

# Unequal Incomes and Partisan Replacement: How Inequality Causes Polarization and Conservatism\*

Ayse Eldes<sup>†</sup>, Nolan McCarty<sup>‡</sup>, Boris Shor<sup>§</sup>, and John Voorheis<sup>¶</sup>

March 5, 2026

## Abstract

Income inequality and political polarization have both increased dramatically in the United States over the last several decades. A recent literature has suggested that these two phenomena may be related and reinforcing: income inequality leads to polarization, and the gridlock induced by polarization reduces the ability of politicians to alleviate rising inequality. Scholars, however, have not credibly identified any underlying causal relationship. Employing a simulated instrumental variables identification strategy with newly available data on polarization in state legislatures and state-level income inequality allows us to obtain credible causal estimates of the effect of inequality on polarization within states. We find that rising income inequality has a large, positive, and significant effect on polarization. Increasing inequality also moves legislatures to the right by replacing moderate Democrats with Republicans. We also explore the extent to which the relationship of inequality varies with respect to campaign finance law, racial heterogeneity, and economic shocks.

Word Count: 9468

**Keywords:** polarization, income inequality, state legislatures, ideology

---

\*This research was supported in part by a grant from the Washington Center for Equitable Growth. We thank Project Votesmart for access to NPAT survey data. The roll call data collection has been supported financially by the Laura and John Arnold Foundation, the Russell Sage Foundation, the Princeton School of Public and International Affairs, the Robert Wood Johnson Scholars in Health Policy program, and NSF Grants SES-1059716 and SES-1060092. Special thanks are due to Michelle Anderson and Peter Koppstein for the data collection efforts. We also thank the following for exemplary research assistance: Steve Rogers, Michael Barber, and Chad Levinson. Party data for Nebraska is included thanks to Seth Masket, who generously provided the informal partisan affiliations for Unicameral legislators. We thank the Sunlight Foundation and the OpenStates project for providing public access to its roll call data for 2011-2014. We also thank seminar participants at Institute for Advanced Study Toulouse, Princeton, UC-Davis, and Stanford GSB.

<sup>†</sup>Department of Politics, Princeton University [aeldes@princeton.edu](mailto:aeldes@princeton.edu)

<sup>‡</sup>Princeton School of Public and International Affairs, Princeton University; [nmccarty@princeton.edu](mailto:nmccarty@princeton.edu)

<sup>§</sup>Department of Political Science & Hobby School of Public Affairs, University of Houston; [bshor@uh.edu](mailto:bshor@uh.edu)

<sup>¶</sup>US Census Bureau; [john.1.voorheis@census.gov](mailto:john.1.voorheis@census.gov)

<sup>||</sup>This work was completed while John Voorheis was affiliated with the University of Oregon, and represents the views of the authors. It does not necessarily reflect the view of the US Census Bureau.

# 1 Introduction

Political polarization is one of the most widely discussed transformations of the American political economy. The ideological distance between the two major political parties has increased substantially since the 1970s.<sup>1</sup> This rise coincides with a dramatic increase in income inequality during the same period. Because in the American constitutional system polarization tends to increase gridlock and to decrease in legislative capacity, political polarization has been blamed for a decline in the ability of governments to respond to the observed increase in inequality. Thus, polarization may contribute to the propagation of inequality over time, even as polarization may itself be partly caused by increases in inequality.

Previous analyses of the potential relationship between income inequality and political polarization have not credibly identified causal effects in either direction, despite the fact that there is a wealth of theory suggesting that there could be a causal relationship. The use of data on state-level income inequality (Voorheis, 2014), state legislative political polarization (Shor and McCarty, 2011, 2022), and an underutilized identification strategy (a variation of Boustan et al. 2013) allows us to identify the causal effect of state income inequality on state legislative political polarization.

We find that income inequality has a statistically significant, positive, and quantitatively large causal effect on political polarization. We find evidence that within-state inequality has a roughly symmetric effect on the mean position of the Democratic and Republican parties. We extend the analysis to consider how income inequality affects the partisan balance of the legislature. We find strong evidence that income inequality shifts the mean ideology within state legislatures to the right and moves average party ideologies away from the center. We also find evidence that inequality increases the share of seats held by Republicans, supporting a mechanism that primarily works as a composition effect. Moderate Democrats are replaced by Republicans, leaving behind a more liberal Democratic party. Our findings suggest that the most conservative wings of the Democratic and Republican parties drive

---

<sup>1</sup>For a review of the evidence on these trends, see McCarty (2019).

these results, changing average chamber ideology to the right through the replacement of moderate Democrats.

We also explore whether institutional changes in campaign finance law create heterogeneity in income inequality's effect on polarization and average chamber ideology. We specifically focus on the *Citizens United v. FEC* Supreme Court decision, which struck down federal and state-level restrictions on independent expenditures. By allowing the effect of income inequality on polarization to systematically vary depending on whether pre-*Citizens United* expenditure restrictions were in place, we can infer whether campaign finance is a causal channel from inequality to rightward shifts in legislatures. We find, consistent with this hypothesis, that removing restrictions on independent expenditures in 2010 increased the effect of income inequality on polarization and average chamber ideology in those states.

We further explore heterogeneity by considering whether the effects of income inequality are contingent on the structural economic change induced by trade shocks, or the varying racial composition of states. We do not detect heterogeneity in the effects of income inequality associated with these factors.

Section 2 briefly surveys the relevant literature on inequality and polarization. Section 3 describes our data and identification strategy for estimating causal effects of state-level income inequality on legislative polarization. We present and discuss our empirical results as well as a series of robustness checks in Section 4. Lastly, Section 5 evaluates whether the effects we identify are heterogeneous with respect to campaign finance law, trade shocks, and racial and ethnic diversity.

## 2 Previous Literature

Scholars generally agree that the U.S. Congress has polarized significantly over the past several decades. Based on the most frequently used measures of congressional polarization, such as those derived from the DW-NOMINATE measures of congressional ideology (Poole

and Rosenthal 1997), the recent rise in polarization began in the late 1970s.

The causes of rising legislative polarization remain far from certain. The literature has been far more successful in ruling out potential causes than in providing a well-supported causal model (Barber and McCarty, 2015; McCarty, 2019). Many journalists and practitioners have argued that features such as gerrymandering, partisan primaries, and under-regulated campaign finance are important contributors to polarization. But little systematic evidence has been produced to support such claims.<sup>2</sup>

In contrast, there is growing evidence that at least some features of our campaign finance system may be connected to increasing polarization. Barber (2016) provides evidence that polarization has grown less in states that limit individual campaign contributions and more in states that limit contributions from corporations and labor unions. Similarly, La Raja and Schaffner (2015) find more polarization in states that limit the role of political parties in campaign finance as candidates in those states are more reliant on ideological individuals and interest groups. Studies of public financing of elections show conflicting evidence of effects on polarization (e.g. Hall (2014) and Masket and Miller (2015)).

Some scholars have explored links between large-scale economic and social change on polarization. A prominent set of hypotheses for rising polarization focuses on the coincident rise of income and wealth inequality since the 1970s.<sup>3</sup>

---

<sup>2</sup>An important role for gerrymandering is undermined both by its inability to explain polarization in the U.S. Senate and by more extensive quantitative analyses. While Theriault (2008) reports small effects of redistricting on polarization, several studies find none. See Mann (2006); McCarty, Poole and Rosenthal (2006, 2009); and Masket, Winburn and Wright (2012). Arguments about the role of partisan primaries in the nomination of extreme candidates have not fared well empirically. Hirano et al. (2010) find that the introduction of a primary had no effect on polarization in the Senate and refute that primaries have become polarizing because turnout has fallen. Additionally, arguments about partisan or “closed” primaries have been rejected by statistical analyses (e.g. McGhee et al. (2014).)

<sup>3</sup>There is considerable debate about the magnitude of the rise of inequality over the past 50 years. Broad-based survey measures of inequality generally such as the ones used in this paper generally show substantial increases over this period (see Congressional Budget Office (2026)). There is less consensus when using tax return based measures to compute top income shares. Piketty and Saez (2003) and Piketty, Saez and Zucman (2018) show dramatic increases in top income shares concentrated at the very highest incomes centiles. But Auten and Splinter (2024) argue that these estimates are sensitive to a variety of assumptions about how to allocate several sources of income not included on tax forms such as underreported individual and business incomes, certain government transfers, and retirement account rollovers. Based on a different set of assumptions and correcting for some errors, Auten and Splinter (2024) find that the increase in the pre-tax Top 1% income share from the 1970s is sharply attenuated and that the after-tax share increased by

Not only have polarization and inequality risen in tandem over the past forty years, but their respective measures declined together during the first part of the 20th Century before leveling off after World War II. McCarty, Poole and Rosenthal (1997) were the first to observe a strong correlation between the time series of income inequality and political polarization. Duca and Saving (2012) and Duca and Saving (2016) extend this time-series-based analysis via a more rigorous treatment of the time-series properties of the data on polarization and inequality. Although these empirical correlations are striking, they do not necessarily indicate whether the correlation represents a causal effect, nor does it reveal in which direction any causality might run.

We must be concerned about reverse causality when polarized legislatures produce policies that promote inequality or gridlock in ways that facilitate further growth of inequality. The second issue with assessments based solely on national time series evidence is omitted variables. The 20th century witnessed any number of large-scale social, political, and economic changes, including the rise and fall of large-scale immigration; Great Depressions and Recessions; wars (popular and unpopular); and civil rights movements. Importantly, many of these trends co-vary with polarization and inequality or represent events at or near the inflection points in the two series. Thus, parsing the specific contribution of inequality is difficult.<sup>4</sup>

Despite the challenges to causal identification faced by observational analyses of the relationship between inequality and polarization, the observed correlation between the two has been recognized as a “stylized fact” about the contemporary American political economy. A small but growing number of theoretical models seek to explain such a relationship. Recent examples include Ma (2014), Feddersen and Gul (2014), and Vlaicu (2016). One implication of the model by Feddersen and Gul (2014) is that income inequality simultaneously moves the mean ideology of a legislature to the right while increasing political polarization.

---

around one percentage point.

<sup>4</sup>Of course, researchers who model the effect of inequality on polarization using only cross-sectional data (e.g. Garand 2010, Gelman, Kenworthy and Su 2010) have a similarly difficult task in attempting to identify causal effects.

Political polarization has been linked to a number of negative policy consequences. Polarization increases gridlock and reduces the ability of legislatures to enact policies (McCarty 2007; Binder 2015) and to modernize statutory frameworks (Mettler 2016). This is especially salient for the political system in the United States, which requires legislation to pass through multiple veto points, a status-quo bias with potentially negative consequences in the faces of changing circumstances. Many states, too, require super-majorities for the passage of certain bills (such as annual state budgets or tax increases). Polarization may thus serve as a mechanism for “political reinforcement” (Barth, Finseraas and Moene 2014). It reduces the capacity of legislators to (a) enact policies which might constrain increases in inequality (e.g. increases in the minimum wage, strengthening union bargaining power), (b) engage in redistribution that reduces inequality in disposable incomes or (c) reform welfare state institutions (Hacker 2005). The positive feedback effect of income inequality on polarization may thus lead to further increases in income inequality. The possibility of such a feedback loop from inequality to polarization to further inequality provides strong motivation for a careful study of these relationships, but also suggests very real empirical challenges for identifying causal effects at any given link in this chain.

The states provide a useful observational setting for studying polarization due to the vast increase in statistical power as opposed to studying one Congress. Until recently, scholars have been unable to measure whether or not similar trends in polarization are present there, given a lack of roll call data and a method to measure ideology on a common scale. Shor and McCarty (2011, 2022), however, have developed measures of state legislator ideology that can be used to measure party positions and polarization over time for the fifty states.

## 3 Data and Identification Strategy

### 3.1 Inequality Data

Until recently, reliable data on income inequality in the United States has been available only at the national level. It is difficult to measure income inequality at sub-national geographies due to censoring of individual incomes in the publicly available micro-data. Such individual-level income micro-data are available from two sources in the United States: tax returns and responses to Census Bureau surveys (chiefly the Current Population Survey or CPS). These data are either censored geographically for privacy purposes (as is the case with the public-use IRS files) or too sparsely distributed to produce credible estimates (as is the case for the CPS for geographies smaller than metropolitan areas). Additionally, the censoring of top incomes by the Census Bureau complicates estimation of inequality even for geographies that have adequate coverage (e.g. states and metropolitan statistical areas.)

Aside from these data quality issues, the different data sources are not equally suited to the calculation of various inequality measures. The IRS data are extremely rich but cover only the population of tax return filers, not the full population of income earners. Filing rates increase with income, so this means that the IRS data are ill-suited to make statements about the entire income distribution.<sup>5</sup> The inequality growth observed when using overall measures like Gini explain why we continue to see an increase in polarization. In our findings, the relevance of those affected by redistributive policies becomes clear – a point that would remain obscured if we did not measure inequality among the general population.

The IRS micro-data also describe a relatively limited definition of income-taxable income accruing to “tax units.” Census Bureau micro-data, on the other hand, are nationally representative samples of the entire US population, not just tax filers. Census Bureau data are therefore better able to recover estimates of income inequality in the population of

---

<sup>5</sup>For this reason the literature that utilizes IRS tax return data (such as Piketty and Saez (2003) and Auten and Splinter (2024)) focuses almost exclusively on top income shares as a measure of income inequality as opposed to indices such as the Gini, Theil and Atkinson which are sensitive to the entire distribution. Tax return data are also sensitive to changes in tax law.

all income earners. Additionally, the Census Bureau micro-data contain rich detail about household structure and non-taxable income sources. It is then possible to use a definition of income (i.e. pre-tax, post-transfer, size-adjusted household income) that more closely aligns with potential consumption than does the income definition in the IRS tax-return data.

A substantial literature has sought to take advantage of the conceptual advantages of Census Bureau income data while addressing its chief drawback—censoring (by the Census Bureau) and under-reporting (by individual respondents) of top incomes. Voorheis (2014) is the first study to provide state-level data on income inequality using Census Bureau micro-data that addresses both censoring and potential under-reporting. This correction is performed by modeling the right tail of the income distribution as following a Generalized Beta II (GB2) distribution. Censored (top coded) incomes and incomes above the 97.5th percentile are replaced by draws from the fitted GB2 distribution in a multiple imputation process. Jenkins et al. (2011) show that this method can closely match inequality trends estimated using uncensored, confidential CPS data, and Voorheis (2014) shows that this method can match the levels and trends in state-level top income shares estimated using public-use IRS data. We have followed the Voorheis method to compute Gini indices through 2022.

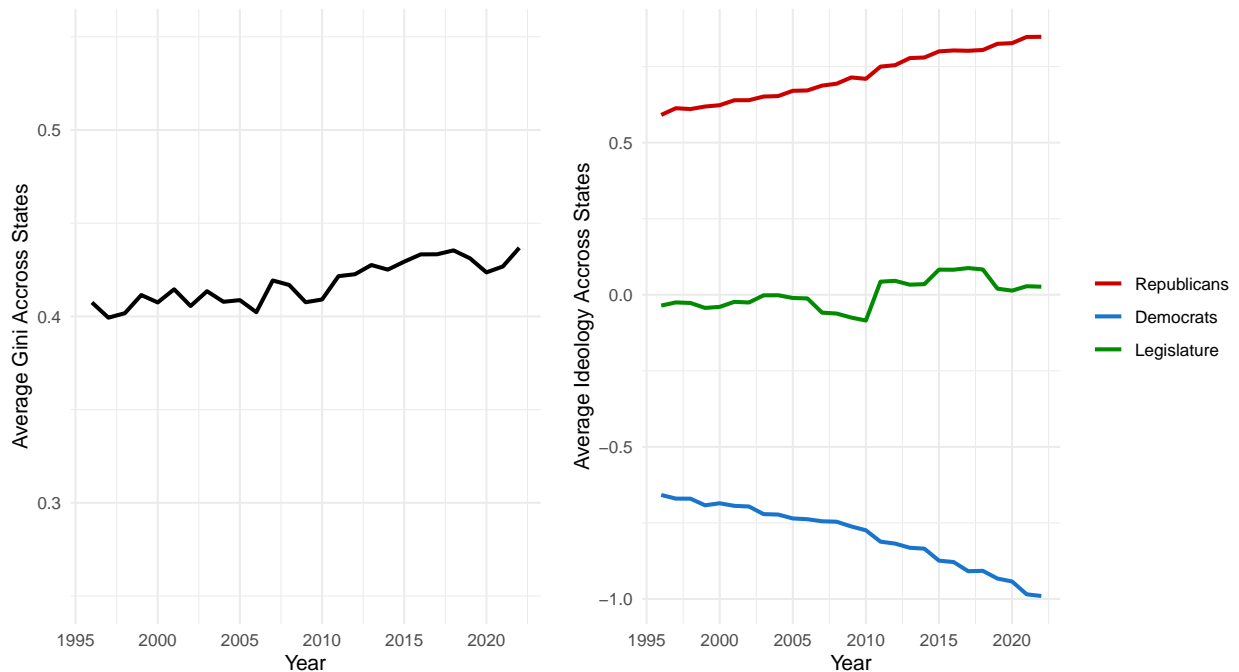
### **3.2 Polarization and Ideology Data**

Empirical spatial models of roll-call votes have become commonplace in political science. These models assume that legislators have symmetric and single-peaked preferences along a latent dimension that is often interpreted as ideological. Under the assumption that legislators vote for their most preferred outcomes, statistical procedures can recover their most-preferred outcome or *ideal point*. Intuitively, legislators who typically vote together have ideal points that are close together, and legislators who rarely vote together have ideal points that are far apart. However, ideal point measures of ideology are only comparable for legislators who vote on a common set of roll calls. This implies that ideology measures can be

estimated only for a single legislative body.<sup>6</sup> Consequently, it has been difficult to develop ideology measures for state legislators that are comparable across states, since legislative agendas differ radically, and there may be no common roll calls.

Shor and McCarty (2011) provide a solution to this problem, however, by using the Project Vote Smart National Political Awareness Test (NPAT), an annual survey of federal and state legislative candidates that has been fielded since the mid-1990s as a way to bridge across legislatures.<sup>7</sup> Using these data, Shor and McCarty (2011) are able to put all state legislators on a common scale and estimate ideology scores for state legislators which are comparable both over time and across states.

Figure 1: Trends in income inequality and political ideology



The updates reported in Shor and McCarty (2022) provide the measures of polarization for this study. The underlying data contains estimates of individual legislator ideal points

<sup>6</sup>Making comparisons over time is facilitated by the overlapping memberships of succeeding legislatures.

<sup>7</sup>The NPAT has subsequently been renamed the “Political Courage Test,” although the survey methodology and questions remain the same. This survey has been used by other scholars to characterize candidate ideology. See, for example, Ansolabehere, Snyder and Stewart (2001*b,a*). More information about the survey is available at: <http://votesmart.org/about/political-courage-test>.

for the vast majority of legislators who held office from 1993-2022. These data are cross-sectional, providing a single average measure capturing the ideology of each legislator which is constant over the course of her legislative career. Consequently, changes in chamber-level ideology are generated only from legislator turnover, and not from changes in individual legislator ideology over time.<sup>8</sup> These data are aggregated to the state-chamber level to produce estimates of the mean ideal point of each party and the overall mean ideology for the chamber. We measure polarization as the difference between the mean ideal points of the Democratic and Republican parties within a legislative chamber.<sup>9</sup>

All states except Nebraska have bicameral legislatures, so we must aggregate scores from two chambers to obtain a state-level measure. We average the polarization measures from the two chambers in each state. Similarly, we use the average, across the two chambers, of the two Democratic and Republican party means to capture asymmetric polarization effects. Figure 1 displays the average Gini coefficient and ideology across all state legislatures from 1996 to 2022 in addition to the within-party ideology scores. As documented by Shor and McCarty (2022), Democrats have moved left somewhat faster than Republicans in state legislatures in this time period.<sup>10</sup>

To complement the two main state-level datasets, we obtain state-level demographic and aggregate economic data from the CPS and Bureau of Economic Analysis (BEA) National Income and Product Accounts (NIPA) tables. These measures include population density, racial composition (the proportions of black and Hispanic residents), education (proportion of the population with a college degree), poverty rates, median income, median age, the proportion of the population under 25 and over 55, and union membership rates. In all, our data covers state-level inequality, state-level polarization, and state demographics for the

---

<sup>8</sup>The assumption that within-legislator movement is small is empirically well documented. See, for example, Poole (2007).

<sup>9</sup>We use means instead of medians as the former are more likely to capture changes at the extreme ends of each party distribution. In Appendix A.4.6, we show that our models return substantively similar results using medians.

<sup>10</sup>Though this may in part be to a composition effect that we discuss below where moderate Democrats have been replaced by conservative Republicans.

period of 1993 to 2022.<sup>11</sup>

### 3.3 Identification Strategy and Empirical Model

Our basic model is

$$Polar_{i,t} = \alpha_i + \beta Gini_{i,t} + \gamma X_{i,t} + \mu_i t + \epsilon_{i,t} \quad (1)$$

where  $\alpha_i$  is a state-fixed effect,  $Gini_{i,t}$  is the income inequality measure for state  $i$  in year  $t$ ,  $X_{i,t}$  is a vector of time-varying state-specific covariates, and  $\mu_i$  is state-specific linear trend.<sup>12</sup> However, in our main specification, we control for any unobserved but non-time-varying heterogeneity by transforming equation 1 into a first-difference model:

$$\Delta Polar_{i,t} = \beta \Delta Gini_{i,t} + \gamma \Delta X_{i,t} + \mu_i + \Delta \epsilon_{i,t} \quad (2)$$

But in the appendix, we also estimate equation 1 directly.<sup>13</sup> Note, however, that estimating either version via OLS will not generally recover the true effect of income inequality on polarization, since there is likely time-varying endogeneity between income inequality and political polarization.

There are three sources of bias that may occur. First, there could be non-random locational sorting of households into more-polarized or less-polarized states based on income. If this residential sorting varies systematically with income, polarization may mechanically

---

<sup>11</sup>The number of observations (651) is less than 750 (50 states and 15 bienniums) due to first differencing, missing ideal point estimations, and off-year elections. Four states have off-year elections, reducing the number of biennial values due to staggered start-years. Our findings are robust to the inclusion of these off-year states and to using a larger dataset with annual observations. Lastly, there is substantial missingness before 1996. Missing values occur when a state does not make roll call vote data available for a particular year. We treat these instances as missing-at-random. We also lose one observation per state due to the first-differencing described below.

<sup>12</sup>The state-specific trends capture the well-known fact that both polarization and inequality are trending upward. We do not, however, want to impose a uniform trend as the growth rates vary significantly across states. In this regard our design is conservative in that it exploits neither cross-state variation in the levels or trends in polarization and inequality.

<sup>13</sup>In the results presented in the main body, we use one observation per state biennial term. We also report the results when we use annual data (as polarization measures can change substantially in off-years due to special elections) in the appendix.

affect state-level income distributions. The direction of this effect is uncertain, however. Whether this process increases or decreases measured inequality over time depends on the relative sizes of the flows at the bottom and top of the income distribution. Second, the causal effect could work in the other direction. More-polarized legislatures, compared to less-polarized legislatures, may enact (or fail to enact) policies that affect the income distribution (either increasing or decreasing inequality). However, such policy effects may be less important in practice at the state level, since almost all tax-and-transfer redistribution occurs at the federal level.

Finally, there may be measurement error bias if income inequality is mismeasured. If any of these effects are present, then the apparent effect of income inequality on polarization revealed by estimating either equation model via OLS is biased. The direction of the bias is uncertain, however. Policy effects inflate the estimates, measurement error bias estimates towards zero, and the net direction of any locational sorting bias is uncertain.

We propose an instrumental variables estimation strategy that is robust to all three sources of bias outlined above. We adapt an instrument proposed by Boustan et al. (2013) and use the GB2 multiple imputation approach from Voorheis (2014) to address censoring and under-reporting in the micro-data. The instrument is constructed by “freezing” the baseline income distribution in each state at some initial year, and then simulating the income distributions for each subsequent year based on nationwide trends in income growth at each decile. This instrument is one example of so-called “Bartik-style” instruments.<sup>14</sup> The identifying variation in this type of instrument comes from the cross-sectional variation in the initial level of income inequality across states. The identifying assumption of this simulated instrumental variables strategy thus amounts to an assertion that the initial level of income inequality is unrelated to subsequent changes in the outcome variable (in our case, political polarization or other measures of ideology).

We construct our instrument as follows. We select the 1990 income distribution as the

---

<sup>14</sup>Baum-Snow and Ferreira (2015) summarizes the theory and practice of using Bartik-style instruments in a variety of settings.

baseline for each state.<sup>15</sup> We estimate average incomes for each centile in this initial year, using the Voorheis (2014) GB2 imputation method. We estimate the growth rates of the average incomes of each centile of the *nationwide* income distribution for each year from 1990 through the end of our estimating sample (2022), again using the GB2 imputation method to calculate average centile incomes for each year.

We then simulate state-level income distributions for each year between 1993 and 2022 as follows. We assign each state centile in the initial year to the matching nationwide centile. We then simulate state-level income distributions for each year by assuming each state centile grows at the matching nationwide centile’s growth rate for that year. Finally, we construct our instrument for income inequality by calculating the Gini coefficient using the simulated centile incomes in each year.

Using the simulated Gini instrument, we can then estimate the effect of income inequality on polarization by two-stage least squares. To strengthen the causal interpretation, we lag both the instrument and the Gini coefficient by one year, ensuring that inequality is measured prior to observed political outcomes. In the appendix, we report placebo tests using contemporaneous and forward-lagged values of the instrument and Gini coefficient, which confirm that our main results are not driven by reverse causality or contemporaneous feedback between inequality and polarization. Our preferred specification is:

$$\text{First Stage: } \Delta Gini_{i,t-1} = \tilde{\mu}_i + \theta \Delta Inst\_Gini_{i,t-1} + \Gamma \Delta X_{i,t} + \nu_{i,t} \quad (3)$$

$$\text{Second Stage: } \Delta Polar_{i,t} = \mu_i + \beta_{2SLS} \Delta \widehat{Gini}_{i,t-1} + \gamma \Delta X_{i,t} + e_{i,t} \quad (4)$$

---

<sup>15</sup>We have also experimented with other years in the range 1988-1992. We settle on 1990 as the baseline year since it produces the strongest instrument (i.e. the instrument with the largest first-stage F-test statistic).

We also estimate models with level data as a robustness check:

$$\text{First Stage: } Gini_{i,t-1} = \tilde{\alpha}_i + \theta Inst\_Gini_{i,t-1} + \Gamma X_{i,t} + \tilde{\mu}_i t + \nu_{i,t} \quad (5)$$

$$\text{Second Stage: } Polar_{i,t} = \alpha_i + \beta_{2SLS} \widehat{Gini}_{i,t-1} + \gamma X_{i,t} + \mu_i t + \epsilon_{i,t} \quad (6)$$

where  $\tilde{\cdot}$  denotes the first stage estimates of the state-effect and state-trend. We can interpret  $\beta_{2SLS}$  as the causal effect of income inequality on polarization.  $X_{it}$  is a vector of time-varying covariates, including state real personal income, the proportion of the state's population that is black or Hispanic, log median income, the proportion of the population with a college degree, population density, the unemployment rate, median age, the proportion of the population over 55 years of age, the proportion of the population under 25 years of age, and the unionization rate. We additionally include the proportion of total state legislators (upper and lower chambers) representing majority-minority districts.<sup>16</sup> This is an important potential confounder, since these districts are both more common in states with high levels of inequality, and are more likely to elect Democrats to the left of the party mean.

Our identification strategy requires that our instrument affects inequality (i.e. instrument relevance) and polarization only through its effect on actual income inequality (i.e. the exclusion restriction). Instrument relevance can be directly tested by performing inference on the first-stage regression. Figure A1 shows a scatter plot of the calculated Gini coefficient for the actual data against the simulated Gini instrument, and Table A8 shows the first-stage estimation results. For the first-differences specifications, the first-stage F-test statistics are well above the rule-of-thumb cutoff of ten, and the first-stage coefficient on the instrument is statistically significant and positive, as expected.

Our instrument is, by design, uncorrelated with any within-state variation over time in political polarization or legislative ideology except through its effect on within-state variation in income inequality. As noted above, the identifying assumption of our empirical strategy

---

<sup>16</sup>We define a district as majority-minority if the proportion of total population who are black or Hispanic is greater than 50%.

amounts to an assumption that initial state income inequality is unrelated to future **changes** in political polarization. This assumption is directly testable. Figure A3 shows scatterplots comparing initial income inequality to subsequent year-to-year changes in the four main outcome variables (polarization, average chamber ideology, average Democratic ideology and average Republican ideology). In all cases, the slope of the line of best fit is close to zero (formally, the slope is not statistically significantly different from zero at conventional levels). Thus, we argue that the identifying assumption of our simulated instrumental variables identification strategy is satisfied, and hence  $\beta_{2SLS}$  can thus be interpreted as the causal effect of income inequality on political polarization (or other measures of ideology).

## 4 Empirical Results

### 4.1 Aggregate State Polarization

We first consider the effect of inequality on state-level polarization, measured by the distance between the mean ideal points of the Democratic and Republican parties. We then disaggregate this effect by examining the influence of inequality on each of the two separate party means. Income inequality may also affect the overall average ideology of the legislature in addition to the distance between parties. We thus consider how inequality might affect the overall mean ideal points of legislative chambers within each state, as well as the partisan balance of legislative chambers, as measured by the proportion of seats held by Republicans in each chamber. As noted earlier, we aggregate across upper and lower state legislative chambers to arrive at single numbers for political polarization and ideology within each state.

Table 1 presents our first main result, showing the effect of income inequality on state polarization using a variety of first difference specifications (e.g. equation 2). The top panel shows the results from our IV model, while the bottom panel shows the results from a naive OLS specification. There are three columns in this table of results (similarly with subsequent

tables). The first column reports a model with no other covariates aside from the state-level Gini coefficient. The second column controls for time-varying economic and demographic factors. The third column include time-varying controls and state fixed effects which capture the state-specific trends in our first-difference model. Each model reports standard errors clustered at the state level to account for arbitrary serial correlation in polarization within states.

Table 1: Effect of Income Inequality on State Legislative Polarization

	(1)	(2)	(3)
<i>IV Results: Gini</i>	1.620*** (0.430)	1.455*** (0.389)	1.542*** (0.390)
<i>OLS Results: Gini</i>	0.071 (0.146)	0.051 (0.134)	0.037 (0.134)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

The IV point estimates for each specification are larger in absolute value than the OLS estimates and more precisely estimated. Our preferred first-differences specification includes state-specific linear trends, shown in column 3. Using this specification, the effect of inequality on polarization is positive and statistically significant at conventional levels. To contextualize the effect size reported in column 3 (1.542), a one within-state standard deviation increase in state income inequality (0.019) would correspond to a session-over-session increase in polarization of 0.029. The average cumulative change in polarization across states over 1993-2022 is about 0.62. Thus, a single onetime standard deviation increase in income inequality could account for about 4.73% of the average total increase in polarization.

## 4.2 Party Means

Our main result suggests that income inequality increases the distance between the mean ideology of the two main political parties within state legislatures. This effect could occur in a number of different ways. Income inequality might move both parties symmetrically away from the political center. Alternately, the effect may be asymmetric, where one party becomes more extreme at a faster rate than the other. To differentiate between these possibilities, we estimate models using the mean ideology of each state party as the dependent variable. By convention, positive values on the ideology scale reflect right-of-center positions and negative values reflect positions that are left of center. A positive coefficient estimate, therefore, implies that inequality moves the party mean to the right and a negative estimated effect implies that inequality moves the party mean to the left. If income inequality moves both parties symmetrically away from the center, we would expect a positive estimated effect on Republican party mean ideology and a negative estimated effect on the Democratic party mean of roughly the same absolute magnitude. On the other hand, if there is an asymmetric effect, then the estimated effect of inequality on ideology for one party will be substantially larger in absolute value.

Table 2 reports the key parameter estimates for the effect of income inequality on Democratic Party mean ideology using first-difference models (the specifications in each column mirror those in Table 1). As expected, the sign of the coefficient is negative which implies that income inequality moves the Democratic party mean to the left. The point estimates indicate that a one within-state standard-deviation (0.019) increase in the change in inequality between years moves Democratic party mean to the left by an additional 0.015, using the IV estimate in column 3. The average *cumulative* change in Democratic party means over 1993-2022 is 0.35 to the left.

Table 3 reports results from first-difference models using the mean ideology of state-level Republican parties as the dependent variable. The IV point estimates of the effect sizes are larger in magnitude than the OLS effect sizes for all specifications and significant. In the

Table 2: Effect of Income Inequality on Democratic Party Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	-0.923*** (0.237)	-0.747*** (0.216)	-0.800*** (0.221)
<i>OLS Results: Gini</i>	-0.215*** (0.077)	-0.157** (0.078)	-0.151* (0.077)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

preferred specification (column 3) the Republican effect is similar in size to the Democratic one, suggesting a somewhat symmetric polarization effect of inequality. A one within-state standard-deviation increase in the change in income inequality (0.019) moves the Republican party mean an additional 0.014 to the right per biennium. To put that in context, the average *cumulative* change in Republican party means from 1993 to 2022 is 0.27.

Table 3: Effect of Income Inequality on Republican Party Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	0.697** (0.302)	0.708** (0.303)	0.741** (0.298)
<i>OLS Results: Gini</i>	-0.143 (0.124)	-0.106 (0.108)	-0.113 (0.106)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

In the Appendix, we also evaluate the effect of inequality on within-party averages using different specifications and alternative measures. In short, the results do not exhibit a

consistent pattern of one party’s average ideology moving faster than the other due to income inequality. The mixed results tells us that income inequality is affecting the parties similarly. Does this mean that the average ideology across the legislature is maintained at the center? We explore this possibility in the next section.

### 4.3 Chamber Means and Partisanship

Income inequality appears to affect the average ideology of parties and the ideological distance between them. But inequality might also affect the average ideology of the legislature overall. To test this possibility, we estimate models using the previous specifications, using the state mean ideology as the dependent variable. State mean ideology is calculated as the weighted average across chambers of the mean ideology score within each chamber, for each state in each year.<sup>17</sup> Table 4 reports estimates for these models. The effects are statistically significant and substantial. There is strong evidence that rising income inequality moves chamber means to the right.

Table 4: Effect of Income Inequality on Average Chamber Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	3.107*** (0.530)	2.391*** (0.547)	2.416*** (0.547)
<i>OLS Results: Gini</i>	0.383** (0.159)	0.282* (0.153)	0.269* (0.153)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

That income inequality causes the mean ideology of state legislative chambers to move to the right may result from three mechanisms. This effect might be the consequence of

<sup>17</sup>See Appendix A.4.6 for results using medians, which produce similar findings.

inequality causing (a) both parties to move to the right, (b) the Republicans to move to the right more than the Democrats move to the left or (c) changes in the partisan balance of the legislature where more Republicans are elected, replacing Democrats. Possibility (a) is ruled out by the previous results where we find that inequality moves Democrats to the *left*. Similarly, possibility (b) is ruled out by the fact that the effect size for movement in Republican ideology is close to that of Democrats. We explore the possibility (c) that income inequality changes the partisan balance of legislatures by increasing the proportion of seats held by Republicans. We estimate similar models as in the previous analysis, now using the proportion of seats in state legislatures held by Republicans as the outcome variable. Table 5 summarizes the results. We find that greater increases in income inequality cause a bigger change in the share of seats held by Republicans. This finding implies a “composition effect,” where increases in income inequality cause Democrats to be replaced by Republicans while simultaneously moving both parties away from the center.

Table 5: Effect of Income Inequality on Republican Seat Share

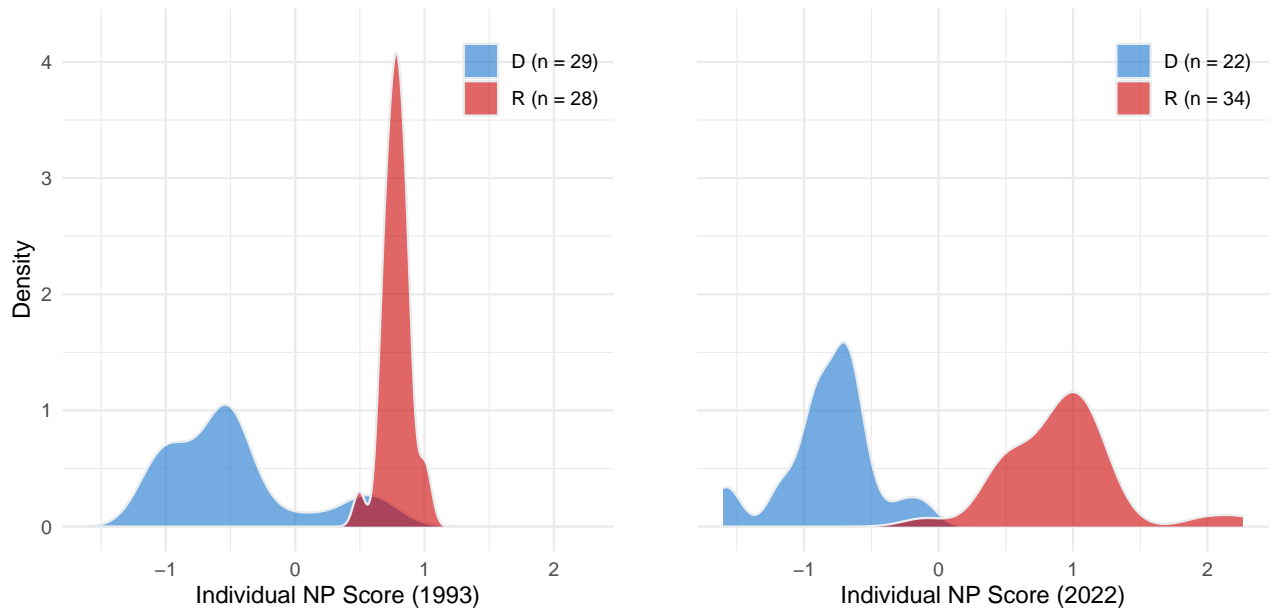
	(1)	(2)	(3)
<i>IV Results: Gini</i>	1.968*** (0.333)	1.479*** (0.353)	1.492*** (0.350)
<i>OLS Results: Gini</i>	0.252** (0.102)	0.166* (0.098)	0.150 (0.100)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

Figure 2 illustrates one example of how the composition effect has taken place. This figure plots a histogram of the ideal points of individual legislators in the Arkansas legislature in 1993 and 2022. In 1993, Democrats exhibited a greater spread in ideal points, with several legislators holding moderate and conservative ideological positions. The contrast to 2022 is

evident: not only are there fewer Democrats, but right-of-center coverage has disappeared. Polarization has increased and average ideology has moved to the right. The detected effect of inequality on average ideology and Republican seat share reveals a mechanism driving polarization. The result is that moderate Democratic seats are replaced by conservative Republicans, moving the overall legislature to the right.

Figure 2: Ideological distribution of the Arkansas legislature in 1993 and 2022



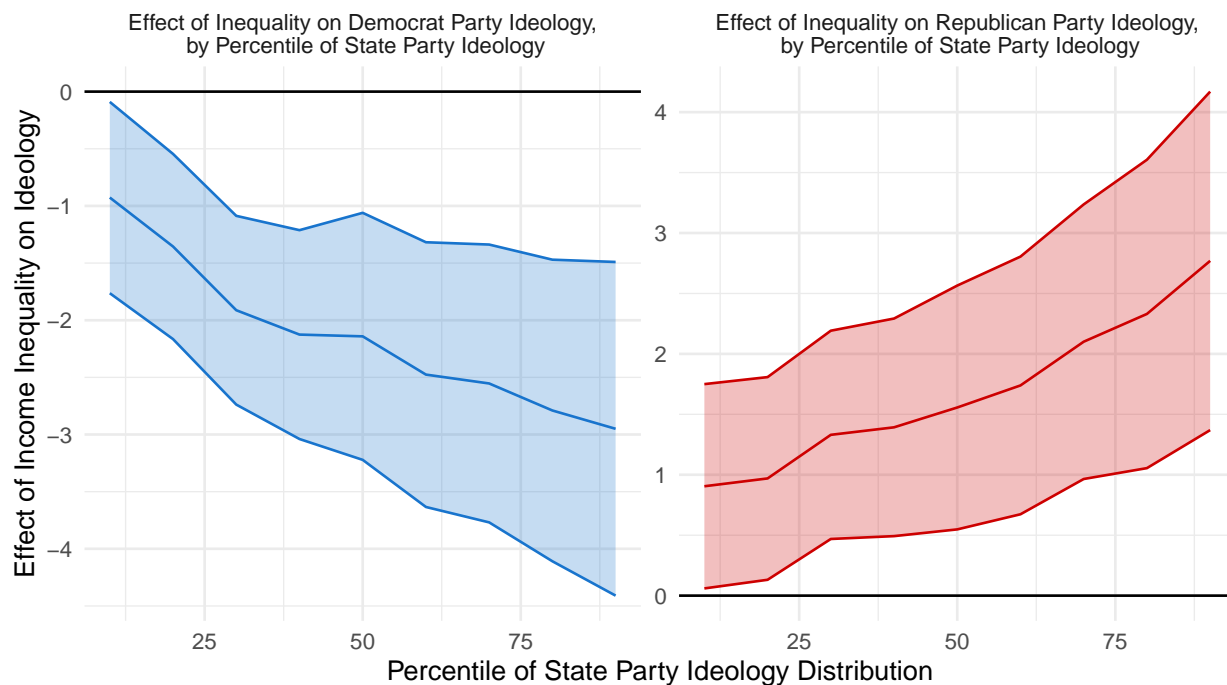
#### 4.4 Party Quantiles and Extremism

One result that has remained constant through a variety of specifications is that income inequality affects the average ideologies of both parties. There is also strong evidence that inequality moves the overall ideology of the legislature to the right. To further investigate these two results, we consider the effect of income inequality on quantiles of the ideology distribution within each party. Because the ideal point scores are ordered by conservatism, the “moderate” wing of the Democratic party is captured by the right tail of the ideology distribution (e.g. the 90th quantile), while the moderate wing of the Republican party is captured by the left tail (e.g the 10th quantile), and the opposite is true for the “extreme”

wings of each party.

Figure 3 summarizes the effect of inequality from models estimated using different quantiles of the party ideology distribution as dependent variables, using our preferred specification. Two striking features are immediately apparent: first, the effect of income inequality on ideology increases with conservatism. This is to say, the effect of inequality is largest for the most conservative (i.e. extreme) wing of Republican parties, and for the most conservative (i.e. moderate) wing of the Democratic parties. In other words, rising income inequality appears to move the extreme wing of the Republican party further to the right, while moving the moderate wing of the Democratic party further to the left. The trend in the Democratic party is consistent with the composition effect, where the most conservative Democrats lose their seats.

Figure 3: Effect of Income Inequality on Party Quantiles (Ordered by Conservatism)



Since the ideal point ideology scores are fixed over time for any individual legislator, the only way that aggregate measures of state polarization can change over time is through the replacement of legislators (either in a normal election, or via retirement, death or party

defection). This fact can be used to rationalize the results on the effect of income inequality on the various aggregate state measures of ideology and political polarization. Rising income inequality moves Republican parties to the right on average (especially for extreme Republicans), and Democratic parties to the left on average (especially for moderate Democrats), but in a way that increases the overall conservatism. These results are all strongly suggestive of an effect of income inequality that produces a smaller and more liberal Democratic party and a larger, more conservative, Republican party.

## **4.5 Heterogeneous Effects: Campaign Finance, Trade Exposure, and Racial Composition**

There are two key findings in our primary analysis. First, income inequality has a causal effect on political polarization across states. Second, inequality contributes to a rightward shift in the average ideology of state legislative chambers. In this section, we examine three potential sources of heterogeneity in these effects. First, we test whether states affected by the Supreme Court’s *Citizens United* ruling exhibit a distinct pattern in the relationship between inequality, polarization, and average chamber ideology. We then assess whether states with higher baseline vulnerability to the China trade shock show a different response to inequality. Finally, we consider whether states that had a larger baseline proportion of minority populations exhibit heterogeneous effects of inequality on polarization and on the average ideology of their legislative chambers.

### **4.5.1 *Citizens United* and Campaign Finance**

One potential mechanism through which rising income inequality might affect the distribution of legislator ideology is through campaign finance, as in Feddersen and Gul (2014). We can explore this potential mechanism by exploiting variation in state-level restrictions on independent campaign expenditures (IE) which were overturned by the *Citizens United* (CU) ruling in 2010. These pre-CU restrictions may have effectively ameliorated the effect

of inequality on polarization and ideology by reducing the ability of top income earners to spend resources on the election of their preferred candidates. We test for the presence of this mechanism with a triple-interaction DiD-IV design by allowing the effect of income inequality to vary systematically with the removal of expenditure limits in 2010. For details on states that imposed independent expenditure restrictions, see Spencer and Wood (2014) and Klumpp, Mialon and Williams (2016). Table A7 lists the state-years with restrictions.

Our treatment indicator,  $LIM_i$  takes the value of 1 if state  $i$  had an independent expenditure limitation prior to *Citizens United* and zero otherwise. To allow the effect of income inequality to vary systematically with the removal of these expenditure restrictions, we then estimate a series of two-stage least squares models which modify the preferred specifications above to allow for the effect of income inequality to vary depending on whether a pre-Citizen's United campaign finance restriction was removed in 2010:

$$\begin{aligned} \text{First Stage 1: } \Delta Gini_{i,t-1} &= \delta_1 LIM_i + \delta_2 Post_t + \delta_3 (LIM_i \times Post_t) \\ &+ \theta_1 \Delta Inst\_Gini_{i,t-1} + \theta_2 (\Delta Inst\_Gini_{i,t-1} \times LIM_i) \\ &+ \theta_3 (\Delta Inst\_Gini_{i,t-1} \times LIM_i \times Post_t) + \Gamma \Delta X_{i,t} + \nu_{i,t}, \end{aligned}$$

$$\begin{aligned} \text{First Stage 2: } \Delta Gini_{i,t-1} \times LIM_i &= \delta_1 LIM_i + \delta_2 Post_t + \delta_3 (LIM_i \times Post_t) \\ &+ \theta_1 \Delta Inst\_Gini_{i,t-1} + \theta_2 (\Delta Inst\_Gini_{i,t-1} \times LIM_i) \\ &+ \theta_3 (\Delta Inst\_Gini_{i,t-1} \times LIM_i \times Post_t) + \Gamma \Delta X_{i,t} + \nu_{i,t}, \end{aligned}$$

$$\begin{aligned} \text{First Stage 3: } \Delta Gini_{i,t-1} \times LIM_i \times Post_t &= \delta_1 LIM_i + \delta_2 Post_t + \delta_3 (LIM_i \times Post_t) \\ &+ \theta_1 \Delta Inst\_Gini_{i,t-1} + \theta_2 (\Delta Inst\_Gini_{i,t-1} \times LIM_i) \\ &+ \theta_3 (\Delta Inst\_Gini_{i,t-1} \times LIM_i \times Post_t) + \Gamma \Delta X_{i,t} + \nu_{i,t}. \end{aligned}$$

$$\begin{aligned}
\Delta Polar_{i,t} = & \delta_1 LIM_i + \delta_2 Post_t + \delta_3 (LIM_i \times Post_t) \\
& + \beta_1 \widehat{\Delta Gini_{i,t}} + \beta_2 \widehat{\Delta Gini_{i,t-1} \times LIM_i} + \beta_3 \widehat{\Delta Gini_{i,t-1} \times LIM_i \times Post_t} \\
& + \gamma \Delta X_{i,t} + e_{i,t}.
\end{aligned}$$

We instrument separately for  $Gini_{i,t-1}$ ,  $Gini_{i,t-1} \times LIM_i$ , and  $Gini_{i,t-1} \times LIM_i \times Post_t$ . In these models, we do not include state or year fixed effects as they are collinear with the treatment indicator,  $LIM_i$ , and the time indicator,  $Post_t$ . The effect of inequality on polarization (or another measure of ideology) for a treated state after *Citizens United* is  $\beta_1 + \beta_2 + \beta_3$ . Here,  $\beta_3$  captures the differential change in the effect of inequality on polarization between states that had their contribution limits removed and those that never imposed such restrictions. This parameter is a difference-in-differences in slopes: the post-minus-pre change in the inequality effect for formerly limited states relative to never limited states. If  $\beta_1 > 0$  and  $\beta_1 + \beta_2 = 0$ , this suggests that before *Citizens United* the independent expenditure limits essentially fully neutralized the effect of inequality on polarization. If  $\beta_3 > 0$  at a significant rate, this means that lifting those limits amplified the effect of inequality on polarization for the affected states.

Tables 6 and 7 summarize the estimation results of several models which allow for heterogeneity of the effect of income inequality based on whether expenditure limits were lifted by *Citizens United*. Each table refers to the results of models using, respectively, political polarization and chamber mean ideology as outcome variables. Additional results are provided in the Appendix for models using Republican party mean ideology, Democratic party mean ideology, and republican seat share as outcome variables. In each table, the first column refers to a model with no covariates and the second with time-varying controls.

Table 6 reports the effect of inequality on polarization by  $LIM_i$ . Although the directions of estimates suggest that having limits reduced the effect of inequality on polarization, these estimates are not statistically significant. Primarily, we are interested in the coefficient estimate for  $Gini \times LIM \times Post$ , which indicates the post-2010 difference in the effect of

Table 6: Citizens United Difference-in-Differences: IV Effect of Inequality on State Legislative Polarization by IE Limits

	(1)	(2)
LIM	-0.016* (0.009)	-0.013 (0.008)
Post	0.023* (0.013)	0.029** (0.013)
LIM $\times$ Post	-0.0003 (0.014)	-0.003 (0.014)
Gini	1.524* (0.922)	1.491** (0.752)
Gini $\times$ LIM	-0.844 (1.023)	-0.954 (0.891)
Gini $\times$ LIM $\times$ Post	0.915 (0.690)	1.207 (0.843)
Other Controls?	No	Yes
$Gini + Gini \times LIM + Gini \times LIM \times Post$	1.59	1.74
$P(Gini + Gini \times LIM + Gini \times LIM \times Post) = 0$	0.001	0
First Stage F	40.94	37.35
Observations	651	651

Note:

\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$

Gini on polarization for states that used to have limits compared to those that never did. The coefficient estimate for  $\text{Gini} \times \text{LIM}$  indicates the difference in effect between limited and unlimited states before *Citizens United*. The left panel of Figure 4 visualizes these estimates.

In our examination of inequality’s relationship with ideology measures in Tables 1 - 4 above, we had found that the strongest effect of ideology was on average chamber ideology. That is also where we find strong evidence of a campaign finance mechanism. Table 7 displays the results of our DiD-IV interaction models with average ideology as the outcome variable. Before *Citizens United*, independent expenditure limits entirely neutralized the effect of inequality on average chamber ideology. Looking at the coefficient estimates using time-varying controls in the second column,  $\beta_1 + \beta_2 \approx 0$  indicates that before *Citizens United*, independent-expenditure limits neutralized the inequality-ideology relationship. After *Citizens United*, affected states exhibit a large and statistically significant increase in the rightward movement of average chamber ideology due to changes in inequality. The right panel in Figure 4 visualizes the confidence intervals for  $\text{Gini} \times \text{LIM}$ ,  $\text{Gini} \times \text{LIM} \text{ times Post}$ , and the total effect of inequality on average ideology ( $\text{Gini} + \text{Gini} \times \text{LIM} + \text{Gini} \times \text{LIM} \times \text{Post}$ ). The results reproduce the pattern in the polarization analysis. Before *Citizens United*, the effect of inequality on average ideology in limited states was less than unlimited states. After *Citizens United*, the effect of inequality on average ideology is greater in previously limited states relative to never-limited states.

In short, lifting independent expenditure limits appears to increase a chamber’s polarization and rightward movement in average ideology due to increases in the growth of inequality. In Appendix A.4.8, we show that party-specific *Citizens United* estimates for Republican and Democrat ideology are directionally consistent with a rightward shift in average chamber ideology, but insignificant. Point estimates indicate that, after *Citizens United*, Republican ideology moved right more and Democratic ideology moved left.

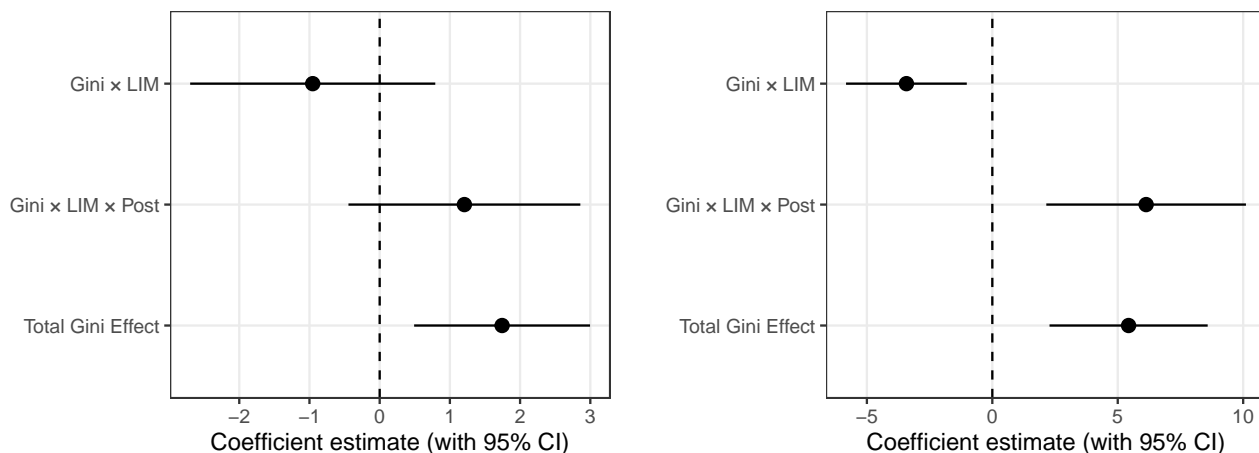
Table 7: Citizens United Difference-in-Differences: IV Effect of Inequality on Average Chamber Ideology by IE Limits

	(1)	(2)
LIM	-0.006 (0.008)	-0.006 (0.007)
Post	-0.004 (0.010)	0.003 (0.010)
LIM $\times$ Post	0.018 (0.012)	0.016 (0.011)
Gini	3.183*** (1.031)	2.725*** (0.858)
Gini $\times$ LIM	-3.636*** (1.258)	-3.424*** (1.228)
Gini $\times$ LIM $\times$ Post	5.905*** (1.742)	6.132*** (2.030)
Other Controls?	No	Yes
$Gini + Gini \times LIM + Gini \times LIM \times Post$	5.45	5.43
$P(Gini + Gini \times LIM + Gini \times LIM \times Post) = 0$	0	0
First Stage F	40.94	37.35
Observations	651	651

Note:

\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$

Figure 4: Estimated effects of income inequality on polarization (left panel) and average chamber ideology (right panel). The top rows plot the pre-Citizens-United inequality effect for states with independent expenditure limits relative to those without limits ( $\text{Gini} \times \text{LIM}$ ). The second rows show the change in that effect after *Citizens United* ( $\text{Gini} \times \text{LIM} \times \text{Post}$ ). The bottom rows plot the confidence interval for the total effect of inequality ( $\text{Gini} + \text{Gini} \times \text{LIM} + \text{Gini} \times \text{LIM} \times \text{Post}$ ). Points represent coefficient estimates and horizontal lines show 95% confidence intervals from the models reported in Tables 6 and 7.



#### 4.5.2 Trade Exposure and Vulnerability to Trade Shock

One potential source of heterogeneity in results is through variation in how much states were vulnerable to economic shock from trade exposure. The literature on the political consequences of the China trade shock shows that regions more vulnerable to import competition experienced substantial economic and social disruption, which in turn reshaped patterns of political representation and polarization (Autor, Dorn and Hanson, 2013; Feigenbaum and Hall, 2015; Autor et al., 2020). Autor et al. (2020) find several polarizing effects of China trade shocks on congressional polarization. Specifically, they find that highly exposed districts were more likely to elect a more extreme congressional representative. Majority white districts were more likely to elect a conservative Republican while majority-minority districts were more likely to elect a progressive Democrat. If trade exposure already produced visible economic winners and losers, the marginal effect of rising income inequality on political behavior may have been attenuated. Conversely, in states less directly affected by trade shocks, inequality itself may have been a more salient driver of distributive conflict

and partisan polarization. We therefore examine whether the causal effect of inequality on legislators’ political behavior differs between states with above-average and below-average baseline exposure to international trade.

To ensure that our measure of trade exposure is predetermined with respect to the period of political outcomes we study, we use state-level trade exposure as of 1990, the period immediately preceding the onset of the China trade shock in the United States. We split states into two samples based on whether they fell above or below the average trade exposure per worker in 1990. This baseline captures structural differences at that time in states’ vulnerability to import competition before changes in manufacturing employment occurred, and thus can be treated as exogenous to subsequent changes in inequality and polarization. We construct a measure of baseline trade exposure using the commuting-zone level data from Autor, Dorn and Hanson (2013). For each state  $s$ , we aggregate commuting-zone exposure using population weights so that exposure reflects the experience of the average worker in the state:

$$TradeExposure_s = \frac{\sum_{c \in s} Exposure_{c,1990} \times PopWeight_{c,1990}}{\sum_{c \in s} PopWeight_{c,1990}}.$$

Because our polarization data begin in 1993, this 1990 measure, calculated from commuter zone exposure rates in Autor, Dorn and Hanson (2013), provides a predetermined proxy for a state’s structural vulnerability to the trade shock. Using 1990 as the pre-period baseline keeps in practice with Autor, Dorn and Hanson (2013).

In Table 8, the first two columns show the estimated effects of income inequality on polarization in states above the 1990 average trade exposure per worker (column one) and below the average (column two). Columns four and five show the estimated effects of income inequality on the average ideology of state chambers. Point estimates for states with low trade exposure are larger than states that had high pre-shock trade exposure. Results using above- and below-median thresholds, reported in the Appendix, show an even starker contrast. However, we cannot reject the null-hypothesis of equal effects. The effect of inequality on polarization in states that had high trade exposure is not statistically differentiable from

Table 8: Effects of Income Inequality by High and Low Pre-Shock Trade Exposure

	<i>Polarization</i>		<i>Average Ideology</i>	
	High	Low	High	Low
<i>IV Results: Gini</i>	1.259** (0.455)	1.891** (0.706)	2.328** (0.870)	2.887*** (0.668)
<i>OLS Results: Gini</i>	0.217 (0.186)	-0.027 (0.212)	0.072 (0.227)	0.576** (0.249)
State FE?	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes
First Stage F	42.59	48.82	42.59	48.82
Observations	262	362	262	362

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

the effect of inequality on polarization in states that had low trade exposure. Therefore, the evidence does not support the claim that vulnerability to trade shocks systematically alters the relationship between income inequality and polarization, but it also does not allow us to rule out moderate differences in effect size.

We also assess whether coefficient estimates for the effect on average chamber ideology statistically significantly different from one another. The coefficient estimates for the two samples are not statistically distinguishable from each other. These findings imply that, although the distributive tensions may have looked differently in these states due to local trade shocks, these shocks did not attenuate or dampen the political effects of income inequality. Taken together with the polarization results, this suggests that while states experienced the China trade shock to varying economic degrees, this variation did not systematically attenuate or amplify the political consequences of rising income inequality. The absence of detectable heterogeneity may reflect genuinely similar political responses across states, or it may reflect limited precision in split-sample IV designs. We leave further exploration of these effects as an opportunity for future research.

### 4.5.3 Effects by Racial Composition of States

An influential scholarly literature has emphasized the partisan realignment on racial issues as an essential cause of contemporary partisan polarization.<sup>18</sup> The arguments range from an enabling one – partisan sorting on race was necessary prior to voter sorting on a broader set of issues (e.g. O’Brian 2024) to a directly causal one – issue positions on other issues especially economic redistribution are largely shaped by racial attitudes (e.g. Gilens 1999 and Alesina, Glaeser and Sacerdote 2001.) Both versions of the argument raise potential concerns about our findings of a causal impact of inequality on political polarization to the extent to which economic inequality may be a product of racial inequality.

We do not think these concerns are significant, however. Beyond the use of our IV strategy, our main results use first-differences and state-specific trends. These specifications should net-out the effects of slower-moving variables such as racial demographics. Second, we control for the racial and ethnic populations in each state-year. These coefficients are small and unreliably estimated as one would expect in a first-difference specification.<sup>19</sup>

While we can rule out that income inequality is proxying for a direct effect of racial diversity, the argument that racial cleavages enable polarization on other issues remains possible. So we address this possibility by looking for whether there is heterogeneity in the effect size of income inequality on our political outcomes related to the racial and ethnic demographics. If racial conflict enables or intensifies partisan economic conflicts, we would expect to see a larger effect of income inequality in the most racially diverse states. Alternatively, it is possible that economic conflict is displaced by racialized ones in diverse states. In that case, we would expect the effect of income inequality on polarization to be smaller in diverse states.

---

<sup>18</sup>The important works here include Carmines and Stimson (1989); Gilens (1999); Alesina, Glaeser and Sacerdote (2001); Schickler (2016); Kuziemko and Washington (2018); Pierson and Schickler (2024); O’Brian (2024).

<sup>19</sup>In the specification of Table 1 column 3, a one-standard deviation increase in Black share of the population is associated with a trivial 0.004 increase in polarization, and the effect is not statistically significant. For Latino share, one standard deviation increase is associated with an also small 0.005 (significant at 0.05 level)

To test these possibilities, we again split states into two samples, this time by whether they fall above or below the 1990 average population proportions of Black residents. The 1990 baseline that serves as a predetermined source of heterogeneity, ensuring that observed heterogeneity reflects preexisting demographic context rather than subsequent demographic change. Results for states that fall above and below the average minority population proportions are displayed in Table 9. The first two columns show estimates for inequality on polarization for each group, and the last two columns show the effect of inequality on polarization and average chamber ideology.

Table 9: Effects of Income Inequality by High and Low Baseline Minority Composition

	<i>Polarization</i>		<i>Average Ideology</i>	
	High	Low	High	Low
<i>IV Results: Gini</i>	1.058** (0.441)	1.825*** (0.582)	1.904** (0.677)	2.630*** (0.773)
<i>OLS Results: Gini</i>	0.153 (0.174)	-0.026 (0.188)	0.465** (0.206)	0.190 (0.207)
State-Specific Linear Trend?	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes
First Stage F	45.64	50.68	45.64	50.68
Observations	224	427	224	427

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

States that had low baseline proportions of Black populations display a stronger estimated effect of income inequality on both polarization and average chamber ideology. This pattern suggests that where racial diversity is lower, inequality may become a clearer cleavage of political conflict, producing greater ideological divergence among legislators. In contrast, in more racially diverse states, distributive tensions may be mediated by identity-based cleavages that diffuse the direct link between inequality and polarization. As before we assume the split samples are independent from each other and conduct a formal hypothesis test to assess whether the estimated effects on polarization and average ideology are statistically

significantly different from one another. The coefficient estimates for the two samples are not statistically distinguishable.

Together, these findings indicate that while point estimates suggest a somewhat stronger effect of inequality on polarization and average chamber ideology in less diverse states, there is no evidence to conclude that the relationships differ significantly across racial compositions. The relationship operates similarly across racial contexts rather than being driven by racial composition.

## 5 Conclusion

We have examined the relationship between income inequality and political polarization in a number of different contexts. By moving to the state level, and by adopting an instrumental variables empirical strategy that discards variation due to non-random sorting across state lines and corrects for measurement error, we are able to recover estimates of the effect of inequality on polarization that are causal. Our results from models examining the effect of income inequality on aggregate polarization within state legislatures align with previous studies of an equivalent national-level relationship (McCarty, Poole and Rosenthal, 2006). We find that within-state income inequality has a significant, positive and quantitatively large effect on within-state legislative political polarization. As parties move away from the center, income inequality also increases the share of Republican seats in state legislatures. The result is a smaller and more extreme Democratic party, a larger more extreme Republican party, and, consequently, an overall movement to the right in the average ideology of chambers. These results are robust to a number of different specifications.

These findings provide an important engagement with studies examining the causes of ideological change and redistributive politics. Meltzer and Richard (1981) famously predict that more inequality should lead to more support for redistribution and therefore left-leaning parties, so why do we find evidence in the opposite direction? Grossmann and Hopkins

(2016)'s characterization of the contemporary American parties helps resolve this puzzle in the context of our findings. In the last two decades, in contrast to a Republican party that has become an ideological movement, the Democratic party has become characterized as a coalition of social groups. Building on this distinction, we argue that Democratic movement to the left is primarily a consequence of coalitional change. As the groups that make up the Democratic coalition shift, so does its average ideology.

Income inequality disadvantages groups at the Democratic coalition's moderate end, like rural and less-educated voters. Because these groups are bound to the party through concrete policy demands rather than ideological commitment Grossmann and Hopkins (2016), the gridlock induced by polarization, which in turn is in part induced by inequality, undermines the basis for their support. The consequence is electoral loss. Seats filled with moderate Democrats flip to being occupied by conservative Republicans and move the overall legislature to the right. This composition effect explains why inequality has similar sized effects on each party even when the within-party paces of ideological change differ.

In studies of Congress, Poole and Rosenthal (1997) and Grossmann and Hopkins (2016) find that Republicans are driving polarization in the post-war period. At the state level, Shor and McCarty (2022) find that Democrats are moving faster from the center than Republicans. Our discovery of a composition effect at the state level uncovers the mechanism for the variation in pace. The parties are almost equally affected by inequality, but legislatures mechanically become more conservative when moderate Democrats are replaced with Republicans.

We also explore three possible sources of heterogeneity in our results. We posit that campaign finance may serve as the causal mechanism linking rising income inequality to rising polarization and rightward movement in the ideology of state legislatures. Our results, when allowing for systematically varying effects based on whether pre-*Citizens United* independent expenditure restrictions were in place suggest that the presence of these restrictions either partially or fully interrupted the effects of income inequality. In state-years with no restric-

tions, the effect of income inequality on average chamber ideology is positive. Our results show the same descriptive story for polarization although the effects are not statistically significant.

Additionally, we explore variation due to vulnerability to local economic shock from trade exposure and differences in the racial composition of states. We test whether states whose workforce had lower exposure to trade exhibited a larger magnitude in the effect of income inequality on both polarization and average chamber ideology. Results do not support this story. We also explore whether our findings differ for states given the proportion of Black constituents in a pre-period baseline. We test whether states with smaller proportions of these minority groups exhibited larger effects of inequality on polarization and average chamber ideology. Similar to trade dynamics, we do not find evidence that states differ in how inequality affects polarization based on racial composition. As with the prior heterogeneity inquiries, racial composition of states and inequality's political effects are worthy of deeper exploration.

## References

Alesina, Alberto, Edward Glaeser and Bruce Sacerdote. 2001. Why Doesn't the US Have a European-Style Welfare System? Technical Report w8524 National Bureau of Economic Research Cambridge, MA: .

**URL:** <http://www.nber.org/papers/w8524.pdf>

Ansolabehere, Stephen, James Snyder and Charles Stewart. 2001a. "Candidate Positioning in U.S. House Elections." *American Journal of Political Science* 45(1):136–159.

Ansolabehere, Stephen, James Snyder and Charles Stewart. 2001b. "The Effects of Party and Preferences on Congressional Roll-Call Voting." *Legislative Studies Quarterly* 26(4):533–572.

Auten, Gerald and David Splinter. 2024. "Income Inequality in the United States: Using Tax data to Measure Long-term Trends." *Journal of Political Economy* 132(7):2179–2227.

Autor, David, David Dorn, Gordon Hanson and Kaveh Majlesi. 2020. "Importing political polarization? The electoral consequences of rising trade exposure." *American Economic Review* 110(10):3139–3183.

Autor, David H, David Dorn and Gordon H Hanson. 2013. "The China Syndrome: Local Labor Market Effects of Import Competition in the United States." *American Economic Review* 103(6):2121–2168.

**URL:** <https://pubs.aeaweb.org/doi/10.1257/aer.103.6.2121>

Barber, Michael J. 2016. "Ideological Donors, Contribution Limits, and the Polarization of American Legislatures." *The Journal of Politics* 78(1):296–310.

**URL:** <http://dx.doi.org/10.1086/683453>

Barber, Michael and Nolan M McCarty. 2015. Causes and Consequences of Polarization. In

- Solutions to Political Polarization in America*, ed. Nathaniel Persily. Cambridge University Press.
- Barth, Erling, Henning Finseraas and Karl O. Moene. 2014. “Political Reinforcement: How Rising Inequality Curbs Manifested Welfare Generosity.” *American Journal of Political Science* pp. n/a–n/a.
- Baum-Snow, Nathaniel and Fernando Ferreira. 2015. Causal Inference in Urban Economics. In *Handbook of Regional and Urban Economics*, vol. 5A, ed. Gilles Duranton, Vernon Henderson and William Strange. Elsevier North-Holland.
- Binder, Sarah. 2015. “The Dysfunctional Congress.” *Annual Review of Political Science* 18:85–101.
- Boustan, Leah, Fernando Ferreira, Hernan Winkler and EM Zolt. 2013. “The Effect of Rising Income Inequality on Taxation and Public Expenditures: Evidence from US Municipalities and School Districts, 1970–2000.” *Review of Economics and . . .* 95(October):1291–1302.
- Carmines, Edward G. and James A. Stimson. 1989. *Issue Evolution: Race and the Transformation of American politics*. Princeton University Press.
- Congressional Budget Office. 2026. The Distribution of Household Income, 2022. Technical Report 61911 Congressional Budget Office Washington, DC: .  
**URL:** <https://www.cbo.gov/publication/61911>
- Duca, John V. and Jason L. Saving. 2012. “Has Income Inequality or Media Fragmentation Increased Political Polarization?” *Federal Reserve of Dallas Working Paper* .
- Duca, John V. and Jason L. Saving. 2016. “Income Inequality and Political Polarization: Time Series Evidence over Nine Decades.” *Review of Income and Wealth* 62(3):445–466.
- Fedderson, Timothy and Faruk Gul. 2014. “Polarization and Income Inequality: A Dynamic Model of Unequal Democracy.” *Working Paper, Northwestern University* .

- Feigenbaum, James J. and Andrew B. Hall. 2015. “How Legislators Respond to Localized Economic Shocks: Evidence from Chinese Import Competition.” *The Journal of Politics* 77(4):1012–1030.  
**URL:** <https://www.journals.uchicago.edu/doi/10.1086/682151>
- Garand, James C. 2010. “Income Inequality, Party Polarization, and Roll-Call Voting in the U.S. Senate.” *The Journal of Politics* 72(04):1109–1128.
- Gelman, Andrew, Lane Kenworthy and Yu-Sung Su. 2010. “Income Inequality and Partisan Voting in the United States.” *Social Science Quarterly* 91(5):1203–1219.
- Gilens, Martin. 1999. *Why Americans hate welfare: race, media, and the politics of anti-poverty policy*. Studies in communication, media, and public opinion Chicago, IL: University of Chicago Press.
- Grossmann, Matt and David A. Hopkins. 2016. *Asymmetric Politics: Ideological Republicans and Group Interest Democrats*. Oxford University Press.  
**URL:** <https://academic.oup.com/book/6600>
- Hacker, Jacob S. 2005. “Policy drift: The hidden politics of US welfare state retrenchment.” *Beyond continuity: Institutional change in advanced political economies* pp. 40–82.
- Hall, Andrew B. 2014. “How the Public Funding of Elections Increases Candidate Polarization.” *Unpublished manuscript, Stanford University, Palo Alto* .
- Hirano, Shigeo, James M. Snyder Jr., Stephen Ansolabehere and John Mark Hanson. 2010. “Primary Elections and Partisan Polarization in U.S. Congressional Elections.” *Quarterly Journal of Political Science* 5(2):169–191.
- Jenkins, Stephen, Richard Burkhauser, Shuaizhang Feng and Jeff Larrimore. 2011. “Measuring inequality using censored data: a multiple-imputation approach to estimation

- and inference.” *Journal of the Royal Statistical Society: Series A (Statistics in Society)* 174(1):63–81.
- Klumpp, Tilman, Hugo Mialon and Michael Williams. 2016. “The Business of American Democracy: Citizens United, Independent Spending, and Elections.” *Journal of Law and Economics* 59.
- Kuziemko, Ilyana and Ebonya Washington. 2018. “Why Did the Democrats Lose the South? Bringing New Data to an Old Debate.” *American Economic Review* 108(10):2830–2867.
- La Raja, Raymond J. and Brian F. Schaffner. 2015. *Campaign Finance and Political Polarization: When Purists Prevail*. University of Michigan Press.
- Ma, Yanlei. 2014. “Income Inequality , Political Polarization and Fiscal Policy Gridlock.” *Working Paper* pp. 1–51.
- Mann, Thomas E. 2006. Polarizing the House of Representatives: How Much Does Gerrymandering Matter? In *Red And Blue Nation?: Characteristics And Causes of America’s Polarized Politics*, ed. Pietro Nivola and David Brady. Washington D.C.: Brookings Press.
- Masket, Seth E., Jonathan Winburn and Gerald C. Wright. 2012. “The Gerrymanderers Are Coming! Legislative Redistricting Won’t Affect Competition or Polarization Much, No Matter Who Does It.” *PS: Political Science* January:39–43.
- Masket, Seth E. and Michael G. Miller. 2015. “Does Public Election Funding Create More Extreme Legislators? Evidence from Arizona and Maine.” *State Politics & Policy Quarterly* 15(1):24–40.
- McCarty, Nolan. 2007. The Policy Effects of Political Polarization. In *Transformations of American Politics*, ed. Paul Pierson and Theda Skocpol. Princeton, NJ: Princeton University Press.

- McCarty, Nolan. 2019. *Polarization: What Everyone Needs to Know*®. Oxford University Press.
- McCarty, Nolan, Keith Poole and Howard Rosenthal. 1997. *Income Redistribution and the Realignment of American Politics*. Washington, DC: AEI Press.
- McCarty, Nolan, Keith T. Poole and Howard Rosenthal. 2006. *Polarized America: The Dance of Ideology and Unequal Riches*. Cambridge: MIT Press.
- McCarty, Nolan, Keith T. Poole and Howard Rosenthal. 2009. “Does Gerrymandering Cause Polarization?” *American Journal of Political Science* 53(3):666–680.
- McGhee, Eric, Seth Masket, Boris Shor, Steven Rogers and Nolan McCarty. 2014. “A Primary Cause of Partisanship? Nomination Systems and Legislator Ideology.” *American Journal of Political Science* 58(2):337–351.
- Meltzer, Allan H. and Scott F. Richard. 1981. “A Rational Theory of the Size of Government.” *Journal of Political Economy* 89(5):914–927.  
**URL:** <http://www.jstor.org/stable/1830813>
- Mettler, Suzanne. 2016. “The Polycscape and the Challenges of Contemporary Politics to Policy Maintenance.” *Perspectives on Politics* 14(02):369–390.
- O’Brian, Neil A. 2024. *The Roots of Polarization: From the Racial Realignment to the Culture Wars*. University of Chicago Press.
- Pierson, Paul and Eric Schickler. 2024. *Partisan Nation: The Dangerous New logic of American Politics in a Nationalized Era*. University of Chicago Press.
- Piketty, Thomas and Emmanuel Saez. 2003. “Income Inequality in the United States, 1913–1998.” *The Quarterly Journal of Economics* 118(1):1–41.

- Piketty, Thomas, Emmanuel Saez and Gabriel Zucman. 2018. "Distributional National Accounts: Methods and Estimates for the United States." *The Quarterly Journal of Economics* 133(2):553–609.
- Poole, Keith and Howard Rosenthal. 1997. *Congress: A Political-Economic History of Roll Call Voting*. New York: Oxford University Press.
- Poole, Keith T. 2007. "Changing Minds? Not in Congress!" *Public Choice* 131(3-4):435–451.
- Schickler, Eric. 2016. "Racial Realignment: The Transformation of American Liberalism, 1932-1965."
- Shor, Boris and Nolan McCarty. 2011. "The Ideological Mapping of American Legislatures." *American Political Science Review* 105(03):530–551.
- Shor, Boris and Nolan McCarty. 2022. "Two Decades of Polarization in American State Legislatures." *Journal of Political Institutions and Political Economy* 3(3–4):343–370.
- Spencer, Douglas M. and Abby Wood. 2014. "Citizens United, States Divided: An Empirical Analysis of Independent Political Spending." *Indiana Law Journal* 89.
- Theriault, Sean M. 2008. *Party Polarization in Congress*. New York: Cambridge University Press.
- Vlaicu, Razvan. 2016. "Inequality, Participation, and Polarization." *Working paper* .
- Voorheis, John. 2014. "State and Metropolitan Area Income Inequality in the United States: Trends and Determinants, 1968–2012." *Working Paper* .

## A Additional Tables and Figures

### A.1 Summary of Key Variables

Because ideal point ideology scores are fixed for an individual legislator, the ideology measures of each legislature changes yearly through the replacement of legislators. An aggregate summary of these ideology measures is given in Table A1. First-differenced variable summaries are given in Table A2.

Table A1: Summary Statistics, Key Variables

Statistic	Mean	St. Dev.	Min	Max
Gini	0.416	0.028	0.331	0.501
Instrument	0.407	0.027	0.344	0.481
Polarization	1.506	0.488	0.237	3.293
Mean Ideology	-0.002	0.396	-0.990	0.911
Dem. Mean	-0.791	0.372	-1.789	0.039
Rep. Mean	0.715	0.325	-0.229	1.613
R Seatshare	0.510	0.172	0.048	0.900

Table A2: First Differenced Summary Statistics, Key Variables

Statistic	Mean	St. Dev.	Min	Max
Gini	0.001	0.019	-0.086	0.087
Instrument	0.001	0.005	-0.015	0.015
Polarization	0.026	0.079	-0.706	1.385
Mean Ideology	0.003	0.065	-0.504	0.713
Dem. Mean	-0.014	0.042	-0.588	0.276
Rep. Mean	0.012	0.053	-0.722	0.797
R Seatshare	0.004	0.047	-0.269	0.739

Table A3 shows summary statistics for state-level demographics, education, and economic measures used as controls in Tables 1-9. To avoid multicollinearity and over-specification, we used a select set of controls: population, minority, union membership, marriage rates, citizenship, age, disposable income, and unemployment rate.

Table A3: Summary Statistics, Covariates

Statistic	Mean	St. Dev.	Min	Max
Median Income	57,025.650	10,295.930	36,574	96,346
Pop. Dens.	192.086	255.005	1.048	1,250.097
Union Membership	0.110	0.057	0.004	0.280
Union Coverage	0.014	0.008	0.000	0.066
Latino	0.102	0.100	0.004	0.502
Black	0.102	0.095	0.000	0.412
Native American	0.015	0.027	0.000	0.176
Asian	0.043	0.081	0.001	0.705
Other Race	0.020	0.029	0.000	0.228
Married	0.441	0.037	0.348	0.558
Divorced	0.136	0.021	0.080	0.197
Native Born	0.909	0.064	0.710	0.994
Noncitizen	0.047	0.033	0.001	0.192
Over 55	0.252	0.045	0.084	0.379
Under 25	0.349	0.033	0.256	0.493
College Degree	0.215	0.061	0.084	0.413
In Poverty	0.126	0.033	0.045	0.255
Attending School/College	0.041	0.013	0.011	0.080
Median Age	36.764	2.747	26	44
Population	6,164,049.000	6,813,651.000	488,167	39,547,996
Unemployment Rate	5.418	1.792	2.200	13.783

Table A4: First Differenced Summary Statistics, Covariates

Statistic	Mean	St. Dev.	Min	Max
Median Income	844.804	2,842.484	-10,652	11,083
Pop. Dens.	1.184	2.003	-6.377	24.720
Union Membership	-0.001	0.018	-0.096	0.099
Union Coverage	-0.0001	0.009	-0.059	0.059
Latino	0.002	0.011	-0.057	0.058
Black	-0.00000	0.007	-0.052	0.057
Native American	-0.00001	0.006	-0.045	0.089
Asian	0.001	0.008	-0.135	0.052
Other Race	0.002	0.012	-0.148	0.182
Married	0.002	0.019	-0.051	0.118
Divorced	-0.001	0.010	-0.056	0.036
Native Born	-0.001	0.009	-0.052	0.038
Noncitizen	0.0004	0.007	-0.030	0.039
Over 55	0.004	0.010	-0.050	0.046
Under 25	-0.002	0.011	-0.054	0.079
College Degree	0.006	0.017	-0.086	0.125
In Poverty	0.004	0.029	-0.068	0.196
Attending School/College	0.001	0.007	-0.027	0.044
Median Age	0.208	0.910	-4	4
Population	50,739.310	88,674.540	-358,140	581,457
Unemployment Rate	-0.048	1.003	-4.200	5.508

Table A5: Within-state standard deviation for key variables (level data)

Variable	Average Within-State SD
Gini	0.018
Instrument	0.008
Polarization	0.182
Mean Ideology	0.099
Dem. Mean	0.096
Rep. Mean	0.107
Rep. Seats	0.095

Table A6: Within-state standard deviation for key variables (differenced data)

Variable	Average Within-State SD
Gini	0.019
Instrument	0.005
Polarization	0.064
Mean Ideology	0.035
Dem. Mean	0.041
Rep. Mean	0.058
Rep. Seats	0.038

Table A7

State	IE Enacted	IE Rescinded
Alabama	1975	2010
Alaska	1996	2010
Arizona	1980	2010
Colorado	2003	2010
Connecticut	1987	2010
Iowa	1975	2010
Kentucky	1974	2010
Massachusetts	1975	2010
Michigan	1976	2010
Minnesota	1988	2010
Montana	1947	2012
New Hampshire	1978	2000
New York	1976	2010
North Carolina	1973	2011
Ohio	2005	2010
Oklahoma	2007	2010
Pennsylvania	1979	2010
Rhode Island	1988	2010
South Dakota	2007	2010
Tennessee	1972	2010
Texas	1987	2010
West Virginia	1908	2010
Wisconsin	1973	2010
Wyoming	1977	2010

## A.2 Measuring Inequality

To correct for top-coding in raw income data from the American Community Survey (ACS) Public Use Microdata Sample (PUMS) files, the right tail of the income distribution for each year in each state is modeled as following a Generalized Beta II distribution. Incomes above the 97.5th percentile are replaced by draws from the fitted distribution. We then construct our state-level Gini measure and instrumental variable. We use a trapezoidal area estimation formula to calculate Gini coefficients for each state-year. For each centile, the change in population share from centile to centile was combined with the average income share to compute the contribution to the total Gini value.

$$\begin{aligned}
 G &= 1 - 2\left(\sum_1^{100} (X_i - X_{i-1})\left(\frac{Y_i + Y_{i-1}}{2}\right)\right) \\
 &= 1 - \left(\sum_1^{100} (X_i - X_{i-1})(Y_i + Y_{i-1})\right),
 \end{aligned} \tag{7}$$

where  $X_i$  is the cumulative population share up to and including centile  $i$  and  $Y_i$  is the cumulative income share up to and including centile  $i$ .

## A.3 Procedure for Constructing Simulated Bartik Instrument

We construct a simulated Gini for each state-year following a shift-share Bartik-style logic that interacts *state inequality fixed at baseline* with *national shifts in the income distribution*. In the following definitions, state is denoted by  $s$ , year is denoted by  $t$ , and centile is denoted by  $c \in \{1, \dots, 100\}$ :

Total income in each centile bin in a state in a given year  $Y_{c,s,t} = \sum_{i \in c} y_{i,s,t}$ .

Number of people in each centile bin in a state in a given year  $N_{c,s,t} = \sum_{i=1}^n \mathbb{1}\{i \in c\}$ .

Population share of each bin relative to all bins in a state in a given year  $p_{c,s,t} = \frac{N_{c,s,t}}{\sum_{j=1}^C N_{j,s,t}}$ .

Income share of each bin relative to all bins  $s_{c,s,t} = \frac{Y_{c,s,t}}{\sum_{j=1}^C Y_{j,s,t}}$ .

The cumulative Lorenz coordinates are

$$X_{c,s,t} = \sum_{j \leq c} p_{j,s,t} \text{ and } Y_{c,s,t} = \sum_{j \leq c} s_{j,s,t},$$

where  $X_{c,s,t}$  is the cumulative population share up to and including centile  $c$  and  $Y_{c,s,t}$  is the cumulative income share up to and including centile  $c$ .

**Step 1: Freeze income distribution at baseline.** We choose baseline year  $\tau = 1990$ . Using national income thresholds of each centile bin in 1990, we assign each household in state  $s$  at  $t = \tau$  to a *national* centile  $c \in \{1, \dots, 100\}$ . The national centile income thresholds come from ordering all incomes and assigning each to a centile bin 1-100 based on where it falls relative to others.

This yields the state's baseline centile population share and income share:

$$p_{c,s}^{\text{base}} \equiv \frac{N_{c,s,\tau}}{\sum_j N_{j,s,\tau}},$$

and its baseline income shares

$$\tilde{s}_{c,s,\tau} \equiv \frac{Y_{c,s,\tau}}{\sum_j Y_{j,s,\tau}}.$$

**Step 2: Measure national annual changes for each centile.** Independently, from the national (pooled-U.S.) microdata, compute the year-to-year change in income share for each centile:

$$\Delta s_{c,t}^{\text{US}} \equiv s_{c,t}^{\text{US}} - s_{c,t-1}^{\text{US}}, \quad t > \tau.$$

**Step 3: Simulate future state income shares using only national trends.** Calculate each state's simulated centile income shares recursively by applying the national changes in each centile to the baseline:

$$\tilde{s}_{c,s,t} = \tilde{s}_{c,s,t-1} + \Delta s_{c,t}^{\text{US}}, \quad \text{for } t > \tau,$$

with  $\tilde{s}_{c,s,\tau}$  from Step 1 and  $\sum_c \tilde{s}_{c,s,t} = 1$  preserved. In each state, each centile's population share remains fixed at the baseline year value  $\tilde{p}_{c,s} \equiv p_{c,s}^{\text{base}}, \quad \forall t$ .

We compute the simulated Gini for each state-year using the simplified calculation in Equation 10 in section A.2.

$$\text{Inst\_Gini}_{s,t} = 1 - \sum_{c=1}^{100} \left( \tilde{X}_{c,s} - \tilde{X}_{c-1,s} \right) \left( \tilde{Y}_{c,s,t} + \tilde{Y}_{c-1,s,t} \right),$$

where  $\tilde{X}_{c,s} = \sum_{j \leq c} \tilde{p}_{j,s}$  is the simulated cumulative population share and  $\tilde{Y}_{c,s,t} = \sum_{j \leq c} \tilde{s}_{j,s,t}$  is the simulated cumulative income share.

$$\begin{aligned} \Delta \text{Inst\_Gini}_{s,t} &= - \sum_{c=1}^{100} \left( \tilde{X}_{c,s} - \tilde{X}_{c-1,s} \right) \left( \left( \tilde{Y}_{c,s,t} + \tilde{Y}_{c-1,s,t} \right) - \left( \tilde{Y}_{c,s,t-1} + \tilde{Y}_{c-1,s,t-1} \right) \right), \\ &= - \sum_{c=1}^{100} \left( \tilde{X}_{c,s} - \tilde{X}_{c-1,s} \right) \left( \Delta s_{c,t}^{US} + \Delta s_{c-1,t}^{US} \right) \end{aligned}$$

The identifying assumption asserts that the level of income inequality in each state in the initial “frozen” year is unrelated to changes in our political ideology measures in subsequent years. Figure A1 and A2 plot the simulated and actual Gini coefficients for each state, and Figure A3 tests the identifying assumption.

Figure A1: First Stage: Simulated vs. Actual State Gini

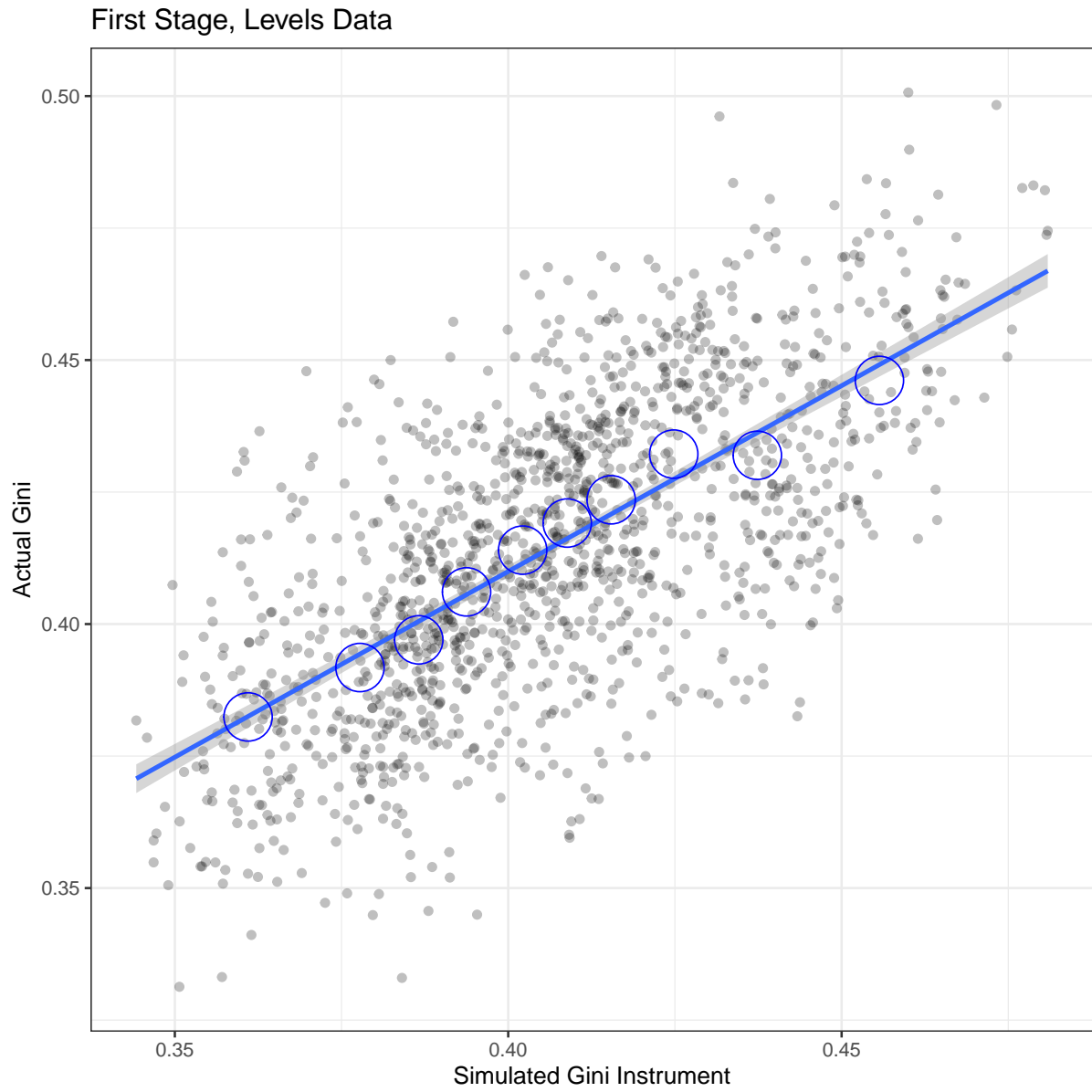


Figure A2: First Stage: Simulated vs. Actual State Gini

First Stage, Differenced Data

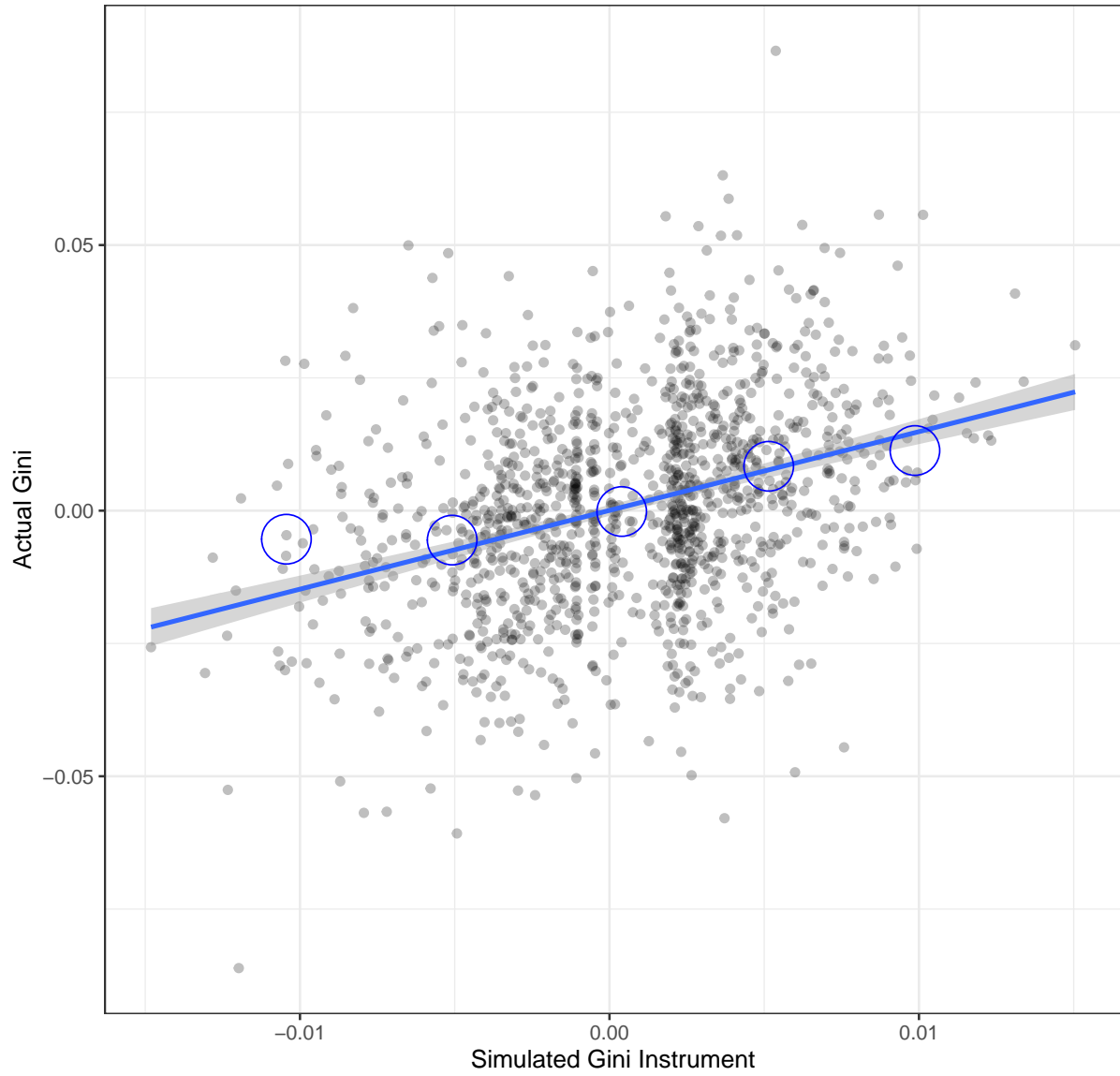


Figure A3: Testing Identifying Assumptions: Initial Inequality is Unrelated to Subsequent Changes in Ideology/Polarization

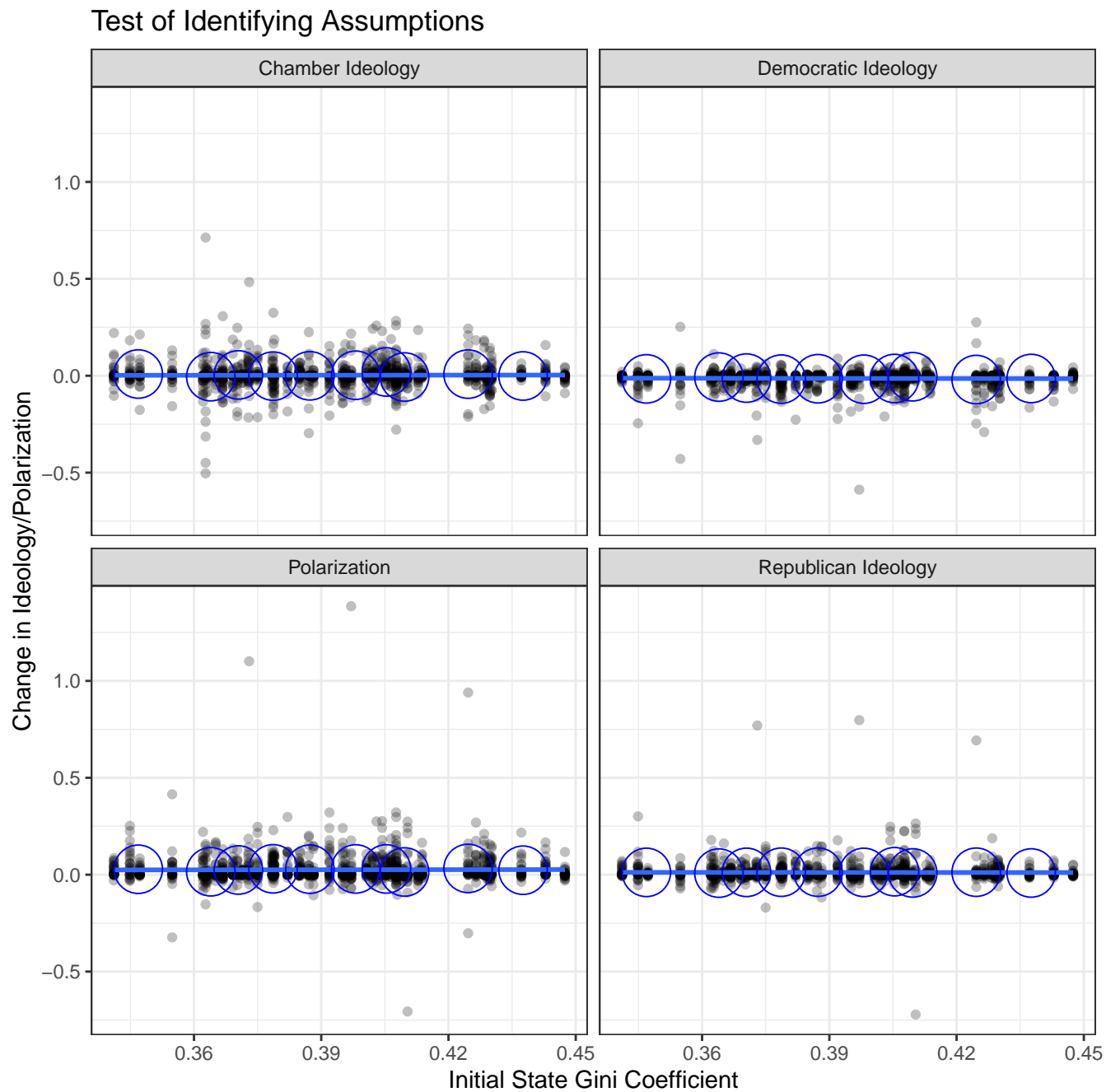


Table A8: First Stage Estimation Results

	<i>Dependent variable:</i>			
	Gini			
	First Diff.		Levels	
	(1)	(2)	(3)	(4)
Population	−0.057 (0.085)	−0.081 (0.091)	−0.005 (0.028)	0.026 (0.067)
Union Mem.	0.038 (0.044)	0.041 (0.044)	0.045* (0.025)	0.007 (0.036)
Perc. Minority-majority	−0.146 (0.111)	−0.160 (0.115)	0.068** (0.028)	−0.057 (0.082)
Black	−0.018 (0.049)	−0.018 (0.050)	0.040 (0.037)	−0.022 (0.040)
Marriage	−0.001 (0.001)	−0.001 (0.001)	−0.001 (0.001)	0.001 (0.001)
Noncitizen	0.000 (0.000)	0.000 (0.000)	0.000*** (0.000)	0.000 (0.000)
Age	0.0003 (0.001)	0.0003 (0.001)	0.0002 (0.0005)	0.0001 (0.0004)
DPI	1.585*** (0.149)	1.593*** (0.148)	0.398*** (0.062)	1.130*** (0.087)
First Stage F	97.43	92.48	140.39	67.79
Controls Variables?	Yes	Yes	Yes	Yes
State FE?	No	Yes	No	Yes
Observations	651	651	703	703

*Note:*

\*p&lt;0.1; \*\*p&lt;0.05; \*\*\*p&lt;0.01

## A.4 Robustness checks

### A.4.1 Polarization Results with Levels Data

The results in Tables 1-5 of the main text estimate a model in first-differences with state-specific trends. The following table shows estimates of the effect of inequality using levels data with state fixed effects as a robustness check of the primary results.

The specifications with levels data are as follows:

$$\text{First Stage: } Gini_{i,t-1} = \tilde{\alpha}_i + \theta Inst\_Gini_{i,t-1} + \Gamma X_{i,t} + \tilde{\mu}_i t + \nu_{i,t} \quad (8)$$

$$\text{Second Stage: } Y_{i,t} = \alpha_i + \beta_{2SLS} \widehat{Gini}_{i,t} + \gamma X_{i,t} + \mu_i t + \epsilon_{i,t} \quad (9)$$

Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2) reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats.

Table A9: Effects of Income Inequality with Levels Data using FE

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results: Gini</i>	2.180*** (0.787)	-1.185** (0.573)	0.995* (0.508)	2.962*** (0.764)	1.375** (0.523)
<i>OLS Results: Gini</i>	0.926** (0.397)	-0.525** (0.238)	0.401 (0.301)	1.172*** (0.371)	0.878*** (0.259)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	80.12	80.12	80.12	80.12	80.12
Observations	703	703	703	703	703

*Note:*

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

### A.4.2 First Differenced Polarization Results with Both Election Years and Non-Election Years

The results in Tables 1-5 of the main text estimate effects using differenced data from election years. The following table shows estimates of the effect of inequality using both election years and non-election years from 1997 to 2022. Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2) reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats.

Table A10: Effects of Income Inequality

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results: Gini</i>	1.048*** (0.311)	-0.510*** (0.162)	0.538** (0.215)	1.034*** (0.232)	0.690*** (0.159)
<i>OLS Results: Gini</i>	0.159 (0.116)	-0.136** (0.064)	0.023 (0.078)	0.211** (0.091)	0.160*** (0.056)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	148.93	148.93	148.93	148.93	148.93
Observations	1,335	1,335	1,335	1,335	1,335

*Note:*

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

### A.4.3 First Differenced Polarization Results with On-Year-Election States

The results in Tables 1-5 of the main text estimate effects using differenced data from all 50 states. The following table shows estimates of the effect of inequality using only states that have on-year elections. Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2)

reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats.

Table A11: Effects of Income Inequality for On-Year-Election States

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results: Gini</i>	1.211*** (0.340)	-0.577*** (0.175)	0.634** (0.242)	1.029*** (0.258)	0.642*** (0.176)
<i>OLS Results: Gini</i>	0.196 (0.126)	-0.153** (0.068)	0.044 (0.084)	0.222** (0.099)	0.158** (0.061)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	131.77	131.77	131.77	131.77	131.77
Observations	1,229	1,229	1,229	1,229	1,229

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

#### A.4.4 Polarization Results with Non-lagged Instrument and Gini

The results in Tables 1-5 of the main text estimate effects by lagging the instrument and Gini measures by one year. To rule out contemporaneous effects, we also check results without lagging. The following table shows estimates of the effect of inequality without lagging these variables. Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2) reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats. With the exception of the estimate for average chamber ideology, we do not see statistically significant effects of income inequality on political outcomes.

Table A12: Effects of Income Inequality

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results: Gini</i>	1.134** (0.533)	-0.675** (0.301)	0.459 (0.447)	2.244*** (0.595)	1.552*** (0.371)
<i>OLS Results: Gini</i>	0.097 (0.120)	-0.089 (0.083)	0.007 (0.085)	0.114 (0.141)	0.173* (0.091)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	84.74	84.74	84.74	84.74	84.74
Observations	651	651	651	651	651

*Note:*

\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$

#### A.4.5 Polarization Results with Once Forward Lagged Instrument and Gini

The results in Tables 1-5 of the main text estimate effects by lagging the instrument and Gini measures by one year. As a placebo test, we instead lag the instrument and Gini coefficient *forward* by one year, linking outcomes in year  $t$  to the values of inequality and its instrument in year  $t + 1$ . For example, the polarization measure for 2009 would be merged to the 2010 value of the Gini and instrument. Future values of the instrument and inequality should not have predictive power for current political outcomes if the causal direction runs from inequality to polarization. The following table shows estimates of the effect of inequality using an instrument and Gini measure that are lagged forward by one year. Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2) reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats. The results below show no statistically significant estimates, providing support that the causal direction indeed runs from inequality to polarization in our main analysis.

Table A13: Effects of Income Inequality

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results:</i> Gini	0.382 (0.474)	0.119 (0.284)	0.500 (0.368)	-0.701* (0.375)	-0.383 (0.234)
<i>OLS Results:</i> Gini	0.016 (0.197)	0.049 (0.110)	0.065 (0.140)	-0.117 (0.222)	0.046 (0.173)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	94.16	94.16	94.16	94.16	94.16
Observations	651	651	651	651	651

*Note:* \*\*\* p<0.01, \*\* p<0.05, \* p<0.1

#### A.4.6 Results Using Medians Instead of Means in Polarization Measure

In the main results of the paper, we measure average party ideologies using the mean rather than median to better capture changes at the extreme ends of each party distribution. Below, we provide results to the main analysis using the median.

Table A14: Effect of Income Inequality on State Legislative Polarization

	(1)	(2)	(3)
<i>IV Results: Gini</i>	1.447*** (0.425)	1.361*** (0.425)	1.456*** (0.428)
<i>OLS Results: Gini</i>	0.048 (0.140)	0.039 (0.143)	0.025 (0.144)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

Table A15: Effect of Income Inequality on Democratic Party Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	-0.768*** (0.267)	-0.646** (0.269)	-0.702** (0.273)
<i>OLS Results: Gini</i>	-0.187** (0.094)	-0.141 (0.099)	-0.129 (0.100)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

Table A16: Effect of Income Inequality on Republican Party Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	0.679** (0.325)	0.715** (0.326)	0.755** (0.325)
<i>OLS Results: Gini</i>	-0.139 (0.099)	-0.102 (0.099)	-0.104 (0.099)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

Table A17: Effect of Income Inequality on Average Chamber Ideology

	(1)	(2)	(3)
<i>IV Results: Gini</i>	3.107*** (0.530)	2.391*** (0.547)	2.416*** (0.547)
<i>OLS Results: Gini</i>	0.383** (0.159)	0.282* (0.153)	0.269* (0.153)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

Table A18: Effect of Income Inequality on Republican Seat Share

	(1)	(2)	(3)
<i>IV Results: Gini</i>	1.968*** (0.333)	1.479*** (0.353)	1.492*** (0.350)
<i>OLS Results: Gini</i>	0.252** (0.102)	0.166* (0.098)	0.150 (0.100)
State-Specific Linear Trend?	No	No	Yes
Other Controls?	No	Yes	Yes
First Stage F	105.5	97.43	92.48
Observations	651	651	651
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1		

#### A.4.7 Polarization Results with Two-Year Lagged Instrument and Gini

The results in Tables 1-5 of the main text estimate effects by lagging the instrument and Gini measures by one year. The following table shows estimates of the effect of inequality using an instrument and Gini measure that are lagged by two years. Each column corresponds to a different outcome variable,  $Y_{i,t}$ . Column (1) reports the estimated effect of income inequality on polarization; column (2) reports results for average Democratic ideology; column (3) for average Republican ideology; column (4) for average chamber ideology; and column (5) for the proportion of Republican seats.

Table A19: Effects of Income Inequality

	<i>Dependent variable:</i>				
	(1)	(2)	(3)	(4)	(5)
<i>IV Results: Gini</i>	0.838 (0.559)	-0.086 (0.259)	0.752* (0.432)	1.648*** (0.469)	0.935*** (0.333)
<i>OLS Results: Gini</i>	0.135 (0.103)	-0.044 (0.060)	0.091 (0.090)	0.290* (0.160)	0.154* (0.084)
State FE?	Yes	Yes	Yes	Yes	Yes
Other Controls?	Yes	Yes	Yes	Yes	Yes
First Stage F	56.8	56.8	56.8	56.8	56.8
Observations	651	651	651	651	651

*Note:*\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ **A.4.8 Independent Expenditure Results for Parties**

The campaign finance analysis in Tables 6-7 of the main text provide estimates for two outcome variables: polarization and average chamber ideology. Here we provide estimates for Republican and Democrat ideology.

Table A20: Citizens United Difference-in-Differences: IV Effect of Inequality on Republican Party Ideology by IE Limits

	(1)	(2)
LIM	-0.010* (0.005)	-0.009 (0.005)
Post	0.005 (0.008)	0.011 (0.007)
LIM × Post	0.001 (0.009)	-0.002 (0.009)
Gini	0.708 (0.725)	0.689 (0.627)
Gini × LIM	-0.146 (0.804)	-0.241 (0.746)
Gini × LIM × Post	0.018 (0.540)	0.327 (0.634)
Other Controls?	No	Yes
$Gini + Gini \times LIM + Gini \times LIM \times Post$	0.58	0.77
$P(Gini + Gini \times LIM + Gini \times LIM \times Post) = 0$	0.046	0.008
First Stage F	40.94	37.35
Observations	651	651

Note:

\*\*\* p<0.01, \*\* p<0.05, \* p<0.1

Table A21: Citizens United Difference-in-Differences: IV Effect of Inequality on Democratic Party Ideology by IE Limits

	(1)	(2)
LIM	0.005 (0.006)	0.005 (0.006)
Post	-0.018** (0.007)	-0.018** (0.007)
LIM $\times$ Post	0.001 (0.009)	0.001 (0.009)
Gini	-0.816** (0.406)	-0.801** (0.338)
Gini $\times$ LIM	0.698 (0.534)	0.712 (0.485)
Gini $\times$ LIM $\times$ Post	-0.897** (0.452)	-0.880* (0.528)
Other Controls?	No	Yes
$Gini + Gini \times LIM + Gini \times LIM \times Post$	-1.01	-0.97
$P(Gini + Gini \times LIM + Gini \times LIM \times Post) = 0$	0.002	0.004
First Stage F	40.94	37.35
Observations	651	651

Note:

\*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$

#### A.4.9 Heterogeneous Effects: Trade Exposure Using Median Threshold

In the main text, we estimate effects separately for states that have an above-average 1990 trade exposure and states that have below-average 1990 trade exposure. We split states using the mean trade exposure measure as the threshold for division. Here, we provide results for the analysis using the median rather than mean.

Table A22: Effects of Income Inequality for States with High Trade Exposure

	(1)	(2)
<i>IV Results: Gini</i>	0.951** (0.410)	3.329*** (0.877)
<i>OLS Results: Gini</i>	0.055 (0.147)	0.187 (0.172)
State-Specific Linear Trend?	Yes	Yes
Other Controls?	Yes	Yes
First Stage F	45.8	45.8
Observations	289	289
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1	

Table A23: Effects of Income Inequality for States with Low Trade Exposure

	(1)	(2)
<i>IV Results: Gini</i>	2.199* (1.216)	4.468*** (1.158)
<i>OLS Results: Gini</i>	-0.045 (0.259)	0.690** (0.319)
State-Specific Linear Trend?	Yes	Yes
Other Controls?	Yes	Yes
First Stage F	26.19	26.19
Observations	286	286
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1	

#### A.4.10 Heterogeneous Effects: Racial Composition Using Median Threshold

In the main text, we use the mean as an average measure of Black population proportions. We estimate effects separately for states with above-average minority compositions and those with below-average minority compositions. Here, we provide a robustness check using the median rather than mean to compare states with high and low minority compositions.

Table A24: Effects of Income Inequality for States with High Minority Composition

	(1)	(2)
<i>IV Results: Gini</i>	1.150*** (0.291)	1.737*** (0.479)
<i>OLS Results: Gini</i>	0.022 (0.152)	0.266* (0.155)
State FE?	Yes	Yes
Other Controls?	Yes	Yes
First Stage F	87.67	87.67
Observations	497	497
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1	

Table A25: Effects of Income Inequality for States with Low Minority Composition

	(1)	(2)
<i>IV Results: Gini</i>	2.215*** (0.582)	2.935*** (0.795)
<i>OLS Results: Gini</i>	0.082 (0.165)	0.289 (0.186)
State FE?	Yes	Yes
Other Controls?	Yes	Yes
First Stage F	53.51	53.51
Observations	493	493
<i>Note:</i>	*** p<0.01, ** p<0.05, * p<0.1	