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INCOME INEQUALITY AND POLITICAL POLARIZATION: TIME SERIES EVIDENCE OVER NINE DECADES

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Rising income inequality and political polarization have led some to hypothesize that the two are causally linked. Properly interpreting such correlations is complicated by multiple factors driving these phenomena, potential feedback between inequality and polarization, measurement issues, and the statistical challenges of modeling non-stationary variables. We find that a more precise measure of inequality (the inverted Pareto–Lorenz coefficient) is more consistently and statistically related to polarization in the short and long runs than the less precise top 1 percent share of income. We find bi-directional causality between polarization and inequality, consistent with theoretical conjecture and less formal evidence in previous studies.

JEL Codes: D31, D63, D72

Keywords: inequality, polarization

1. INTRODUCTION

Income inequality and political polarization have risen in the U.S., as illustrated in Figure 1 and documented by a host of inequality studies (e.g., Atkinson *et al.*, 2011) and tracked by Poole and Rosenthal's (1997, 2007) index of polarization in the U.S. Congress and the partisan thermometer ratings from the American National Election Studies (Prior, 2007). The coincidence of and controversy surrounding these trends has led some to hypothesize that increased inequality and political polarization are linked, such as Bartels (2008), Feddersen and Gul (2013), McCarty *et al.* (2002, 2006, 2013), and Stiglitz (2012). Understanding the factors that shape, and are shaped by, greater polarization can be important for assessing the macro-economic and political-economic prospects for the U.S. and possibly other countries. Indeed, the economic repercussions of increased polarization of Congress were evident in the 2011 political impasse over the Treasury's debt ceiling, which raised fears that the U.S. might default and prompted Standard &

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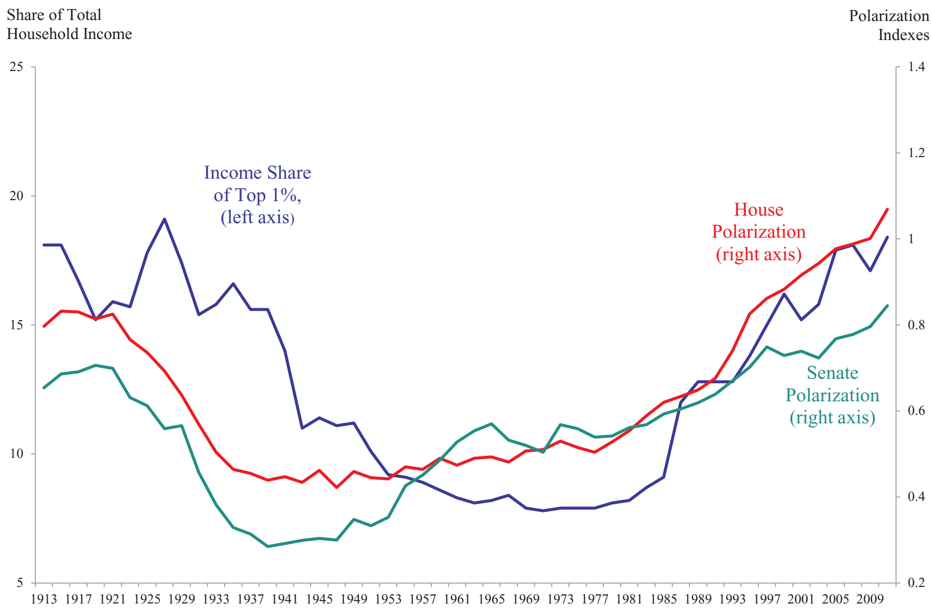

Figure 1. Political Polarization and the Income Share of the Top 1 Percent

Sources: Updates of Poole and Rosenthal's (1997, 2007) DWNominate Scores and inequality data of Piketty and Saez (2006), and authors' calculations converting the latter into corresponding biennial values.

Poor's (2011) to downgrade the credit rating of Treasury debt in September 2011.[1] And the more recent shutdown of the federal government in October 2013 directly lowered annual GDP growth in 2013q4 by nearly 1 percentage point.

While some political science researchers have studied polarization, most of their evidence relies on cross-sectional patterns, which limits our ability to understand and test potential key factors that may have shifted polarization and income inequality over time. So far, studies have mentioned correlations between income inequality and political polarization, but interpreting correlations is challenged by the multiple factors that drive each of these phenomena, potential feedback between inequality and polarization, and the statistical challenges of modeling non-stationary variables. In short, without careful analysis, it is difficult to determine what is driving what and why. This paper partly addresses this gap and contributes to the literature with time series tests of whether greater income inequality temporally leads to increased polarization and whether higher polarization temporally leads to greater income inequality.

Using a post-World War II (WWII) sample, Duca and Saving (2012a) found that media fragmentation had a stronger statistical association with polarization than did income inequality.[2] Furthermore, there are different potential channels

[1]For more on polarization and macroeconomic implications, see Mian *et al.* (2012); see also Bernanke (2012) and Yellen (2012), who make reference to both direct and indirect uncertainty effects of fiscal policy.

[2]For more on the effects of the rise of cable TV, see Baum and Kernell (1999) and Prior (2007).

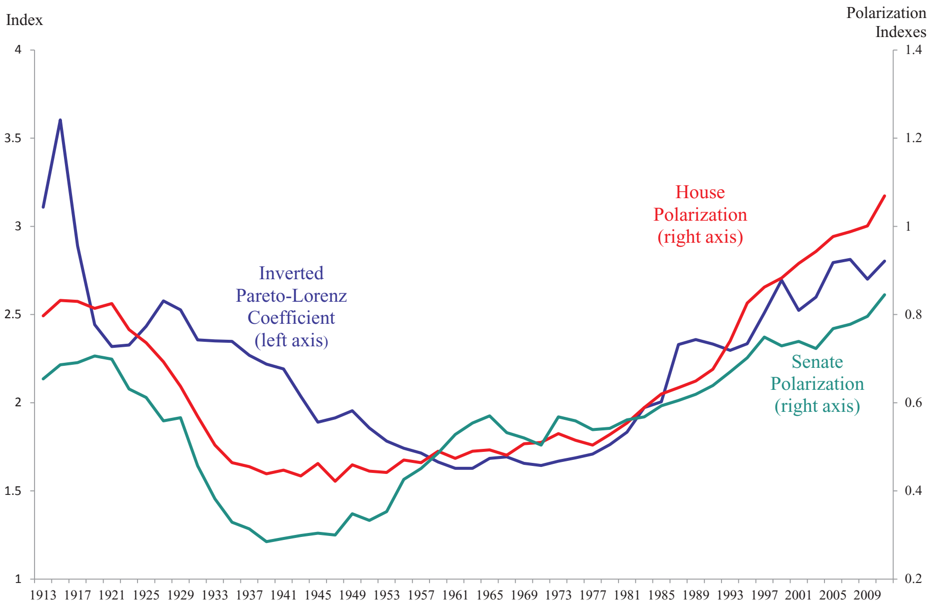

Figure 2. Political Polarization and a Broad Measure of Income Inequality

Sources: Updates of Poole and Rosenthal's (1997, 2007) DWNominate Scores and inequality data of Piketty and Saez (2006), and authors' calculations converting the latter into corresponding biennial values.

for how inequality may induce greater polarization, as well as different measures of inequality. When one goes beyond simple income shares as a measure of inequality, a more consistent and complicated statistical relationship between inequality and polarization emerges. This post-WWII finding is reflected when comparing Figure 1, which plots polarization and biennial averages of the income share of the top 1 percent of families, with Figure 2, which charts a more accurate gauge of income inequality, the inverted Pareto–Lorenz coefficient, which measures income inequality within the top 10 percent of families. Juxtaposing Figures 1 and 2 suggests that the more accurate measure of inequality (the inverted Pareto–Lorenz coefficient) is more consistently correlated with polarization indexes than the income share of the top 1 percent.

Duca and Saving (2012a) also found post-WWII evidence of bi-directional feedback between a precise measure of inequality (the Gini coefficient, available since 1947) and polarization, ostensibly reflecting the impacts of inequality on polarization and how shifts in federal policies may affect inequality (e.g., the New Deal until the Reagan revolution of the 1980s). In Figure 2, it appears that Senate polarization falls ahead of the broad inequality measure in the 1930s, while the latter rises slightly ahead of polarization since the early 1980s. Owing to limited data on their media fragmentation proxy variable, however, their study only covers a post-WWII sample, a limitation that makes it difficult to interpret the few low frequency trends in polarization, which fell in the 1930s, stayed low until the 1970s, and recently returned to pre-Depression era levels (McCarty *et al.*, 2013).

The current study analyzes time series data starting in 1913, when income statistics began with the start of the modern federal income tax, and carefully distinguishes between various measures of income inequality potentially relevant to the polarization debate. Our work provides rigorous time series evidence of bidirectional causality between polarization and inequality, consistent with theory and less formal evidence in studies by Feddersen and Gul (2013) and McCarty *et al*. (2002, 2006, 2013), among others.

To present these findings, Section 2 briefly reviews the selected measures of income inequality and factors influencing inequality, and then measures of political polarization and possible influences affecting it. Section 3 reviews possible interactions between the two types of data, as well as testable hypotheses. The fourth section presents and reviews the time series evidence on the statistical relationship between income inequality and political polarization. The conclusion provides some interpretation and perspective on the empirical results.

2. Possible Factors Influencing Income Inequality and Political Polarization

This section briefly reviews the empirical measures of income inequality and political polarization, and then discusses how each are influenced and driven by a variety of possible factors. Their endogenous response to outside influences has implications for both analyzing their time series relationship with one another and for interpreting the time series results.

2.1. *Measuring Income Inequality Over the Past 100 Years*

In this study, we use two measures of income inequality from Piketty and Saez (2006) that are available since 1913: the income share of the top 1 percent of families (*Top1%*) and the inverted Pareto–Lorenz coefficient (*IPL*, a term coined by Atkinson, 2003).[3] The latter is considered a more precise measure of income inequality for two reasons. First, as Atkinson *et al*. (2011) point out, the information content of arbitrary income shares like the top 1 percent share is subject to temporary income shocks and distortions that can be affected by the particular, arbitrary choice of a threshold level of income.[4] In contrast, the inverted Pareto–Lorenz coefficient used here describes how unequal the share of income is within a segment of the income distribution—specifically, within the top 10 percent of income. As a result, the inverted Pareto–Lorenz coefficient for families in the top 10 percent of the income distribution (Piketty and Saez 2006) is less noisy than the top 1 percent income share, as shown in Figure 3, and reflected in a notably smaller

[3]Partly because Piketty and Saez estimate *IPL* from data that exclude capital gains, the two inequality variables used are based on income excluding capital gains. Another reason is that in regressions not shown in the tables, fits were higher in models that used the top 1 percent share when income excluded capital gains, perhaps reflecting that the series inclusive of capital gains is noisier.

[4]Using measurements based on the Pareto distribution benefits from that distribution's property that the ratio of average income of those with incomes above a threshold y^h to the threshold level y^h does not depend on y^h. This ratio, β, equals $\alpha/(\alpha - 1)$, where α is a coefficient describing the cumulative distribution of income (y) used by Pareto: $1 - F(y) = (k/y)^{\alpha}$ ($k > 0$, $\alpha > 1$), with a corresponding density of income function $f(y) = \alpha k^{\alpha}/y^{(1-\alpha)}$. A lower level of α implies a more unequal distribution of income, and implies a higher level of β.

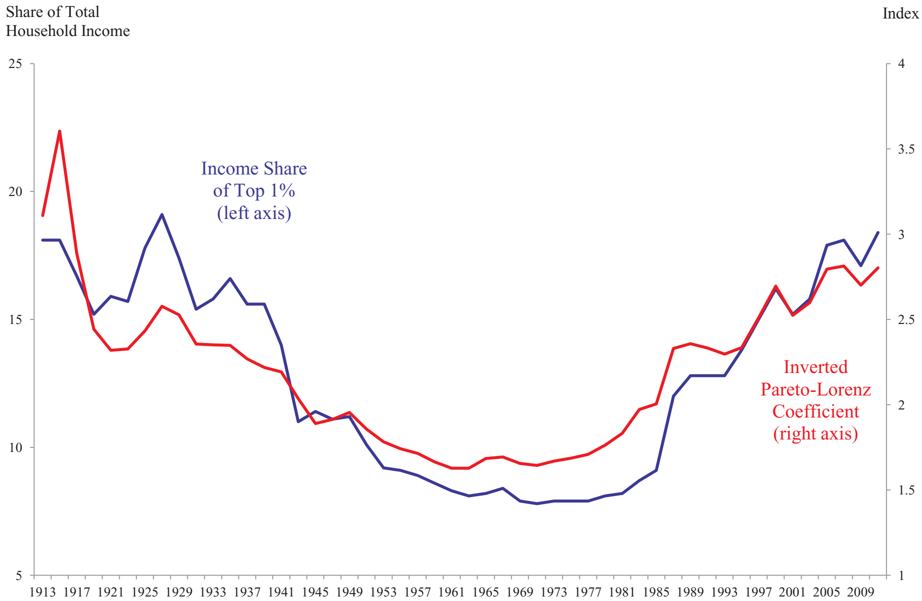

Figure 3. After WWI Top 1 percent Income Share Noisier than Inverted Pareto–Lorenz Coefficient

Sources: Updates of Poole and Rosenthal's (1997, 2007) DWNominate Scores and inequality data of Piketty and Saez (2006), and authors' calculations converting the latter into corresponding biennial values.

coefficient of variation (0.21 versus 0.30, respectively). This feature makes it easier to statistically identify long-run relationships using cointegration analysis, and likely for this reason, a significant long-run relationship between polarization and inequality is more consistently found using the less noisy inverted Pareto–Lorenz coefficient, as shown later.

The second reason to prefer the inverted Pareto–Lorenz coefficient is that Piketty and Saez construct it using more accurate data than are available for the top 1 percent income share. They construct the *IPL* using income data only from actual income tax returns on high income families, whereas the top 1 percent income share variable compares accurately measured income of the top 1 percent with pre-WWII estimates of total income derived from several sources. Because the U.S. income tax was levied only on high incomes before WWII, we have more consistent data on family income for upper-income households and the distribution of income within high income families than we have on family income for those in other parts of the income distribution. This allows the inverted Pareto–Lorenz coefficient for the upper 10 percent of incomes to be estimated more directly than either the top 1 percent income share or the more conceptually sound Gini coefficient, which is not considered here because it requires data on the entire income distribution that are unavailable before 1947. For this reason, neither the U.S. Census nor Piketty and Saez provide pre-1947 estimates of the U.S. Gini coefficient. It is reassuring, however, that the inverted Pareto–Lorenz coefficient

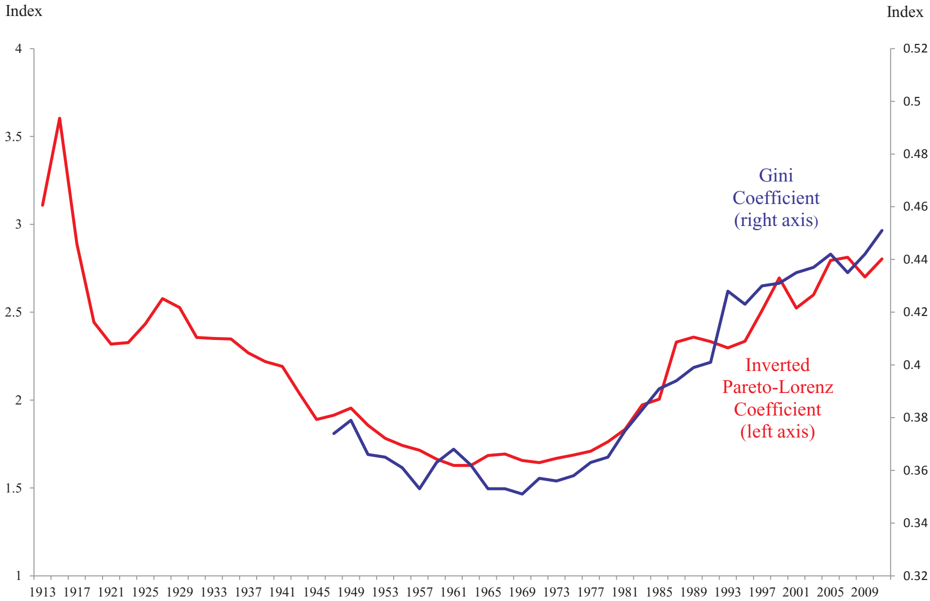

Figure 4. Post-WWII Gini and Inverted Pareto–Lorenz Coefficients Are Highly Correlated

Sources: Inequality data of Piketty and Saez (2006), and authors' calculations converting the latter into corresponding biennial values.

moves closely with the Gini coefficient (Figure 4) in the post-WWII era for which reliable Gini data are available.

2.2. *Measuring Political Polarization Over the Past 100 Years*

The measures of political polarization we analyze are the Poole and Rosenthal (1997, 2007) indexes of polarization in the House (*PolarH*) and Senate (*PolarS*). Unlike interest-group measures that consider only certain votes and are often geared toward finding certain results, the DW-Nominate scores from Poole and Rosenthal consider all votes without regard to partisan considerations. The polarization index for each is based on estimates of legislator ideal points from a spatial representation or mapping of legislator preferences along a scale between –1 and 1, where it is assumed that legislators' votes reflect their underlying preferences with some allowance for individual preferences to evolve. Using different criteria, one can construct an ideal point for legislators from each of the parties, and then calculate how much the average position of Democratic and Republican legislators has diverged at different points in time.[5] This polarization index tracks the ordinal or relative (not cardinal or absolute) preferences of the parties over time. We use the polarization indexes based on the primary dimension of differences that Poole

[5]For a non-technical description of the Poole and Rosenthal methodology, see McCarty (2010).

and Rosenthal track, which they argue corresponds to the role of the government in the economy in the modern sense of the terms liberal, moderate, and conservative.

Our time series analysis is applied to data covering the period 1913–2013, when data on both political polarization in congressional voting and income inequality are available. Including the pre-WWII period helps us better identify long-run relationships for two reasons. First, the time series is longer, which by adding more biennial readings makes cointegration analysis more applicable and relevant. Second, a post-WWII analysis is open to the concern that the data show a delayed upward trend that might be correlated with any factor having a similar trend. In contrast, recent polarization readings are in a range similar to those of the 1920s. If inequality is related to political polarization, then both series should move in similar ranges in recent decades as they did in the 1920s. Indeed, the inverted Pareto–Lorenz coefficient moves in a similar range in these two periods, and this is consistent with cointegration estimates reported below, which indicate a consistent long-run relationship between the two variables over the pre- and post-WWII periods. Three measures of U.S. income inequality are highly correlated since 1947 and likely track the same underlying trend in income inequality. Of these measures, the most accurate and complete one, the Gini coefficient, is available only since 1947. This leaves only two that are available since 1913. Of these, the inverted Pareto–Lorenz coefficient is preferable to the top 1 percent income share on both theoretical and empirical grounds as reflected in the analysis of Atkinson *et al.* (2011) and the relative coefficients of variation of these two variables.

2.3. *What Factors Influence Income Inequality and Political Polarization?*

Increased inequality and political polarization have been linked to several sources in the empirical literature, depicted by the uppermost and lowermost boxes of the flow chart in Figure 5. Greater inequality has arisen from changes in technology (the left-most middle box), which have generally reduced the returns to less-skilled labor and raised the skill/education premium (see, e.g., Lemieux, 2006; Goldin and Katz, 2007; Atkinson *et al.*, 2011). Additionally, technological changes, particularly since the 1970s, have also contributed to a fragmentation of visual media linked to the rise of cable TV (Baum and Kernell, 1999; Duca and Saving, 2012a, 2012b). This may contribute to increased political polarization, depicted in the lowest box of Figure 5, either through the effect of less news viewership because of more non-news entertainment alternatives (Prior, 2005, 2007) or through a "silo" effect of TV viewers self-sorting into watching news from biased sources that reinforce viewer priors (e.g., Sunstein, 2007; Iyengar and Hahn, 2009; Campante and Hojman, 2010; Gul and Pesendorfer, 2012).

Other, less technologically driven changes—listed in the far right box—have also contributed to higher inequality and political polarization. Shifts in demographic patterns, such as the assortative matching of people into pairs of highly educated couples have contributed to a less even distribution of income across families (see Fernandez and Rogerson, 2001) and have reduced a sense of common political interests, possibly contributing to greater polarization (see, e.g., Mann and Ornstein, 2012). In recent decades there has even been a shift to more income

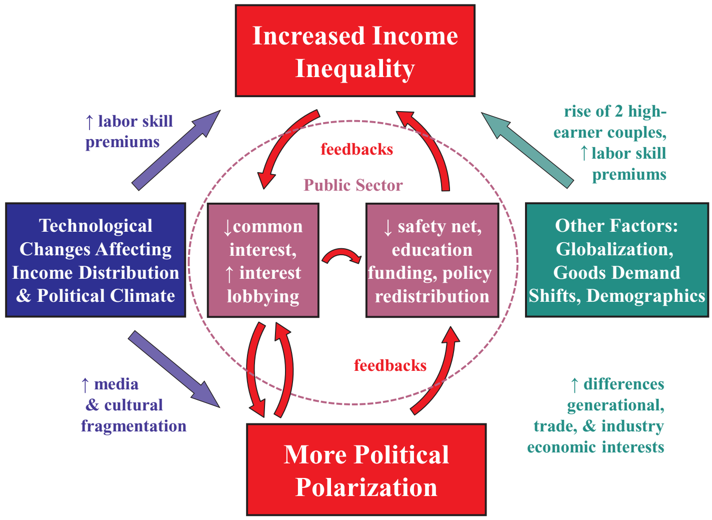

Figure 5. Influences on Income Inequality and Political Polarization

segregation across neighborhoods (Watson, 2009; Taylor and Fry, 2012) that may further reduce common interests. Other factors, such as the rise of globalization and shifts in goods demand toward new products (e.g., high tech), have also been linked to wider skill premiums and increased inequality (Dreher and Gaston, 2008; Cozzi and Impullitti, 2010). The rise of globalization has also been linked to a divergence in interests and voting behavior across socio-economic groups (Weck-Hannemann, 2001), which may induce greater polarization.[6]

As stressed by Atkinson (2003), among others, changes in public policy have also contributed to increased income inequality. As depicted by the second from the right middle box in Figure 5, the shift to a smaller safety net in the 1980s and 1990s and a somewhat less progressive income tax may have been factors,[7] as well as limits on federal and state government support for higher education amid a rising share of young Americans attending college. To some extent, this may reflect a feedback from greater political polarization, where feedback effects between polarization and income inequality are depicted with curved flow arrows in Figure 5. A reduced sense of common interest increases political polarization, which feeds back onto income inequality via less voter and legislative support for both income redistribution and higher education subsidies. The role played by factors affecting voter preferences is consistent with Poole and Rosenthal's (1997) finding that shifts in political polarization are less statistically linked to individual legislators changing their voting behavior ("conversions") and are more statistically linked to replacing members of Congress.

[6]For more on the various factors influencing the degree of political polarization, see McCarty *et al.* (2002, 2006, 2013), Poole and Rosenthal (1984), and Rosenthal (2004).

[7]Doerrenberg and Peichl (2012) find that changes in government spending, but not taxation, had significantly lowered inequality in a panel of OECD nations. OECD (2011, p. 270) found little change in the net impact of U.S. benefits and taxes on inequality between 1979 and 2004 comparing disposable and before tax income.

Taken together, these factors and interconnections have two important implications that relate to our work here. One is that income inequality and political polarization are not exogenous variables that are likely to be trendless or stationary. Several factors affect each or both of them, complicating how to interpret statistical relationships between them. Second, there are plausible bi-directional feedbacks, implying that inequality and political polarization may not be statistically exogenous to each other.

3. Estimating the Relationship Between Income Inequality and Political Polarization

This section describes the variables at our disposal to measure income inequality and political polarization, sets out the specifications we employ, and provides our statistical results.

3.1. *The Long-Run Variables Tracking Polarization and Income Inequality*

We consider four long-run variables available since 1913. These are the biennial Poole and Rosenthal indexes of polarization in the House and Senate (*PolarH* and *PolarS*, respectively) based on the revised methodology used to construct the latest estimates inclusive of the 2011 Congress and the biennial averages of Piketty and Saez's (2006) annual measures of the income share of the top 1 percent of families and the inverted Pareto–Lorenz (*IPL*) coefficient. The *IPL* is constructed based on income exclusive of capital gains and for comparability we use the Piketty and Saez top 1 percent income share series that omits capital gains (this series is more closely linked to polarization likely because the series inclusive of capital gains is noisier).

As Rosenthal (2004) notes, the House polarization index is more volatile than its Senate counterpart. Both are integrated of order 1, meaning that the first differences of each polarization index are stationary—as is the case for both inequality measures. Cointegration techniques for estimating long-run relationships between the levels of variables are suitable for I(1) variables. Accordingly, we estimate cointegration models of House and Senate polarization. One appeal of modeling the House is that all members are up for election every two years, whereas the composition of the Senate reflects members selected over three different elections. On the other hand, unlike the House, Senate elections are not potentially affected by gerrymandering.[8]

3.2. *Empirical Approach to Testing for Long-Run Relationships*

Cointegration analysis is also amenable to testing whether right-hand side variables are exogenous to the dependent variable, providing evidence on whether income inequality drives political polarization and/or the reverse. We use vector-error correction models (VECMs) to jointly estimate the long-run relationship between two variables, Y_1 and Y_2, in a cointegrating vector and short-run effects in first difference equations, respectively:

[8]McCarty *et al.* (2009) find that gerrymandering had little effect on polarization in the U.S. House.

$$Y_1 = \alpha_0 + \alpha_1 Y_2 \tag{1}$$

$$\Delta Y_1 = \beta_1[Y_1 - \alpha_0 + \alpha_1 Y_2]_{t-1} + \sum_{i=1} \gamma_i \Delta Y_{1t-i} + \sum \delta_i \Delta Y_{2t-i} + \lambda_1 X_t + \varepsilon_{1t}$$

$$\Delta \ln(Y_2) = \beta_1[\ln(Y_1) - \alpha_0 + \alpha_1 \ln(Y_2)]_{t-1} + \sum_{i=1} \gamma_i \Delta \ln(Y_2)_{t-i} + \sum \delta_i \Delta \ln(Y_1)_{t-i} + \lambda_2 X_t + \varepsilon_{2t}$$

where the lag length of first difference endogenous variables is selected to yield a cointegrating relationship which minimizes the Schwartz Information Criterion (SIC), X is a vector of exogenous factors, ε_{it} are residuals, and λ_i, γ_i, and $\delta_{i\ v}$ are row vectors of coefficients.

We experimented with several short-run variables used in past papers, including war deaths (with or without a dummy for the draft era), midterm congressional elections, the election of a new or re-election of an incumbent president (*Pres2nd* = 1 for the first Congress following a president's re-election), whether a president was being considered for impeachment (*Impeach* = 1 for the 1973 and 1997 congresses), and first or second terms of a president. None were consistently statistically significant. Lacking a strong theoretical argument for their inclusion, none are included in any of the specifications reported in the tables.

Some of our specifications do include a variable for the rise and fall of the New Deal, which a large political science literature treats as a "realigning political era" and may therefore merit special consideration (Campbell *et al.*, 1960). A variety of idiosyncratic factors including FDR's unique political gifts, the shared suffering of the Depression, and the perception that a once-in-a-lifetime opportunity to change the country's political center-of-gravity had been reached led an unusually broad set of voters to coalesce around New Deal policies (Converse, 1976).[9] This resulted in a temporary reduction in the ideological distance between Democrats and Republicans.[10,11] To control for a possible negative effect of the New Deal coalition on polarization, we added a short-run variable, *NewDeal*, which equals 1 for the 1931 to 1951 congresses, to span all the congresses during the FDR and Truman presidencies, plus the Congress elected in 1930, when the developing Great Depression contributed to large Republican losses in House and Senate races in the North.[12]

[9]These changes occurred through a combination of conversion, mobilization, and other factors. See Erikson and Tedin (1981) and Andersen (1979) for a fuller explanation of these points.

[10]This characterization is consistent with evidence from Achen and Bartels (2008) that the political realignment of the 1930s owed less to a shift in ideology among voters and more to their "retrospective" reaction in holding the Republican Party responsible for the Great Depression. They argue that "the great partisan realignment of this period was largely due to accumulation of myopic retrospections" (p. 7).

[11]Reduced polarization could have also arisen because the New Deal created a coalition of limited government and pro-segregationist southern Democrats with more interventionist, pro-civil rights northern progressives. Widening the dispersion in voting within the Democratic congressional delegation on economic matters and the role of the federal government may have narrowed the ideological distance between Democrats and Republicans.

[12]Using a slightly earlier endpoint for the New Deal variable produces qualitatively similar results.

Another potential realignment effect that may affect polarization is the recent rise of the "Tea Party," which has pressured Republicans in Congress to vote in ways that differentiate them more distinctly from their Democratic counterparts (Abramowitz, 2011). While it is too early to assess whether the Tea Party phenomenon will have as lasting an impact as that of the New Deal,[13] its effect over the last few years has been found to increase turnout among conservative portions of the electorate and drive moderates of both parties out of Congress, exacerbating polarization beyond what it would otherwise have been (Skocpol and Williamson, 2013). The variable *TeaParty* equals 1 for the 2011 Congress and 0 otherwise, and its inclusion essentially tests whether the 2011 polarization readings are outliers from other information and factors included in the models tested. Indeed, including both *NewDeal* and *TeaParty* can be interpreted as controlling for political realignment effects beyond those statistically reflected in the income inequality variables used (though we are careful to report results with and without these controls).

3.3. *Tests of Whether Inequality and Polarization Are Related in the Long Run*

Cointegration tests for polarization in the House and Senate using different sets of inequality and short-run variables are reported in Tables 1 and 2, respectively. In each table, Models 1–3 use *Top1%* and Models 4–6 use the inverted Pareto–Lorenz coefficient to track income inequality. In each table, Models 1 and 4 omit short-run variables outside of lagged first differences in polarization and inequality, while Models 2 and 5 include *NewDeal* and Models 3 and 6 include both political realignment variables *NewDeal* and *TeaParty*. In the long-run vectors, inequality is dated at time $t-1$ to reflect the economic conditions prevailing when the Congress was elected that votes in the time t biennial period. Much tighter long-run relationships and much higher corrected R-square statistics (10 percent and over 20 percent higher for corresponding House and Senate models, respectively) were obtained using this dating than if inequality at time t were used. This increased performance is consistent with the view that economic conditions at the time of elections affect the composition, perceived mandate, and behavior of the Congress over its term in office.[14]

The VECMs estimated use data spanning the 1913–2011 congresses. For all the House models, a lag length of 5 was used—the length needed to obtain a unique, significant cointegrating variable and/or minimize the SIC statistic and, if possible, also yield clean model residuals using the VECLM statistics on lags $t-1$ through $t-6$. Except for Model 1 (for which a lag length of 5 was selected), in Senate Models 2–6 a lag length of 6 was used, which was the lag length needed to obtain a unique, significant cointegrating variable and/or minimize the SIC statistic and, if possible, also yield clean model residuals using the VECLM statistics on lags $t-1$ through $t-6$.

[13]Even after a possible realignment plays itself out, researchers often fail to agree on whether it, in fact, represented a realignment, with these uncertainties magnified in the midst of the phenomenon (see Rosenof, 2003, for more on this issue).

[14]In addition to this plausible interpretation, another possible contributing factor is that aggregate inequality data are available with a lag of one to two years, and are not contemporaneously available in the information sets of voters.

TABLE 1

Biennial Models of Political Polarization in the U.S. House of Representatives Using Two Income Inequality Measures (1913–2011 Congresses spanning votes over 1913–2012)

Equilibrium Long-Run Relationship: $PolarH_t = \lambda_0 + \lambda_1 Inequality_{t-1}$

	$Top1\%_{t-1}$			IPL_{t-1}		
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant	–0.2472	–0.2478	–0.3014	0.0449	0.0999	–0.0556
$Inequality_{t-1}$	0.0293**	0.0292**	0.0248**	0.3084**	0.3346**	0.3135**
	(4.97)	(4.08)	(4.81)	(5.73)	(5.89)	(7.21)
Eig. (1 vec.)	0.262	0.199	0.371	0.285	0.272	0.416
Eig. (2 vec.)	0.018	0.003	0.001	0.015	0.001	0.000
Trace (1 v.)	13.81^{+}	9.71	20.28**	15.06^{+}	13.64^{+}	23.15**
Trace (2 v.)	0.76	0.15	0.36	0.65	0.00	0.01
Max-Eig (1 v.)	13.04^{+}	9.56	19.92**	14.41*	13.63^{+}	23.15**
Max-Eig (2 v.)	0.76	0.15	0.36	0.65	0.00	0.01
Cointegrated?	Yes$^{+,+}$	No	Yes**,**	Yes$^{+,*}$	Yes$^{+,+}$	Yes**,**

Short-Run Models: $\Delta PolarH_t = \alpha_0 + \alpha_1(EC)_{t-1} + \beta_i\Delta(PolarH)_{t-i} + \theta_i\Delta(\Delta Inequality_t)_{t-1-i} + \delta Y_t$

Sample	1927–2011	1927–2011	1927–2011	1927–2011	1927–2011	1927–2011
Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant	0.0033	0.0067	0.0051	0.0042	0.0096^{+}	0.0091^{+}
	(0.80)	(1.07)	(0.93)	(1.14)	(1.71)	(1.78)
EC_{t-1}	–0.096	–0.134	–0.190*	–0.121*	–0.162*	–0.192**
	(–1.29)	(–1.47)	(–2.45)	(–2.14)	(–2.41)	(–3.12)
$NewDeal_t$		–0.013	–0.015		–0.019	–0.022
		(–0.73)	(–1.01)		(–1.25)	(–1.61)
$TeaParty_t$			0.080**			0.058*
			(2.94)			(2.42)
$\Delta PolarH_{t-1}$	0.337^{+}	0.332^{+}	0.385*	0.277^{+}	0.251	0.262^{+}
	(1.83)	(1.79)	(2.32)	(1.71)	(1.53)	(1.72)
$\Delta PolarH_{t-2}$	0.497**	0.493**	0.492**	0.482**	0.454**	0.463**
	(2.72)	(2.68)	(3.00)	(3.01)	(2.81)	(3.08)
$\Delta PolarH_{t-3}$	–0.016	0.013	0.018	–0.082	–0.033	–0.036
	(–0.08)	(0.06)	(0.09)	(–0.47)	(–0.18)	(–0.22)
$\Delta PolarH_{t-4}$	0.017	0.006	0.058	0.045	0.014	0.024
	(0.09)	(0.03)	(0.32)	(0.28)	(0.09)	(0.16)
$\Delta PolarH_{t-5}$	0.400^{+}	0.401^{+}	0.426*	0.396*	0.368*	0.408*
	(1.91)	(1.90)	(2.27)	(2.34)	(2.17)	(2.57)
$\Delta Inequality_{t-2}$	0.000	–0.000	0.002	0.001	–0.010	–0.002
	(0.06)	(–0.02)	(0.52)	(0.01)	(–0.26)	(–0.06)
$\Delta Inequality_{t-3}$	0.003	0.004	0.025	0.021	0.015	–0.006
	(0.72)	(0.81)	(0.60)	(0.53)	(0.38)	(–0.17)
$\Delta Inequality_{t-4}$	0.001	0.001	0.002	0.082*	0.080*	0.065^{+}
	(0.25)	(0.30)	(0.43)	(2.04)	(2.00)	(1.72)
$\Delta Inequality_{t-5}$	0.002	0.003	0.007	–0.025	–0.015	0.002
	(0.46)	(0.65)	(1.50)	(–0.85)	(–0.47)	(0.07)
$\Delta Inequality_{t-6}$	0.002	0.004	0.003	0.063**	0.062**	0.052*
	(0.50)	(0.76)	(0.72)	(2.76)	(2.73)	(2.44)
Adjusted R^2	0.397	0.387	0.513	0.558	0.560	0.621
S.E.	0.0259	0.0261	0.0233	0.0222	0.0221	0.0205
VECLM(1)	2.19	1.69	1.20	7.43	8.48	6.38
VECLM(6)	2.24	2.51	0.97	1.79	2.25	2.00
F-Statistic: $NewDeal = TeaParty = 0$			4.69*			3.58*

TABLE 1 (*continued*)

Augmented Dickey–Fuller Unit Root Tests Using Schwartz Information Criterion (1913–2011 Congresses, covering 1913–2012)					
	Level (SIC lag)			Level (SIC lag)	
PolarH	–0.0743	(0)	Δ *PolarH*	–5.9245**	(0)
PolarS	–1.2093	(0)	Δ *PolarS*	–5.3137**	(0)
IPL	–1.5693	(1)	Δ *IPL*	–5.1137**	(0)
Top1%	–0.2959	(0)	Δ *Top1%*	–6.1188**	(1)

Notes: $^{+}$, * and ** denote 90%, 95%, and 99% significance levels, respectively. t-statistics in parentheses. Lag lengths of 5 yielded the strongest evidence for unique, significant vectors. The significance level of VECLM statistics accounts for size of the vector. Lag lengths for unit root tests are based on the Schwartz Information Criterion (SIC).

In cases when a statistically significant and unique cointegrating vector could not be identified, the lag length yielded was chosen to minimize the SIC statistic. In the models, inequality is lagged by an extra period—doing so yielded notably better model fits. This timing may reflect that the state of inequality leading up to a congressional election is linked to the polarization in voting behavior of members during the ensuing legislative session, consistent with Poole and Rosenthal's (1997) finding that changes in polarization are dominated by congressional turnover. The estimation allowed for possible time trends in the long-run variables without an independent time effect in the vector not attributable to measured factors. The lagged first differences shorten the estimation period to 1927–2011 for the House and to 1929–2011 for the Senate in all but one model (Senate Model 1, where the sample is from 1927 to 2011, reflecting a lag length of 5).

In two of the three models of House polarization shown in the upper panel of Table 1, a significant and unique cointegrating vector using *Top1%* could be identified, with a significant vector not found in Model 2 that only added *NewDeal* as a political realignment variable. The same pattern using *Top1%* arises in models of Senate polarization in the upper-panel of Table 2. In both tables, marginally significant cointegrating vectors could be identified for the models that omitted both political realignment variables *NewDeal* and *TeaParty*, whereas highly significant vectors could be identified in models that included both of these variables. These findings are implied by the eigenvalue and trace statistics rejecting the null hypothesis of no significant long-run relationship and by tests that accepted the null hypothesis of no more than 1 unique vector for Models 1 and 3 in each table. By itself, this finding implies that there is a statistically significant and consistent relationship between polarization and *Top1%* when both political realignment terms are included, while in their absence, there was only marginally significant evidence of a long-run cointegrating relationship. This pattern, coupled with the joint significance of the two realignment variables, implies that inequality is not the only factor that notably affects political polarization in congressional voting behavior.

Another condition for an error-correction model to be valid is that the lagged error-correction term in the short-run model (lower panel of Table 1) should be statistically significant and negative, implying that the time t change in polarization tends to be negative if in the previous time period the actual level of

TABLE 2

Biennial Models of Political Polarization in the U.S. Senate Using Two Income Inequality Measures (1913–2011 Congresses spanning votes over 1913–2012)

Equilibrium Long-Run Relationship: $PolarH_t = \lambda_0 + \lambda_1 Inequality_{t-1}$

	$Top1\%_{t-1}$			IPL_{t-1}		
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant	0.4171	0.2657	0.4016	0.2233	0.0975	0.2083
$Inequality_{t-1}$	0.0092	0.0217*	0.0104^{+}	0.1450*	0.2051**	0.1522**
	(1.32)	(2.56)	(1.73)	(2.07)	(3.51)	(3.10)
Eig. (1 vec.)	0.276	0.240	0.339	0.266	0.372	0.484
Eig. (2 vec.)	0.000	0.000	0.000	0.033	0.008	0.009
Trace (1 v.)	13.90^{+}	11.53	17.40*	14.40^{+}	19.92**	28.11**
Trace (2 v.)	0.01	0.00	0.00	1.39	0.35	0.36
Max-Eig (1 v.)	13.90^{+}	11.53	17.40*	13.01^{+}	19.57**	27.76**
Max-Eig (2 v.)	0.01	0.00	0.00	1.39	0.35	0.36
Cointegrated?	Yes$^{+,+}$	No	Yes*,*	Yes$^{+,+}$	Yes**,**	Yes**,**

Short-Run Models: $\Delta PolarS_t = \alpha_0 + \alpha_1(EC)_{t-1} + \beta_i\Delta(PolarS)_{t-i} + \theta_i\Delta(Inequality)_{t-1-i} + \delta Y_t$

Sample	1927–2011	1929–2011	1929–2011	1929–2011	1929–2011	1929–2011
Variable	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Constant	0.0038	0.0299**	0.0241*	–0.0081	0.0211**	0.0184**
	(0.74)	(3.64)	(3.34)	(–1.60)	(3.85)	(2.88)
EC_{t-1}	–0.023	–0.193*	–0.264**	–0.149*	–0.206**	–0.245**
	(–0.27)	(–2.38)	(–2.84)	(–2.42)	(–3.02)	(–3.98)
$NewDeal_t$		–0.089**	–0.076**		–0.055*	–0.052**
		(–3.50)	(–3.55)		(–2.60)	(–2.73)
$TeaParty_t$			0.080*			0.061*
			(2.66)			(2.16)
$\Delta PolarS_{t-1}$	0.272	0.416*	0.427**	0.344*	0.295*	0.268*
	(1.48)	(2.57)	(2.84)	(2.46)	(2.25)	(2.16)
$\Delta PolarS_{t-2}$	0.308^{+}	0.119	0.176	0.287^{+}	0.193	0.222^{+}
	(1.79)	(0.75)	(1.19)	(1.97)	(1.38)	(1.68)
$\Delta PolarS_{t-3}$	–0.030	–0.105	–0.160	0.058	0.040	0.040
	(–0.16)	(–0.66)	(–1.09)	(0.41)	(0.31)	(0.32)
$\Delta PolarS_{t-4}$	0.222	0.258	0.315*	0.179	0.118	0.134
	(1.21)	(1.62)	(2.11)	(1.30)	(0.90)	(1.08)
$\Delta PolarS_{t-5}$	0.093	–0.134	–0.073	0.106	–0.003	0.025
	(0.05)	(–0.84)	(–0.49)	(0.74)	(–0.02)	(0.19)
$\Delta PolarS_{t-6}$		–0.110	0.010	0.167	0.065	0.137
		(–0.69)	(0.07)	(1.16)	(0.47)	(1.01)
$\Delta Inequality_{t-2}$	0.007	–0.008	–0.024	–0.062	–0.084*	–0.057
	(0.12)	(–1.60)	(–0.48)	(–1.34)	(–2.00)	(–1.39)
$\Delta Inequality_{t-3}$	–0.004	0.001	0.001	0.059	0.026	0.025
	(–0.68)	(0.11)	(0.23)	(1.25)	(0.60)	(0.58)
$\Delta Inequality_{t-4}$	0.041	–0.001	0.003	0.095^{+}	0.069	0.076^{+}
	(0.65)	(–0.26)	(0.60)	(1.90)	(1.51)	(1.73)
$\Delta Inequality_{t-5}$	0.001	0.008	0.014*	0.042	0.044	0.098^{+}
	(0.24)	(1.64)	(2.56)	(0.86)	(0.97)	(1.99)
$\Delta Inequality_{t-6}$	–0.006	–0.005	–0.005	–0.011	–0.004	–0.032
	(–0.97)	(–1.04)	(–1.06)	(–0.31)	(–0.13)	(–0.98)
$\Delta Inequality_{t-7}$		0.015**	0.016**	0.125**	0.090**	0.098**
		(2.95)	(3.35)	(4.85)	(3.46)	(3.97)
Adjusted R^2	0.123	0.414	0.500	0.500	0.569	0.619
S.E.	0.0336	0.0269	0.0248	0.0248	0.0230	0.0217
VECLM(1)	10.37*	3.48	4.49	0.58	2.22	2.27
VECLM(6)	0.71	5.10	6.34	1.50	0.36	0.42
F-Statistic: *NewDeal* = *TeaParty* = 0			10.78**			5.76**

Notes: $^{+}$, * and **denote 90%, 95%, and 99% significance levels, respectively. t-statistics in parentheses. Lag lengths of 6 yielded the strongest evidence for unique, significant vectors using models 2–6, while a lag length of 5 worked best in finding a unique vector in model 1. Using a lag length of 6 in model 1 a unique cointegrating vector that was at least marginally significant could not be found. The significance level of VECLM statistics accounts for size of the vector.

polarization was above its equilibrium level. For this reason, it is important to estimate a long-run equilibrium relationship and to assess whether it helps explain short-run changes.[15] In models using *Top1%* the error-correction term is significant at the 95 percent confidence level only in House Model 3 and Senate Models 2 and 3. This suggests mixed evidence that long run deviations of polarization from the equilibrium levels implied by *Top1%* may help explain short-run changes in polarization.

In each table, Models 4–6 correspond to Models 1–3 except that they replace *Top1%* with the inverted Pareto–Lorenz measure of income inequality (*IPL*). In contrast to those using *Top1%*, a marginally significant or significant unique cointegrating vector could be identified in every model using *IPL*. In addition, the long-run inequality coefficient is more highly significant in the corresponding House polarization models when *IPL* is used. In each of the three *IPL* models of House polarization, the error-correction term is significant, especially in Model 4 that omits political realignment variables and Model 6 that includes both. A significant and negatively signed error-correction coefficient is important because it means that polarization changes in period t tend to reduce the previous gap between the levels of polarization and its estimated long-run equilibrium. Otherwise, the estimated long-run relationship provides no information on short-run changes in polarization.

Even more noteworthy, in Models 4–6 of Senate polarization the long-run *IPL* coefficient is always statistically significant, whereas *Top1%* is insignificant in Model 1 and is only marginally significant in Model 3. In the Senate Models 4–6 (Table 2) that use *IPL*, a statistically significant and negative error-correction coefficient was estimated in each case, whereas in Models 1–3 using *Top1%*, the error-correction term was marginally significant when both political realignment variables were present in Model 3, with the error-correction term only significant in Model 2 that adds only *NewDeal*.

The estimated long-run coefficients for the best fitting House and Senate models use the more accurate measure (*IPL*) rather than *Top1%* to track inequality and include the jointly significant short-run political realignment variables, whose inclusion raises the significance of the estimated cointegrating relationship when comparing Models 4 and 6 for the House and Senate. These vectors, from Model 6 in Tables 1 and 2, imply that equilibrium polarization rises with income inequality as tracked by the inverted Pareto–Lorenz coefficient (*IPL*):

$$\text{(2)} \qquad \text{House:} \quad PolarH = -0.056 + \underset{(7.21)}{0.314}\, IPL^{**} \quad \text{(Model 6, Table 1)},$$

$$\text{(3)} \qquad \text{Senate:} \quad PolarS = 0.208 + \underset{(3.10)}{0.152}\, IPL^{**} \quad \text{(Model 6, Table 6)},$$

where t-statistics are in parentheses and ** denotes significance at the 99 percent confidence level.

[15]Also, the estimated short-run models reflect that short-run movements stem from changes in relevant variables and the lagged adjustment of actual levels of a variable to its long-run equilibrium.

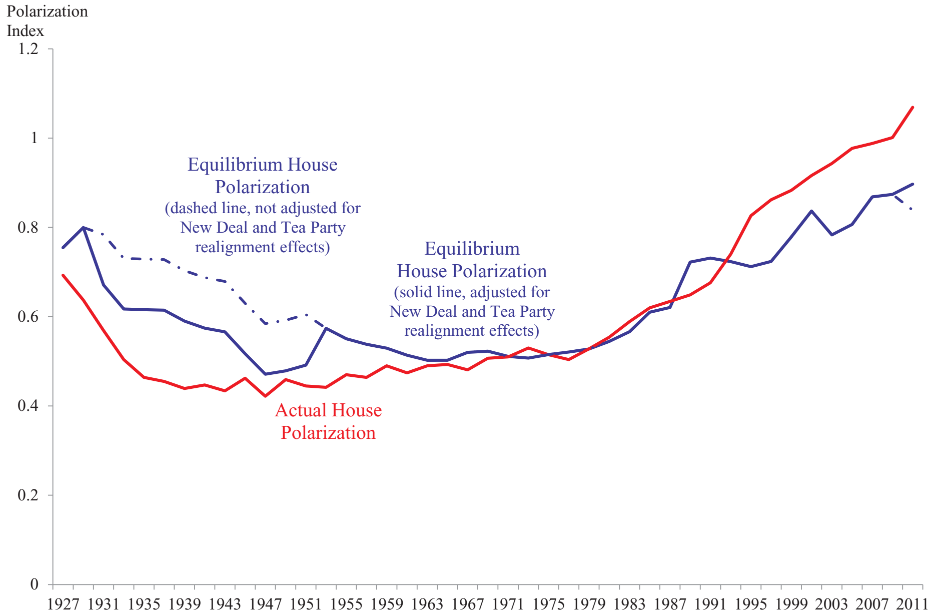

Figure 6. Inequality-Based Equilibrium Estimates Track House Polarization Trends Well

Sources: Updates of Poole and Rosenthal's (1997, 2007) DWNominate Scores and authors' calculations from results for model 6 in Table 1.

The respective House and Senate polarization indexes in Figures 6 and 7 weakly line up with the long-run equilibrium values from equations (2) and (3), which are adjusted for the constant in the short-run equation but *not* for the more persistent estimated effects of the New Deal realignment shift variable and the estimated one-off short-term Tea-Party effect. The simple equilibrium relationships from these equations miss the U-shaped drops in polarization coinciding with the forming and fraying of the New Deal coalition. This can be viewed as a medium term effect spanning the 10 congresses covered by the New Deal 0–1 variable. To adjust for this medium term effect, we add to the respective equilibrium values the dummy multiplied by its estimated short-run coefficient divided by the corresponding estimated speed of adjustment. The resulting equilibrium estimates line up well with actual polarization indexes in Figures 6 and 7, in contrast to estimates lacking this realignment effect. Nevertheless, there is no clear pattern of the implied equilibrium levels moving before the actual polarization readings. This suggests there may be bidirectional feedback between inequality and polarization, an issue examined in Section 4.

3.4. *Tests of Whether Inequality and Polarization Are Related in the Short Run*

Results for the corresponding short-run models of changes in House and Senate polarization are in the lower panels of Tables 1 and 2, respectively. Consistent with longer-run results mentioned above, model fits also reflect that the inverted Pareto–Lorenz coefficient contains notably more information for modeling polarization than the *Top1%* income share. Among the House results in Table 1, the

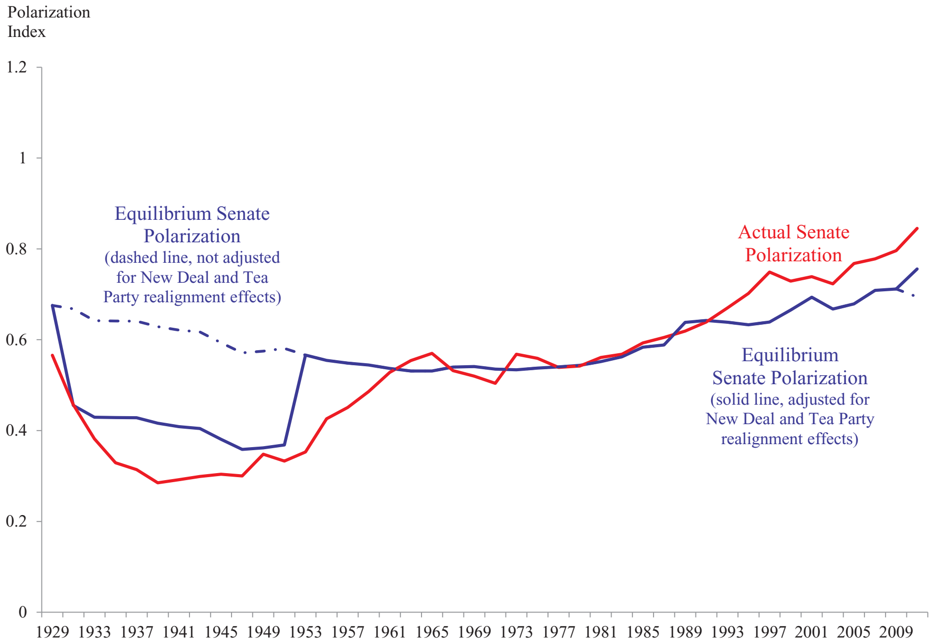

Figure 7. Inequality-Based Equilibrium Estimates Track Senate Polarization Trends Well

Sources: Updates of Poole and Rosenthal's (1997, 2007) DWNominate Scores and authors' calculations from results for model 6 in Table 2.

models using *IPL* account for 11–17 percent more of the variance in polarization than the corresponding models that use the *Top1%* income share variable (Models 4 versus 1, 5 versus 2, and 6 versus 3). The improvement in fit is also notable for modeling polarization in the Senate, where corresponding models using a lag length of 6 account for 11–38 percent more of the variance than their *Top1%* counterparts.

The better fit of corresponding models that use *IPL* reflects that *IPL* contains more information not only in the short run through lagged first difference effects, but also through long-run relationships as implied by generally more significant error-correction terms than those in corresponding models using the *Top1%* share measure of inequality.

Recall that the statistically significant, negative coefficients on the error-correction terms in each of the six models that use *IPL* imply that polarization falls in a congressional session if, in the prior session, polarization had been above the long-run equilibrium implied by its relationship with income inequality. In terms of yielding stronger long-run relationships that have information for short-run movements in polarization and of yielding better fitting models of short-run changes in polarization, models that use the more carefully and rigorously gauged *IPL* outperform and are preferable to corresponding models using the *Top1%* share. Hence, for both the House and Senate, among models omitting extra short-run controls for political realignment (Models 1 and 4 in each of Tables 1 and 2), Model 4 is preferable to Model 1, while among models including political realignment controls, Model 6 is preferable to Model 3.

Both the New Deal and Tea Party political realignment variables are statistically significant, with expected negative and positive signs, respectively. One interesting pattern is that while the qualitative results using *IPL* are unaffected by the inclusion of these two political realignment variables, the model fits are notably higher when both are present. This not only reflects their joint statistical significance, but also that the error-correction coefficients (speeds of adjustment) are higher and more significant because their presence helps better identify the long-run relationships between polarization and income inequality. For example, in Model 4 of the House index in Table 1, the magnitude of the error- or equilibrium-correction coefficient implies that about 12 percent of the gap between actual and equilibrium polarization is eliminated in the following congressional session. In Model 6 of Table 1, which is identical, except that it includes *NewDeal* and *TeaParty*, the speed is higher at 19 percent. The impact of including the two realignment variables on the estimated speed of adjustment is much larger for polarization in the Senate. The estimated speed of adjustment is 15 percent per session in Model 4 in Table 2, which omits the two realignment variables, versus a much faster 25 percent per session when *NewDeal* and *TeaParty* are added in Model 6. The improvement in the estimated speed of adjustment also likely reflects that including both realignment terms picks up omitted short-run influences.

In terms of the specification yielding evidence of the strongest long-run relationship and best fit in explaining short-run changes, Model 6 is the preferred one. That said, there may be a perceived subjective element to the definitions of short-run variables like *New Deal* and *Tea Party*, and Model 4 is the preferred specification among the two models that omit them (Models 1 and 4). Nevertheless, coefficients from these models estimated over a sample that omits the New Deal and Tea Party periods were similar to those of Model 6 for the House and Senate.[16]

It is also reassuring that when comparing Models 4 and 6 in the upper-panel of each table, the inclusion of realignment variables raises the significance of the overall estimated cointegrating relationship and that of the long-run coefficient on *IPL*, but hardly affects the estimated size of the long-run constant and the coefficients on *IPL*. This robustness of long-run coefficient estimates, coupled with faster estimated speeds of adjustment toward long-run equilibrium in the short run and notably improved model performance in tracking short-run changes, suggests that inclusion of the realignment variables tracks effects on polarization other than those directly attributable to income inequality.

4. Is Income Inequality Exogenous to Polarization? Is There Bidirectional Causation?

As discussed in Section 2, both income inequality and political polarization are likely affected by several variables in the long run. Moreover, the long-run causation need not just be from income inequality to political polarization mainly because inequality has been found to be affected by shifts in public policy that may reflect the effects of polarization on government spending, taxes, and trade policy. Indeed, bi-directional causality is a central theme of much of the work on political

[16]In this shorter sample, highly significant and unique cointegrating vectors were found.

polarization (see, e.g., McCarty *et al.*, 2006, 2013). Partly to examine empirically whether causation between these two variables is bi-directional, the models presented earlier were estimated as a vector error-correction model containing separate equations for changes in polarization and income inequality, which were regressed on the same error-correction term along with the same lags of changes in the long-run variables and same set of short-run variable(s). If the error-correction term is significant in the model of polarization but is insignificant in the model of inequality, then formal econometric evidence indicates that income inequality is "weakly exogenous" to polarization, as discussed in Urbain (1992) and which Granger and Lin (1995) would have described as evidence that income inequality is caused, in a long-run sense, by political polarization. If the error-correction term is significant in both VECM component equations of changes in polarization and inequality, then there is evidence of bidirectional causality.

As reported in Table 3, income inequality "causes" polarization in each of the three House models and in each Senate model that uses the superior *IPL* measure of inequality. In all six IPL models (Models 4–6 in Tables 1 and 2), polarization "causes inequality." The joint significance of the two political realignment variables (reported for Models 3 and 6 in Tables 1 and 2) implies that the Model 6 results indicating bidirectional causality are the most relevant. For those with doubts about the endogeneity of the realignment variables, the results from Model 4, which omits

TABLE 3
WEAK EXOGENEITY TESTS

A. Testing Whether *Polarization* is Weakly Exogenous to the *IPL* Inequality Variable

Estimate Short-Run Model: $\Delta(Polar)_t = \alpha_0 + \alpha_1(EC)_{t-1} + \beta_i\Delta(Polar)_{t-i} + \theta_i\Delta(IPL)_{t-i} + \delta Y_t$

Test whether α_1 is equal to zero: for the House, rejected in Models 4–6; and for the Senate, rejected in Models 4–6.

Variable	Model 4	Model 5	Model 6
House Polarization			
EC_{t-1}	−0.121*	−0.162*	−0.192**
	(−2.14)	(−2.41)	(−3.12)
Senate Polarization			
EC_{t-1}	−0.149*	−0.206**	−0.245**
	(−2.42)	(−3.29)	(−3.98)

B. Testing Whether the *IPL* Inequality Variable is Weakly Exogenous to *Polarization*

Estimate Short-Run Model: $\Delta(IPL)_t = \alpha_0 + \alpha_1(EC)_{t-1} + \beta_i\Delta(Polar)_{t-i} + \theta_i\Delta(IPL)_{t-i-1} + \delta Y_t$

Test whether α_1 is equal to zero: for the House, rejected in Models 4–6; and for the Senate, rejected in Models 4–6.

Variable	Model 4	Model 5	Model 6
House Polarization			
EC_{t-1}	0.647**	0.751**	0.704**
	(2.77)	(2.80)	(3.15)
Senate Polarization			
EC_{t-1}	0.506*	0.582*	0.663*
	(2.22)	(2.34)	(2.58)

Notes: $^{+}$, * and **denote 90%, 95%, and 99% significance levels, respectively. t-statistics in parentheses.

both realignment variables, are most relevant. In those parsimonious models of House and Senate polarization, there is strong evidence of bidirectional causality.[17]

In other models that replace the inverted Pareto–Lorenz coefficient (*IPL*) with the top 1 percent income share (not shown to conserve space), the evidence is stronger for causality running from polarization to inequality and less strong for causality running from inequality to polarization. This may reflect either that the top 1 percent share is more prone to measurement error or that special interest lobbying is both encouraged in a highly polarized environment and tends to favor the extremely rich, as tracked by the income share of the top 1 percent. In general, the statistical evidence suggests bidirectional effects or important feedbacks between income inequality and polarization that are illustrated in Figure 5. This is consistent with the view that increased inequality makes it more difficult to achieve political consensus, either through undermining a sense of common interest or through fostering more rent-seeking, and that polarization undermines support for redistributive policies, thereby inducing greater inequality.

5. Conclusion

Using appropriate time-series techniques, this paper examines the statistical relationships between income inequality and the congressional polarization indexes of Poole and Rosenthal, both of which have trended up sharply in recent decades. While public discussion of a possible relationship has focused on tracking income inequality with the income share of the top earning 1 percent of families, this measure is noisy and may not accurately track inequality as well as other measures, such as the inverted Pareto–Lorenz coefficient, as argued by Atkinson *et al.* (2011). Indeed, absent controls for New Deal and Tea Party political realignment effects, information from the long-run relationship between income inequality and political polarization does not consistently add statistically significant information about short-run changes in House and Senate polarization if inequality is measured by the top 1 percent income share. However, regardless of whether both of these controls for realignment are included, inequality adds more information about polarization using the more accurate gauge of income inequality, the inverted Pareto–Lorenz coefficient. When this more precise measure is used, there is stronger evidence that income inequality is cointegrated with polarization and short-run changes in polarization are better explained. These results reflect the advantage of using more accurate gauges of income inequality, as stressed by Atkinson (2003) and Atkinson *et al.* (2011), that match the care and precision taken by Poole and Rosenthal (1997, 2007) in measuring political polarization.

A second, important contribution of this paper is in documenting the existence of long-run, bi-directional feedback effects between income inequality and political polarization. This finding is consistent with the view that both income inequality and political polarization are endogenous variables that feedback onto each other (McCarty *et al.*, 2002, 2006, 2009, 2013; Bartels, 2008). This reinforces

[17]If inequality and polarization enter the cointegrating vector in Model 6 with the same $t-1$ lag, significant bidirectional causality is found for both House and Senate polarization. The same obtains for estimating Model 3 using data from the pre-Tea Party sample of 1913–2009.

concerns in some quarters that increased income inequality: (1) weakens the perception of shared destiny and thereby spawns political polarization; (2) induces more special interest rent-seeking by the very wealthy through concentrating the gains from lobbying; and/or (3) fosters a perception that one's political opponents are working against the national interest, which limits support for social-insurance programs that encompass all segments of American society.

Aside from these two contributions to the empirical literature, this study's results have other political-economic implications. Until other structural changes affecting income inequality or inducing political reform or realignment occur that reduce discord between the political parties, our findings suggest that a high degree of polarization in congressional voting is likely to persist. In this plausible scenario, continued political polarization could have major economic ramifications, even beyond short-run economic effects, such as those from the shutdown of the federal government in late 2013. For example, the lack of political consensus for addressing the U.S.'s long-run fiscal challenges was the main rationale mentioned by Standard & Poor's (2011) when it downgraded the credit rating of U.S. Treasury debt. The long-term budgetary challenges posed by entitlement programs have been exacerbated by the rise in national debt during the Great Recession, with no clear indication of how or whether legislators will bridge their differences to address them—and whether a highly polarized electorate will even allow them to do so.

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