Corporate Capture of Congress in Carbon Politics: Evidence from Roll Call Votes*

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This Draft: September 2023

*We would like to thank Heitor Almeida, Yao Chen, Jack Cooney, Tatyana Deryugina, Umit Gurun, Jie He, Xuewen Liu, Nolan Miller, Brett Myers, Jeff Netter, Enrico Onali, Neslihan Ozkan, Joshua Pollet, Jacopo Ponticelli, Yiming Qian, Julian Reif, Maximilian Rohrer, Dragon Tang, Grzegorz Trojanowski, Steven Utke, Roman Wang, Tiancheng Yu, and conference and seminar participants at the 2023 Midwest Finance Association meetings, Hong Kong University, Texas Tech University, University of Bristol, University of Connecticut, University of Exeter, University of Georgia, and University of Illinois at Urbana-Champaign for helpful comments and suggestions. Yifu Zhou provided excellent research assistance. We retain responsibility for any remaining errors.

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Abstract

How special interests shape legislative decisions is a fundamental question in political economy. In this paper, we examine the influence of carbon-emitting corporations on legislative voting behavior on climate bills and the impact of the votes on firm value. Using a comprehensive sample of votes on contested climate bills in the US House of Representatives and Senate, we find that politicians with high carbon dependency, i.e., those whose campaigns received more contributions from carbon-emitting firms, are more likely to cast climate-skeptic votes. This relation is stronger when politicians face greater electoral pressure. Using the narrow defeat of incumbent politicians to generate plausibly exogenous shocks to their elected peers' carbon dependency, we find that the relation is likely causal. We further find evidence that carbon-emitting firms benefit from their connected politicians casting climate-skeptic votes.

JEL Classification: G18, P16, D72, Q54, Q58

Keywords: Special interests, climate change legislation, roll call voting, campaign contributions, political connection

1 Introduction

The influence of campaign contributions by special interest groups on legislative votes has been a subject of much debate. Some contend that campaign contributions allow special interests to purchase the support of legislators, while others argue that the primary goal of such contributions is to influence election outcomes rather than directly sway legislative decision-making. Despite the intense debate on this topic, evidence on the effects of campaign contributions on congressional voting behavior remains scarce.

In this paper, we examine the influence of contributions by carbon-emitting economic interests on congressional votes on climate change legislation. We focus on climate bills because of two reasons. First, climate change caused by greenhouse gas emissions is one of the biggest environmental challenges the world faces. Since carbon emissions generate significant negative externalities, firms seeking to maximize profits have little incentive to curb emissions because they do not bear the full cost of their actions. Thus, government intervention, e.g., through a carbon tax or cap-and-trade program, can play a crucial role in combating climate change.[1] The effectiveness of government intervention, however, could be undermined by special interests. Since climate legislations can have far-reaching effects on firm value and performance, corporations have strong incentives to influence the political process.

Second, climate-related bills provide a setting where one can readily assess the impact of a reasonably large set of legislations on individual firms. Specifically, the classification of climate bills as either "pro-climate" or "anti-climate" by the League of Conservation Voters (LCV) provides a convenient way to determine whether a bill contributes to alleviating climate change. As carbon-emitting firms are typically adversely affected by climate-friendly

[1] In a survey of finance academics, professionals, and regulators, Stroebel and Wurgler (2021) find that regulatory risk is viewed as the top climate risk to businesses and investors over the next five years. Similarly, Krüger, Sautner, and Starks (2020) show that regulatory climate risks have already become important concerns for a majority of institutional investors surveyed.

policies and tend to benefit from climate-unfriendly ones, carbon emissions can serve as a useful metric to assess the positions of firms on a given climate bill.

Corporations that stand to lose from more climate-friendly legislations have spent heavily on influencing the political process. For example, according to the Center for Responsive Politics, the amount of campaign contributions by oil and gas companies to federal candidates during the 2019-2020 election cycle is more than eight times higher than that of renewable energy companies ($139.3 million vs. $14.2 million). While there is speculation in the popular press that political donations by high emission companies have allowed them to gain undue influence over elected politicians and outcomes on climate legislations (e.g., *The Economist*, 2009; *New York Times*, 2009, 2010, 2021),[2] there is no systematic evidence that campaign contributions by carbon-emitting corporate interests influenced politicians' roll call votes on climate bills.

The purpose of our paper is to investigate how corporations influence legislative voting behavior on climate bills and the impact of the votes on firm value. If campaign donations enable firms to influence politicians' roll call votes, politicians whose campaign is backed to a greater extent by carbon-emitting firms should be more likely to cast climate-skeptic votes in Congress (i.e., voting against climate-friendly bills and voting for climate-unfriendly bills). Since roll call votes can directly impact the legislative outcome, carbon-emitting firms should benefit from climate-skeptic votes cast by politicians to whom they make campaign contributions.

We use a sample of 37,443 roll call votes cast on bills related to climate change in the US House of Representatives and Senate during the period from 2011 to 2021 to examine the

[2] For example, commenting on campaign contributions made by oil and gas companies and electric utilities, a *New York Times* (2009) article quotes the director of a watchdog group as stating that "what making campaign contributions provides you [...] is enhanced access with members of Congress. It doesn't guarantee outcomes but it increases your odds of being able to influence the outcomes." As another example, in a 2007 interview on CBS, when asked why the federal government had not acted on climate change, then-Senator John McCain responded that "utility companies and the petroleum companies and other special interests [...] are the ones that have blocked progress in the Congress of the United States and the administration."

influence of corporate interests in the politics of climate change. We focus on 126 climate bills that pass or fail within ±10% around the passing threshold, because individual legislators are more likely to be pivotal on contested bills. Consistent with a wide partisan divide on the issue of climate change, Republican lawmakers are much more likely to cast climate-skeptic votes than their Democratic counterparts; the average likelihood is 93.2% for Republicans, as compared to 3.0% for Democrats.

We obtain data on greenhouse gas emissions from the US Environmental Protection Agency's Greenhouse Gas Reporting Program (GHGRP), which requires facilities that emit greenhouse gases and suppliers of fuels and industrial gases to report their emissions starting from 2010. We use contribution-weighted emissions to measure the "carbon dependency" of a politician, i.e., the extent to which the politician's election campaign is backed by carbon-emitting firms. Specifically, for a politician at a given point in time, carbon dependency is computed as a weighted average of scaled GHG emissions of firms that contributed to the politician's election campaigns in the past, weighted by the dollar value of the contributions.

We show that carbon dependency positively predicts the likelihood of a politician casting a climate-skeptic vote. This result holds after controlling for high-dimensional fixed effects, such as bill fixed effects, bill × political party fixed effects, state × year fixed effects, and legislator fixed effects, thereby eliminating many potential sources of omitted variable bias that can confound inferences. For example, our bill × political party fixed effects mitigate the concern that party-line voting on legislative bills drives our results. Also, time-varying economic fundamentals and/or public opinions in a state cannot explain our results either, because our state × year fixed effects absorb observed and unobserved time-varying heterogeneity across lawmakers from different states. In terms of economic magnitude: a one standard deviation increase in carbon dependency is associated with an increase of 1.1 to 2.2 percentage points in the likelihood of casting a climate-skeptic vote, which is economically nontrivial considering that 13.5% and 19.8% of the climate bills in our sample have a margin of victory within 1.1 and 2.2 percentage points, respectively.

Consistent with the idea that politicians become more responsive to corporations' campaign contributions when facing greater electoral pressure, we find that the positive relation between carbon dependency and the likelihood of casting a climate-skeptic vote becomes stronger when the lawmaker's winning margin in the most recent election is small and when the lawmaker is from a swing district/state.

While the inclusion of a large set of fixed effects enables us to rule out alternative interpretations based on bill-specific factors, time-varying characteristics of states, and time-invariant factors that are specific to individual lawmakers, it remains possible that omitted variables, e.g., time-varying factors that are specific to lawmakers, drive both carbon dependency and voting decisions. To address this potential endogeneity concern, we use narrowly defeated incumbent politicians to generate plausibly exogenous shocks to their elected peers' carbon dependency.

When an incumbent focal politician loses reelection, corporations that contributed to her election campaign suffer a loss in political capital and hence have to make up for the loss by securing the support of like-minded elected politicians (peer politicians). As a result, relative to corporations that contributed to a focal politician that narrowly wins, those that contributed to a focal politician that narrowly loses will direct their contributions to the focal politician's elected peers, which could generate plausibly exogenous changes in the carbon dependency of the peers.

Importantly, peer politicians of the defeated focal politician may exhibit differential exposures to the shock, with some positively exposed and some negatively exposed. If a peer politician is funded to a lesser extent by carbon money than the losing focal politician before the close race, the peer is likely to receive disproportionately more carbon money afterward when the losing politician's donors direct their contributions to peers. In this case, the peer is positively exposed to the shock in the sense that he or she is likely to experience an increase in carbon dependency afterward due to post-election "makeup" contributions made by the losing politician's donors. On the other hand, if a peer politician is funded to a greater extent by carbon money than the losing focal politician before the close race, the peer is negatively exposed because he or she is likely to receive disproportionately more non-carbon money and experience a decrease in carbon dependency afterward.

We use a triple-difference specification to test this prediction. The first difference is between treated politicians ($Treated$), i.e., peers of a focal politician that loses reelection bid in a close race, and control politicians, i.e., peers of a focal politician that wins reelection in a close race. The second difference is from before to after the close-race event ($Post$). The third difference is across peer politicians with varying levels of exposure to the shock ($Exposure$), which is captured by the difference in the pre-event carbon dependency between the focal politician and the peers. The triple-difference term ($Treated \times Post \times Exposure$) captures how a close loss of a focal politician changes the peer politicians' tendency to cast climate-skeptic votes when the peers are more positively exposed to the shock (i.e., likely to receive disproportionately more campaign contributions from carbon interests). If shocks to carbon dependency induced by focal politicians' narrow losses influence peer politicians' roll call voting behavior, the coefficient on the triple-difference term should be positive and significant.

We find evidence consistent with the idea that changes in carbon dependency induced by close races influence politicians' roll call voting behavior. Specifically, compared to the peers of a focal politician who wins reelection in a close race, the peers of a focal politician who loses become more likely to cast climate-skeptic votes after the close-race event than before the event when their pre-event carbon dependency is lower than that of the focal politician (and hence their carbon dependency is likely to increase after the event). The economic magnitude of this effect is also significant. For instance, the coefficient on the triple-difference term suggests that a one standard deviation higher exposure to the shock is associated with a 0.9 to 1.2 percentage points higher likelihood of casting a climate-skeptic vote for treated politicians, relative to control politicians, after the close race than before. This result is obtained after controlling for a host of fixed effects, including close-race event fixed effects, bill $\times$ political party fixed effects, and state $\times$ year fixed effects. This evidence provides support for a causal interpretation of the relation between campaign contributions and legislative votes.

Given the results on the influence of corporate interests on politicians' roll call votes, a natural question is how much corporations benefit from this influence. To explore this question, we construct a climate-skeptic vote share measure for each firm-bill pair by computing a weighted average climate-skeptic vote share among politicians who received campaign contributions from the firm, weighted by the dollar value of the contributions. If firms benefit from their connected politicians voting in their favor, carbon-emitting firms should earn higher abnormal stock returns when their connected politicians disproportionately cast climate-skeptic votes. We find evidence consistent with this prediction. Specifically, we find that a one standard deviation increase in scaled emission, combined with a one standard deviation increase in the climate-skeptic vote share, is associated with an increase of 11.6 to 15.8 basis points in the CAR around the vote on a climate bill. Furthermore, the higher abnormal stock return associated with carbon-emitting firms when their connected politicians vote in their favor is driven entirely by bills that climate skeptics win. This result provides suggestive evidence that firms benefit from influencing the voting behavior of politicians and sheds light on the underlying motives behind corporations' political contributions.

Our paper is related to three strands of the literature, with the first being that on climate regulation and its impact on financial markets. Climate regulation is crucially important to the transition to a low-carbon economy. Recent studies show that regulatory risk from climate change is viewed as the most important short-run risk by investors and academics in the finance community (Krüger, Sautner, and Starks, 2020; Stroebel and Wurgler, 2021) and that it can have significant impacts on security pricing (e.g., Chava, 2014; Ilhan, Sautner, and Vilkov, 2019; Ramelli, Wagner, Zeckhauser, and Ziegler, 2021; Bolton and Kacperczyk, 2021, 2022; Seltzer, Starks, and Zhu, 2022). The climate policymaking process, however, could be captured by carbon-emitting economic interests. Our paper is the first in the literature to provide evidence that carbon-emitting corporate interests influence climate legislation. Our results suggest that corporations most affected by climate legislation use their influence to actively shape legislative decision-making on climate bills.

Second, our paper adds to the literature on congressional voting, which has provided mixed evidence on whether campaign donations by corporations influence legislators' roll call votes. Early studies find little systematic evidence that campaign contributions impact congress members' voting behavior, measured using voting indexes or scores at the politician level.[3] For example, Ansolabehere, Figueiredo, and Snyder (2003) survey about 40 studies that examine the relation between PAC contributions and roll call voting behavior and conclude that "PAC contributions show relatively few effects on voting behavior." More recent studies show a positive correlation between contributions and votes (e.g., Stratmann, 2002; Mian, Sufi, and Trebbi, 2010). However, since campaign contributions and legislators' votes could be driven by "the same underlying factors" (Stratmann, 2002), these studies generally caution against interpreting the correlation as evidence that money buys votes. Our paper contributes to this literature by using plausibly exogenous shocks to campaign contributions to identify the effect.

Third, our paper connects to the literature on the value of political connections. A number of studies examine the influences of firms' political activities and connections, such as

[3] An incomplete list of early studies that examine this question includes Silberman and Durden (1976), Chappell (1982), Kau, Keenan, and Rubin (1982), Peltzman (1984), Hall and Wayman (1990), and Bronars and Lott (1997). These studies generally conduct analyses at the politician level, rather than at the politician-vote level, which is what we do in this paper.

campaign contributions, lobbying, and politically connected corporate executives and board members, on firm outcomes.[4] Our paper contributes to this literature in two important ways. First, instead of looking at firm outcomes, we focus on politicians' actions, specifically their roll call votes, which allows us to shed light on a specific channel through which politicians affect firm value. Second, our identification strategy based on close-call elections and the granular data of roll call votes allow us to provide causal evidence on the impact of political connections.

Our study highlights carbon-emitting economic interests as an important force in shaping climate policymaking, which might provide a partial explanation for the limited progress on climate reforms in the U.S. despite rising public support. For example, a survey conducted by Pew Research Center in 2019 reveals that two-thirds of U.S. adults believe that the federal government is doing too little to reduce the effects of global climate change. While climate reforms are in the public interest, they tend to impose significant costs on carbon-emitting firms' private interests. Our results suggest that carbon interests' influence on the political process can be a barrier to government action on addressing climate change.

The rest of the paper is organized as follows. Section 2 describes sample selection and reports summary statistics. Section 3 presents results on the relation between carbon dependency and legislative voting behavior. Section 4 presents results on the valuation effect of votes on climate bills, and Section 5 concludes.

[4] See, e.g., Lenway, Morck, and Yeung (1996), Kroszner and Stratmann (1998), Fisman (2001), Johnson and Mitton (2003), Ansolabehere, Snyder, and Ueda (2004), Khwaja and Mian (2005), Faccio (2006), Faccio, Masulis, and McConnell (2006), Jayachandran (2006), Ferguson and Voth. (2008), Claessens, Feijen, and Laeven (2008), Faccio and Parsley (2009), Goldman, Rocholl, and So (2009), Cooper, Gulen, and Ovtchinnikov (2010), Yu and Yu (2011), Duchin and Sosyura (2012), Blanes i Vidal et al. (2012), Adelino and Dinc (2014), Akey (2015), Acemoglu et al. (2016), Faccio and Hsu (2017), Brown and Huang (2020), and Faccio and Zingales (2022).

2 Data and Sample

We obtain congressional voting records on environmental legislations for the period between 2011 and 2021 from the League of Conservation Voters (LCV). We choose 2011 as the starting year of our analysis because data on greenhouse gas emissions, which we measure in the year before the vote on a bill, become available starting from 2010. We focus on climate bills that pass or fail within ±10% around the passing threshold, because individual legislators are more likely to be pivotal on contested bills.[5] We consider a bill as related to climate change if LCV lists "climate" or "climate change" as one of the issues addressed by the bill or if the description of the bill contains key words related to climate change, including "climate change," "global warming," "greenhouse gas," and "methane." There are a total of 126 contested votes on climate bills during the sample period. Panel A of Table 1 shows the distribution of climate bills by year and chamber. The number of climate bills in a year ranges from six to 18. The House of Representatives and the Senate vote on 75 and 51 climate bills, respectively, during the sample period.

LCV classifies each environmental related bill as "anti-environment" or "pro-environment". As an example of an anti-environment bill, consider H.R. 2042, the Ratepayer Protection Act of 2015, which would allow states to opt out of carbon pollution standards for fossil fuel-fired power plants. Voting in favor of this bill constitutes a climate-skeptic vote, whereas voting against it represents a climate-supportive vote. The bill passed the House on June 24, 2015, which marks a win for climate skeptics. As an example of a pro-environment bill, consider S.J. Res. 14, which seeks to repeal the Environmental Protection Agency's rollback of the 2016 methane emission standards in the oil and gas industry and allow the Agency to implement stronger measures against methane pollution. In this case, voting in favor of the bill is a climate-supportive vote, while voting against it is a climate-skeptic vote. The

[5] During our sample period, a large majority (80.3%) of climate bills pass or fail within ±10% around the passing threshold.

amendment passed the Senate on April 28, 2021, which represents a win for climate activists (i.e., a climate-supportive win or a climate-skeptic loss).

Panel B of Table 1 shows that 64.3% of the bills are classified as climate-skeptic and the outcome of 61.1% of the bills is a climate-skeptic win. These aggregate statistics, however, mask a wide partisan divide on the issue of climate change. To shed light on this, we partition the sample into three periods, corresponding to Democratic control of both houses of Congress (2021), Republican control of both houses (2015 through 2018), and split control by the two parties (2011 through 2014, 2019, and 2020). Panel B of Table 1 shows that during Democratic control of both houses of Congress, only 16.7% of the climate bills being voted on are climate-skeptic bills and climate skeptics prevail on only 5.6% of the climate bills. In contrast, the corresponding numbers during Republican control of both houses are 75.0% and 88.5%. The period with split control of Congress has numbers falling between those for the periods with unified control of Congress. This pattern is consistent with the observation that Democrats favor policies that mitigate climate change, whereas Republicans have largely resisted such policies.

[Insert Table 1 about here]

Our main sample of roll call votes contains 37,443 votes cast on these 126 climate bills. We construct an indicator, *Climate-skeptic vote*, which equals one if the legislator votes for a climate-skeptic bill or votes against a climate-supportive bill and zero otherwise. Panel A of Table 2 shows that the variable has a mean of 50.5% and a standard deviation of 50.0%. Consistent with the wide partisan divide, Republican lawmakers are much more likely to cast climate-skeptic votes than their Democratic counterparts. The average likelihood is 93.2% for Republicans, as compared to 3.0% for Democrats. Importantly, there is considerable variation in the tendency to cast a climate-skeptic vote among Democratic lawmakers and among Republican ones. The *Climate-skeptic vote* indicator for Democrats has a standard deviation of 0.170, and that for Republicans has a standard deviation of 0.251. In untabulated results, the indicator has a within-legislator (across-legislator) standard deviation of 0.138 (0.143) among Democrats and 0.227 (0.126) among Republicans, suggesting that within-legislator variability is as important as across-legislator variability for lawmakers from both parties.

We obtain data on greenhouse gas emissions from the US Environmental Protection Agency's Greenhouse Gas Reporting Program (GHGRP), which requires facilities that emit greenhouse gases and suppliers of fuels and industrial gases to report their emissions starting from 2010. GHGRP provides facility-level emissions information in metric tons of CO_2 equivalent in terms of global warming potential. The program covers about 85 to 90% of total U.S. greenhouse gas emissions. We aggregate facility-level emission data to the company level.

We obtain data on campaign contributions from the Center for Responsive Politics (CRP) and merge the CRP data with the GHGRP data on greenhouse gas emissions by company name. We focus on hard money Political Action Committee (PAC) contributions (following Cooper, Gulen, and Ovtchinnikov, 2010). Since firms are subject to mandatory reporting of emissions and campaign contributions regardless of whether they are publicly-traded or privately-owned, the merged data include both public firms and private ones.

We construct a new measure, referred to as "carbon dependency," to gauge the extent to which a legislator's election campaign is backed by carbon-emitting firms. For a politician at a given point in time, carbon dependency is computed as a weighted average of the most recent GHG emission shares of firms that contributed to the politician's election campaigns over the preceding 60 months, with weights assigned based on the dollar amount of the contributions. The emission share of a firm is the reported carbon equivalent emissions of the firm in a year divided by the total reported carbon equivalent emissions under GHGRP in that year. The carbon dependency measure quantifies a politician's ties to carbon-emitting firms.

Panel A of Table 2 shows that carbon dependency has a mean of 0.042% and a standard deviation of 0.036%. Consistent with the view that Republican lawmakers are backed to a greater extent by carbon interests than their Democratic counterparts, the mean carbon dependency is 0.060% for Republicans, as compared to 0.022% for Democrats.

Panel A of Table 2 also shows summary statistics of other variables. *District emission share* is the reported carbon equivalent emissions of facilities in a lawmaker's congressional district (for representatives) or state (for senators) in a year as a fraction of total reported carbon equivalent emissions under GHGRP in that year. Republican lawmakers' districts/states tend to have a higher emission share than those of Democrats; the mean *District emission share* is 0.605% for Republicans, as compared to 0.321% for Democrats, suggesting that districts/states that are more dependent on carbon-emitting industries are more likely to elect Republicans. *Tenure* is the number of years since the legislator assumed office. The average tenure is about 11 years.

We construct a panel of firm-bills to examine the valuation impact of votes on contested climate bills. For each contested climate bill, we consider all firms with common stocks traded on the NYSE, Nasdaq, and AMEX. The panel consists of 78,081 firm-bill pairs. For each firm-bill pair in our sample, we compute a weighted average climate-skeptic vote share on the bill among legislators whose campaigns received PAC contributions from the firm, weighted by the dollar value of the contributions. If all politicians backed by a firm cast climate-skeptic (climate-supportive) votes on a bill, the climate-skeptic vote share takes a value of one (zero). Panel B of Table 2 shows that the climate-skeptic vote share has a mean of 0.605 and a standard deviation of 0.253. Climate skeptics win the vote on 62.3% of the bills in the panel of firm-bills.

To capture firms' exposure to climate bills, we define *Emission share* as a firm's reported carbon equivalent emissions in the year before the vote on a bill as a fraction of total reported carbon equivalent emissions under GHGRP in that year. Panel B of Table 2 shows that *Emission share* has a mean of 0.072% and a standard deviation of 0.320%. We use cumulative abnormal returns (CARs) from five days before to five days after the date of the vote on a climate bill to measure valuation effects. We compute both market-adjusted and DGTW characteristics-adjusted abnormal returns. The market-adjusted and DGTW characteristics-adjusted CARs have a mean of −0.174% and 0.114%, respectively.

Panel B of Table 2 also reports summary statistics on firm characteristics, including total assets, market-to-book ratio, prior stock return, leverage, capital expenditures, S&P 500 membership, institutional ownership, and the number of analysts covering the firm. We measure these firm characteristics using the most recent information publicly available to investors at the time when a given climate bill is voted on.

[Insert Table 2 about here]

3 Carbon Dependency and Roll Call Votes

In this section, we first run OLS regressions to examine the relation between a politician's carbon dependency in campaign financing and the likelihood of casting climate-skeptic votes. We then use the narrow defeat of an incumbent politician as a shock to the carbon dependency of her elected peers to identify the effect of carbon dependency on politicians' roll call voting behavior.

3.1 Baseline results

If campaign contributions influence legislative decision making, politicians with greater carbon dependency should be more likely to vote in favor of carbon-emitting corporate interests. To the extent that climate-unfriendly bills tend to benefit firms with higher carbon emissions and climate-friendly bills tend to harm them,[6] politicians with higher carbon dependency should be more likely to cast climate-skeptic votes, i.e., vote for climate-unfriendly bills and against climate-friendly ones.

We run OLS regressions using the sample of roll call votes to examine the relation between carbon dependency and the tendency to cast climate-skeptic votes. The main regression takes the following form,

$$\textit{Climate-skeptic vote}_{l,b} = \alpha_b + \beta \times \textit{Carbon dependency}_{l,t-1} + \gamma \times \mathbf{X}_{l,t-1} + \varepsilon_{l,b}, \qquad (1)$$

where $\textit{Climate-skeptic vote}_{l,b}$ is an indicator for whether legislator l casts a climate-skeptic vote on bill b; α_b is bill fixed effects; $\textit{Carbon dependency}_{l,t}$ is legislator l's carbon dependency measured in month $t-1$, i.e., the month before bill b is put to a vote; $\mathbf{X}_{l,t-1}$ consists of legislator characteristics measured in month $t-1$, including the carbon emission share of the district/state that the legislator represents, political party indicators (using independents as the base group), and tenure. We additionally control for bill $\times$ political party fixed effects, state $\times$ year fixed effects, and legislator fixed effects in some regressions. We compute standard errors using two-way clustering by bill and by legislator to allow for arbitrary within-bill and within-legislator correlation in residuals.

[6] We use stock returns around the vote on climate bills to test this assumption. If carbon-emitting firms on average gain from climate-unfriendly bills and lose from climate-friendly ones, their stock price should react positively to a climate-skeptic win on a bill (i.e., the passage of a climate-unfriendly bill or the rejection of a climate-friendly bill). We construct a panel of firm-bills and regress abnormal stock returns around the vote on a climate bill on an indicator for a climate-skeptic win, firms' carbon emissions, an interaction between the two variables, and firm controls as well as a host of fixed effects. Appendix Table A1 shows that the coefficient on the interaction term is positive and significant in all specifications, suggesting that carbon-emtting firms generally benefit from climate-unfriendly bills and suffer from climate-friendly ones.

Table 3 presents the regression results. Column 1 includes bill fixed effects, and column 2 further includes state × year fixed effects. With bill fixed effects, identification comes from variation across legislators with different degrees of carbon dependency within a given bill. These fixed effects control for any observed and unobserved heterogeneity across bills (e.g., issues being voted upon and the public opinion about the bill), allowing us to rule out differences in bills as potential explanations for our results. With state × year fixed effects, identification comes from variation in carbon dependency across legislators within a state-year. This specification allows us to control for time-varying climate, economic, social, and political considerations specific to a state that could impact legislators' voting decisions. For example, the state × year fixed effects effectively capture the impact of climate-related events specific to a state-year, such as natural disasters and extreme weather conditions, thereby isolating the influence of carbon dependency on legislators' voting decisions from confounding effects related to such events. Besides state × year fixed effects, column 3 includes bill × political party fixed effects, which mitigate the concern that party-line voting on legislative bills drives our results. All three columns show that the coefficient of *Carbon dependency* is positive and significant at the 1% level, indicating that legislators with greater carbon dependency in campaign funding are more likely to cast climate-skeptic votes. The economic magnitude is nontrivial: a one standard deviation increase in carbon dependency is associated with an increase of 1.9 to 2.2 percentage points in the likelihood of casting a climate-skeptic vote. Since 18.3% and 19.8% of the climate bills in our sample have a margin of victory within 1.9 and 2.2 percentage points, respectively, these results suggest that the change in the tendency to cast a climate-skeptic vote induced by a one standard deviation change in carbon dependency, when aggregated across legislators, can flip the result of a substantial number of climate bills.

Column 4 of Table 3 further includes legislator fixed effects, forcing identification to come from variation in a given legislator's carbon dependency over time. The coefficient of *Carbon* *dependency* remains positive and significant. A one standard deviation increase in carbon dependency is associated with an increase of 1.1 percentage points in the likelihood of casting a climate-skeptic vote, which is again economically nontrivial given that 13.5% of the climate bills have a margin of victory within 1.1 percentage points. The reduction in magnitude with the inclusion of legislator fixed effects is expected, because across-legislator variation in carbon dependency could be correlated with factors that influence voting behavior on climate bills. For example, legislators who are skeptical of climate change are likely to cast climate-skeptic votes in Congress and, at the same time, attract a disproportionate amount of campaign contributions from carbon-emitting firms. In this case, it is legislators' preferences on climate policies that drive both carbon dependency and voting behavior. To the extent that lawmakers' preferences are relatively stable over a legislator's tenure, the inclusion of legislator fixed effects helps to mitigate the concern that the observed results are driven by such preferences.

Table 3 also shows that Republican legislators are significantly more likely to cast climate-skeptic votes than Independents and that Democratic legislators do not differ significantly from Independents in their tendency to cast climate-skeptic votes. This partisan divide is consistent with the univariate result reported in Panel A of Table 2. Also, the coefficient of $Tenure$ is negative and generally marginally significant, suggesting that lawmakers become less likely to take a climate-skeptic stance as their tenure increases.

Interestingly, the coefficient on *District emission share* is positive but statistically insignificant. In untabulated analysis, the coefficient becomes significantly positive when we drop the political party indicators from the regression in the first two columns of Table 3. These results suggest that districts with higher levels of carbon emissions are more likely to elect legislators who are skeptical of climate change, which tend to be Republicans, because these legislators might be seen as better able to protect the jobs of constituents who work in carbon-emitting industries.

[Insert Table 3 about here]

3.2 Controlling for public opinions

Lawmakers may respond to their constituents' concerns by aligning their votes with public opinions on climate change. As Howe, Mildenberger, Marlon, and Leiserowitz (2015) show, there is substantial geographic variation of beliefs toward climate change. Lawmakers may cast climate-skeptic votes when their constituents are more skeptical about climate change. To control for this, we use data from Yale Climate Opinion Maps (YCOM) to measure public opinions on climate change. YCOM is a project developed by the Yale Program on Climate Change Communication. It is based on a nationally representative survey, which aims to capture the beliefs, attitudes, and perceptions of Americans towards climate change. The survey was conducted biennially from 2014 through 2018 and switched to being conducted annually from 2019 onward. We use their congressional district-level and state-level data for 2014, 2016, 2018, 2019, and 2020, and link them to our sample of roll call votes by US representatives and senators, respectively. For each vote, we use the most recent YCOM data available at the time of the vote. Thus, the sample used in this analysis spans the period from 2015 to 2021.

YCOM provides a comprehensive view of public opinions on climate change. We use the percentage of adults who think global warming is happening as a measure for public opinions on climate change, although using other measures, such as that of adults who think global warming will harm people in the US and that of adults who support regulating CO_2 as a pollutant, yields qualitatively similar results. We add the public opinion measure as a control in Eq. (1) and report the results in Table 4. The coefficient on the measure is negative and significant in the first three columns, suggesting that politicians are less likely to cast climate-skeptic votes when their constituents are aware of and concerned about climate change. However, the coefficient becomes insignificant in column 4 when we introduce legislator fixed effects into the regression. Thus, the relation between constituents' beliefs and legislators' votes is largely driven by legislator-specific effects. This result suggests that constituents who are more concerned about climate change are more likely to elect legislators who already share their views, rather than the preferences of the constituents influencing the legislators' votes.

Importantly, the coefficient on *Carbon dependency* remains positive and significant after controlling for public opinions on climate change, suggesting that the relation between campaign contributions and voting behavior is not simply driven by constituents' preferences.

[Insert Table 4 about here]

3.3 Cross-sectional tests

Politicians may become more responsive to corporations' campaign contributions when facing greater electoral pressure. We use two proxies to capture electoral pressure. The first is an indicator for whether the lawmaker's winning margin in the most recent election is less than 20 percentage points (*Non-landslide*). Politicians who have won non-landslide elections are more likely to be concerned about losing their seat, and hence may be more likely to heed the demands of their donors in order to increase their chances of reelection.

The second proxy for electoral pressure is an indicator for whether the lawmaker is from a swing district/state. We define a congressional district as a swing district if the winning candidates for the House of Representatives in that district in the past five years do not come from the same party. Similarly, a state is considered as a swing state if the winning candidates for the Senate in that state in the past ten years do not come from the same party. In swing districts and states, the outcome of an election is more uncertain, which may prompt politicians to be more responsive to corporate donors in order to secure reelection.

We add one of the two measures of electoral pressure and its interaction terms with all explanatory variables to Eq. (1). Table 5 reports the regression results. The coefficients on the interaction terms combining *Carbon dependency* and the electoral pressure measures are positive and significant in six out of eight specifications, suggesting that greater electoral competition makes politicians more responsive to campaign contributions in their roll call votes. The only two specifications where the coefficients on the interaction terms lose statistical significance are those in columns 4 and 8 where legislator fixed effects are included, although the magnitude of the coefficients remains economically large.

[Insert Table 5 about here]

3.4 Identification

While the inclusion of a large set of fixed effects enables us to rule out alternative interpretations based on bill-specific factors, time-varying characteristics of states, and time-invariant factors that are specific to individual lawmakers, a remaining identification concern is that time-varying politician-level omitted variables drive both carbon dependency and voting decisions. For example, a big shale discovery in a congressional district is likely to increase the likelihood that the congressperson in that district casts climate-skeptic votes, and at the same time the congressperson is likely to receive more campaign contributions from carbon-emitting firms. To address this potential endogeneity concern, we use the narrow defeat of an incumbent politician to generate plausibly exogenous shocks to her elected peers' carbon dependency.[7]

[7] Lee, Moretti, and Butler (2004) and Lee (2008) provide evidence that the outcome of close congressional elections is akin to random assignment in the narrow range around the 50% vote share threshold and that the districts where a party narrowly wins or loses have similar pre-determined characteristics. A few studies have used close-race elections as identification strategies (e.g., Lee, Moretti, and Butler, 2004; Lee, 2008; Akey, 2015).

When an incumbent focal politician loses reelection, corporations that contributed to her election campaign suffer a loss in political capital and hence have to make up for the loss by securing the support of like-minded elected politicians (peer politicians). As a result, relative to the peers of a focal politician that narrowly wins reelection, the peers of a focal politician that narrowly loses are likely to receive "makeup" contributions from corporations that contributed to the focal politician, which could generate plausibly exogenous changes in the carbon dependency of these peers.

To illustrate, suppose Representative L narrowly loses her reelection bid, whereas Representative W narrowly wins. Companies contributing to Representative L's election campaign, relative to those contributing to Representative W's campaign, will increase their contributions to similar-minded elected peer congresspersons to make up for the loss in political connection. Thus, compared to Representative W's peers, Representative L's peers will experience a plausibly exogenous shock to their carbon dependency, which varies with their pre-event level of carbon dependency relative to that of Representative L. As a result of post-election "makeup" contributions made by Representative L's donors to the peer politicians, the carbon dependency of these peers would become more similar to that of Representative L. This implies that peers with a lower (higher) pre-event carbon dependency than Representative L would experience an increase (decrease) in carbon dependency afterward. If changes in carbon dependency influence votes, then positively exposed peer politicians should become more likely to cast climate-skeptic votes after the close-race event than negatively exposed ones. In other words, because donors to a defeated focal politician shift their contributions to the focal politician's elected peers, the peers would become more similar to the focal politician in terms of carbon dependency and hence the likelihood of casting climate-skeptic votes.[8]

[8]If elected peer politicians heed the lessons of the defeat of the focal politician, they might move *away* from the focal politician's stance on climate change. Thus, a learning effect would bias against finding our results.

Figure 1 provides an illustration of how the defeat of a focal politician affects peer politicians. Representatives T_1 and T_2 are elected congresspersons in the same party and same state as Representative L, so these two politicians are peers of Representative L and are treated politicians. Representatives C_1 and C_2 are elected congresspersons in the same party and same state as Representative W, so they are peers of Representative W and are control politicians. If Representative L's narrow defeat prompts the donors to step up support of Representatives T_1 and T_2, the carbon dependency of these two treated politicians should move towards that of Representative L after the event, which in turn should affect their likelihood of casting climate-skeptic votes. Specifically, suppose Representative T_1 is funded to a lesser extent by carbon money than Representative L before the event, as indicated by a lighter brown color. Thus, Representative T_1 is likely to receive disproportionately more carbon money after the event than before, which could lead to an increase in her tendency to cast climate-skeptic votes. Representative T_2, on the other hand, is funded to a greater extent by carbon money than the focal politician before the event, as indicated by a darker brown color. As a result of the narrow defeat of the focal politician, Representative T_2 is likely to receive disproportionately more non-carbon money afterward and become less likely to cast climate-skeptic votes. In contrast, the carbon dependency and likelihood of casting climate-skeptic votes of the control politicians, i.e., Representatives C_1 and C_2, should not exhibit such changes.

We use a triple-difference specification to test this prediction. The first difference is between treated politicians ($Treated$), i.e., peers of a focal politician that loses reelection bid in a close race, and control politicians, i.e., peers of a focal politician that wins reelection in a close race. The second difference is from before to after the close-race event ($Post$). The third difference is between peer politicians that are differentially exposed to the shock ($Exposure$), which is captured by the difference in the pre-event carbon dependency between the focal politician and the peers. The triple-difference term ($Treated \times Post \times Exposure$)

captures how a close loss of a focal politician changes the peer politicians' tendency to cast climate-skeptic votes when the peers are more positively exposed to the shock (i.e., likely to receive disproportionately more campaign contributions from carbon interests). If shocks to carbon dependency induced by focal politicians' narrow losses influence peer politicians' roll call voting behavior, the coefficient on the triple-difference term should be positive and significant.

We define peer politicians as elected congresspersons that are in the same party, same state, and same chamber as the focal politician. As Table 2 shows, Democratic and Republican lawmakers have very different views about climate change, suggesting that party affiliation is an important consideration when corporations make their campaign contribution decisions. Geography is another important consideration, because politicians systematically favor local firms (e.g., Faccio and Parsley, 2009). Thus, following the defeat of an incumbent focal politician, firms that contributed to the politician's election campaign are likely to increase their contributions to same-party, same-state elected politicians in the same chamber as the focal politician to make up for the loss in political capital.

To focus on congressional races for which the outcome is likely to be exogenous, we use a sample of close races, defined as those for which the winning margin, i.e., the difference between the percentages of votes received by the winner and the runner-up, is within ±2.5%.[9] We require that the outcome of the close races be a narrow defeat or a narrow win for an incumbent legislator. We exclude a narrow defeat (win) of an incumbent legislator if it coincides with a close win (defeat) of a same-state same-party candidate in the same chamber, because in such cases the shocks to political connection tend to cancel out for connected firms.

[9] The choice of the winning margin within ±2.5% reflects a trade-off between selecting a bandwidth large enough to have a reasonable sample size to detect the effects, yet small enough to avoid incorporating races for which the outcome is likely to be endogenous to the characteristics of the politicians. Appendix Table A2 shows that the results are qualitatively similar when using alternative cutoffs, i.e., ±3% and ±2%.

We focus on close races involving incumbent representatives rather than senators because the distribution of campaign contributions is different across the two types of races. In close races involving incumbent representatives, PAC donors are more likely to focus their contributions on the incumbent, which means that they are more likely to make post-race "make-up" contributions to like-minded peer politicians when the incumbent loses. In contrast, in senate races, PAC donors often donate to both candidates, which reduces the need for them to make post-race "make-up" contributions. In our sample of close races involving incumbent representatives, only 2.9% of the donors to the incumbent also contribute to the opponent in the race. This is in contrast to close races involving incumbent senators, where 94.3% of the donors to the incumbent also contribute to the opponent.

For a close race to be included in our sample, we further require that the focal politician involved in the race have at least one peer politician, defined as one who is in the same party and state as the focal politician, and that the peer politician cast votes on climate bills both in the pre- and post-election periods. The final sample includes 28 close races involving 319 peer politicians.

We construct a panel of roll call votes by peer politicians during a two-year window before and a two-year window after the close-race events. We run the following triple-difference regression,

$$\begin{aligned} \textit{Climate-skeptic vote}_{e,l,b} = \alpha_e + \alpha_b + \gamma_1 \times \textit{Treated}_e \times \textit{Post}_{e,b} \times \textit{Exposure}_{e,l} \\ + \gamma_2 \times \textit{Treated}_e \times \textit{Exposure}_{e,l} + \gamma_3 \times \textit{Post}_{e,b} \times \textit{Exposure}_{e,l} \\ + \gamma_4 \times \textit{Pass}_e \times \textit{Post}_{e,b} + \gamma_5 \times \textit{Post}_{e,b} + \gamma_6 \times \textit{Exposure}_{e,l} + \varepsilon_{e,l,b}, \end{aligned} \tag{2}$$

where $\textit{Climate-skeptic vote}_{e,l,b}$ is the indicator for climate-skeptic vote on bill b cast by legislator l, which is a peer of the focal politician in close-race event e; α_e is event fixed effects; α_b is bill fixed effects; $\textit{Treated}_e$ is an indicator that equals one if the focal politician in event e loses a close race and zero if he or she wins one; $Post_{e,b}$ is an indicator that equals one if bill b is in the two-year period after event e and zero if it is in the two-year period before the event; $Exposure_{e,l}$ is the difference in carbon dependency between the focal politician in event e and legislator l immediately before the close-race event. A more positive value of $Exposure_{e,l}$ indicates that legislator l is less carbon dependent than the focal politician in close-race event e before the race, and hence the narrow defeat of the focal politician is likely to lead to an increase in the carbon dependency of legislator l after the race. We compute standard errors using three-way clustering by close-race event, by bill, and by legislator to allow for arbitrary correlation in residuals at different levels. The triple-difference term ($Treated \times Post \times Exposure$) captures how the narrow defeat of a focal politician changes the peer politicians' tendency to cast climate-skeptic votes when the peers are likely to receive more campaign contributions from carbon interests. If shocks to carbon dependency induced by focal politicians' narrow losses influence peer politicians' roll call voting behavior, the coefficient on the triple-difference term, i.e., γ_1, should be positive and significant.

Panel A of Table 6 reports the results. We start with a specification that includes bill fixed effects and close-race event fixed effects and progressively saturate the specification with bill $\times$ political party fixed effects, state $\times$ year fixed effects, and legislator fixed effects. The coefficient on the triple-difference term is positive and significant at conventional levels in all specifications, suggesting that shocks to carbon dependency influence politicians' roll call voting behavior. Thus, relative to the peers of a focal politician that wins reelection in a close race, the peers of a focal politician that loses become more likely to cast climate-skeptic votes after the close-race event than before the event when the peers have lower pre-event carbon dependency than the focal politician (and hence are likely to experience an increase in carbon dependency after the event). The economic magnitude of this effect is also significant. For instance, the coefficient on the triple-difference term implies an increase of 0.9 to 1.2 percentage points in the likelihood of casting climate-skeptic votes for a one standard deviation increase in exposure (about 0.033%) for treated politicians, relative to control ones, after the close-race event than before.

The premise of our identification strategy is that exposure to treatment should positively predict changes in treated politicians' carbon dependency after the close-race event. To test this, we construct a panel of peer politicians with two observations for each peer politician in each close-race event, one before the event and the other after the event. We run the following triple-difference regression at the politician level,

$$\begin{aligned} Carbon\ dependency_{e,l,t} = \alpha_e &+ \delta_1 \times Treated_e \times Post_{e,t} \times Exposure_{e,l} \\ &+ \delta_2 \times Treated_e \times Exposure_{e,l} + \delta_3 \times Post_{e,t} \times Exposure_{e,l} \\ &+ \delta_4 \times Pass_e \times Post_{e,t} + \delta_5 \times Post_{e,t} + \delta_6 \times Exposure_{e,l} + \varepsilon_{e,l,t}, \quad (3) \end{aligned}$$

where $Carbon\ dependency_{e,l,t}$ is the carbon dependency of legislator l, which is a peer of the focal politician in close-race event e, at time t (either immediately before the event or two years after the event); α_e is event fixed effects; $Treated_e$ is an indicator that equals one if the focal politician in event e loses a close race and zero if he or she wins one; $Post_{e,t}$ is an indicator that equals one if time t is after event e and zero if it is before the event; $Exposure_{e,l}$ is the difference in carbon dependency between the focal politician in event e and legislator l before the close-race event. We compute standard errors using two-way clustering by close-race event and by legislator to allow for arbitrary correlation in residuals within an event and within a legislator. The triple-difference term ($Treated \times Post \times Exposure$) captures how the narrow defeat of a focal politician changes a peer politician's carbon dependency when the peer politician is more positively exposed to treatment, i.e., the pre-event carbon dependency of the peer politician is less than that of the focal politician. If a focal politician's narrow defeat prompts the donors to step up support of elected peer politicians and hence move the peer politicians' carbon dependency towards that of the focal politician, the coefficient on the triple-difference term, i.e., δ_1, should be positive and significant.

Panel B of Table 6 reports the results. We again progressively saturate the specification with more stringent fixed effects, including event fixed effects, political party × year fixed effects, state × year fixed effects, and legislator fixed effects. The coefficient on the triple-difference term is positive and significant at the 1% level in all specifications, suggesting that relative to the peers of a focal politician that wins reelection in a close race, the peers of a focal politician that loses experience an increase in carbon dependency after the close-race event than before the event when the peers are more positively exposed to treatment. The economic magnitude of this effect is also significant. For instance, the coefficient on the triple-difference term implies an increase of 0.012 to 0.013 percentage points in carbon dependency for a one standard deviation increase in exposure (about 0.031%) for treated politicians, relative to control ones, after the close-race event than before. These numbers are economically meaningful given that carbon dependency has a mean of 0.030% and a standard deviation of 0.039% in the triple-difference sample. These results suggest that the narrow defeat of a focal politician moves the peer politicians' carbon dependency towards that of the focal politician, providing support for the premise of our identification strategy.

[Insert Table 6 about here]

The identification strategy assumes that absent close-race events for focal politicians, peer politicians do not systematically change their voting behavior on climate bills. To verify the parallel-trends assumption, we conduct falsification tests by assuming that the close-race events occur two years prior to the actual dates. We repeat the regression in Eq. (2) using the votes by peer politicians in the two years before the pseudo-events and the two years after. The results, reported in Panel A of Table 7, show that the coefficient on the triple-interaction term, $Treated \times$ "$Post$" $\times Exposure$, is insignificant in all columns, suggesting that peer politicians of a focal politician that narrowly loses and those of a focal politician that narrowly wins exhibit similar patterns in voting behavior in the four years prior to the actual close-race events. These results suggest that the parallel trends assumption is likely to hold in our setting.

For completeness, we also test whether peer politicians' carbon dependency moves towards that of the focal politician before the close-race events. We repeat the regression in Eq. (3) using peer politicians' carbon dependency around the pseudo-events. The results, reported in Panel B of Table 7, show that the coefficient on the triple-interaction term, $Treated \times$"$Post$"$\times Exposure$, is insignificant in all columns, suggesting that peer politicians do not exhibit systematic changes in carbon dependency as predicted by the contribution reallocation effect in the four years prior to the actual close-race events.

[Insert Table 7 about here]

4 Stock Returns Around Votes on Climate Bills

Given the results on the influence of corporate interests on politicians' roll call votes, a natural question is how much corporations benefit from their influence on politicians. To explore this question, we construct a climate-skeptic vote share measure for each firm-bill pair by computing a weighted average climate-skeptic vote share among politicians that received campaign contributions from the firm, weighted by the dollar value of the contributions (as described in Section 2). If firms benefit from their connected politicians voting in their favor, carbon-emitting firms should earn higher abnormal stock returns when their connected politicians disproportionately cast climate-skeptic votes.[10]

[10] Abnormal stock returns could arise around the vote on a climate bill because of two main reasons. First, the bill can have a direct impact on firm value. That is, a climate-skeptic win on a bill can negatively affect the cash flows of high-emission firms relative to those of low-emission firms, and a climate-supportive win may have the opposite effect. Second, the outcome of the bill contains information about the likelihood

We run the following difference-in-differences regression using the panel of firm-bills to evaluate the valuation impact of connected politicians' votes,

$$CAR_{f,b} = \alpha_b + \kappa_1 \times Emission_{f,t-1} \times \textit{Skeptic vote share}_{f,b} + \kappa_2 \times Emission_{f,t-1} + \kappa_3 \times \textit{Skeptic vote share}_{f,b} + \kappa \times \mathbf{X}_{f,t-1} + \varepsilon_{f,b}, \quad (4)$$

where $CAR_{f,b}$ is the cumulative abnormal stock return (CAR) of firm f around the vote on bill b; α_b is bill fixed effects; $Emission_{f,t-1}$ is firm f's reported carbon equivalent emissions as a fraction of total reported carbon equivalent emissions under GHGRP in year $t-1$, i.e., the year before the vote on bill b; $\textit{Skeptic vote share}_{f,b}$ is a weighted average climate-skeptic vote share on bill b among politicians that received campaign contributions from firm f, weighted by the dollar value of the contributions; $\mathbf{X}_{f,t-1}$ is a set of firm characteristics measured before the vote. We use market-adjusted and DGTW characteristic-adjusted abnormal stock returns to measure the stock price reaction to climate bills. We use an 11-day window centered around the date of the vote on a climate bill (i.e., five days before to five days after the vote), because uncertainty over the outcome of a bill may be resolved before the bill reaches a floor vote and it may take some time for the market to fully incorporate the information about the impact of a bill. We compute standard errors using two-way clustering by bill and by industry to allow for arbitrary correlation in residuals within a bill and within an industry. The difference-in-differences term ($Emission \times \textit{Skeptic vote share}$) captures how stock market reactions to the vote on a bill vary with firms' carbon emissions and their connected politicians' climate-skeptic votes. If carbon-emitting firms benefit from climate-skeptic votes cast by their connected politicians, the coefficient on the difference-in-differences term should be positive and significant.

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of future legislative actions on climate change. For example, the passage of a climate bill may indicate increased legislative support for climate-related issues, thereby resulting in an upward revision of the market's expectation about future climate legislation.

The first three columns in both panels of Table 8 show the results. Consistent with the prediction that carbon-emitting firms benefit from climate-skeptic votes cast by their connected politicians, the coefficient on the difference-in-differences term is positive and significant at the 1% level in all three specifications. A one standard deviation increase in scaled emissions (.0032), combined with a one standard deviation increase in connected politicians' climate-skeptic vote share (.253), is associated with an increase of 13.5 to 15.8 and 11.6 to 14.6 basis points, respectively, in the market-adjusted and characteristic-adjusted CAR around the vote on a climate bill.

The positive valuation effect of connected politicians' climate-skeptic votes on carbon-emitting firms should be particularly pronounced when the outcome of the vote is in these firms' favor, i.e., when it is a climate-skeptic win. To test this, we add interaction terms between an indicator for a climate-skeptic win outcome for bill b, $Skeptic\ win_b$, and the main independent variables in Eq. (4) to test a triple-difference effect.[11] The last three columns in both panels of Table 8 report the results of the triple-difference model. The coefficient on the triple-interaction term ($Emission \times Skeptic\ vote\ share \times Skeptic\ win$) is positive and significant whereas that on the double-interaction term ($Emission \times Skeptic\ vote\ share$) is insignificant, suggesting that the positive abnormal stock return associated with carbon-emitting firms when their connected politicians cast climate-skeptic votes is driven entirely by bills on which climate skeptics win the vote. In terms of economic magnitude, a one standard deviation increase in scaled emissions, combined with a one standard deviation increase in connected politicians' climate-skeptic vote share, is associated with an increase of 19.2 to 26.9 and 16.3 to 22.7 basis points, respectively, in the market-adjusted and characteristic-adjusted CAR around the vote on a climate bill when the outcome is a climate-skeptic win than when it is a climate-supportive win. This result suggests that carbon-emitting firms benefit from their connected politicians casting climate-skeptic votes. Thus, influencing

[11] It is not necessary to include the *Skeptic win* indicator by itself in the regression because of the inclusion of bill fixed effects.

legislative outcomes can be an important channel through which corporations derive value from political connections.

[Insert Table 8 about here]

5 Conclusion

How special interests shape legislative decisions is a fundamental question in political economy. In this paper, we examine the influence of corporations on climate change legislation, which plays an important role in addressing the negative externalities associated with greenhouse gas emissions. Using a comprehensive sample of $37,443$ votes cast on 126 contested climate bills in the US House of Representatives and Senate during the period from 2011 to 2021, we find that politicians with greater carbon dependency, i.e., those whose campaigns receive more contributions from carbon-emitting firms, are more likely to cast climate-skeptic votes. In economic terms, the results suggest that the change in the tendency to cast a climate-skeptic vote induced by a one standard deviation change in carbon dependency, when aggregated across legislators, can flip the result of about 13.5% to 19.8% of the climate votes in our sample depending on the specification.

This relation continues to hold after including high-dimensional fixed effects, such as bill fixed effects, bill $\times$ political party fixed effects, state $\times$ year fixed effects, and legislator fixed effects, thereby eliminating many potential sources of omitted variable bias that can confound inferences. We also find evidence suggesting that this relation is stronger when politicians face greater electoral pressure, which is consistent with electoral competition making politicians more responsive to campaign contributions. Exploiting narrowly defeated focal politicians as plausibly exogenous shocks to their elected peers' carbon dependency, we find that the positive relation between carbon dependency and the likelihood of casting climate-skeptic votes is likely causal.

We also show that carbon-emitting firms earn higher abnormal stock returns around the vote on a climate bill when their connected politicians disproportionately cast climate-skeptic votes on the bill. Furthermore, the higher abnormal stock return associated with carbon-emitting firms when their connected politicians vote in their favor is driven entirely by bills on which climate skeptics win the vote. These results provide suggestive evidence that firms benefit from influencing the voting behavior of politicians.

Overall, our evidence highlights the influence of corporate interests on climate policy. The results that campaign contributions by carbon-emitting firms enable these firms to tilt climate legislation in their favor suggest that entrenched special interests can use their political influence to undermine the role of government in addressing climate change.

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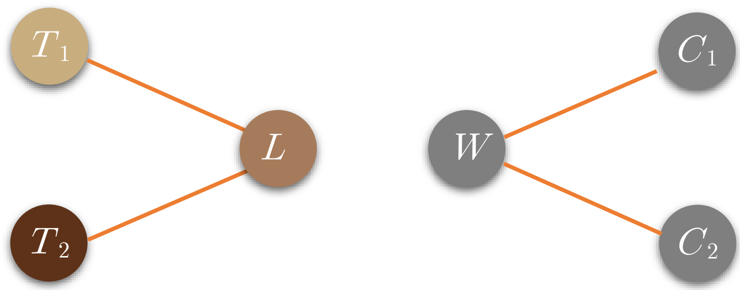

Figure 1. Illustration of the triple-difference test

This figure provides an illustration of the effects of the defeat of a focal politician on peer politicians. Representative L is a congressperson that narrowly loses her reelection bid. Representatives T_1 and T_2 are elected congresspersons in the same party and same state as Representative L, so these two politicians are treated politicians. Representative W narrowly wins her reelection bid; Representatives C_1 and C_2 are elected congresspersons in the same party and same state as Representative W and are control politicians. If Representative L's narrow defeat prompts the donors to step up support of Representatives T_1 and T_2, the carbon dependency of these two treated politicians should move towards that of Representative L after the close-race event, which in turn should affect their likelihood of casting climate-skeptic votes. Specifically, suppose Representative T_1 is funded to a lesser extent by carbon money than Representative L before the event, as indicated by a lighter brown color. Thus, Representative T_1 is likely to receive disproportionately more carbon money after the event than before, which could lead to an increase in her tendency to cast climate-skeptic votes. Representative T_2, on the other hand, is funded to a greater extent by carbon money than the focal politician before the event, as indicated by a darker brown color. As a result of the narrow defeat of the focal politician, Representative T_2 is likely to receive disproportionately more non-carbon money afterwards and become less likely to cast climate-skeptic votes. In contrast, the carbon dependency and likelihood of casting climate-skeptic votes of the control politicians, i.e., Representatives C_1 and C_2, should not exhibit such changes.

Table 1: Summary statistics on climate bills

This table reports summary statistics for the sample of contested climate bills during the period from 2011 through 2021. Contested climate bills are those that pass or fail within ±10% around the majority threshold. Panel A reports the distribution of the bills by year and chamber. Panel B reports the characteristics of climate bills, including the fraction of climate bills that are classified as climate-skeptic and the fraction of climate bills on which the vote outcome is a climate-skeptic win. We split our sample into three periods, corresponding to Democratic control of both houses of Congress (2021), split control by the two parties (2011 through 2014, 2019, and 2020), and Republican control of both houses (2015 through 2018).

Panel A: Distribution of climate bills by year and house

Year	All	House	Senate
2011	6	3	3
2012	9	5	4
2013	11	7	4
2014	10	9	1
2015	15	7	8
2016	10	8	2
2017	14	7	7
2018	13	6	7
2019	13	9	4
2020	7	4	3
2021	18	10	8
All years	126	75	51

Panel B: Characteristics of climate bills

	No. of climate bills	% climate-skeptic bills	% climate-skeptic win
Full sample	126	64.3%	61.1%
Subsamples by party control of Congress:			
Democratic control	18	16.7%	5.6%
Split control	56	69.6%	53.6%
Republican control	52	75.0%	88.5%

Table 2: Summary statistics on roll call votes and firm-bills

Panel A reports summary statistics on the sample of roll call votes. *Climate-skeptic vote* is an indicator that equals one if the legislator casts a climate-skeptic vote on a bill (i.e., voting against a climate friendly bill or voting for a climate unfriendly bill) and zero otherwise. *Carbon dependency* is a weighted average of the most recent GHG emission shares of firms that contributed to the politician's election campaigns in the past 60 months, weighted by the dollar value of the contributions. The emission share of a firm is the reported carbon equivalent emissions of the firm in a year divided by the total reported carbon equivalent emissions under GHGRP in that year. *District emission share* is the reported carbon equivalent emissions of facilities in a lawmaker's congressional district (for representatives) or state (for senators) in a year as a fraction of total reported carbon equivalent emissions under GHGRP in that year. *Tenure* is the number of years since the legislator assumed office. *Democrat* and *GOP* are indicators for Democratic and Republican lawmakers, respectively. Panel B reports summary statistics on the sample of firm-bills. *Skeptic vote share* is a weighted average climate-skeptic vote share on the bill among legislators whose campaigns received PAC contributions from the firm, weighted by the dollar value of the contributions. *Skeptic win* is an indicator that equals one for climate bills on which the vote outcome is in favor of climate obstructionists (i.e., the passage of a climate-skeptic bill or the rejection of a climate-skeptic bill) and zero otherwise. *Emission share* is the reported carbon equivalent emissions of the firm in the year before the vote on a bill as a fraction of total reported carbon equivalent emissions under GHGRP in that year. $CAR_M[-5, +5]$ is the cumulative market-adjusted abnormal return of the firm from five days before to five days after the date of the vote on a climate bill. $CAR_{DGTW}[-5, +5]$ is the cumulative DGTW characteristics-adjusted abnormal return of the firm from five days before to five days after the date of the vote on a climate bill. *Total assets* is the book value of total assets. *M/B* is market-to-book ratio. *Prior stock return* is the market-adjusted stock return in the prior 12 months. *Leverage* is book leverage ratio. *CapEx* is capital expenditures over assets. *S&P* 500 is an indicator that equals one if the firm is included in the S&P 500 index and zero otherwise. *Institutional ownership* is the fractional ownership by institutional investors in the firm. *Number of analysts* is the number of analysts that issue earnings forecasts for the firm. We measure firm characteristics using information publicly available before the month when a given bill is put to a vote for. For each variable in each sample, we report the mean, standard deviation, 25th, 50th, and 75th percentiles.

Panel A: The sample of roll call votes

	Mean	S.D.	25th	50th	75th
Full sample (N = 37,443)					
Climate-skeptic vote	0.505	0.500	0.000	1.000	1.000
Carbon dependency (%)	0.042	0.036	0.013	0.034	0.059
District emission share (%)	0.470	1.111	0.030	0.129	0.411
Tenure	10.749	9.086	4.000	8.000	16.000
Democrat	0.471	0.499	0.000	0.000	1.000
GOP	0.526	0.499	0.000	1.000	1.000
Democrats (N = 17,638)					
Climate-skeptic vote	0.030	0.170	0.000	0.000	0.000
Carbon dependency (%)	0.022	0.024	0.005	0.014	0.029
District emission share (%)	0.321	0.752	0.017	0.071	0.249
Tenure	12.227	9.481	4.000	10.000	19.000
GOP (N = 19,703)					
Climate-skeptic vote	0.932	0.251	1.000	1.000	1.000
Carbon dependency (%)	0.060	0.036	0.036	0.052	0.077
District emission share (%)	0.605	1.341	0.066	0.203	0.589
Tenure	9.434	8.519	3.000	7.000	13.000

Panel B: The sample of firm-bill pairs

	Mean	S.D.	25th	50th	75th
Main variables					
Skeptic vote share	0.605	0.253	0.455	0.597	0.792
Skeptic win	0.623	0.485	0.000	1.000	1.000
Emission share (%)	0.072	0.320	0.000	0.000	0.001
$CAR_M[-5, +5]$ (%)	-0.174	6.783	-3.120	-0.084	2.851
$CAR_{DGTW}[-5, +5]$ (%)	0.114	6.087	-2.587	0.128	2.819
Firm characteristics					
Total assets ($ billions)	51.327	192.614	2.949	9.452	32.114
Market-to-book	1.761	1.163	1.096	1.387	1.941
Prior stock return	0.030	0.353	-0.136	0.021	0.179
Leverage	0.637	0.232	0.484	0.618	0.800
CapEx	0.042	0.046	0.011	0.029	0.058
S&P 500	0.386	0.487	0.000	0.000	1.000
Institutional ownership	0.643	0.335	0.526	0.757	0.879
Number of analysts	15.903	10.980	7.000	15.000	22.000

Table 3: Carbon dependency and the likelihood of casting climate-skeptic votes

This table presents regression analysis of the relation between carbon dependency and the likelihood of casting climate-skeptic votes. The unit of observation is a congressperson's vote on a climate bill. All variables are defined in Table 2. Numbers in parentheses are t-statistics using standard errors two-way clustered by bill and by legislator. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Dependent variable =	*Climate-skeptic Vote*			
	(1)	(2)	(3)	(4)
Carbon dependency	0.622	0.548	0.516	0.316
	(4.84)***	(4.65)***	(4.45)***	(2.09)**
District emission share	0.006	0.003	0.003	0.009
	(1.58)	(0.92)	(1.00)	(1.22)
Tenure	-0.001	-0.001	-0.001	-0.001
	(1.80)*	(1.78)*	(1.73)*	(1.21)
Democrat	0.004	-0.057		
	(0.20)	(1.17)		
GOP	0.879	0.796		
	(41.44)***	(15.88)***		
Bill FEs	Yes	Yes	Absorbed	Absorbed
State × year FEs	No	Yes	Yes	Yes
Bill × party FEs	No	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	37,443	37,443	37,443	37,441
Adjusted R-squared	0.82	0.83	0.83	0.86

Table 4: Carbon dependency and the likelihood of casting climate-skeptic votes: Controlling for public opinions

This table presents regression analysis of the relation between carbon dependency and the likelihood of casting climate-skeptic votes controlling for public opinions on climate change. We measure public opinions using data from Yale Climate Opinion Maps (YCOM). *Global warming happening* is the percentage of adults who think global warming is happening at the congressional district level for representatives and at the state level for senators. Since YCOM data is only available starting from 2014 and we lag this variable, the sample used in this analysis covers the period from 2015 to 2021. All other variables are defined in Table 2. The unit of observation is a congressperson's vote on a climate bill. Numbers in parentheses are t-statistics using standard errors two-way clustered by bill and by legislator. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Dependent variable =	*Climate-skeptic Vote*			
	(1)	(2)	(3)	(4)
Carbon dependency	0.331	0.341	0.333	0.462
	(2.87)***	(2.92)***	(2.85)***	(2.36)**
Global warming happening	-0.579	-0.438	-0.431	-0.098
	(5.36)***	(4.01)***	(3.80)***	(0.53)
District emission share	0.004	0.002	0.002	0.010
	(1.48)	(0.72)	(0.63)	(0.89)
Tenure	-0.000	-0.000	-0.000	-0.003
	(0.34)	(0.59)	(0.61)	(0.87)
Democrat	-0.014	-0.081		
	(0.59)	(1.52)		
GOP	0.837	0.765		
	(31.72)***	(13.57)***		
Bill FEs	Yes	Yes	Absorbed	Absorbed
State × year FEs	No	Yes	Yes	Yes
Bill × party FEs	No	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	25,874	25,874	25,874	25,874
Adjusted R-squared	0.84	0.85	0.85	0.88

Table 5: Carbon dependency in campaign fundings and the likelihood of casting climate-skeptic votes: The role of close-call elections and swing districts

This table presents regression analysis of the relation between carbon dependency and the likelihood of casting climate-skeptic votes conditional on electoral pressure. We use two proxies to capture electoral pressure. *Non-landslide* is an indicator that equals one if the lawmaker's winning margin in the most recent election is within 20 percentage points and zero otherwise. *Swing district/state* is an indicator that equals one if the lawmaker is from a swing district/state. We define a congressional district as a swing district if the winning candidates for the House of Representatives in that district in the past five years do not come from the same party. Similarly, a state is considered as a swing state if the winning candidates for the Senate in that state in the past ten years do not come from the same party. All other variables are defined in Table 2. Numbers in parentheses are *t*-statistics using standard errors two-way clustered by bill and by legislator. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Dependent variable =	*Climate-skeptic Vote*							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Carbon dependency × *Non-landslide*	0.924	0.562	0.581	0.313				
	(3.83)***	(2.47)**	(2.63)***	(0.77)				
Carbon dependency × *Swing district/state*					0.956	0.769	0.808	0.577
					(2.38)**	(1.96)*	(2.06)**	(1.06)
Carbon dependency	0.361	0.333	0.312	0.340	0.493	0.448	0.430	0.201
	(3.05)***	(3.08)***	(2.94)***	(1.86)*	(4.06)***	(3.90)***	(3.84)***	(1.07)
Control variables and interaction terms	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Bill FEs	Yes	Yes	Absorbed	Absorbed	Yes	Yes	Absorbed	Absorbed
State × year FEs	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Bill × party FEs	No	No	Yes	Yes	No	No	Yes	Yes
Legislator FEs	No	No	No	Yes	No	No	No	Yes
Observations	37,443	37,438	37,394	37,389	37,443	37,443	37,443	37,441
Adjusted R-squared	0.82	0.83	0.84	0.86	0.82	0.83	0.83	0.86

Table 6: Triple-difference analysis of the effects of narrow defeats of incumbent focal politicians on their peers

This table presents the results of the triple-difference tests of the impact of focal politicians' close-race events on their peers. The dependent variable in Panel A is an indicator for whether the peer politician casts a climate-skeptic vote, and that in Panel B is the carbon dependency of the peer politician. *Treated* is an indicator that equals one if the focal politician loses reelection in a close race and zero if the focal politician wins. *Post* is an indicator that equals one if the bill is in the two-year period after the close-race event and zero if it is in the two-year period before the event. *Exposure* is the difference in carbon dependency between the focal politician and the peer politician in question before the close-race event. A positive value of *Exposure* indicates that the focal politician is more carbon dependent than the peer politician. Numbers in parentheses are t-statistics using clustered standard errors. The standard errors are three-way clustered by close-race event, by bill, and by legislator in Panel A, and two-way clustered by close-race event and by legislator in Panel B. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Panel A: The effect of focal politicians losing a close race on peer politicians' votes

Dependent variable =	*Climate-skeptic Vote*			
	(1)	(2)	(3)	(4)
Treated × *Post* × *Exposure*	0.363	0.276	0.276	0.314
	(4.51)***	(3.03)***	(3.08)***	(1.83)*
Treated × *Post*	0.022			
	(1.98)*			
Treated × *Exposure*	-0.877	-0.833	-0.833	1.078
	(1.45)	(1.43)	(1.43)	(1.07)
Post × *Exposure*	-0.193	-0.162	-0.162	-0.164
	(4.23)***	(3.17)***	(2.68)**	(5.07)***
Exposure	-0.225	-0.243	-0.243	-0.330
	(1.23)	(1.29)	(1.28)	(0.57)
Close-race event FEs	Yes	Yes	Yes	Yes
Bill FEs	Yes	Yes	Absorbed	Absorbed
State × year FEs	No	Yes	Yes	Yes
Bill × party FEs	No	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	8,523	8,523	8,523	8,523
Adjusted R-squared	0.86	0.86	0.86	0.88

Panel B: The effect of focal politicians losing a close race on peer politicians' carbon dependency

Dependent variable =	*Carbon dependency*			
	(1)	(2)	(3)	(4)
Treated × Post × Exposure	0.407	0.390	0.390	0.390
	(3.30)***	(3.11)***	(3.10)***	(3.10)***
Treated × Post	0.007			
	(1.62)			
Treated × Exposure	-0.089	-0.081	-0.081	0.093
	(1.94)*	(2.04)**	(2.03)**	(1.04)
Post × Exposure	0.140	0.158	0.158	0.158
	(2.17)**	(2.52)**	(2.52)**	(2.52)**
Exposure	-0.908	-0.917	-0.917	-0.689
	(21.71)***	(22.98)***	(22.92)***	(9.71)***
Close-race event FEs	Yes	Yes	Yes	Yes
State × year FEs	No	Yes	Yes	Yes
Party × year FEs	Yes	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	786	778	778	778
Adjusted R-squared	0.82	0.82	0.82	0.85

Table 7: Triple-difference analysis of the effect of campaign contribution composition changes induced by close races: Falsification tests

This table presents the results of falsification tests assuming that the close-race events occur 24 months before the actual date. The dependent variable in Panel A is an indicator for whether the peer politician casts a climate-skeptic vote, and that in Panel B is the carbon dependency of the peer politician. *Treated* is an indicator that equals one if the focal politician loses reelection in a close race and zero if the focal politician wins. "*Post*" is an indicator that equals one if the bill is in the two-year period after the pseudo-event and zero if it is in the two-year period before the event. *Exposure* is the difference in carbon dependency between the focal politician and the peer politician in question before the close-race event. A positive value of *Exposure* indicates that the focal politician is more carbon dependent than the peer politician. Numbers in parentheses are t-statistics using clustered standard errors. The standard errors are three-way clustered by close-race event, by bill, and by legislator in Panel A, and two-way clustered by close-race event and by legislator in Panel B. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Panel A: The effect of focal politicians losing a pseudo close race on peer politicians' votes

Dependent variable =	*Climate-skeptic Vote*			
	(1)	(2)	(3)	(4)
Treated × "*Post*" × *Exposure*	-0.641	-0.561	-0.561	-0.378
	(1.62)	(1.36)	(1.36)	(1.15)
Treated × "*Post*"	-0.023			
	(1.16)			
Treated × *Exposure*	-0.057	-0.097	-0.097	1.600
	(0.13)	(0.24)	(0.24)	(1.61)
"*Post*" × *Exposure*	0.042	-0.038	-0.038	0.047
	(0.34)	(0.38)	(0.37)	(1.43)
Exposure	-0.284	-0.244	-0.244	-0.399
	(2.12)**	(2.20)**	(2.20)**	(0.69)
Close-race event FEs	Yes	Yes	Yes	Yes
Bill FEs	Yes	Yes	Absorbed	Absorbed
State × year FEs	No	Yes	Yes	Yes
Bill × party FEs	No	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	6,078	6,078	6,078	6,078
Adjusted R-squared	0.89	0.89	0.89	0.91

Panel B: The effect of focal politicians losing a pseudo close race on peer politicians' carbon dependency

Dependent variable =	*Carbon dependency*			
	(1)	(2)	(3)	(4)
Treated × "*Post*" × *Exposure*	-0.069	-0.154	-0.154	-0.154
	(0.54)	(1.25)	(1.25)	(1.25)
Treated × "*Post*"	-0.006			
	(1.00)			
Treated × *Exposure*	-0.036	0.006	0.006	-0.013
	(0.25)	(0.04)	(0.04)	(0.10)
"*Post*" × *Exposure*	-0.160	-0.143	-0.143	-0.143
	(3.19)***	(3.16)***	(3.15)***	(3.15)***
Exposure	-0.810	-0.819	-0.819	-1.030
	(15.22)***	(15.00)***	(14.96)***	(9.91)***
Close-race event FEs	Yes	Yes	Yes	Yes
State × year FEs	No	Yes	Yes	Yes
Party × year FEs	Yes	No	Yes	Yes
Legislator FEs	No	No	No	Yes
Observations	646	644	644	644
Adjusted R-squared	0.87	0.87	0.87	0.88

Table 8: Abnormal stock returns around votes on climate bills

This table presents the regression analysis of abnormal stock returns around votes on climate bills conditional on firms' carbon emissions, connected politicians' votes, and vote outcome. The unit of observation is a firm-bill pair. The dependent variable in Panel A is market-adjusted CARs from five days before to five days after the date of the vote on a climate bill, and that in Panel B is DGTW characteristics-adjusted CARs. All variables are defined in Table 2. Numbers in parentheses are t-statistics using standard errors two-way clustered by bill and by industry. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Panel A: Market-adjusted CAR

Dependent variable =	$CAR_M[-5, +5]$					
	(1)	(2)	(3)	(4)	(5)	(6)
Emission × Skeptic vote share × Skeptic win				2.375	2.545	3.379
				(3.27)***	(3.25)***	(4.68)***
Emission × Skeptic vote share	1.672	1.949	1.690	0.243	0.475	-0.238
	(2.82)***	(3.85)***	(2.69)***	(0.31)	(0.67)	(0.28)
Emission × Skeptic win				-1.217	-1.388	-1.868
				(2.42)**	(2.47)**	(3.81)***
Skeptic vote × Skeptic win				-0.008	-0.008	-0.006
				(1.67)	(1.60)	(1.27)
Emission	-1.143	-1.360	-1.595	-0.448	-0.601	-0.613
	(2.91)***	(3.71)***	(3.36)***	(0.88)	(1.18)	(1.09)
Skeptic vote share	-0.003	0.001	0.000	0.002	0.005	0.004
	(1.03)	(0.34)	(0.14)	(0.43)	(1.68)*	(1.09)
Firm size		-0.000	-0.012		-0.000	-0.012
		(0.22)	(4.18)***		(0.16)	(4.20)***
Market-to-book		0.001	-0.003		0.001	-0.003
		(0.68)	(2.74)***		(0.71)	(2.74)***
Prior stock returns		0.004	-0.003		0.004	-0.003
		(0.87)	(0.60)		(0.86)	(0.61)
Leverage		0.000	0.012		0.000	0.012
		(0.17)	(2.19)**		(0.19)	(2.20)**
CapEx		-0.025	-0.043		-0.025	-0.043
		(1.49)	(1.66)		(1.49)	(1.64)
S&P 500		0.002	-0.000		0.002	-0.000
		(1.04)	(0.07)		(0.98)	(0.15)
Institutional ownership		-0.002	-0.009		-0.002	-0.009
		(1.02)	(2.35)**		(0.99)	(2.33)**
Log(1+*Number of analysts*)		0.000	-0.000		0.000	-0.000
		(0.27)	(0.10)		(0.28)	(0.13)
Bill FEs	Yes	Yes	Yes	Yes	Yes	Yes
Industry FEs	No	Yes	No	No	Yes	No
State × year FEs	No	No	Yes	No	No	Yes
Firm FEs	No	No	Yes	No	No	Yes
Observations	75,249	75,249	75,241	75,249	75,249	75,241
Adjusted R-squared	0.06	0.06	0.10	0.06	0.06	0.10

Panel B: DGTW Characteristics-adjusted CAR

Dependent variable =	$CAR_{DGTW}[-5, +5]$					
	(1)	(2)	(3)	(4)	(5)	(6)
Emission × Skeptic vote share × Skeptic win				2.012	2.241	2.874
				(4.04)***	(4.56)***	(4.72)***
Emission × Skeptic vote share	1.437	1.797	1.444	0.250	0.533	-0.081
	(3.11)***	(3.58)***	(2.44)**	(0.47)	(0.89)	(0.13)
Emission × Skeptic win				-1.072	-1.279	-1.662
				(3.42)***	(4.08)***	(4.11)***
Skeptic vote × Skeptic win				-0.009	-0.008	-0.007
				(2.07)**	(1.93)*	(1.57)
Emission share	-1.075	-1.322	-1.454	-0.476	-0.642	-0.642
	(3.68)***	(4.17)***	(2.87)***	(1.48)	(1.80)*	(1.28)
Skeptic vote share	-0.002	0.000	0.001	0.004	0.005	0.005
	(0.75)	(0.33)	(0.62)	(1.23)	(1.81)*	(1.54)
Firm size		-0.001	-0.010		-0.001	-0.010
		(3.04)***	(4.61)***		(2.97)***	(4.64)***
Market-to-book		-0.000	-0.004		-0.000	-0.004
		(0.60)	(3.74)***		(0.57)	(3.73)***
Prior stock returns		0.002	-0.004		0.002	-0.004
		(0.63)	(1.05)		(0.63)	(1.06)
Leverage		0.003	0.019		0.003	0.019
		(1.28)	(4.08)***		(1.29)	(4.03)***
CapEx		-0.020	-0.038		-0.020	-0.038
		(1.24)	(1.56)		(1.24)	(1.55)
S&P 500		0.002	0.003		0.002	0.003
		(2.07)**	(1.25)		(1.97)*	(1.15)
Institutional ownership		-0.002	-0.008		-0.002	-0.008
		(1.04)	(2.31)**		(1.02)	(2.30)**
Log(1+*Number of analysts*)		-0.000	-0.000		-0.000	-0.001
		(0.36)	(0.51)		(0.33)	(0.55)
Bill FEs	No	Yes	Yes	No	Yes	Yes
Industry FEs	No	Yes	No	No	Yes	No
State × year FEs	No	No	Yes	No	No	Yes
Firm FEs	No	No	Yes	No	No	Yes
Observations	71,636	71,636	71,628	71,636	71,636	71,628
Adjusted R-squared	0.01	0.01	0.05	0.00	0.01	0.05

Appendix Table A1: Abnormal stock returns around votes on climate bills

This table presents the regression analysis of abnormal stock returns around votes on climate bills conditional on firms' carbon emissions and vote outcome. The unit of observation is a firm-bill pair. The dependent variable in the first three columns is market-adjusted CARs from five days before to five days after the date of the vote on a climate bill, and that in the last three columns is DGTW characteristics-adjusted CARs. All variables are defined in Table 2. Numbers in parentheses are *t*-statistics using standard errors two-way clustered by bill and by industry. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Dependent variable =	$CAR_M[-5, +5]$			$CAR_{DGTW}[-5, +5]$		
	(1)	(2)	(3)	(4)	(5)	(6)
Emission share × *Skeptic win*	0.371	0.325	0.394	0.271	0.234	0.269
	(3.47)***	(3.06)***	(2.95)***	(2.61)**	(2.25)**	(2.28)**
Emission share	-0.294	-0.280	-0.833	-0.311	-0.285	-0.742
	(5.65)***	(2.00)*	(2.63)**	(4.02)***	(2.06)**	(2.71)***
Firm size		-0.000	-0.012		-0.001	-0.010
		(0.25)	(4.19)***		(3.06)***	(4.62)***
Market-to-book		0.001	-0.003		-0.000	-0.004
		(0.66)	(2.74)***		(0.61)	(3.75)***
Prior stock returns		0.004	-0.003		0.002	-0.004
		(0.88)	(0.60)		(0.65)	(1.05)
Leverage		0.000	0.012		0.003	0.019
		(0.11)	(2.20)**		(1.21)	(4.15)***
CapEx		-0.026	-0.043		-0.020	-0.037
		(1.46)	(1.66)		(1.31)	(1.57)
S&P 500		0.002	-0.000		0.002	0.003
		(1.00)	(0.10)		(2.04)**	(1.23)
Institutional ownership		-0.003	-0.009		-0.003	-0.008
		(1.05)	(2.34)**		(1.07)	(2.30)**
Log(1+*Number of analysts*)		0.000	-0.000		-0.000	-0.001
		(0.28)	(0.13)		(0.34)	(0.54)
Bill FEs	Yes	Yes	Yes	Yes	Yes	Yes
Industry FEs	No	Yes	Yes	No	Yes	Yes
State × year FEs	No	No	Yes	No	No	Yes
Firm FEs	No	No	Yes	No	No	Yes
Observations	75,249	75,249	75,241	71,636	71,636	71,628
Adjusted R-squared	0.06	0.06	0.10	0.01	0.01	0.05

Appendix Table A2: Triple-difference analysis of the effects of narrow defeats of incumbent focal politicians on their peers: Alternative cutoffs for close-race events

This table presents the results of the triple-difference tests of the impact of focal politicians' close-race events on their peers using alternative cutoffs for close-race events, i.e., the winning margin being within ±3% and ±2%. The dependent variable in Panel A is an indicator for whether the peer politician casts a climate-skeptic vote, and that in Panel B is the carbon dependency of the peer politician. The first two columns in each panel report the results using the ±3% sample, and the last two columns report the results using the ±2% sample. We use the same specifications as those in the last two columns of Table 6, which include the most stringent set of fixed effects. *Treated* is an indicator that equals one if the focal politician loses reelection in a close race and zero if the focal politician wins. *Post* is an indicator that equals one if the bill is in the two-year period after the close-race event and zero if it is in the two-year period before the event. *Exposure* is the difference in carbon dependency between the focal politician and the peer politician in question before the close-race event. A positive value of *Exposure* indicates that the focal politician is more carbon dependent than the peer politician. Numbers in parentheses are *t*-statistics using clustered standard errors. The standard errors are three-way clustered by close-race event, by bill, and by legislator in Panel A, and two-way clustered by close-race event and by legislator in Panel B. Significance at the 10% (*), 5% (**), or 1% (***) level is indicated.

Panel A: The effect of focal politicians losing a close race on peer politicians' votes

Dependent variable =	*Climate-skeptic Vote*			
	[−3%, +3%]		[−2%, +2%]	
	(1)	(2)	(3)	(4)
Treated × *Post* × *Exposure*	0.181	0.197	0.519	0.602
	(2.14)**	(1.78)*	(4.73)***	(2.29)**
Treated × *Post*	0.000	0.000	0.000	0.000
	(0.00)	(0.00)	(0.00)	(0.00)
Treated × *Exposure*	-0.919	1.187	-0.865	1.076
	(1.54)	(1.16)	(1.10)	(1.18)
Post × *Exposure*	-0.087	-0.087	-0.153	-0.158
	(1.10)	(1.42)	(2.69)**	(3.83)***
Exposure	-0.323	-0.364	-0.243	-0.392
	(1.60)	(0.76)	(1.28)	(0.81)
Close-race event FEs	Yes	Yes	Yes	Yes
Bill FEs	Absorbed	Absorbed	Absorbed	Absorbed
State × year FEs	Yes	Yes	Yes	Yes
Bill × party FEs	Yes	Yes	Yes	Yes
Legislator FEs	No	Yes	No	Yes
Observations	10,685	10,685	6,897	6,897
Adjusted R-squared	0.85	0.88	0.87	0.89

Panel B: The effect of focal politicians losing a close race on peer politicians' carbon dependency

Dependent variable =	*Carbon dependency*			
	[−3%, +3%]		[−2%, +2%]	
	(1)	(2)	(3)	(4)
Treated × *Post* × *Exposure*	0.456	0.456	0.541	0.541
	(4.87)***	(4.88)***	(3.62)***	(3.62)***
Treated × *Exposure*	-0.073	-0.005	-0.083	0.095
	(2.17)**	(0.05)	(2.04)*	(1.02)
Post × *Exposure*	0.154	0.154	0.150	0.150
	(2.49)**	(2.59)**	(2.40)**	(2.40)**
Exposure	-0.925	-0.695	-0.917	-0.657
	(27.31)***	(8.48)***	(22.50)***	(9.37)***
Close-race event FEs	Yes	Yes	Yes	Yes
State × year FEs	Yes	Yes	Yes	Yes
Party × year FEs	Yes	Yes	Yes	Yes
Legislator FEs	No	Yes	No	Yes
Observations	970	970	658	658
Adjusted R-squared	0.82	0.86	0.81	0.84