Political Control Over Redistricting and the Partisan Balance in Congress

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We estimate the impact of a political party's legal ability to unilaterally redistrict Congressional seats upon partisan seat share allocations in the U.S. House of Representatives. Controlling for stateXdecade and year effects, we find an 8.2 percentage point increase in the Republican House seat share in the three elections following Republican control over redistricting in the past two decades. We find no significant or sizable effect for Democrats. The effects over the past five decades in aggregate are smaller and insignificant for both parties. In the past two decades, these effects are sizable though not pivotal for Congressional control. *Keywords: Gerrymandering, redistricting, voting*

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"The partisan gerrymanders in these cases deprived citizens of the most fundamental of their constitutional rights: the rights to participate equally in the political process, to join with others to advance political beliefs, and to choose their political representatives." - Justice Elena Kagan, Dissent, Rucho et al. v. Common Cause et al.

"Extreme partisan gerrymandering is a real problem for our democracy" - Justice Brett Kavanaugh

I. Introduction

"Representatives and direct Taxes shall be apportioned among the several States which may be included within this Union, according to their respective Numbers, which shall be determined by adding to the whole Number of free Persons, including those bound to Service for a Term of Years, and excluding Indians not taxed, three fifths of all other Persons. The actual Enumeration shall be made within three Years after the first Meeting of the Congress of the United States, and within every subsequent Term of ten Years, in such Manner as they shall by Law direct." - Article 1 Section 2, U.S. Constitution

Geography-based democratic representation either leads over time to unequal representation or allows for alterations to geographical districts. In the United States, redistricting in order to maintain balance in the House of Representatives is mandated by Article 1 Section 2 of the U.S. constitution. Every ten years, the U.S. government is required to undertake a census of the population and use it as a basis for redistricting in order to ensure equal representation. However, this process potentially allows for politicians to redraw political boundaries in order to affect partisan control over both federal and state legislatures.

Partisan interest in redistricting became apparent as early as 1812, when the Massachusetts State Senate redrew electoral boundaries, as mandated by the U.S. constitution. The redrawn districts benefited the Democratic-Republican party over the opposition Federalist party. Governor Elbridge Gerry of Massachusetts, a Democratic-Republican, signed the redistricting bill into law though he personally lamented the highly partisan process. The redistricting resulted in some oddly shaped districts. One Federalist newspaper, the Boston Gazette, noted that one of the state senate districts looked like a salamander. The Gazette coined the term "Gerry-mander" in a political cartoon from March, 1812.

In recent years, there has been increased concern over whether or not redistricting leads to gerrymandering: i.e. whether or not parties redistrict in order to increase their share of legislative seats. A body of theoretical work shows that self-interested political parties will redistrict by *cracking* opposition districts with a narrow majority and *packing* the opposition into lop-sided districts in an attempt to increase own-party seat share (Shotts, 2001; Gilligan and Matsusaka, 1999). More recent work points out that packing is beneficial but when a party is uncertain of partisan leanings and voter turnout, it is usually not optimal to crack (Friedman and Holden, 2008).

There also has been an empirical literature on the impact of redistricting. Much of the best empirical work to date simulates counterfactual maps subject to legal or norms-based constraints such as requirements that the district be connected and that it be compact (Chen, Rodden et al., 2013; Chen and Rodden, 2015; Stephanopoulos and McGhee, 2015). This literature then computes probabilities of an outcome at least as partisan as the actual outcome. Though this work tends to find that in many states few alternative ways of drawing districts yield greater imbalance in the relationship between voting behavior and representation, it is nonetheless possible that these districts which confer partisan advantage to one party or another reflect natural geographical boundaries or reflect natural political communities. Some of this work suggests that partisan advantage due to asymmetries in clustering across districts is due to the clustering of like-minded individuals (Chen, Rodden et al., 2013) rather than the intentional design of parties. A large literature has noted increased political sorting over time (Bishop, 2009; Hopkins, 2017; Kaplan and Sullivan, 2018). Nonetheless, a recent paper (Jeong and Shenoy, 2016) shows that parties engineer narrow state government victories in the election before redistricting.

Instead of looking at maps relative to counterfactual potential maps, we look at whether federal seat shares increase in the direction of a party when that party has legal control over the redistricting process. To do so, in contrast with most of the prior literature, we provide comprehensive evidence on the prevalence of partisan gerrymandering over 50 years of American history, across parties, and heterogeneously by the number of seats in a state.

We first develop a measure of the amount of redistricting as the fraction of a state which changes districts. We show empirically that almost all redistricting happens once a decade by legislatures in power in years that end in 1. We then estimate the impact of the ability of one party to pass a redistricting bill without votes from the opposition party on that party's fraction of seats in Congress in subsequent elections.

Controlling for year effects as well as state X decade effects, we find a statistically significant positive impact of 4.7 percentage points of Republican legal control on the Republican seat share in Congress in the subsequent election. The effect is 9.1 percentage points restricted to the past two decades. The average effect over the next three federal elections is positive but not statistically significant for the full five decade sample but increases in size to 8.2 percentage points and becomes statistically significant at conventional levels when we restrict our data to the past two decades. In contrast, we do not find statistically significant effects or sizes as large with Democratic control except for large state delegations restricted to the past two decades. Our estimates for the impact of Republican control over redistricting are relatively stable regardless of state delegation size. However, we do find statistically significant increases in the Democratic seat share following Democratic control over redistricting only when we restrict to the past two decades and only in states with more than five seats. Overall, simple back of the envelope calculations suggest that partisan redistricting can on net account for less than 5% of the gap between Republicans and Democrats in the House during each of the 1970s, 1980s and 1990s. However, the same calculations show that it can account for 57% of the gap in the 2000s and 51% in the 2010s.

In the next section, we further discuss the relevant literature. In section 3, we discuss important institutional features of the U.S. redistricting process. In section 4, we discuss our empirical methods. In section 5, we give an overview of the data we use for our estimation. In section 6, we present our main results. In section 7, we perform an exercise in which we compute aggregate impacts of the rights to redistrict upon the partisan balance in Congress. Finally, in section 8, we conclude.

II. Literature Review

As we noted, most of the literature on redistricting has looked cross-sectionally at the relationship between vote shares and seat shares to infer the impact of partisan bias in the construction of districts. Two recent contributions to the literature on redistricting also provide panel rather than cross-sectional evidence on the impact of redistricting. A recent book, (McGann et al., 2016), looks at partisan differences in concentration of of partisanship by district. They use the skewness of the distribution of the Democratic two-party vote-shares across districts to measure partisan bias at a state level. This is similar to the efficiency gap measure developed by (Stephanopoulos and McGhee, 2015) in that it is a measure of concentration of a party's vote. They find that the distribution of Democratic two party vote-shares is right skewed; in other words, there are many more very Democratic districts than there are very Republican districts. Of course, this is possibly due to Democrats living in more homogeneous communities (i.e. urban areas). Thus, they also do look at the change in the skewness of the distribution of Democratic two-party vote shares. They find that Democratic two-party skewness increased after the 2010 redistricting cycle.

The paper which is most similar to ours is (Stephanopoulos, 2017). He estimates the impact of unified control on the efficiency gap, a measure of a party's 'wasted' votes within a Congressional district using a panel of states. He regresses the efficiency gap on state and time fixed effects and a small number of state-level demographic characteristics. Relative to (Stephanopoulos, 2017), we have a number of differences. First, we estimate the impact upon Congressional seat shares which is of more interest; additionally, estimating effects upon seat shares allows us to simulate counterfactual seat allocations in Congress by party in the absence of any party having legal control. Second, we code the legal requirements for control by a party over the redistricting process going back to 1971. The correlation coefficient between our legal control variable and the unified control variable used in (Stephanopoulos, 2017) in redistricting years is 0.66 - decently positively correlated but far from 1. Since we measure the legal ability of a party to control redistricting, the fact that their estimates are smaller in magnitude relative to ours is consistent with an impact of measurement error. Third, their two-way fixed effects estimation strategy is subject to aggregation bias in the presence of cohort-heterogeneity across cohorts. We find substantial time-varying heterogeneity in our estimated effects which we demonstrate does in fact cause substantial bias in the estimated coefficient. Fourth, we show dynamic evidence of the effect. We show that it occurs precisely in redistricting years and lasts for at least two more Congressional elections. We also do not find a substantial trend in seat shares prior to redistricting. Finally, we separate our effect by parties and find large differences across the parties. Our results are generally consistent with the asymmetric polarization literature in political science and in economics (Gentzkow, 2016; Grossmann and Hopkins, 2016; McCarty, Poole and Rosenthal, 2016).

III. Institutional Background

The use of district boundary creation to influence elections goes back to the period before the Constitution when the Articles of Confederation was law. Patrick Henry, along with other anti-Federalists, purportedly altered Virginia's 5^{th} Congressional District in an attempt to prevent the strong Federalist, James Madison, from returning to Congress (Labunski, 2006). In 1789 (the following year), the states ratified decadal redistricting. Since the adoption of the Constitution, redistricting has happened within the first three years of the decade in almost all states in every decade with the exception of the 1920s.[1]

The process of redrawing districts happens in two phases. In the first phase, reapportionment, the U.S. Congress uses the Census data and by January 25^{th} of the year following the completion of the Census, assigns numbers of seats in the House of Representatives to each state. Though there are multiple possible methods to apportion seats, Congress uses the Hamilton-Hill method which minimizes deviations in number of representatives per person across states. Though the first Congress had 105 members and an average of approximately 33,000 individuals per representative, the size of Congress grew over time until it was capped in 1911 to 435 (after the addition of Arizona and New Mexico to the United States in 1912). This cap was reauthorized in 1929 and has been in place continuously since then except for a temporary increase in 1959 when Alaska and Hawaii joined the United States and the number of representatives rose temporarily to 437.

After reapportionment, the states are notified of the number of representatives that they are apportioned. The federal government historically has given individual states wide latitude to redistrict as they see fit. After a sequence of Supreme Court decisions in the 1960s (*Baker v. Carr*, *Wesberry v. Sanders*, and *Reynolds*

[1] In the 1920s, reapportionment would have led to a shift of 11 seats away from rural areas towards urban areas which had grown in size dramatically due to immigration. A coalition of representatives from rural areas made sure that the reapportionment was blocked until 1929 (Anderson, 2015).

v. Sims,), states have been required to equalize the number of people in each district. Though reapportionment results in relative balance across states in representation in the House of Representatives, individual states created districts with a high degree of population imbalance. For example, one district in Tennessee represented 2,340 and another in the same state represented 42,298 people represented. The worst example of representational imbalance was in the Nevada state legislature where one district contained 568 voters and another approximately 127,000. In the early 1960s, the Warren court handed down three rulings. First, in 1962, *Baker v. Carr* established that redistricting was subject to judicial review. Then, in 1964, *Wesberry v. Sanders* mandated equal population in federal Congressional districts. *Reynolds v. Sims* also in 1964, then extended equal representation to state legislative districts. In subsequent decisions (*Karcher v. Daggett*, 1983; *Vieth v. Jubelirer*, 2003), the Supreme Court clarified that Congressional Districts should be exactly equal in size to the degree possible whereas for state legislative districts deviations of up to 10% across districts would be allowed (*Brown v. Thomson*, 1983) (Ansolabehere and Snyder, 2008).

In addition, the Supreme Court has also decided that as long as districts are sufficiently compact, redistricting in order to create majority minority Congressional districts is legal but other racially-based reasons are not (Thornburg v. Gingles, 1986; Shawn v. Reno, 1993; Miller v. Johnson, 1995). However, the courts have been more reluctant to disallow redistricting for partisan gain (Davis v. Bandemer, 1986; Vieth v. Jubelirer, 2004). In Vieth v. Jubelirer (2004), Supreme Court Justice Anthony Kennedy made the argument that in principle the Supreme Court could intervene to prevent partisan gerrymandering but that there was no empirical metric with which the Supreme Court could intervene. He stated, "That no such standard has emerged in this case should not be taken to prove that none will emerge in the future." In 2019, the Supreme Court decided in Rucho v. Common Cause that the Supreme Court did not have the authority to intervene in order to limit partisan redistricting. However, court battles are ongoing at the state level.

States differ in their redistricting laws and processes. Seven states do not redistrict federal Congressional boundaries because they only have one federal representative: Alaska, Delaware, Montana, North Dakota, South Dakota, Vermont and Wyoming. We drop these states from our main analysis as they do not participate in redistricting. Moreover, since these districts have 100% partisan seat shares unless they elect an independent, including these high variance observations decreases our precision.

We drop Nebraska because since 1934, Nebraska has not allowed political parties to operate at the state level. Thus, it is difficult to tell whether or not Democrats or Republicans have control over legislative bodies and thus whether one party has legal control over the redistricting process[2].

[2]Nebraska also became the only state to have a unicameral state legislature with the passage of the same 1934 law

Over the fifty years which our data set encompasses, 12 states have used a commission to draw maps and implement redistricting in at least some decades. The composition of the redistricting commissions varies by state but many attempt to appoint a bipartisan commission by balancing the number of Democratic appointees and Republican appointees to the commission (Arizona, California, Hawaii, Idaho, Maine, Montana, New Jersey and Washington). [3] Three other states (New York, Ohio, and Rhode Island) use commissions that do not have partisan balance. In Iowa, non-partisan staff draw maps every decade and the maps are then sent for approval to the legislature and the governor. Commissions vary from state to state in the extent of their redistricting powers and restrictions. We treat all state-decades with a commission as being of a singular "commission" type. We include these states in our benchmark specification but also show robustness to dropping them as well as recoding them.

In the rest of the states, the legislatures or legislature and governor draw the maps. When one party has a trifecta, they have the capacity to draw the maps and pass a redistricting bill without help from the other party. In Connecticut and North Carolina, control over both chambers of the legislature is sufficient because the governor cannot veto the redistricting bill.

We estimate the impact of the legal ability of a party to redistrict upon subsequent federal partisan seat shares. This depends upon the laws in the state, the number of seats, the distribution of legislative seats across both chambers and usually the control of the governorship. We use variation across decades and across states to estimate the impact of partisan control over redistricting on partisan seat shares. The details of the laws and how they vary by states over time are explained in greater detail in the Data Appendix. In Figure 1, we color code states by decade with blue for Democratic control, red for Republican control, and gray if neither party had control. We code neither party as having control if the government was divided and required approval by all chambers plus the governor, if the legislature was divided and only the legislature was required to pass a redistricting bill, or if redistricting was delegated to a commission. In the 1970s through 1990s, the Democrats had a much higher share of states with legal control. However, in the 2000s, control was largely balanced across parties and in the 2010s, the Republicans maintained partisan control in a substantially higher fraction of states.

IV. Empirical Methods

In this section, we present the empirical methods that we will use to estimate our main effects. In our main specification, we regress an outcome variable $O_{s,y}$, the Republican House of Representative seat share, on a measure of partisan legal control conditional upon stateXdecade and year effects; $DemControl$ and $RepControl$ are dummy variables which take on a value of 1 if either the

[3] Montana uses the commission in years in our data set where they have more than one Representative.

Democrats or the Republicans respectively have the legal ability to pass a redistricting bill solely on votes from their own party; and $\gamma_{s,d}$ and δ_y are stateXdecade and year fixed effects respectively.

We include one lead and 3 lags of our main treatment variables: $DemControl$ and $RepControl$. These together with the contemporaneous effect saturate the decade and allows us to trace the dynamic path of our effect over three elections in addition to checking for a pre-trend. Our main specification is thus given by:

$$\begin{aligned} O_{s,d,y} = \alpha + \sum_{k=-1}^{3} \beta_k^D L_k.DemControl_{s,y} \\ + \sum_{k=-1}^{3} \beta_k^R L_k.RepControl_{s,y} + \gamma_{s,d} + \delta_y + \epsilon \end{aligned} \tag{1}$$

We perform tests of the difference in average outcome in the three elections after those ending in 2 (i.e. those ending in 4, 6, and 8) relative to the contemporaneous vote share:

$$\begin{aligned} \frac{\beta_4^D + \beta_6^D + \beta_8^D}{3} - \beta_2^D = 0 \\ \frac{\beta_4^R + \beta_6^R + \beta_8^R}{3} - \beta_2^R = 0 \end{aligned} \tag{2}$$

We additionally show as robustness a two-way fixed effects model with state and year effects. Our main specification with stateXdecade rather than state effects is more restrictive in order to improve upon identification of the average effect of partisan control upon partisan seat shares. A recent econometric literature (Abraham and Sun, 2018; de Chaisemartin and D'Haultfoeuille, 2019; Goodman-Bacon, 2018) notes that two way fixed estimators do not properly aggregate an average of cohort-specific treatment effects in the presence of heterogeneous treatment effects by cohort. (Abraham and Sun, 2018), for example, show that in simple models where each cross-sectional unit is treated once and permanently, the two way fixed effects estimator can be represented as a weighted sum of cohort-specific treatment effects. Since early adopters are largely used as controls for late adopters, their treatment effects are usually weighted negatively. In fact, (Abraham and Sun, 2018) show that, for this reason, two way fixed effects estimates can lie outside of the convex-hull of cohort-specific treatment effects.

In our context, we show in the results section that estimates increase over time and are thus likely to be biased upwards, affirming our choice of main specification. We also assume that all treatment of partisan control happens based upon the state legislatures and governors in power in a year ending in one. In a very small fraction of cases, redistricting is done by legislators not in power in years ending in one. To the degree this is the case, our ITT estimates are likely attenuated ATE estimates by introducing measurement error in treatment. We show, however, that almost all redistricting is done by state legislatures in power in years ending in one. In return, however, all treated units are treated at the same time within a decade and thus the issues of concern for (Abraham and Sun, 2018; de Chaisemartin and D'Haultfoeuille, 2019; Goodman-Bacon, 2018) are not of concern for us.

We are further concerned that there are long run trends in partisanship, even within decades, due to realignment of the party system particularly after the passage of the 1964 Civil Rights Act (Abramowitz and Saunders, 1998; Gentzkow et al., 2016; Jensen et al., 2012; Kuziemko and Washington, 2018; Schickler, 2016). We therefore use, as a robustness check, an additional very taxing specification; we add in state specific linear time trends in the outcome variable:

$$
\begin{aligned}
O_{s,d,y} = \alpha + \sum_{k=-1}^{3} \beta_k^D L_k.DemControl_{s,y} \\
+ \sum_{k=-1}^{3} \beta_k^R L_k.RepControl_{s,y} + \gamma_{s,d} + \mu_s y + \delta_y + \epsilon
\end{aligned}
\tag{3}
$$

where $\mu_s y$ is a state-specific linear time trend.

Finally, we cluster all of our results by state. We do this for two reasons. First, clustering accounts for serial correlation within states over-time. Since popularity of parties and individual politicians persists over time, it is important to allow for serial correlation in state partisan seat shares over time within a state. Second, our data is heteroskedastic. Variances in state delegation shares are substantially higher in smaller states for mechanical reasons. However, since delegation sizes do not change much over our fifty year time period, errors are largely heteroskedastic within a state over time. Clustering at the state level accounts both for the patterns of serial correlation and the patterns of heteroskedasticity present in the data.

V. Data

A. *Vote Shares and Seat Shares*

Our main independent variable is legal control by a party over redistricting. We compute this using state partisan control data from Klarner et al. (2013). We obtain this back to 1969. This data is available through 2011. From 2012 onward we collect state partisan control data from the National Conference of State Legislatures' legislative partisan composition tables. We thus have a balanced panel of states from 1969 through 2018.

B. *Legal Control and Unified Control*

In order to determine whether a party controls the redistricting process in any given state-decade, we collect data on how redistricting is conducted. For each state, we collect the state's statutory and constitutional rules for the redistricting process, including any changes to the rules over time. We code each state-decade as one of (1.) Single district state, (2.) Legislature + Governor state, (3.) Legislature only state, or (4.) Commission. If the state has a commission, we furthermore classify it as an advisory commission if it merely provides a recommendation or a statutory commission if it has legal authority to pass a redistricting plan. We classify each commission as partisan or non-partisan depending upon whether a majority of commission members can be in a partisan manner. In our main specification, we treat all commission types as selected in a non-partisan manner. We use further sub-classifications of commission types in Table 4.

For 2000 onwards, data on rules comes from Justin Levitt's website. For the pre-2000 period, we employed a team of undergraduates to collect documents from individual state legislatures and from the National Conference of State Legislatures. We present and document our main treatment variable in the Data Appendix.

As a robustness check, we also use the data on legal control from (Friedman and Holden, 2009). This data goes from 1969 through 2004. It was assembled by Friedman and Holden based upon prior work by Cox and Katz (ICPSR 6311) and subsequent work by Gary Jacobson. Richard Holden graciously provided us with the data.

C. *Measuring the Extent of Redistricting*

In our main specification, we assume that redistricting is in effect for Congresses elected in years ending in 2. To document that this is in fact the case, we create a quantitative measure of the extent of redistricting. We use ARC-GIS to geocode every map from every state for every Congress between 1969 and 2018. When there is a change between two Congresses, we compute the geographical overlap between each pre-existing Congressional district and each new district. For each pre-existing district, we assign to it a unique new district with which it has maximum geographical overlap. We then sum over all pre-existing districts and compute the fraction of overlap as a share of all land. Thus, we compute the change in land area in every district[4]. Our results are presented in Figure 2. We show our measure separately for Congresses elected in years ending in 2 and for those elected in years not ending in 2. Moreover, we do this for both the full sample as well as the more recent sample incorporating only the past two decades. Both over the full sample and in the recent sample, almost all state-decades in

[4]We computed this measure based upon land area rather than population since Census Tracts, which are population-based, were only introduced across the entire United States for the 2000 census.

our sample do redistrict. Approximately 30% shift between 1% and 10% of their land; approximately 25% shift between 10% and 20% of their land; well over 90% of state-decades shift less than 40% of their land and all shift less than 50% of their land. By contrast, in other years, almost 100% of states have no change in district boundaries.

In Table 1, we also show results from regressions of our quantitative redistricting measure on partisan legal control. First, we compute our measure at the stateXdecade level controlling for decade and state fixed effects. Second, we show estimates at the level of a stateXyear with year and stateXdecade fixed effects. The coefficients from the decadal results are very similar to the coefficients for election years ending in 2 for specifications at the level of a Congress. The results over five decades show 6.3% more redistricted land when Republicans have control relative to no party having legal control; these results are significant with a 6% p-value. The coefficient for Democrats, by contrast, is 1.4% and statistically insignificant at conventional levels. In the past two decades, the coefficient for Republicans is larger. It is 9.8% and is statistically significant at less than a 1% level. The coefficient for Democrats is also larger; it is 4.5% but not statistically significant at conventional levels.

Of course, it is possible that Republicans have legal control and are more dominant in more rural areas where larger shifts in land do not substantively translate into larger shifts in population. Therefore, we also show the same results with a population-based rather than land-based measure of district change. These estimates show the percentage of people rather than land who switch districts as a result of redistricting. We show these population-based results in Column 6. Due to data constraints, we only show results for the past two decades. Overall, the estimates are of similar magnitude though slightly less precise.

VI. Main Results

We now present our main results. Our variation comes from comparing seat shares across state-decades where a party has legal control to other states where it does not, controlling for year effects. It is crucial for our results that we have enough instances of legal control. In Figure 3, we show four histograms which display the number of instances of partisan control separately for each party by size of state delegation. We show histograms for both Democrats and Republicans for the full sample as well as separately restricted to the past two decades. In fact, Republicans only had 12 instances of partisan control in the first three decades of our sample.

Overall, we include 212 state-decades in our sample. Out of these, we find 62 instances of Democratic control over redistricting. In contrast, we find 32 instances of Republican control. This is due to the dominance of Democrats in the earlier portion of our sample. In fact, when we restrict to the past two decades, we see 20 instances of Republican partisan control but only 14 of Democratic control. In part, this more recent Republican dominance is due to historic losses of control by the Democratic party in the 2010 elections.

In general, the size distribution of states skews slightly larger for Republican control than it does for Democratic control. This may seem somewhat surprising. However, many of the larger Democrat-dominated states either have had Republican governors (California, New York) during redistricting or have used redistricting commissions (California, New Jersey, Washington).[5]

Our main estimates are presented in Table 2. The results are split into two panels: one for the effect of Republican control and one for the effect of Democratic control. The coefficients are jointly estimated in a single regression for a given column across panels. Different columns represent different regressions, estimated with Equation(1). The first four columns show estimates from the full sample and the second four columns show estimates restricted to the past two decades. For each sample, we show results for states with more than one representative, more than two representatives, more than five representatives and more than 10 representatives respectively.[6] We do not ever include single representative states because they do not redistrict.

In the first election after redistricting, we find a 4.7 percentage point increase in a state delegation's Republican seat share following Republican control over redistricting. The effect sizes rise to 8.0 percentage points as we raise state size restrictions. These results are all statistically significant with at least a 95% level; effect sizes for recent decades are larger. In columns 5-8, we present estimates restricted to the past two decades, these estimates are larger in magnitude. In the maximal sample of states with more than one representative, Republican control leads to an 9.1 percentage point increase in a state delegation's Republican share following Republican redistricting control. Size restrictions lower coefficients but all estimates lie at 6.5 percentage points or above and are all statistically significant with a 0.06 p-value or less.[7]

In the second-to-bottom row of the panel, we average effects over the three elections following redistricting. The full sample estimates show a statistically insignificant positive coefficient of 2.8 percentage points for Republican control. Independent of size restrictions, we do not see statistically significant increases in the Republican share of a state's delegation following Republican-controlled redistricting. However, in the past two decades, we do see large and statistically significant average increase in the Republican seat share of a delegation following Republican redistricting. Our benchmark estimate yields an 8.2 percentage point increase on average over the following three elections. These estimates are statistically significant at conventional levels. Due to concerns over cluster size, we also present Wild-Cluster bootstrapped p-values in brackets for average effects; results are similar. Estimates are 6.3 percentage points in states with more than

[5]California moves to a commission system during our study period.

[6]We compute the number of representatives in a state-decade as given by the number of representatives in years ending in 3 - i.e. just after federal reapportionment.

[7]Note that the t-statistic for the 10 or more size restriction is above 2 but because of the substantial degrees of freedom correction, the p-value is 0.06.

two seats and 7.7 percentage points restricted to both states with more than five as well as more than ten seats respectively. The averages do not include effects for estimates beyond three elections because the other two elections in a decade are the baseline election (years ending in zero) and the first lead (years ending in eight).

In contrast to the effects we find of Republican control, we do not find statistically significant or large decreases in Republican seat share after Democratic-controlled redistricting. This is true whether we estimate over the full fifty year sample or whether we restrict to recent years. The reduction in the Republican seat share even in the initial election after Democratic control over redistricting is 1.8 for the full sample and 0.3 in the recent period. We do see large, statistically significant effects only in the recent period when looking at large Democratic states. Restricting to the eight instances in the past two decades with Democratic control in states with more than five seats, we find an average decrease in the Republican seat share of 8.9 percentage points; restricting to states with more than ten seats, we find a decrease of 12.6 percentage points. These estimates are statistically significant with a 95% level of confidence using conventional standard errors. Using the wild cluster bootstrap, both have a p-value less than 0.1 and the more than five restriction yields a p-value less than 0.05. Though these large estimates are estimated off of a small number of treatments and only a few coefficients are statistically significant, they do reflect evidence of partisan redistricting in large Democrat-controlled states in recent years.

In the rows entitled "Control X Election Ending in 8", we show the coefficients on partisan legal control for the election prior to the one in which the government with legal control was elected. The leads are small and statistically insignificant with the exception of our estimates restricted to states with more than five seats for our Republican control estimates and our main sample for Democratic control. The small number of statistically significant leads are consistent [8] with what one would expect by random chance.

In Table 3, we show the robustness of our results to alternative specifications and data. Our estimates are largely robust. In column 1, we repeat our baseline estimates. In column 2, we replace our legal control variable with unified control as our main treatment variable. [9] In column 3, we use the treatment variable from the Friedman-Holden data Friedman and Holden (2009). The Friedman-Holden data end in 2004 and thus we extend their cutoff year past 2004 using their method of classifying partisan redistricting.[10] The estimates for the impact of Republican control over the full sample are slightly more than double those in the main sample but remain statistically indistinguishable from zero at conventional levels. In column 4, we control linearly for the statewide vote share for House

[8] Three of sixteen estimates are statistically significant at conventional levels.

[9] Unified control is usually used in the political science literature to look at the impact of control over redistricting because of the costs of collecting the legal control variable.

[10] We do not report estimates using the Friedman-Holden data for the 2000+ time period since their actual data only covers 20% of the time period.

of Representatives races to account for time-varying political preferences of the electorate. In other words, we control linearly for the seat-share/vote-share map. We do this because we are concerned that legal control may be endogenous to partisan preference shocks at the state level. Our estimates decline slightly to 6.9 percentage points and statistical significance falls slightly to just below the 95% level of confidence. We do not see any sizable or statistically significant estimates for Democrats. For Republicans, with the exception of the Friedman-Holden estimates, our estimates lie within 1.3 percentage points of our baseline estimates.

In column 5, we present estimates in which we drop commission state-decades from our sample. The coefficient on Republican control for the recent sample rises by 3.6 percentage points when we drop states with an electoral commission for all specifications; in particular, the Republican effect in the recent sample becomes statistically significant at the 10% level. This potentially suggests that commission states redistrict in a slightly more partisan manner than states with divided government. However, the differences are not large. However, the other estimates including the estimates restriction to the prior two decades are within one percentage point of the benchmark estimates.

In column 6, we show our two way fixed effects estimates. The estimates are substantially larger though still statistically insignificant at conventional levels for Democrats. The effect size increase by 50% for Republican control in the recent period and by 300% in the full sample. Since effect sizes for Republicans are increasing over time, this is exactly what we would expect given the recent work in the new panel effects estimation literature. The differences between the two-way fixed effects design and the baseline state-decade fixed effects and year effects design precisely validate the need for our baseline design.

We then show results for our very taxing specification state-specific linear trends baseline model. We estimate this model out of concerns that state specific trends such as the realignment of the parties may induce a correlation both between legal control and increases in the dominant party's seat share. These could even happen within decades. Our estimates become less precise, likely due to over-fitting given the limited degrees of freedom. However, the estimates remain remarkably similar given the large number of covariates added. All estimates are within 1.4 percentage points of our baseline estimates.

Finally, we reclassify partisan legal control as non-partisan where the same party also had legal control in the prior decade. Since parties with ongoing legal control may not gerrymander much if their redistricting goals have already been achieved, we expect our coefficients to increase in magnitude, potentially by a sizable amount. In most cases of legal control over redistricting in our five decade sample, that same party did not have legal control in the prior decade. In column 8, we estimate on this sample of new legal control and we precisely find that all coefficients are larger in magnitude as well as of the same sign. The increase in the effect for Republicans is particularly large. Thus, our results are consistent with a time-invariant effect of new Republican control; the difference between the earlier and later time periods then is due to the greater prevalence of new Republican control in the recent period.

We also show robustness to our definition of legal control and present the results in Table 4. We consider two different types of robustness. In columns 2-3, we classify types of commissions and allow different ways of defining commissions as instances of partisan legal control. In the last three columns, we reclassify instances of a unified legislature with a governor of the opposite party as instances of partisan legal control. We do this because legislatures often can over-ride gubernatorial vetos with a strong enough majority. Moreover, legislatures can threaten governors with veto over-riding on unrelated legislation. Since veto thresholds can vary by type of bill, we consider different levels of minimal legislative majority in columns 4-6 as instances of partisan legal control by the legislature in the cases of divided government with a unified legislature.

Returning to commissions, we define commission types and then discuss our different codings for the purposes of our robustness checks. A partisan commission is a commission which can be appointed with a net partisan balance. These commissions are often appointed by the Governor or the majority leaders of the state legislatures. Some commissions are appointed in a non-partisan or a bi-partisan (i.e. balanced in partisanship) manner.[11] An advisory commission draws maps and submits them to the legislature and governor for legislative and gubernatorial approval. However, advisory commissions have no legal authority to redistrict. A non-advisory commission draw maps and the maps are automatically accepted as law. Non-advisory commissions do not need gubernatorial or legislative approval. We consider redefining commissions as being party-controlled depending upon how they are selected and whether they have the legal ability to directly implement the maps that they draw.

In the first column, we present our baseline results. Then, We consider redefining advisory commissions as instances of partisan legal control when they are appointed in a partisan manner (column 2). We also estimate a model (column 3) where we redefine legal control to include commissions when they are are merely advisory (i.e. they do not have the ability to directly implement the maps they draw).[12] Column 2 shows estimates of partisan legal control with net partisan-appointed commissions redefined as instances of partisan legal control. The estimates barely change from our baseline estimates. In column 3, we show

[11]For example, in some states, commissions are composed of five members, one member appointed by each of the majority and minority leaders in each of the two chambers. The fifth member is then appointed by a majority of the four directly appointed members.

[12]We do not show estimates based upon samples where we redefine combinations of advisory/non-advisory and partisan/non-partisan appointment as partisan appointment because all non-advisory commissions are also appointed in a bi-partisan or non-partisan manner. Thus, redefining advisory commissions as partisan legal control when government is unified is the same as redefining only advisory commissions with partisan appointment under the same circumstances. Also, redefining partisan appointment state-decades as instances of partisan legal control is akin to redefining partisan appointment with advisory commissions. Since, in all of these cases, results are identical, we limit redefinitions based solely upon changing advisory commissions as well as partisan-appointed commissions individually.

estimates with advisory commissions reassigned as partisan in the case of unified control.[13] All estimates show an increased positive impact on the Republican seat share, mostly by approximately two percentage points. These results suggest that advisory commissions do not necessarily reduce partisan redistricting though the differences in the estimates are due to a small number of state-decades and results are only sizable for the case of Republican unified control.

In the last three columns of Table 4, We also redefine legal control as partisan when the governor is a different party from the legislature but the legislature is unified and has over 60% (column 5) and over 66.7% (column 6) majority in both houses respectively. We do this because this may give the legislature the ability to pass a redistricting bill over a Gubernatorial veto. For Connecticut and Maine, which have legislative redistricting thresholds of 60% to pass a bill, regardless of gubernatorial approval, we redefine partisan control to those levels (column 4). None of these changes make a substantial difference in our estimates. Across all definitions, the variation in estimates are similarly small. We thus conclude from Table 4 that our estimates do not depend upon our particular definition of legal control.

One issue with our estimates is that observations are naturally heteroskedastic. A state-decade with two seats (i.e. Maine over the full sample) inherently changes the seat share by 0.5 when one seat changes party. However, a state-decade with 53 seats (i.e. California in the 2000s and 2010s) changes the seat share by less than 0.02 when one seat changes party. We thus re-estimate Tables 3 and 4, weighting them by delegation size. These estimates appear in Tables A1 and A2 in the Online Appendix respectively. We notice that the estimates are largely similar; however the standard errors are approximately 20% lower on average. The reduction in the standard errors is consistent with a reduction in heteroskedasticity.

Estimates of Democratic control do increase in Table A1 relative to Table 3 since larger states are more heavily weighted and effect sizes are larger for Democrats in larger states during recent period.[14] We note that since we include our baseline estimates in Tables 3 and 4, these estimates also appear in the Online Appendix Tables. Table A2, showing the weighted version of the legal control definition robustness table yields relatively similar estimates to Table 4. Also, estimates are similar across the different alternative definitions of legal control.

We additionally perform four quasi-placebos for the full sample as well as for the most recent two decades. For each sample period, these placebos lead to four placebo coefficients each for Republicans and for Democrats. In Table 5, we re-estimate effects as if redistricting were done by the state governments in power

[13]All the advisory commission states require passage of redistricting by a majority of each legislative chambers and the signature of the Governor.

[14]Additionally, unified control estimates in the recent period become substantially smaller and statistically indistinguishable from zero. Since one of the contributions of our paper is to code and use legal control, we see the robustness of our legal control measure as opposed to the traditional unified control measure as further validation of the benefits of using legal control.

in years ending in 3, 5, 7, and 9 respectively. These are not actual placebos. If legal control in the treatment year persists, we may be picking up actual lags and leads of treated effects. Thus, it is all the more striking that we find only one statistically significant coefficient for Democrats or Republicans out of the eight quasi-placebo coefficients in the full sample and none in the 2000+ sample. Moreover, the true effect estimated for Republicans in the 2000+ sample is the largest in magnitude of the 20 coefficients. That would happen by random chance if all the estimates were independent (as noted above, they are not) with a probability of 5%. Also 15 out of the 20 estimates are less than half the size of our estimate of Republican control in the past two decades.

Overall, our robustness and heterogeneity tests find substantial support for partisan gerrymandering by the Republican party in recent decades and also for the Democratic party in recent decades only for large states.

VII. Aggregate Effects

We have so far estimated the impact of party legal control over redistricting on subsequent seat shares. What has the been the impact of this upon the aggregate balance in the House of Representatives? We now translate our estimates of average seat share impacts by party into aggregate partisan effects and compare them in size to partisan seat margins.

We do this by year and party. In particular, we use decade and party specific estimates for each year for both Democrats and Republicans and compute implied seat share changes, rounding to the nearest seat. We then multiply by the number of treated states and the average number of seats in each treated state. We also note when the changes would have resulted in a shift in the balance of the House of Representatives. Analytically, we compute:

(4)

$$\Delta DemSeats_{y,d} = 2XInt\left[\beta_y^D N_{d,y}^D I\left(Control_d^D\right)\right] - Int\left[\beta_y^R N_{d,y}^R I\left(Control_d^R\right)\right]$$

where $\beta_{y,d}^P$ is the effect of party P control on a state's seat share fraction for party P in year$y$ and decade d, $N_{d,y}^P$ is the number of seats in that state's delegation in decade and year, and $I\left(Control_d^P\right)$ is a dummy which takes on a value of 1 if party P has legal control over redistricting in decade d.

We show the results of these computations in Table 6. Overall, we find little evidence of a sizable shift in partisan balance in the House of Representatives until the 2000s. Before the 2000s, net effects are no more than five seats. In the 2000s, we compute that seats shifted by 12 seats towards the Republicans and in the 2010s, we see a shift of 27 seats towards the Republican party. The reason for the small net effects through most of the past 50 years but much larger recent effects is due to a combination of two factors. First, the effect of partisan control upon seat shares has increased over time. We can see this in the difference between the effects for the 2000+ period and in the full period. We can also this by looking at the differences in partisan control in the 1970s and in 2000s. Second, state legislatures have shifted from overall Democratic dominance to overall Republican dominance. This is partly due to realignment and the shift of the South of the United States to the Republican party as well as to the poor performance of the Democratic party in the 2010 election which were critical for redistricting. This has been consequential because this switch in dominance has been from a party with a low impact of legal control on seat shares to one with a high impact of legal control on seat shares. For example, in the 2000s, the Democrats had legal control in two more states than the Republican party. Moreover, the average delegation size in both Republican and in Democratic legal control states was 13. However, because of the greater impact of Republican legal control, our estimates imply a net shift of 12 seats to the Republican party.

Interestingly, we do not find that greater imbalance in legal control over redistricting plays an important role. In fact, imbalances were much larger earlier in the sample. In the 1970s, Democrats held control in 8 more seats than the Republicans; in the 1980s, this increased to 13 and in the 1990s, it increased to 15. In the past two decades, precisely when legal control has become more consequential for net partisan balance, these gaps have fallen to 2 in the 2000s and 8 in favor of the Republicans in the 2010s.

Overall, we find an increased role for legal control over redistricting in recent decades. This increased impact is driven by two main factors: (1.) the rise in the effect of legal control and (2.) the switch from the lower impact Democratic party to the higher impact Republican party.

VIII. Conclusion

In this paper, we have shown that parties sometimes act in their own political interest by reshaping districts to increase their party's representation in Congress when they have the power to do so. The estimated size of the effects are large. In the past two decades, Republican control over redistricting has led to an increase of 8.2 percentage points in the average of a state delegation's Republican seat share in the subsequent three elections. We do not, however, find a similar effect of Democratic control except for a small number of large, Democratic states.

It may not be surprising that parties manipulate vote aggregation to benefit themselves. However, there are reasons why they might not. First, there may be a moral sense of fairness in political competition which may restrain parties from engaging in manipulative behavior. Second, parties in non-competitive environments may not feel the need to gerrymander. Third, parties in competitive states may worry about future retribution. Fourth, parties may limit themselves for fear of incurring court involvement in redistricting.[15] Unfortunately, we cannot dis-

[15]The reasons behind why partisan legal control sometimes leads to partisan redistricting and some-

tinguish between these different motives to the degree they exist. However, they provide unanswered questions for future research. Finally, though currently there is not enough sample size to look at the impact of independent commissions using our methodology, given the increasing numbers of parties who have switched to independent or bipartisan commissions, future research on their efficacy would be of great interest.

times does not are both interesting and amenable to empirical analysis but beyond the scope of the current paper.

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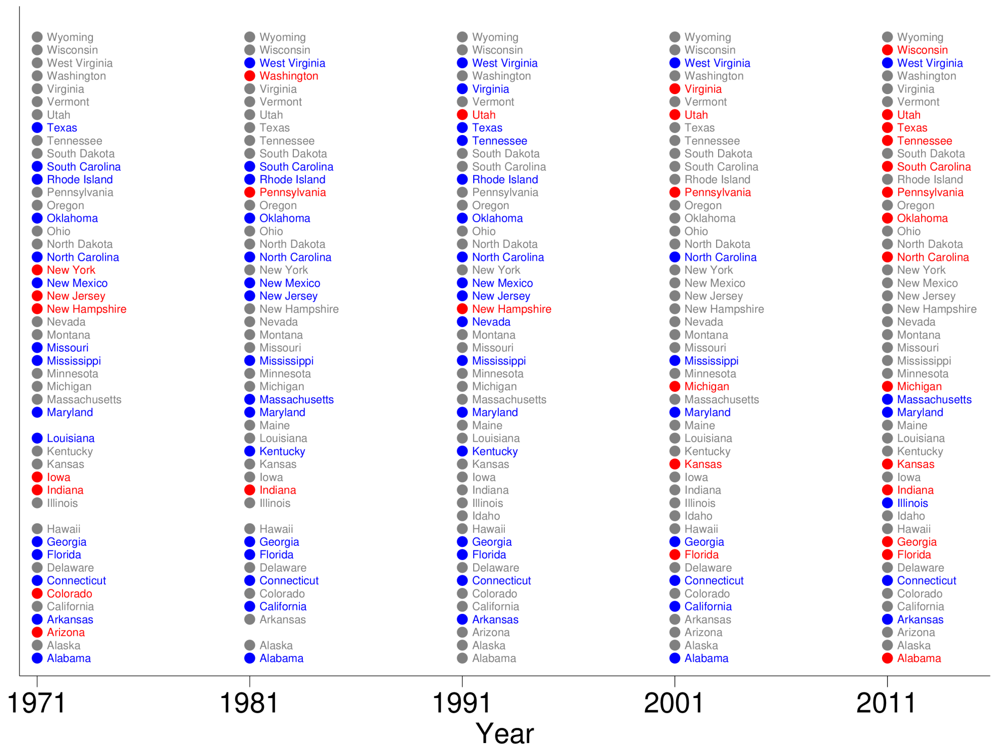
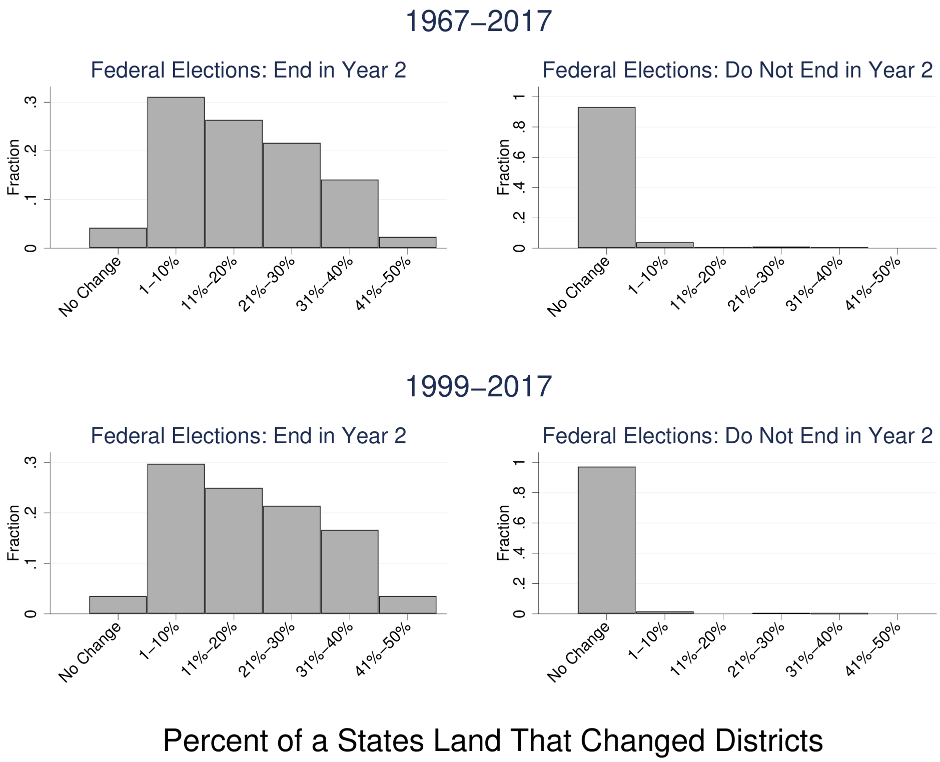
IX. Figures and Tables

FIGURE 1. STATES WITH LEGAL CONTROL



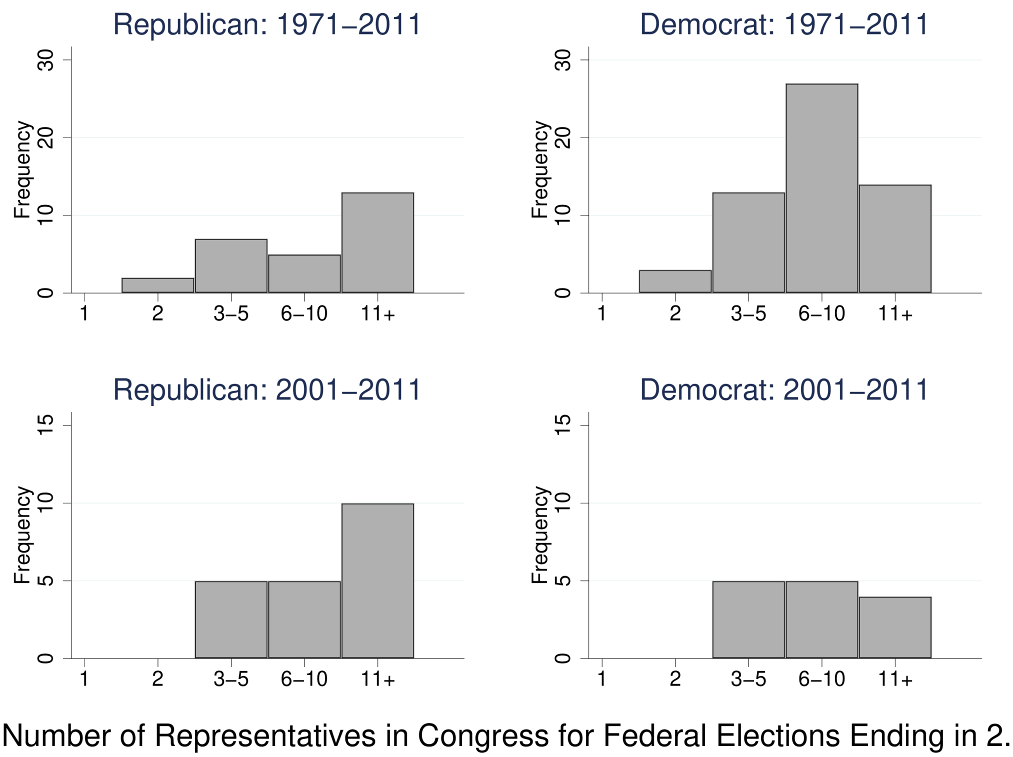
Blue dots indicate Democrat legal control over redistricting in a state-decade; red dots indicate Republican legal control; grey dots indicate no legal control.

Figure 2. Timing of Redistricting



Graphs show the distribution of percentages of land whose district changed in a state since the prior federal election. The top panel shows changes over the entire sample period. The bottom panel restricts the sample to the 1999-2017 time period. Graphs on the left show the distribution of changes for federal elections ending in 2; graphs on the right show the distribution of changes for federal elections ending in all other years.

Figure 3. Distribution of State Size with Legal Control



Graphs show distributions of delegation sizes under partisan control of redistricting. The top panel shows distributions over the entire sample period. The bottom panel restricts to the 1999-2017 time period. Graphs on the left show the size distribution under Republican control. Graphs on the right shows the size distribution under Democratic control.

Table 1—District Changes from the Prior Election Period

	(1)	(2)	(3)	(4)	(5)	(6)
Republican Control: Effect on District Change						
Control x Election Ending in 2	0.070**	0.063*	0.097**	0.084*	0.098**	0.084*
	(0.029)	(0.031)	(0.040)	(0.043)	(0.040)	(0.043)
Control x Election Ending in 4		-0.012			-0.005	-0.006
		(0.009)			(0.005)	(0.006)
Control x Election Ending in 6		-0.009			0.001	0.000
		(0.008)			(0.002)	(0.002)
Control x Election Ending in 8		-0.005			0.006	0.005
		(0.013)			(0.004)	(0.004)
Democrat Control: Effect on District Change						
Control x Election Ending in 2	0.009	0.014	0.040	0.017	0.045	0.021
	(0.022)	(0.024)	(0.040)	(0.044)	(0.042)	(0.046)
Control x Election Ending in 4		0.002			-0.010	-0.014
		(0.008)			(0.008)	(0.010)
Control x Election Ending in 6		0.005			0.015	0.016
		(0.010)			(0.020)	(0.023)
Control x Election Ending in 8		0.013			0.016	0.016
		(0.010)			(0.016)	(0.017)
Sample	All	All	2000+	2000+	2000+	2000+
Outcome Basis	Land	Land	Land	Pop	Land	Pop
Number of Observations	1060	1060	420	420	420	420
R2	0.641	0.642	0.667	0.681	0.668	0.683'

Each column display coefficients from a single regression. The dependent variable is the fraction of land within a state changing districts since the prior decade in Columns 1 and 3, the fraction of land within a state changing districts since the prior election in Columns 2 and 5, the fraction of population within a state changing districts since the prior election in Column 4 and the fraction of the population within a state changing districts since the prior election in Column 6. The treatment variable is legal control in the years ending 1. State-year level regressions are conditional upon state-decade and year fixed effects. State-decade level regressions are conditional upon decade and state fixed effects. Columns 1-2 estimate over the full sample. Columns 3-6 estimate over the 2000+ sample.

Table 2—Main Specification

	All Years				2000 Onward			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Republican Control: Effect on Republican Representative Seat Share								
Control x Election Ending in 2	0.047**	0.048**	0.063**	0.080***	0.091**	0.065***	0.076**	0.078*
	(0.023)	(0.023)	(0.027)	(0.024)	(0.035)	(0.024)	(0.028)	(0.038)
Control x Election Ending in 4	0.002	0.024	0.022	0.040	0.059*	0.054*	0.071**	0.075
	(0.043)	(0.039)	(0.040)	(0.050)	(0.034)	(0.029)	(0.029)	(0.048)
Control x Election Ending in 6	0.034	0.058	0.044	0.064	0.094	0.071*	0.084**	0.078
	(0.048)	(0.046)	(0.038)	(0.047)	(0.058)	(0.039)	(0.039)	(0.059)
Average Effect	0.028	0.043	0.043	0.061	0.082**	0.063**	0.077**	0.077
	(0.035)	(0.032)	(0.034)	(0.039)	(0.040)	(0.028)	(0.031)	(0.046)
	[0.462]	[0.196]	[0.232]	[0.113]	[0.044]	[0.050]	[0.033]	[0.116]
Control x Election Ending in 8	-0.000	-0.008	-0.048**	-0.014	0.022	0.002	-0.047*	-0.010
	(0.041)	(0.040)	(0.020)	(0.021)	(0.045)	(0.039)	(0.026)	(0.033)
Democrat Control: Effect on Republican Representative Seat Share								
Control x Election Ending in 2	-0.018	-0.009	-0.014	-0.027	-0.003	-0.022	-0.096**	-0.127**
	(0.023)	(0.021)	(0.017)	(0.023)	(0.045)	(0.041)	(0.040)	(0.042)
Control x Election Ending in 4	-0.010	0.020	0.030	0.024	-0.015	-0.019	-0.105**	-0.131**
	(0.031)	(0.026)	(0.024)	(0.029)	(0.050)	(0.049)	(0.039)	(0.053)
Control x Election Ending in 6	0.016	0.050	0.049	0.028	0.027	0.004	-0.066	-0.119**
	(0.033)	(0.031)	(0.029)	(0.039)	(0.066)	(0.052)	(0.041)	(0.051)
Average Effect	-0.004	0.021	0.022	0.008	0.003	-0.013	-0.089**	-0.126**
	(0.025)	(0.023)	(0.019)	(0.028)	(0.049)	(0.043)	(0.038)	(0.044)
	[0.846]	[0.370]	[0.270]	[0.783]	[0.942]	[0.747]	[0.069]	[0.035]
Control x Election Ending in 8	-0.055**	-0.041	-0.015	-0.016	-0.035	-0.050	0.024	0.008
	(0.026)	(0.026)	(0.018)	(0.016)	(0.051)	(0.050)	(0.025)	(0.031)
Sample	All	All	All	All	2000+	2000+	2000+	2000+
Size Restriction >	1	2	5	10	1	2	5	10
Republican Treatments	32	30	22	14	20	20	15	10
Democrat Treatments	62	57	42	18	14	14	8	4
Number of Observations	1060	900	675	295	420	370	260	120
R2	0.778	0.840	0.851	0.846	0.869	0.877	0.922	0.856

Each column presents coefficients from a single regression. Each observation is a state-year. Treatment is unilateral legal control of a political party over redistricting. The dependent variable is the state's Republican share of seats in federal House of Representatives. All specifications include state-decade and year fixed effects. Columns 1-4 use data from 1967-2017. Columns 5-8 use data from 1999-2017. Size restrictions along columns restrict the sample to states with more than the listed number of representatives. Rows show estimates of the effect of legal control. Average effect reports the average of the effects for elections for federal elections ending in 2, 4, and 6 respectively. Average effects represent the mean of the coefficients on elections year ending in 2, 4, and 6. Control X Election Ending in 8 represents the coefficient on control for elections in the year 8 from the previous decade. Standard errors clustered by state are in parentheses. P-values from the Wild Cluster bootstrap are presented in square brackets for average effects.

Table 3—Robustness and Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All Years								
Rep Average Effect	0.028	0.032	0.062	0.032	0.032	0.113***	0.036	0.096***
	(0.035)	(0.030)	(0.038)	(0.032)	(0.038)	(0.031)	(0.041)	(0.033)
Dem Average Effect	-0.004	0.000	0.007	-0.014	-0.006	-0.038	-0.007	-0.019
	(0.025)	(0.025)	(0.023)	(0.021)	(0.030)	(0.037)	(0.030)	(0.047)
Number of Observations	1060	1060	1060	1060	885	1060	1060	1060
R2	0.778	0.778	0.778	0.838	0.750	0.495	0.794	0.778
2000 and Onward								
Rep Average Effect	0.082**	0.072*		0.069*	0.108*	0.121***	0.096*	0.086**
	(0.040)	(0.042)		(0.036)	(0.056)	(0.036)	(0.051)	(0.038)
Dem Average Effect	0.003	0.002		-0.007	0.022	0.049	-0.005	0.025
	(0.049)	(0.049)		(0.040)	(0.059)	(0.073)	(0.066)	(0.068)
Number of Observations	420	420	420	420	315	420	420	420
R2	0.869	0.871	0.867	0.885	0.828	0.753	0.903	0.869
Legal Control	X							
Unified Control		X						
Holden Data			X					
Vote Share Control				X				
Exclude Commissions					X			
2-Way FE						X		
State Linear Trends							X	
Change in Control								X

Each column within a panel presents the average of coefficients from a single regression. The top panel presents coefficients over the full sample from 1969-2017. The bottom panel presents coefficients from a restricted panel spanning 1999-2017. All specifications include state-decade and year fixed effects. Treatment is unilateral legal control of a political party over redistricting. The dependent variable is the state's Republican share of seats in the federal House of Representatives. Rows show estimates of the average effect of legal control by a party on seat-shares in federal elections ending in 2, 4, and 6. Column 2 replaces legal control with unified control. Column 3 replaces our legal control variable with a similar measure due to Holden-Friedman. Column 4 adds in statewide Republican vote share in elections for the House of Representatives as a control. Column 5 excludes states with electoral commissions from the sample. Column 6 replaces state-decade fixed effects with state fixed effects. Column 7 adds state-specific linear time trends to the baseline model in Column 1. Column 8 drops decades with legal control where the same party had legal control in the prior decade. Standard errors clustered by state are in parentheses.

Table 4—Varying Definitions of Legal Control

	(1)	(2)	(3)	(4)	(5)	(6)
All Years						
Rep Average Effect	0.028	0.034	0.051*	0.031	0.026	0.023
	(0.035)	(0.034)	(0.029)	(0.035)	(0.032)	(0.035)
Dem Average Effect	-0.004	0.009	0.013	0.008	-0.018	-0.022
	(0.025)	(0.025)	(0.025)	(0.026)	(0.025)	(0.024)
Number of Observations	1060	1060	1060	1060	1060	1060
R2	0.778	0.779	0.779	0.779	0.777	0.778
2000 and Onward						
All Years						
Rep Average Effect	0.082**	0.083**	0.104**	0.081**	0.082**	0.076**
	(0.040)	(0.039))	(0.039)	(0.039)	(0.037)	(0.036)
Dem Average Effect	0.003	0.006	0.013	0.002	0.001	-0.040
	(0.049)	(0.025)	(0.049)	(0.055)	(0.025)	(0.035)
Number of Observations	420	420	420	420	420	420
R2	0.869	0.870	0.871	0.870	0.869	0.870
Baseline Legal Control	X					
Partisan Appointed Commissions Included in Treatment Pool		X				
Advisory Commissions Included in Treatment Pool			X			
Super Majority ME, CT				X		
Super Majority 60%					X	
Super Majority 66%						X

Each column within a panel presents the average of coefficients from a single regression. The top panel presents coeffcieints over the full sample from 1969-2017. The bottom panel presents coefficients from a restricted panel spanning 1999-2017. All specifications include state-decade and year fixed effects. Treatment is unilateral legal control of a political party over redistricting. The dependent variable is the state's Republican share of seats in federal House of Representatives. Rows show estimates of the average effect of legal control by a party on seat-shares in federal elections ending in 2, 4, 6. In column 2, partisan legal control includes states with commmisions with net partisan appointment. In column 3, partisan legal control treatment includes states with advisory commisions with unified legislative and gubernatorial control. In column 4, partisan legal is modified to reflect the 60% thresholds used for redistricting bills in Connecticut and Maine. In column 5, partisan legal control treatment includes states with a unified legislature and divided governor where both legislative chambers are above a 60% threshold. In column 6, partisan legal control treatment includes states with a unified legislature and divided governor where both legislative chambers are above a 66% threshold.

Table 5—Year of Control Placebos

	(1)	(2)	(3)	(4)	(5)
All Years					
Republican Control Average Effect	0.028	-0.031	-0.006	-0.032	-0.057*
	(0.035)	(0.038)	(0.023)	(0.033)	(0.034)
Democrat Control Average Effect	-0.004	-0.002	0.040	0.011	0.042
	(0.025)	(0.024)	(0.033)	(0.027)	(0.025)
Number of Observations	1060	1016	1060	1060	1016
R2	0.778	0.788	0.827	0.844	0.809
2000 and Onward					
Republican Average Effect	0.082**	-0.065	0.048	0.043	-0.041
	(0.040)	(0.076)	(0.035)	(0.031)	(0.043)
Democrat Average Effect	0.003	-0.030	-0.042	-0.012	0.021
	(0.049)	(0.048)	(0.094)	(0.070)	(0.045)
Number of Observations	420	378	420	420	378
R2	0.869	0.863	0.890	0.880	0.877 '
Year of Control	1	3	5	7	9

Each column within a panel presents averages of coefficients from a single regression. Each observations is a state-year. Treatment is unilateral legal control of a political party over redistricting in the year ending in the Year of Control listed at the bottom of the table. Column 1 shows baseline results and columns 2-5 show placebos. The dependent variable is the Republican share of seats from the state. All specifications include state-decade and year fixed effects. The top panel presents estimates over the full sample from 1969-2017. The bottom panel presents estimates from the restricted sample from 1999-2017. Standard errors clustered by state are in parentheses.

Table 6—Aggregate Partisan Effects by Decade

	1970s	1980s	1990s	2000s	2010s
States with Dem Control	15	16	17	8	6
Average Seats, Dem Control	8	10	9	13	8
States with Rep Control	7	3	2	6	14
Average Seats, Rep Control	12	14	3	13	12
Seat Share Effect: Dems	1	1	1	1	0
Seat Share Effect: Reps	5	2	0	13	27
Net Effect	4	1	1	12	27
Average Margin	95	86	62	21	53
Net Effect as % of Avg Margin	4%	1%	2%	57%	51%

Each column presents numbers for a particular decade. States with Dem Control and States with Rep Control show the number of states with Democratic and Republican legal control in the decade respectively. Average Seats is the average number of seats after redistricting in states with Democratic and Republican legal control respectively. Seat Share Effect presents a back-of-the-envelope computation of the gross number of seats gained from legal control over redistricting, broken down by party. Net effect is the absolute value of the net change in seats as a result of redistricting. Average margin is the average of the absolute value of the difference between Republican seats and Democratic seats in the Congresses elected in the years ending with 2, 4 and 6 in the decade.

Table A.1—Robustness and Heterogeneity Using Analytical Weights

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All Years								
Rep Average Effect	0.041	0.025	0.055	0.032	0.066*	0.120***	0.031	0.079**
	(0.027)	(0.029)	(0.037)	(0.020)	(0.034)	(0.030)	(0.032)	(0.033)
Dem Average Effect	0.013	-0.005	0.016	-0.012	0.018	-0.062**	0.005	-0.034
	(0.017)	(0.024)	(0.019)	(0.016)	(0.022)	(0.024)	(0.025)	(0.020)
Number of Observations	1060	1060	1060	1060	794	1060	1060	1060
R2	0.824	0.839	0.823	0.870	0.831	0.519	0.838	0.824
2000 and Onward								
Rep Average Effect	0.068*	0.022		0.057**	0.060	0.130***	0.061	0.086**
	(0.034)	(0.061)		(0.026)	(0.039)	(0.028)	(0.074)	(0.037)
Dem Average Effect	-0.043	-0.072		-0.037	-0.055	0.023	-0.020	-0.015
	(0.036)	(0.052)		(0.033)	(0.037)	(0.043)	(0.063)	(0.042)
Number of Observations	420	420	420	420	315	420	420	420
R2	0.880	0.913	0.877	0.901	0.867	0.788	0.909	0.880
Legal Control	X							
Unified Control		X						
Holden Data			X					
Vote Share Control				X				
Exclude Commissions					X			
2-Way FE						X		
State Linear Trends							X	
Change in Control								X

Each column within a panel presents the average of coefficients from a single regression. The top panel presents coefficients over the full sample from 1969-2017. The bottom panel presents coefficients from a restricted panel spanning 1999-2017. All specifications include state-decade and year fixed effects. Treatment is unilateral legal control of a political party over redistricting. The dependent variable is the state's Republican share of seats in the federal House of Representatives. Rows show estimates of the average effect of legal control by a party on seat-shares in federal elections ending in 2, 4, and 6. Column 2 replaces legal control with unified control. Column 3 replaces our legal control variable with a similar measure due to Holden-Friedman. Column 4 adds in statewide Republican vote share in elections for the House of Representatives as a control. Column 5 excludes states with electoral commissions from the sample. Column 6 replaces state-decade fixed effects with state fixed effects. Column 7 adds state-specific linear time trends to the baseline model in Column 1. Column 8 drops decades with legal control where the same party had legal control in the prior decade. Standard errors clustered by state are in parentheses.

Table A.2—Varying Definitions of Legal Control Using Analytical Weights

	(1)	(2)	(3)	(4)	(5)	(6)
All Years						
Rep Average Effect	0.041	0.047*	0.052**	0.041	0.035	0.038
	(0.027)	(0.026)	(0.025)	(0.026)	(0.025)	(0.026)
Dem Average Effect	0.013	0.018	0.019	0.015	-0.005	0.005
	(0.017)	(0.018)	(0.018)	(0.018)	(0.013)	(0.014)
Number of Observations	1060	1060	1060	1060	1060	1060
R2	0.824	0.825	0.826	0.825	0.823	0.824
2000 and Onward						
All Years						
Rep Average Effect	0.068*	0.073**	0.077**	0.068**	0.077***	0.069**
	(0.034)	(0.031))	(0.031)	(0.033)	(0.028)	(0.031)
Dem Average Effect	-0.043	-0.037	-0.036	-0.046	-0.031*	-0.058*
	(0.036)	(0.018)	(0.036)	(0.038)	(0.018)	(0.032)
Number of Observations	420	420	420	420	420	420
R2	0.880	0.882	0.883	0.881	0.889	0.882
Legal Control	X					
Partisan Appointed Commissions Included in Treatment Pool		X				
Advisory Commissions Included in Treatment Pool			X			
Super Majority ME, CT				X		
Super Majority 60%					X	
Super Majority 66%						X

Each column within a panel presents the average of coefficients from a single regression. The top panel presents coeffcieints over the full sample from 1969-2017. The bottom panel presents coefficients from a restricted panel spanning 1999-2017. All specifications include state-decade and year fixed effects. Treatment is unilateral legal control of a political party over redistricting. The dependent variable is the state's Republican share of seats in federal House of Representatives. Rows show estimates of the average effect of legal control by a party on seat-shares in federal elections ending in 2, 4, 6. In column 2, partisan legal control includes states with commmisions with net partisan appointment. In column 3, partisan legal control treatment includes states with advisory commisions with unified legislative and gubernatorial control. In column 4, partisan legal is modified to reflect the 60% thresholds used for redistricting bills in Connecticut and Maine. In column 5, partisan legal control treatment includes states with a unified legislature and divided governor where both legislative chambers are above a 60% threshold. In column 6, partisan legal control treatment includes states with a unified legislature and divided governor where both legislative chambers are above a 66% threshold.

X. Data Appendix

We compile a novel data set on the legal rules that states use to create Congressional district lines from 1968 to 2012. We coded types of legal systems for redistricting across states over 5 decades. We grouped state-decades into one of six categories: (1.) Single district states not eligible for redistricting, (2.) States where redistricting bills are passed by state legislatures and are not subject to a Gubernatorial veto, (3.) States where redistricting bills are passed by state legislatures but where the Governor has veto rights, (4.) States where potentially-partisan advisory commissions (i.e. commissions that are not appointed in a bi-partisan or non-partisan manner) draw the maps but the legislature needs to pass a redistricting bill in order for it to become law, (5.) States where advisory commissions, appointed in a non-partisan or balanced partisan manner, draw the maps but the legislature needs to pass a redistricting bill in order for it to become law, and (6.) States with an independent commission which is appointed in a non-partisan or bi-partisan manner and which has the legal authority to implement a redistricting plan without legislative or gubernatorial approval.

In the 2000+ time period, we rely upon descriptions from Justin Levitt's website: `https://redistricting.lls.edu/2010districts.php`. In the pre-2000 period, we rely upon a combination of sources. First, the National Conference of State Legislatures has documented all historical commissions. Second, we rely upon state legislative documents for each non-single-district state. Third, we rely on law.justia.com. Finally, we also make use of academic articles in some cases. Our sources are documented in greater detail in: `https://docs.google.com/spreadsheets/d/1nZuugxJe09PfCHVIsLyXjGx5cnlKTVnvtDYivDNFdiM/edit?usp=sharing`.

In this document, we point out general patterns, a few anomalies and coding decisions. Most states are of the legislative + gubernatorial veto type. Only Connecticut and North Carolina do not allow for a Gubernatorial veto. In addition, two states, Connecticut and Maine, set a 2/3 majority threshold for passage of a redistricting bill. Five states are one-district states throughout the five-period decade spanning our data. Two others, Montana and South Dakota, start as 2-district states and change to a 1-district state during our time span, while Nevada starts as a 1-district state and eventually reaches 4-districts in our time span. Some states transition to commission states during the time period spanned by our data. However, no states revert from a commission back to legislative redistricting. Montana does transition from a commission state to a 1-district state. For our main specification, we code any state with a commission of any type (4, 5 or 6) as not having legal control by either party. We show robustness to redefining commissions of type 4 as under partisan legal control depending upon the composition of the state legislature in Appendix Table 1.

For all states, we estimate an intention to treat estimate. Thus, we code based upon the law for the decade that was in place in years ending in 1 when redistricting normally happens. Hawaii, in 1968, passed a constitutional amendment which called for redistricting in 1969, 1973 and then every ten years starting in 1981. It also called for a commission system as of 1973. We thus code Hawaii in the 1970s as a commission state.

Table D.A1—State-Level Congressional Redistricting Laws By Decade

State	1970s	1980s	1990s	2000s	2010s
Alabama	3	3	3	3	3
Alaska	1	1	1	1	1
Arizona	3	3	3	6	6
Arkansas	3	3	3	3	3
California	3	3	3	3	6
Colorado	3	3	3	3	3
Connecticut	2	2	2	2	2
Delaware	1	1	1	1	1
Florida	3	3	3	3	3
Georgia	3	3	3	3	3
Hawaii	6	6	6	6	6
Idaho	3	3	3	6	6
Illinois	3	3	3	3	3
Indiana	3	3	3	3	3
Iowa	3	5	5	5	5
Kansas	3	3	3	3	3
Kentucky	3	3	3	3	3
Louisiana	3	3	3	3	3
Maine	3	5	5	5	5
Maryland	3	3	3	3	3
Massachusetts	3	3	3	3	3
Michigan	3	3	3	3	3
Minnesota	3	3	3	3	3
Mississippi	3	3	3	3	3
Missouri	3	3	3	3	3
Montana	6	6	1	1	1
Nebraska	3	3	3	3	3
Nevada	1	3	3	3	3
New Hampshire	3	3	3	3	3
New Jersey	3	3	3	6	6
New Mexico	3	3	3	3	3
New York	3	4	4	4	4
North Carolina	2	2	2	2	2
North Dakota	1	1	1	1	1
Ohio	3	3	4	4	4
Oklahoma	3	3	3	3	3
Oregon	3	3	3	3	3
Pennsylvania	3	3	3	3	3
Rhode Island	3	3	3	3	4
South Carolina	3	3	3	3	3
South Dakota	3	1	1	1	1
Tennessee	3	3	3	3	3
Texas	3	3	3	3	3
Utah	3	3	3	3	3
Vermont	1	1	1	1	1
Virginia	3	3	3	3	3
Washington	3	3	6	6	6
West Virginia	3	3	3	3	3
Wisconsin	3	3	3	3	3
Wyoming	1	1	1	1	1

Note: Numbers represent different legal systems for redistricting: 1: Single District - The state was apportioned a single congressional district and thus there was no need for districting. 2: Legislature Only - The State Legislature has full control over the redistricting process with no possibility of a Gubernatorial veto. 3: Legislature and Governor: The State Legislature is in charge of developing a Congressional Redistricting plan but the Governor has veto rights. 4: Advisory Commission: An advisory commission draws redistricting maps and presents them to the legislature for passage. Advisory commissions of this type are appointed in a manner that lacks partisan balance. 5: Non-Partisan Advisory Commission: An advisory commission which is appointed in a non-partisan manner or on a bi-partisan basis so as to maintain partisan balance on the commissions. 6: Independent Commission - Independent commissions are appointed on a non-partisan basis and have the legal authority to draw and implement a redistricting plan without gubernatorial or legislative approval. For the 2000s and 2010s redistricting cycles data was collected from a website by Justin Levitt. For the 1980s and 1990s cycles the majority of the data came from court cases whose summaries were aggregated by the National Conference of State Legislatures website. The full documentation of the cases were then examined, often via law.justia.com. For the 1970s redistricting cycle, a variety of sources were used. The primary ones were state specific sites either documenting the history of redistricting in the state or documenting historical state constitutional amendments as well as a paper on the 1970s redistricting cycle in which the processes were characterize