Incremental Democracy: The Policy Effects of Partisan Control of State Government

Devin Caughey, Massachusetts Institute of Technology
Christopher Warshaw, George Washington University
Yiqing Xu, University of California, San Diego

How much does it matter whether Democrats or Republicans control the government? Unless the two parties converge completely, election outcomes should have some impact on policy, but the existing evidence for policy effects of party control is surprisingly weak and inconsistent. We bring clarity to this question, using regression-discontinuity and dynamic panel analyses to estimate the effects of party control of state legislatures and governorships on a new annual measure of state policy liberalism. We find that throughout the 1936–2014 period, electing Democrats has led to more liberal policies, but that in recent decades the policy effects of party control have approximately doubled in magnitude. We present evidence that this increase is at least partially explained by the ideological divergence of the parties' office holders and electoral coalitions. At the same time, we also show that party effects remain substantively modest, paling relative to policy differences across states.

In November 1948, the Ohio Democratic Party gained control of state government for the first time in 10 years. With the popular Frank Lausche at the top of their ticket, the Democrats defeated the incumbent Republican governor and won majorities in both houses of the legislature. During their two years of unified control, however, Ohio Democrats failed to pass any major new liberal policies. In fact, Governor Lausche, a fiscal conservative who had defeated a more liberal candidate in the Democratic primary, proposed a budget cutting state expenditures, and the liberal initiatives he did support, such as a ban on racial discrimination in employment, failed to make it through the Democratic legislature (Chen 2009, 165, 273; *Time* 1956; Usher 1994). Six decades later, in 2012, North Carolina Republicans experienced a similar triumph with the election of Governor Pat McCrory, who completed the GOP takeover of the state initiated two years earlier with its capture of the legislature. Unlike Ohio Democrats in 1948, North Carolina Republicans took advantage of their newfound control by passing a flood of conservative legislation: cutting unemployment insurance, repealing the estate tax, "flattening" the income tax, relaxing gun laws, and tightening restrictions on abortion (Fausset 2014).

These two cases, Ohio in 1948 and North Carolina in 2012, suggest very different conclusions about the policy effects of party control of state government. Does electing Democrats rather than Republicans have only an incremental, or perhaps nonexistent, impact on state policies, or does it result in dramatic policy shifts that leapfrog over the median voter? The scholarly literature exhibits surprisingly little consensus on this question. Many classic studies of state politics emphasize the exceedingly weak or even negative cross-sectional association between state policy liberalism and Democratic control of state offices, suggesting that electoral pressure to converge on the median voter may be so strong as to all but eliminate differences between Democrats and Republicans (e.g., Erikson, Wright, and McIver 1993; Hofferbert 1966). More recent studies employing panel or regression-discontinuity (RD) designs have uncovered partisan policy effects but typically

Devin Caughey (caughey@mit.edu) is an associate professor at the Massachusetts Institute of Technology in Cambridge, Massachusetts. Christopher Warshaw (warshaw@gwu.edu) is an assistant professor at George Washington University in Washington, DC. Yiqing Xu (yiqingxu@ucsd.edu) is an assistant professor at the University of California, San Diego in San Diego, California.

We are grateful for support from the School of Humanities, Arts, and Social Sciences at MIT. Data and supporting materials necessary to reproduce the numerical results in the paper are available in the *JOP* Dataverse (https://dataverse.harvard.edu/dataverse/jop). An online appendix with supplementary material is available at http://dx.doi.org/10.1086/692669.

The Journal of Politics, volume 79, number 4. Published online August 4, 2017. http://dx.doi.org/10.1086/692669

only for certain offices, on some policies, in a subset of states, or under particular conditions (e.g., Alt and Lowry 1994; Besley and Case 2003; Kousser 2002; Leigh 2008).

Combining multiple research designs, a long historical perspective, and a wealth of new data, we offer clearer answers to the question of partisan effects on policy. We improve upon existing research in three major ways. First, we use a much more comprehensive policy measure, the policy liberalism scale developed by Caughey and Warshaw (2016), which is estimated from a data set of nearly 150 distinct policies covering each year between 1936 and 2014. Second, we use more credible identification strategies. Specifically, we estimate the effects of Democratic governors and state legislatures using two designs: an electoral RD design, which exploits variation in party control induced by very close elections, and dynamic panel analysis, which exploits year-specific partisan variation within states. These designs enable us to isolate the causal effects of partisan control from other time-varying determinants of state policy, such as changes in public opinion. Third, we examine whether party effects have grown over time, and whether this growth is related to partisan polarization at the mass and elite levels.

We find that partisan effects on state policy have indeed increased substantially over the past eight decades, with the growth concentrated in the last quarter century. Between the 1930s and 1980s, the partisan composition of state governments had little causal impact on the ideological orientation of state policies. Since the 1980s, however, partisan effects have grown dramatically. We find little indication that this growth differed by region or was driven by the anomalously Democratic partisanship of the formerly "solid" South. We do, however, find robust support for the hypothesis that partisan polarization has increased partisan effects on policy. Specifically, we find greater policy effects when and where Democratic and Republican identifiers diverge more in their policy views and where roll-call voting in the state legislature is more polarized by party.

Notwithstanding their dramatic growth, the substantive magnitude of partisan effects on policy should not be exaggerated. Even today, for example, electing a Democratic rather than Republican governor only has an incremental effect on policy.[1] It should be expected to increase monthly welfare payments by only $1–$2 per recipient and to increase by just half a percentage point the proportion of policies on which a state has the liberal policy option. These effects are small relative to policy differences across states and also relative to partisan differences in legislative voting records. Our findings thus partially assuage the normative concern that partisan polarization has resulted in a "leapfrog democracy" of wide policy swings and poor congruence with citizens' preferences (Bafumi and Herron 2010; see also Lax and Phillips 2011; Poole and Rosenthal 1984).

The remainder of this article is organized as follows. We first discuss the substantive and theoretical background for our inquiry. We then turn to empirics, beginning with a description of our annual measure of state policy liberalism. Next, we estimate the policy effects of Democratic governors and state legislatures using RD and dynamic panel analyses. We also document the strong relationship between the growth in party effects on policy and partisan polarization. The final section discusses the implications of our results.

SUBSTANTIVE BACKGROUND

Although the relationship between state policies and the partisanship of state officials is a long-standing focus of the state politics literature, there is no consensus regarding the causal effects of partisan control on state policy. Most classic studies find little association between states' policies and the partisanship of their officials.[2] After controlling for public opinion, some studies even find Democratic party control and liberal policies to be negatively correlated across states (e.g., Erikson et al. 1993; Lax and Phillips 2011).

These cross-sectional studies, however, are hampered by two important methodological limitations. First, they lack a credible identification strategy. As a result, their findings about the effect of party control on policy could be biased by any number of omitted variables that are correlated with partisan control of government (economic conditions, mass or elite policy preferences, etc.). Second, their findings are all based on a single slice of time and sometimes a single policy area. As a result, it is hard to know whether each study's results are generalizable to other time periods or policy areas.

A smaller literature has used panel data to examine policy effects using more credible causal identification strategies. Most studies, including those with strong designs, find that in general partisan control of the governorship does not substantially affect policy. Besley and Case (2003), for example,

1. We are using the term "incremental" in a more general way than the incrementalism literature in public administration (e.g., Lindblom 1959). Most importantly, our explanation for why policy change is incremental is not based on the cognitive or informational limitations of decision-makers.

2. Hofferbert (1966), for example, finds "no significant relationship" between "the party in power and public policy" on welfare issues. Winters (1976) finds that party control of state government makes "little or no difference" for tax burdens and spending. Hanson (1984) finds no significant effects of party control on Medicaid programs, while Plotnick and Winters (1985) find no effect of party control on AFDC benefits.

estimate a two-way fixed-effects model of four state policy indicators and find a mix of liberal, conservative, and indeterminate effects of Democratic governors. Studies that employ electoral RD designs to examine the policy effects of governors find similarly ambiguous and contingent effects. For instance, Fredriksson, Wang, and Warren (2013) find that reelectable Democratic governors increase taxes, but term-limited ones decrease them. Similarly, Leigh (2008) examines a total of eight policy indicators and finds significant effects on just one, leading him to conclude that governors "behave in a fairly nonideological manner" (256). Evidence that party control of the state legislature influences policy outcomes is more consistent but hardly universal. Panel studies have found that control of the legislature influences some policies, such as civil rights, tax burdens, and welfare benefits (e.g., Besley and Case 2003; Chen 2007; Reed 2006), but has no effect on others (e.g., Konisky 2007). Each of these studies, however, focuses on only a handful of policies. It is thus hard to know what to make of their mixed and ambiguous results. Moreover, it is difficult to assess whether their results generalize to the larger policy agenda.

In sum, the state politics literature exhibits little agreement regarding the policy effects of partisan control of state government (see appendix section A.1 for a more comprehensive summary of the previous literature that further demonstrates this point; appendix available online). On the whole, these studies have found "weak and oftentimes conditional" evidence that party control affects state policies (Kousser and Phillips 2009, 70). In the sections that follow, we bring clarity to this debate with both new theory and new evidence on the policy effects of the partisan composition of state government.

THEORETICAL FRAMEWORK

Like many other works on state politics, our basic theoretical framework is a model of two-party competition over a one-dimensional policy space.[3] In a perfectly Downsian world, in which electorally motivated parties adopt the positions of the median voter, party control of state offices has no effect on state policies. Only if the parties diverge from the median voter do partisan policy effects—counterfactual differences in policy liberalism under Democratic versus Republican control—actually emerge.

Given that candidates cannot perfectly predict election outcomes and often care about influencing policy in addition to winning office, we should in general expect some degree of ideological divergence between the two parties (Grofman 2004; Roemer 2001, 72). In fact, as Gerring (1998) shows, national party conflict has had a strong ideological component throughout US history, with the parties' current ideological orientations dating back to 1928 for the Republicans and 1952 for the Democrats. Within states, Democratic senators (Poole and Rosenthal 1984), candidates and activists (Erikson, Wright, and McIver 1989), and state legislators (Shor and McCarty 2011) take more liberal policy positions than their Republican counterparts. Within-state partisan divergence on economic issues extends back to the New Deal realignment, if not before, but even on racial issues, where the national parties took longer to sort out ideologically, Democrats have been more liberal than same-state Republicans since the 1940s (Feinstein and Schickler 2008).

Given this evidence for partisan divergence, the more interesting question is not whether partisan effects exist but how large they are. If centripetal pressures dominate, then the parties in each state will converge closely on the state's median voter and differ only modestly in their policy platforms. Policy effects will be further attenuated by the limitations imposed by the minority party and other constraints on the majority party's capacity to implement their preferred policies (e.g., Alesina, Londregan, and Rosenthal 1993). Governors, for example, cannot simply implement their ideal points but rather must compromise with a legislature in which the opposing party probably has at least some influence. Such limitations on Democrats' and Republicans' desire and capacity to implement divergent policies lead us to the expectation that policy effects should generally be small relative to, say, the policy variation across states.

Nevertheless, there are also good reasons to expect partisan effects on state policy to have increased over the period we examine. At the national level, Democratic and Republican officials have become increasingly ideologically polarized, especially since the 1970s (McCarty, Poole, and Rosenthal 2006). Policy conflict between the national parties has become increasingly aligned with what is now defined as "liberalism" and "conservatism" (Noel 2014). Whether due to true polarization (Abramowitz 2010) or partisan sorting (Fiorina and Abrams 2008), the mass public has followed suit, increasing the ideological distance between the parties' electoral coalitions (Hill and Tausanovitch 2016). As formal theorists have long noted, ideological divergence between parties' primary electorates increases the electoral incentives for party nominees to diverge from the median voter (e.g., Adams and Merrill 2008). Moreover, if candidates are drawn from the set of party identifiers, their own sincere policy views should become more extreme as well (e.g., Cadigan and Janeba 2002; Thom-

3. Caughey and Warshaw (2016) show that throughout this period, cross-state policy variation was primarily structured by a single latent dimension, and modeling state policies as a function of two or more latent dimensions does little to improve the model's predictive accuracy.

sen 2014). Mass polarization between the parties has thus reinforced and exacerbated elite polarization (Jacobson 2012), resulting in larger policy effects of the partisan composition of government.[4] Indeed, some scholars have warned that polarization has become so extreme that representatives now "leapfrog" over the median voter, leading to wide swings between liberal and conservative policy outcomes incongruent with the preferences of the median voter (Bafumi and Herron 2010; Lax and Phillips 2011; Poole and Rosenthal 1984).

In sum, these theoretical results and empirical trends give rise to several expectations. On one hand, the centripetal pull of electoral competition and the limitations on officials' capacity to fully implement their policy preferences lead to the expectation that policy effects will be modest, at least relative to policy differences between states. On the other hand, given the growth of partisan polarization, partisan effects on policy are likely to be larger now than in the past. To the extent that this growth has been driven by the diverging policy preferences of Democratic and Republican officials (as opposed to, say, increases in the majority party's control over state policy), we should also expect policy effects to be larger where Democratic and Republican politicians are more ideologically polarized. Finally, if elite polarization is rooted in ideological divergence between the parties' electoral coalitions, we should expect the magnitude of policy effects to be correlated with the extent of mass polarization. We assess these hypotheses below, but first we describe our strategy for measuring the dependent variable in our analysis: state policy liberalism.

AN ANNUAL MEASURE OF STATE POLICY LIBERALISM

Studies of state policy generally employ one of two measurement strategies: they either analyze one or more policy-specific indicators, or they construct composite measures intended to summarize the general orientation of state policies (Jacoby and Schneider 2014, 568). Each approach has advantages and disadvantages. An important benefit of policy-specific indicators is that they yield easily interpretable measures and causal estimates. When the concept of interest is the overall orientation of state policies, however, individual policies are often inadequate. A state's minimum wage, for example, is at best a partial indicator of the liberalism of its economic policies, let alone its policies in other domains.[5] Another downside of focusing solely on continuous policies such as taxes and expenditures is that it ignores categorical policies like the abortion restrictions enacted by North Carolina Republicans after the 2012 election. Finally, relying on a few noisy policy indicators leads to a substantial loss of statistical power. The combination of multiple outcome variables and low statistical power can easily lead to inferential errors about effect magnitudes because only a few unusually large point estimates will pop out as significant (Gelman, Hill, and Yajima 2012). It is thus unsurprising that studies focusing on individual policies have typically found significant (sometimes large) partisan effects on a few policies but null results for many others. For the same reasons, studies of city policies have often found similar patterns of results (e.g., Ferreira and Gyourko 2009; Gerber and Hopkins 2011).

To address these problems, many studies of state policy rely on indices, factor scores, or other holistic summaries of the liberalism of state policies (e.g., Erikson et al. 1993; Hofferbert 1966). Such composite measures substantially reduce measurement error and thus increase statistical power if, as seems reasonable with state policies, the indicators on which they are based tap into a single latent variable (Ansolabehere, Rodden, and Snyder 2008). In addition, composite measures of policy liberalism often come closer to capturing the outcome of interest, which is usually not a specific policy domain but rather the overall ideological orientation of state policies. A major disadvantage of the composite approach, however, has been the difficulty of constructing time-varying measures of state policy liberalism. Because of this, all existing analyses of the determinants of state policy liberalism employ cross-sectional designs inimical to credible causal inferences.

In our analysis, we utilize the dynamic measure of state policy liberalism recently developed by Caughey and Warshaw (2016), who use a data set of nearly 150 policies to estimate a policy liberalism score for each state in each year between 1936 and 2014. The policy liberalism scores are estimated using a dynamic Bayesian factor-analytic model for mixed

4. Other factors too have probably contributed to an increase in partisan effects on policy. For example, policy effects in state legislatures should depend on the degree to which the majority party can use its control to skew policy outcomes away from the median legislator in the chamber (e.g., Cox, Kousser, and McCubbins 2010). Over the past half century, there is a variety of evidence that the two parties in Congress have leveraged their greater homogeneity into strong formal mechanisms of party discipline and control, enhancing the majority's influence over policy making (Aldrich and Rohde 2000). Given state legislatures have polarized too (Shor and McCarty 2011), it is plausible that party power has increased there as well (but see Mooney [2013] who finds no evidence that the formal powers of state speakers have increased since 1981). Another contributing factor is the decline in the nonpolicy benefits of holding office as patronage-oriented machines have been replaced by an activist base of issue-oriented "amateurs" (Wilson 1962). Since candidates should adopt more moderate (and thus electorally appealing) policy positions to the extent that they value holding office in itself (Calvert 1985), the decline of patronage politics has probably contributed to ideological divergence as well.

5. Adcock and Collier (2001) call this a failure of content validation.

data, which allows the inclusion of both continuous and ordinal indicators of state policy (over 80% of the variables in the policy data set are ordinal, mainly dichotomous).[6] The policy data set that Caughey and Warshaw used to estimate these scores was designed to include all politically salient state policy outputs on which comparable data are available for at least five years.[7] The data cover a wide range of policy areas, including social welfare (e.g., Aid to Families with Dependent Children [AFDC]/Temporary Assistance for Needy Families [TANF] benefit levels), taxation, labor (e.g., right to work), civil rights (e.g., fair housing laws), women's rights (e.g., jury service for women), morals legislation (e.g., antisodomy laws), family planning (e.g., ban on partial birth abortion), the environment (e.g., state endangered species acts), religion (e.g., public schools allowed to post Ten Commandments), criminal justice (e.g., death penalty), and drugs (e.g., marijuana decriminalization). Despite the diversity of policies, Caughey and Warshaw (2016) find little evidence that policy variation across states is multidimensional, and they report that the global measure correlates highly with domain-specific indexes of policy liberalism. Data on at least 43 different policies are available in every year, enough to estimate policy liberalism quite precisely.[8]

Table 1 provides a sense of how policy liberalism corresponds to substantive differences across states in 1950 and 2010. Mississippi and Massachusetts, which bookend the policy liberalism scale throughout the period, are included for both years; the other three states in each year were chosen because their policy liberalism differ from each other by about one standard deviation.[9] The second column indicates the percentage of dichotomous policies on which the state had the liberal option.[10] In a typical year, a one-unit change in policy liberalism corresponds to a 14-point increase in a state's percentage of liberal policies. The next four columns provide examples of highly discriminating dichotomous policies of varying "difficulty," and the rightmost column provides an example of a continuous policy, average monthly AFDC/TANF benefits per recipient family.[11]

Figure 1 plots the policy liberalism time series of every state between 1936 and 2014, with light and dark loess lines for states with Democratic and Republican governors, respectively. Strikingly, until the end of the twentieth century states with Democratic governors actually had more conservative policies than Republican-controlled states (the patterns for state legislatures are similar). The figure thus confirms the classic finding of a weakly negative relationship between state policy liberalism and Democratic control. Since 2000, however, party control has become aligned with state politics, and the gap in policy liberalism between Democratic- and Republican-controlled states has rapidly widened. The realignment of the South is only partly responsible for this shift, for even the non–South Republican states were at least as liberal as Democratic ones until the late 1990s. Whether the increasing correlation between party control of government and policy is causal—and not simply the result of a better match between ideology and partisanship—is the subject of the empirical analyses in the next section.

EMPIRICAL ANALYSIS OF POLICY EFFECTS

Evaluating policy divergence between the parties requires isolating the policy effects of partisan composition from other determinants of state policy; otherwise, partisan effect estimates will be biased. The public's ideological mood, for example, may affect policy not only through partisan turnover but also through the anticipatory responsiveness of incumbents (Stimson, MacKuen, and Erikson 1995), introducing spurious correlation into naive estimates of partisan effects. In order to isolate the policy effects of partisan composition per se, we rely on two identification strategies. The first is an RD design, which exploits the exogenous variation in party control induced by narrowly decided elections. Intuitively, extremely close elections may be thought of as coin flips that randomly install one party's candidate into office, independent of all other policy determinants. Our second identification strategy is a dynamic panel analysis, which exploits overtime variation within states while controlling for national trends and states' recent history of policy liberalism. We use the RD design to establish our basic findings and then follow up with dynamic panel analysis, whose greater statistical efficiency allows us to examine these findings with greater nuance and precision.

6. The model is dynamic in that policy liberalism is estimated separately in each year, and the policy-specific intercepts (or "difficulties") are allowed to drift over time. This has the effect of dampening shifts that are common to all states. If, instead, the intercepts are held constant, the policies of all states are estimated to have become substantially more liberal, especially before the 1980s. The precise structure of the item parameters in the policy model do not significantly affect our results, however, since our estimation strategies net out shifts in policy liberalism common to all states.

7. Unlike many studies, the data set explicitly excludes social outcomes (e.g., incarceration or infant-mortality rates) as well as more fundamental government institutions (e.g., legislative term limits).

8. For further details on the policy liberalism measure, see sections A.2–A.3 of the appendix and Caughey and Warshaw (2016).

9. The policy liberalism scores have zero-mean and unit-variance across state-years. In a typical year, the cross-sectional standard deviation is around 0.9.

10. There are 41 dichotomous policies available in 1950 and 45 in 2010.

11. The welfare benefits are expressed in 2012 dollars and are adjusted for cost-of-living differences.

Table 1. Illustrative Policies of Selected States, 1950 and 2010

	Policy Liberalism	Percent Liberal	1950 Women on Juries	1950 Labor Anti-Injunction	1950 Housing Aid	1950 Fair Employment Commission	AFDC Benefit
Mississippi	−1.35	28	0	0	0	0	$460
Delaware	−.94	30	1	0	0	0	$642
Montana	.05	44	1	1	0	0	$838
Wisconsin	.93	56	1	1	1	0	$1,028
Massachusetts	1.33	62	1	1	1	1	$1,036

	Policy Liberalism	Percent Liberal	2010 Corporal Punishment Ban	2010 Prevailing Wage Law	2010 Medicaid Abortion	2010 Greenhouse Gas Cap	TANF Benefit
Mississippi	−2.29	17	0	0	0	0	$253
Virginia	−.89	33	1	0	0	0	$262
Nevada	−.13	45	1	1	0	0	$304
Minnesota	1.13	66	1	1	1	0	$323
Massachusetts	2.02	77	1	1	1	1	$352

Note. AFDC = Aid to Families with Dependent Children; TANF = Temporary Assistance for Needy Families.

Regression-discontinuity analysis

Electoral RD designs exploit the fact that a sharp electoral threshold, 50% of the two-party vote share, determines which party controls a given office (Lee 2008; Pettersson-Lidbom 2008). The validity of the RD design hinges on the assumption that only the winning candidate—and not the distribution of units' potential outcomes—changes discontinuously at the threshold. Unlike US House elections, where incumbents appear to have an advantage in very close elections (Caughey and Sekhon 2011), our analysis of state legislative and gubernatorial elections uncovers no statistically significant pretreatment discontinuities. Following Calonico, Cattaneo, and Titiunik (2014a, 2014b), we estimate both pre- and post-treatment discontinuities with local linear regression, using a bandwidth chosen to minimize mean square error (MSE) and adjusting confidence intervals to account for bias in the local-linear estimator.

RD for governor. Consistent with Eggers et al. (2015) and Folke and Snyder (2012), we find no significant discontinuities in the partisan composition of the state government at the time of the gubernatorial election (appendix section A.4, table A3). The only worrisome covariate is contemporaneous Policy Liberalism, which is somewhat higher where the Democrat barely won. The imbalance disappears, however, when Policy Liberalism is converted to a first difference.[12] In light of the better balance on first-differenced Policy Liberalism as well as for increased statistical efficiency, we estimate treatment effects on changes in Policy Liberalism rather than on levels.

Figure 2 illustrates the estimation of the policy effects of Democratic governors (relative to Republican governors) using the electoral RD design. The dependent variable is change in Policy Liberalism between the year of the governor's election and the governor's first year in office. On average, barely electing a Democratic governor is estimated to increase change in Policy Liberalism by about 0.03. Consistent with our expectations, this estimate is quite small relative to the variation in Policy Liberalism across states. Even the largest plausible one-year effect, which the confidence interval suggests is around 0.07, is less than one-tenth the cross-sectional standard deviation of Policy Liberalism.[13] Substantively, an effect of this size

12. The imbalance also disappears if we residualize Policy Liberalism using a regression with lagged dependent variables. Lee and Lemieux (2010, 331–33) suggest residualizing or differencing the dependent variable in RD designs as a way to increase statistical efficiency.

13. The point estimates are larger if Policy Liberalism itself is the dependent variable, but they are statistically significant only if Policy Liberalism is residualized using two-way fixed-effects. Adding lagged dependent variables to the residualizing regression yields point estimates very close to the

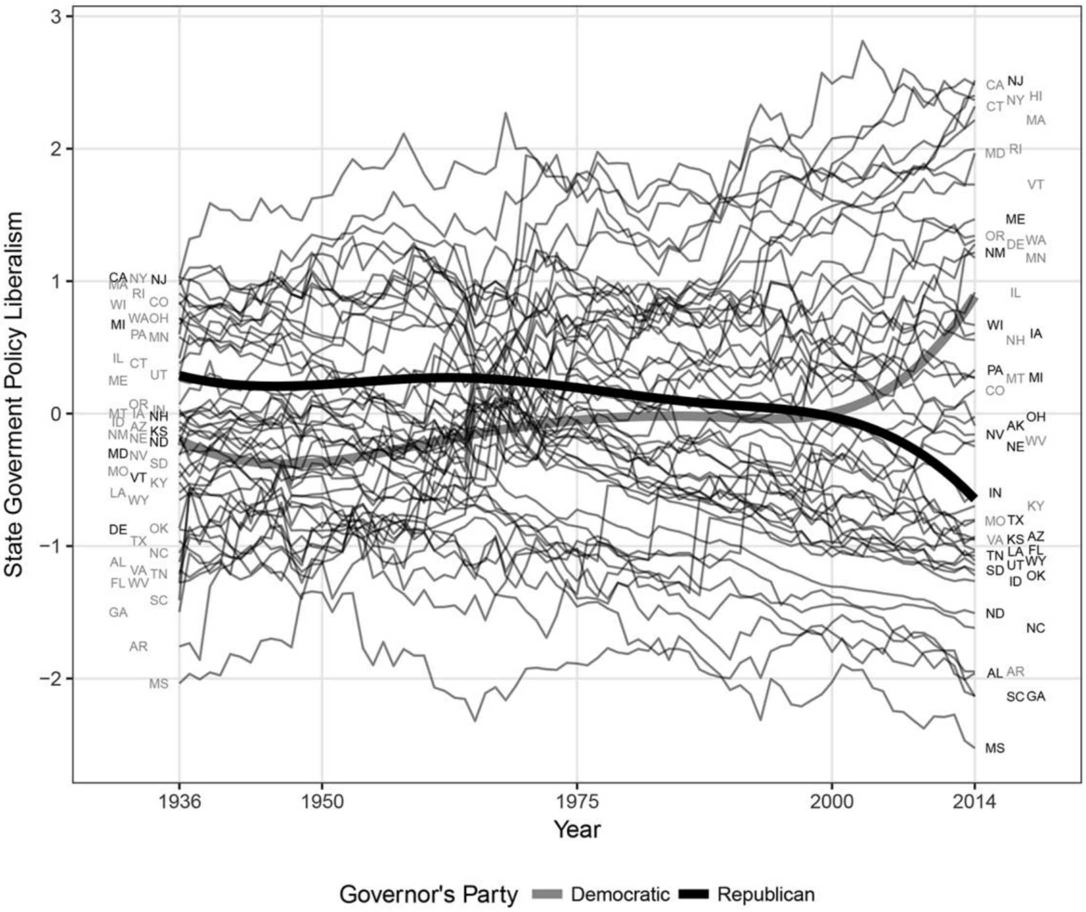

Figure 1. Yearly state Policy Liberalism, 1936–2014. Light and dark loess lines indicate the average Policy Liberalism of states with Democratic and Republican governors, respectively.

corresponds to about a one-point increase in a state's percentage of liberal policies.

Moreover, as figure 3 indicates, there is little solid evidence that policy effects cumulate over time. The effect after two years is only a bit larger than the one-year effect, and the effects after three and four years are essentially the same magnitude as the first year, though less precisely estimated. It thus appears that any effect of electing a Democratic governor is accomplished by the governor's second year in office. One possible reason for this lack of cumulation is that winning a gubernatorial election typically causes a party to lose seats in the next state legislative election (Folke and Snyder 2012), which could in turn lead to countervailing policy shifts. Indeed, voters' desire to counterbalance gubernatorial policy effects by electing a legislature of the opposing party may be a primary mechanism for such midterm slumps (Alesina et al. 1993).[14]

These local average treatment effect (LATE) estimates, however, conceal substantial temporal heterogeneity in the effect of partisan control. Mirroring the cross-sectional correlations plotted in figure 1, the policy consequences of electing a Democratic governor have grown markedly, particularly in recent decades. These changes are visualized in figure 4, which plots the evolution of gubernatorial policy effects over time. Each point and confidence interval in this plot corresponds to the gubernatorial RD estimate in a two-decade window. That is, the leftmost point is the estimated effect on one-year policy change for the period 1936–56, and the rightmost one is the same estimate for 1994–2014. This figure shows that through the 1970s, Democratic governors had essentially no estimated effect on Policy Liberalism. The magnitude of the

estimates for change in Policy Liberalism but a little more precisely estimated. Given this fact and the pretreatment differences in lagged Policy Liberalism we have the most confidence in the estimates with change in Policy Liberalism as the dependent variable.

14. Note that some governors have two-year terms and others have four-year terms. However, we see no difference in the cumulation of policy effects across states with different term lengths.

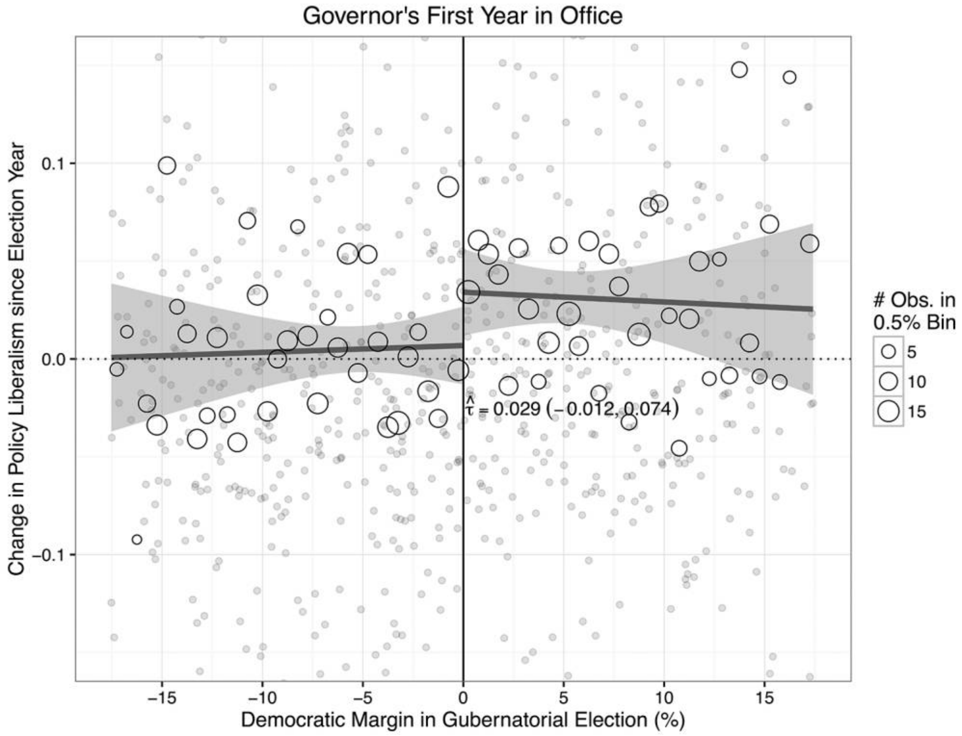

Figure 2. RD estimate of the effect of electing a Democratic governor on change in policy liberalism after the governor's first year in office. Estimates are based on triangular-kernel local linear regression, with MSE-optimal bandwidths and robust confidence intervals calculated by *rdrobust* (Calonico et al. 2014a). Open circles are averages in 0.5% bins. Shaded 95% confidence intervals are based on conventional standard errors.

estimates jumps up in the 1969–89 window but not until a decade later do the estimates become unambiguously positive. Between 1980 and 2014, the estimates hover around 0.06—approximately double the LATE estimate for the whole 1936–2014 period.

RD for State House. Descriptively, the cross-sectional relationship between policy liberalism and Democratic control of the state house and senate looks very similar to what figure 1 shows for governor: negative until around 1975, then nonexistent until the end of the twentieth century, when a strong positive association quickly emerged. However, this growing association in recent years could be due to an increase in the effect of public opinion or other changes in the political environment. Therefore, as we did for governors, we apply an RD design to estimate the causal effects of barely electing a Democratic majority in the state house (the lower chamber of the state legislature).[15] Because majority control of the legislature is a function of many elections rather than just one, however, we must construct a more complex assignment variable than in the gubernatorial RD.

The specific approach we follow is the multidimensional RD (MRD) design described by Feigenbaum, Fouirnaies, and Hall (forthcoming), which combines information from multiple close legislative elections.[16] The assignment variable they suggest is the Euclidean distance between a vector of district-level electoral results and the electoral results required for majority status. The first step in constructing this variable is to determine the number of seats (m) short of majority status the minority party is after a given election.[17] Then, obtain the Euclidean distance from majority status by summing the squares of the margins in the minority party's m closest losses in that election and taking the square root. Multiply this measure by -1 if the Democrats are in the minority. For example, if the Democrats are $m = 2$ seats short of a majority and the margins in their two closest losses are 3% and 4%, respectively, then the value of the assignment variable is $-1 \times (3^2 + 4^2)^{1/2} = -5$. Using data from Klarner

15. We do not examine the state senate because typically only a portion of senate seats are up for election in a given year.

16. For related multidimensional approaches to RD, see Folke (2014) and Reardon and Robinson (2012). An alternative design would be to use Democratic seat share as the assignment variable rather than a function of electoral results. We explored this design and found that it yields poor balance on important covariates, suggesting that seat share is too discrete and manipulable to be used as an RD assignment variable.

17. We estimate majority status based on the two-party seat share.

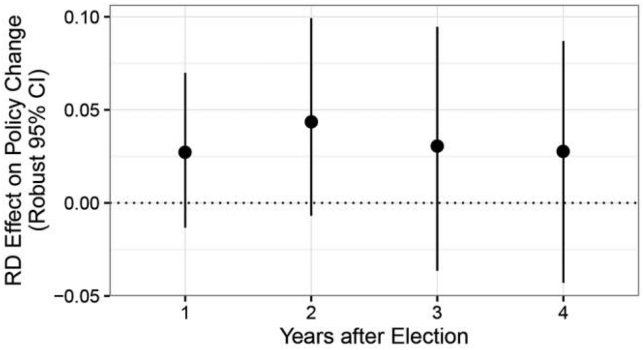

Figure 3. RD estimates of the effect of electing a Democratic governor, one to four years after the election.

et al. (2013), we are able to implement the multidimensional RD design for state house elections between 1968 and 2012.[18] None of the covariates exhibit statistically significant discontinuities, though the estimates of imbalance are somewhat less precise than in the gubernatorial RD (appendix section A.4, table A4).

Figure 5 plots the RD estimates of the policy effects of narrowly elected Democratic house majorities. Overall, the results are similar to those for governor. Narrowly electing a Democratic house majority causes a 0.05 increase in Policy Liberalism change after one year but no additional increase in the second year. Beyond the second year, these effects dissipate even more sharply than for governors. Indeed, the point estimate four years after the election is slightly negative, indicating that the positive effects of the first year are wiped out by the fourth year. As with governors, this could be the result of the endogenous political response to policy changes in the first two years, a possibility supported by the fact that narrowly winning a legislative majority decreases a party's probability of controlling the legislature in the future (Feigenbaum et al. Forthcoming). Finally, figure 4 shows that like gubernatorial policy effects, legislative policy effects have also grown over time. From a baseline of essentially zero, the one-year effect of electing a Democratic house has gradually climbed, reaching 0.08 by the end of the period and showing no signs of slowing.

Dynamic panel analysis

Given its transparent and testable identifying assumptions, the RD design is an appealing mode of causal inference, but its emphasis on observations near the RD threshold restricts the effective sample size. Thus, to increase statistical power we complement and extend the RD analyses reported above with an analysis that exploits within-state partisan variation in the full panel of state-years.

The crucial identifying assumption in the panel analysis is that the statistical model characterizes the counterfactual outcome each state would have exhibited under a different treatment assignment (i.e., a governor of the opposite party).[19] If unobserved confounding across states were constant across time and year-specific shocks affected all states equally, then the effect of a Democratic governor would be identified under a two-way fixed-effect (FE) model. This model, which is used by Besley and Case (2003) and others, assumes that the timing of shifts in party control is uncorrelated with time-varying state-specific determinants of policy liberalism (Angrist and Pischke 2009, 243–44). Unfortunately, given that ideological trends in state politics are likely to affect both partisan fortunes and policy outcomes, this assumption is unlikely to hold in this application.[20] We therefore estimate dynamic panel models with two-way FEs and lagged values of our dependent variable (Beck and Katz 2011):

$$y_{it} = \delta Gov_{it} + \beta Maj_{it}^{H} + \gamma Maj_{it}^{S} + \sum_{l=1}^{L} \rho_l y_{i,t-l} + \alpha_i + \xi_t + \epsilon_{it}, \tag{1}$$

where Gov_{it} indicates a Democratic governor; Maj_{it}^{H} indicates a Democratic house majority; Maj_{it}^{S} indicates a Democratic senate majority; $y_{i,\,t-l}$ is state i's policy liberalism l years before t; ρ_l is the coefficient on the l-th lag; and α_i and ξ_t are state- and year-specific intercepts, respectively.[21] All of the panel results reported in this article are qualitatively robust to alternative estimation strategies.[22]

Table 2 shows the results from the dynamic panel analysis. We first report gubernatorial estimates based on the conventional two-way FE model without LDVs in column (1).

18. Since multimember house districts cause complications for the design, state-years with multimember districts are dropped from the analysis. We also drop Nebraska, which has a nonpartisan legislature.

19. For details see appendix section A.5.

20. Another concern with the two-way FE model is that lagged dependent variables (LDVs) are potential confounders. This is because state policies change incrementally, and thus are highly correlated over time; meanwhile, policy outcomes could also affect the partisan composition of state government.

21. The FE-LDV estimator of δ in equation (1) is biased (Nickell 1981), but when the number of time periods is large, as it is in our case, the bias is a minor concern (Beck and Katz 2011; Gaibulloev, Sandler, and Sul 2014). Nonstationarity is also not a problem in our application (see appendix section A.6).

22. We explored a variety of alternative strategies to account for time-varying confounding, including state-specific time trends and a latent factor approach to interactive fixed effects (e.g., Bai 2009; Gaibulloev et al. 2014; Xu 2017). For details, see appendix section A.8. All diagnostic criteria indicate, however, that linear, quadratic, or even cubic time trends do not account for the dynamics of policy liberalism as well as LDVs do, and that latent factors are not necessary once LDVs are included.

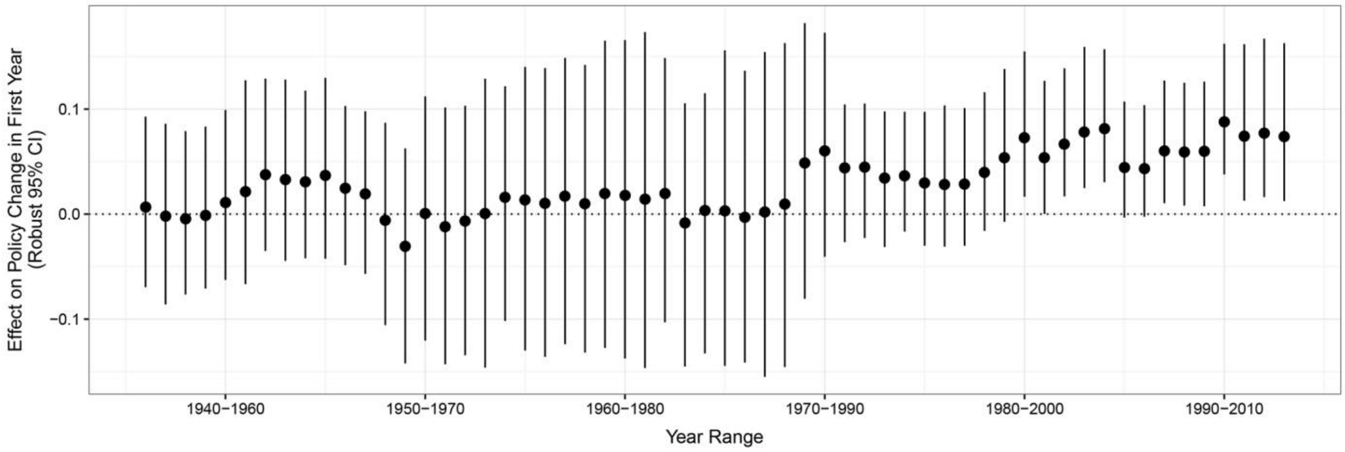

Figure 4. Changes in gubernatorial policy effects across the 1936–2014 period

These (implausible) two-way FE estimates suggest that relative to Republicans,[23] Democratic governors increase state Policy Liberalism by about 0.07, and that Democratic control of the state house and senate increases it by 0.17 and 0.27, respectively. The estimates shrink dramatically, however, if we control for LDVs. Column (2) reports the results from our preferred baseline specification, a FE-LDV model with two lagged terms, as specified by equation (1) with $L = 2$.[24] Under this specification, the estimated immediate effects of a Democratic governor, Democratic control of the house, and Democratic control of the senate are 0.01, 0.03, and 0.02, respectively.[25] All three estimates remain highly statistically significant, but the point estimates are an order of magnitude smaller than the FE model. This strongly suggests that FEs alone do not adequately account for within-state trends in Policy Liberalism and are likely to overestimate policy effects (for further evidence on this point, see appendix section A.8).

It is important to note that the effect of a Democratic legislative majority has a different interpretation in the dynamic panel analysis than in the RD analysis. In the RD design, the estimand is the LATE of electing a bare Democratic majority rather than a bare Republican majority. In the dynamic panel analysis, however, the estimand conflates the effect of chamber control per se with that of seat share since the party in control typically has more than a bare majority. This conceptual difference notwithstanding, the estimates for majority control barely change if we control for seat share because share has little independent association with Policy Liberalism (appendix section A.10). Indeed, for both state house and governor, the panel estimate are somewhat smaller than (though statistically indistinguishable from) the corresponding RD estimate, suggesting that parties receive little additional policy benefit if they win control by a larger-than-bare margin. Table 2 also explores the possibility that the policy effects of one institution depend on party control of other institutions. We might expect, for example, that capturing the governorship yields greater policy benefits if the same party also controls both houses of the legislature. However, there is no clear evidence of positive interaction effects between the coefficients.[26]

Next, we examine whether the results differ between the South and non-South. As column (4) of table 2 shows, the results for the non-South are substantively similar to (and statistically indistinguishable from) those for the whole sample. This makes sense because both the RD and dynamic panel analyses implicitly place greater weight on competitive states (those with closer elections and more alternation in party control), and until recently state politics in the South was dominated by the Democratic party. Due to the lack of partisan variation in Southern states, the estimates for the South are very imprecise, and none are distinguishable from zero.

Partisan polarization and the growth in party effects on policy

We saw in figures 4 and 6 that partisan effects on policy have grown markedly, especially in the last quarter century. What has driven these increases? One obvious potential culprit is

23. Among the 3,630 state year observations, only 29 have independents as governors. Dropping these observations does not change our main finding at all.

24. The gubernatorial estimate remain very stable if we control for more than two LDVs; see appendix section A.9.

25. In a dynamic panel model, a treatment will affect not only the contemporaneous outcome but also outcomes in future periods through the channel of the LDVs. The effect on the contemporaneous outcome is often called the "immediate" effect.

26. Appendix section A.7 shows a graph of these interactions.

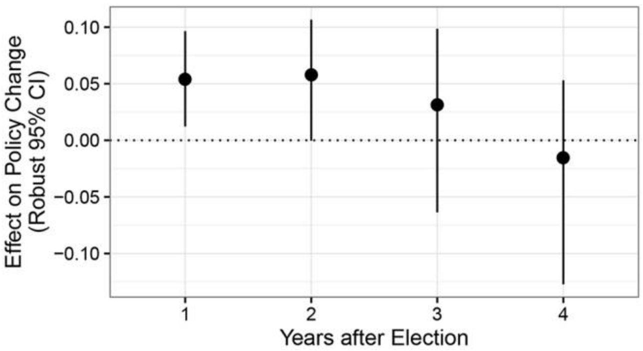

Figure 5. RD estimates of the effect of electing a Democratic state house, one to four years after the election. RD = regression discontinuity.

polarization in the policy preferences (whether sincere or induced) between Democratic and Republican candidates and office holders, which is well documented among members of Congress and other national politicians (e.g., Layman, Carsey, and Horowitz 2006). If, as seems likely, the policy positions of state-level politicians have also diverged by party, we should expect them to pursue increasingly distinct policies in office, thus increasing partisan effects on policy. Moreover, to the extent that government officials are responsive to their partisan subconstituencies, we should also expect elite polarization—and thus partisan effects on policy—to be larger where the policy preferences of Democrats and Republicans in the public diverge more (Adams and Merrill 2008; Clinton 2006; Jacobson 2012).

Preliminary evidence for this last point is provided by figure 7, which plots the cross-sectional relationship between elite and mass partisan divergence. We measure Elite Divergence (vertical axis) as the ideological distance between the median Democrat and median Republican in the state legislature, which Shor and McCarty (2011) have estimated annually since 1993. Analogously, we measure Mass Divergence (horizontal axis) as the ideological distance between the average Democrat and average Republican identifier in the state public, using the estimates of mass-level economic policy liberalism developed by Caughey, Dunham, and Warshaw (2016). This measure, available for each state in each year

Table 2. Policy Effects of Democratic Control of the Governorship, State House, and State Senate

	Policy Liberalism t				
	Full Sample			Non-South	South
Outcome Variable	(1)	(2)	(3)	(4)	(5)
Democratic governor	**.065**	**.012**	**.014**	**.010**	.022
	(.031)	(.004)	(.007)	(.004)	(.012)
Democratic house majority	**.165**	**.030**	**.043**	**.032**	.014
	(.051)	(.006)	(.013)	(.007)	(.011)
Democratic senate majority	**.271**	**.021**	.008	**.021**	−.023
	(.058)	(.006)	(.012)	(.006)	(.011)
Democratic house majority × Democratic senate majority			−.002		
			(.017)		
Democratic governor × Democratic house majority			−**.032**		
			(.016)		
Democratic governor × Democratic senate majority			.009		
			(.015)		
Democratic governor × Democratic house majority × Democratic senate majority			.025		
			(.021)		
State and year FEs	✓	✓	✓	✓	✓
Policy liberalism $t-1$		✓	✓	✓	✓
Policy liberalism $t-2$		✓	✓	✓	✓
Observations	3,678	3,586	3,586	2,749	837
States	49	49	49	38	11
R^2	.871	.988	.988	.983	.947

Note. Standard errors produced by block bootstraps (clustered at the state level) of 1,000 times are in parentheses. Nebraska is not included in the sample. Coefficients statistically significant at the 5% level are in bold font. FE = fixed effect.

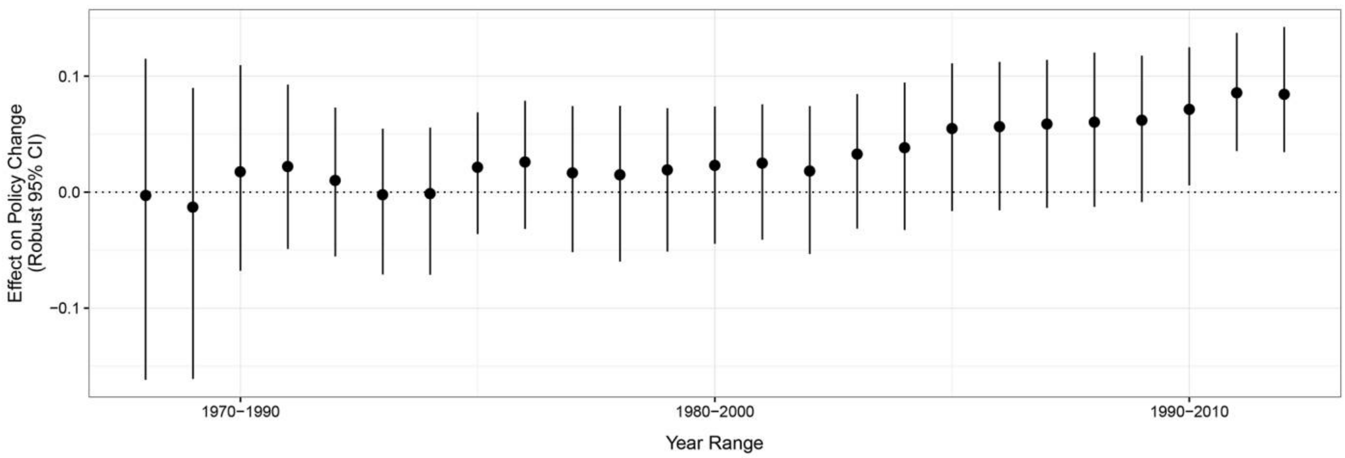

Figure 6. Changes in legislative policy effects across the 1968–2012 period

between 1946 and 2014, was derived from a dynamic group-level item-response model of over 800,000 survey respondents' preferences on economic issues (Caughey and Warshaw 2015).[27] Plotting the within-state averages of both measures over the 1993–2014 period, figure 7 shows that although their correlation is not perfect ($r = 0.5$), states with greater Mass Divergence clearly tend to have more polarized state legislatures.

Next, we examine whether partisan effects on policy also tend to be larger where Mass and Elite Divergence is greater. To simplify the analysis, we create a modified version of the panel model in equation (1) that includes a variable indicating the proportion of government offices/chambers (i.e., governorship, state house, and state senate) controlled by the Democratic Party.[28] The first column of table 3 reports the results of a specification that interacts this Democratic Control variable with indicators for three time periods: 1936–1968, 1969–1991, and 1992–2014.[29] Consistent with the RD estimates in figures 4 and 6, the coefficient estimates indicate that the effect of Democratic Control was roughly constant in the first two periods but doubled in magnitude after 1991. As column (2) shows, the results are qualitatively identical if we restrict the analysis to the years for which mass partisan divergence is available (1947–2014). If we also interact Democratic Control with lagged Mass Divergence, however, the former's interaction with the post-1992 dummy is reduced to insignificance. This suggests, though hardly proves, that era indicators may simply be proxying for changes in Mass Divergence over time.

Ideally we would conduct the same analysis for Elite Divergence, but the Shor-McCarty state legislative ideal points do not extend before 1993. Nevertheless, we can still examine whether Elite Divergence moderates the effect of Democratic Control in the post-1993 period. The answer, provided by column (4), is a clear yes. The coefficient estimate for the interaction of Democratic Control and lagged Elite Divergence indicates that the former's effect increases by 0.05 for

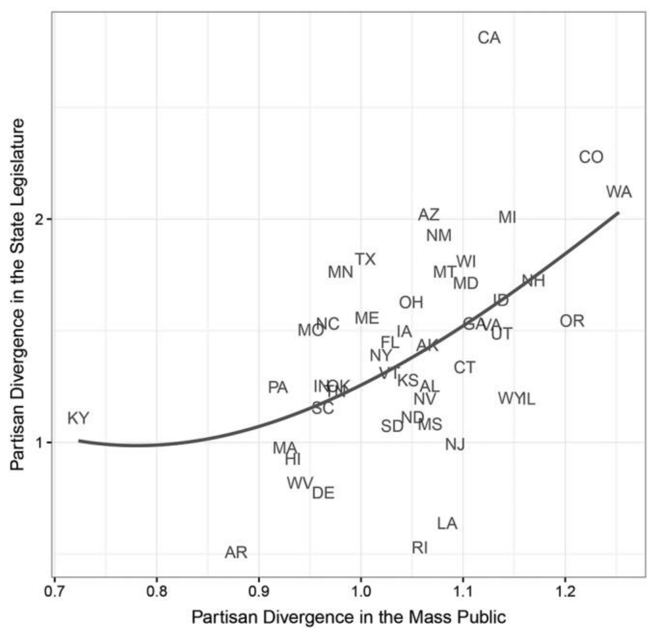

Figure 7. Relationship between mass partisan divergence (Caughey et al. 2016) and elite partisan divergence (Shor and McCarty 2011), averaged within states across the 1993–2014 period. The fitted line is a three-knot natural spline.

27. See appendix section A.11 for a more comprehensive description of the measure of opinion divergence between Democrats and Republicans in each state.

28. The linearity assumption implied by the use of this index seems reasonable in light of the roughly additive effects of different offices reported in table 2.

29. We defined the eras in this way because they divide the years that our measure of mass partisan divergence is available into three equal parts.

Table 3. Moderators of Partisan Effects on Policy

	Policy Liberalism t				
Outcome Variable	(1)	(2)	(3)	(4)	(5)
Democratic control	**.055**	**.063**	.042	−.024	−.089
	(.018)	(.022)	(.026)	(.046)	(.058)
Democratic control × era 1969–1991	−.014	−.020	−.027	NA	NA
	(.022)	(.024)	(.024)		
Democratic control × era 1992–2014	**.066**	**.061**	.029	NA	NA
	(.022)	(.026)	(.026)		
Mass divergence$_{t-1}$			−.025		−.005
			(.015)		(.068)
Democratic control × mass divergence$_{t-1}$			**.015**		**.028**
			(.008)		(.013)
Elite divergence$_{t-1}$				−.027	−.020
				(.018)	(.019)
Democratic control × elite divergence$_{t-1}$				**.049**	**.038**
				(.014)	(.013)
Years covered	1936–2014	1947–2014	1947–2014	1994–2014	1994–2014
State and year FEs	✓	✓	✓	✓	✓
State × era FEs	✓	✓	✓	NA	NA
Policy liberalism $t-1$	✓	✓	✓	✓	✓
Policy liberalism $t-2$	✓	✓	✓	✓	✓
Observations	3,586	3,182	3,182	812	812
States	49	49	49	49	49
R^2	.989	.989	.989	.995	.995

Note. Standard errors produced by block bootstraps (clustered at the state level) of 1,000 times are in parentheses. Nebraska is not included in the sample. Coefficients statistically significant at the 5% level are in bold font. Measures of mass partisan divergence and elite partisan divergence are rescaled based on their standard deviations during the period of 1994–2014. FE = fixed effect; NA = not applicable.

every standard deviation increase in the latter.[30] This result persists even if Democratic Control is also interacted with dummies for state and year, which indicates that the moderating effect of Elite Divergence is not driven by national time trends in partisan policy effects or by durable state differences in these effects. It is interesting to note that the implied effect of Democratic Control when the party medians in the legislature are equal is essentially 0, as one would expect if candidates converged on the same policy positions.[31] Finally, the rightmost column of table 3 demonstrates that both Mass Divergence and Elite Divergence continue to moderate Democratic Control when they are included in the same model. This suggests that Mass Divergence may lead to or proxy for ideological differences between Democratic and Republican candidates that are not fully captured by roll-call patterns in the state legislature.

Taken together, the evidence presented in this section corroborates the hypothesis that the magnitude of party effects is a function of the ideological distance between candidates of different parties. More tentatively, they also suggest that the size of policy effects may be influenced by the mass public as well, whether through electoral pressures to cater to more-or-less extreme primary electorates or some other mechanism. Given that partisan divergence has increased at both the mass and elite levels (Caughey et al. 2016; Hill and Tausanovitch 2016), these results thus provide a potential explanation for the growth of partisan effects on state policy.

DISCUSSION AND IMPLICATIONS

Commenting on state politics around 1980, Erikson et al. observed that Democratic and Republican parties in each state "respond to state opinion—perhaps even to the point of

30. We tested the validity of the multiplicative interaction models using diagnostic tools proposed by Hainmueller, Mummolo, and Xu (2016). Both the overlap and linearity assumptions appear to be valid.

31. No state is estimated to have no elite divergence, but some get quite close. The least polarized state-year is Arkansas in 1993, whose Elite Divergence score is 0.4 (the average score across state-years is 3).

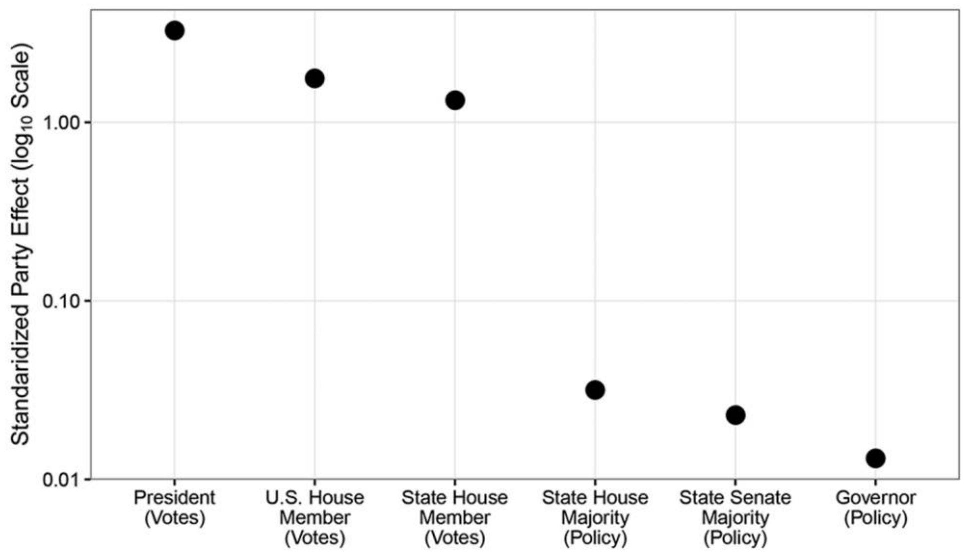

Figure 8. Position effects and policy effects. The left three quantities are counterfactual differences in roll-call ideal points between Republicans and Democrats occupying the same office. The right three are analogous estimated effects of party control on state policy liberalism. For comparability, each of the estimates is standardized by the cross-sectional standard deviation of the dependent variable. The vertical axis is on the $\log_{10}$ scale, so each line represents an effect ten times larger than the line below it.

enacting similar policies when in … control" (1993, 121). Based on an analysis spanning eight decades, we come to similar conclusions about statehouse democracy at that time. Before the 1990s, electing Democratic rather than Republican governors and legislatures generally had small effects on the liberalism of state policies. Since Erikson et al.'s seminal analysis, however, partisan effects have grown rapidly, and electing Democrats now has an unambiguously positive impact on policy liberalism. In other words, the parties have increasingly diverged in the policies they implement in office, a trend that seems at least partly attributable to the growing ideological gap between the parties' candidates and electoral coalitions.

The substantive magnitude of contemporary policy effects, however, should not be overstated. In 2010, for example, Democratic governors, houses, and senates are each estimated to increase policy liberalism by around 0.04 per year. As table 1 suggests, an effect of this size would be expected to increase a state's percentage of liberal policies by a small amount, on the order of 0.5%. Or, to take an important welfare policy, it would increase average monthly TANF benefits per recipient family by a little over \$1.[32] The substantive magnitude of partisan effects on policy also pales relative to the cross-sectional differences between states. The estimated policy effect of a switch in unified party control in recent decades is one-tenth the size of the typical difference between states, suggesting that many decades of Republican governors and legislatures would be required to make the policies of Massachusetts as conservative as those of Mississippi.[33]

As a final point of comparison, consider the focus of most research on partisan polarization: the difference between candidates' policy positions, as measured by their roll-call records, campaign platforms, or financial supporters (e.g., Ansolabehere, Snyder, and Stewart 2001; Bonica 2014; Poole and Rosenthal 1984). We call such differences "position effects." Numerous studies have found that party affiliation is by far the most powerful predictor of politicians' policy positions, at both the national and the state level (e.g., Shor and McCarty 2011). Figure 8 confirms this finding, showing that there is a difference of one to four standard deviations in the ideal points of otherwise similar presidents, US House members, and state house members from opposing parties.[34]

32. Calculated based on the linear association between policy liberalism and TANF benefits in 2010.

33. Of course, this hypothetical comparison glosses over two complications. First, Massachusetts Republicans are less conservative than Mississippi Republicans, so party effects may differ across states (see Erikson et al. [1993], however, for evidence that the within-state divergence of the parties does not vary strongly with state liberalism). The second complication is that the comparison ignores any endogenous political response to changes in policy liberalism. We have both theoretical (e.g., Alesina and Rosenthal 1995) and empirical (e.g., Folke and Snyder 2012) reasons to believe that voters will respond to rightward (leftward) changes in state policy by electing more Democrats (Republicans) to state office.

34. The ideal point measure for the US House and president is DW-NOMINATE (Poole and Rosenthal 2007). The House estimate based on an RD design (estimates based on two-way fixed effects or any other estimator are very similar); the president estimate is simply the raw difference between

By contrast, analogously standardized policy effects are nearly two orders of magnitude smaller.[35] Of course, the two sets of quantities are not fully comparable—some are defined at the individual level, others at the level of the office or body—and standardizing the estimates does not necessarily put them on the same scale as each other, let alone the same scale as citizens. But the vast differences in magnitude between position and policy effects cannot help but cast a very different light on partisan polarization. In particular, they call into question the concern that alternation in party control leads to "wide swings in policy" that "do not well represent the interests of middle-of-the-road voters" (Poole and Rosenthal 1984, 1061). Whether due to status quo bias, the necessity of compromise, or the realities of policy making as opposed to symbolic position taking, the effects of party control appear much less dramatic by the metric of actual policy outcomes. Even if the policy positions of politicians from different parties "leapfrog" over the citizens they represent (Bafumi and Herron 2010), partisan control of government has only incremental effects on policy outcomes. In short, Democrats and Republicans may disagree consistently and even violently, but the actual policy consequences of these disagreements are far less dramatic.

ACKNOWLEDGMENTS

We thank participants at the 2014 MPSA Conference and seminar participants at MIT, Rochester, Yale, and Duke for feedback on previous versions of this manuscript. We are grateful for feedback on earlier drafts of this manuscript from Thad Kousser, Jens Hainmueller, Andy Hall, Danny Hidalgo, Dan Hopkins, Chris Tausanovitch, and Eric Schickler. We appreciate the research assistance of James Dunham, Melissa Meek, Kelly Alexander, Aneesh Anand, Tiffany Chung, Emma Frank, Joseff Kolman, Mathew Peterson, Charlotte Swasey, Lauren Ullmann, and Amy Wickett. We also appreciate the willingness of Carl Klarner to generously share data. All mistakes, however, are our own.

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Democratic and Republican president-years since 1936. The figure for the state house is based on the matching estimate of intradistrict partisan divergence in ideal points reported in table 2 of Shor and McCarty (2011, 548).

35. These are the estimates reported in column (2) of table 2, divided by the standard deviation of policy liberalism across states in a typical year.

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