Has Income Inequality or Media Fragmentation Increased Political Polarization?

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Abstract

The increasing polarization of Congressional voting patterns has been attributed to factors including generational shifts, economic conditions, increased media fragmentation, and greater income inequality. The first of these factors is difficult to test with time series data owing to the low frequency of generational shifts, while the tendency of business cycles to reverse suggests that economic cycles are unable to account for long-term shifts in polarization. This leaves two main possible long-run drivers: the increasingly fragmented state of American media as stressed by Prior (2005, 2007) and Duca and Saving (2012a), and increased income inequality, as emphasized by McCarty, Poole, and Rosenthal (2006, forthcoming) and Stiglitz (2012).

Using statistical techniques suitable for analyzing variables with shifting long-run averages we find evidence indicating that media fragmentation has played a more important role than inequality, at least as tracked by available data and measures. Periods when the share of Americans with access to cable or satellite TV has risen are followed by upward shifts in polarization. Furthermore, our results suggest that the polarization arising from media fragmentation or inequality may make it more difficult to achieve the political consensus needed to address major challenges, such as the long-run fiscal imbalances facing the United States.

JEL Codes: D72, G11

Keywords: income inequality, media fragmentation, political polarization

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American politics has become increasingly polarized in recent years, as reflected in indicators such as Poole and Rosenthal's (1997, 2007) index of polarization in the U.S. House of Representatives and Senate (Figure 1) or the partisan thermometer ratings from the American National Election Studies (Prior, 2007). Polarization has been shown to have a variety of negative social ramifications as people turn from addressing shared problems and instead "bowl alone" (Putnam, 2000). This in turn makes it more difficult to address pressing policy matters, sometimes with serious economic consequences.

A prominent example is the "fiscal cliff" issue facing the U.S. economy in late 2012 and early 2013. Without the approval of new legislation, sizable increases in taxes and cuts in government spending could occur in early 2013 that are large enough to threaten the U.S. economic recovery (see Bernanke, 2012, and Congressional Budget Office, 2012).While virtually all observers agree that going over the cliff would take the U.S. economy into recession, the increasingly polarized fiscal policy environment has made finding a solution more problematic than in years past.

And the fiscal cliff is not simply a one-time anomaly, having been immediately preceded by the 2011 political impasse over raising the Treasury's debt ceiling, which had raised fears of a U.S. default arising from failure to make timely debt payments and prompted Standard & Poors (2011) to downgrade the credit rating of U.S. government debt in summer 2011. While the current economic environment has temporarily forestalled the interest-rate increases that often occur around such a downgrade, an inability to address the nation's fiscal imbalancesraises two types of threats. In the short-run, not addressing the fiscal cliff could tip the U.S. economy back into recession. Over a longer horizion, a prolonged inability to address the long-term fiscal challenges facing the U.S. would likely lead to higher future government borrowing costs after the interest rate effects of the recession and Euro-debt crisis abate. Through such channels, polarization can potentially affect the economy's short-run and long-run growth paths in addition to having other sociological and political impacts.

Increased partisanship in American politics has been attributed to shifting generational attitudes (Strauss and Howe, 1991), cyclical economic conditions (Gelman, et al., 2010; and Pontusson and Rueda, 2008), greater income inequality (McCarty, Poole, and Rosenthal 2006, forthcoming), and an increasingly fragmented state of media (e.g., Davis and Owen, 2008; Iyengar and Hahn, 2009; Jamieson and Cappella, 2008, Jones, 2001, and Prior, 2005 and 2007). The first of these factors is difficult to test with time series data owing to the low frequency of generational shifts, while the tendency of business cycles to reverse implies that economic cycles cannot account for long-term shifts in polarization. This leaves two main testable long-run drivers: the increasingly fragmented state of American media and greater income inequality.

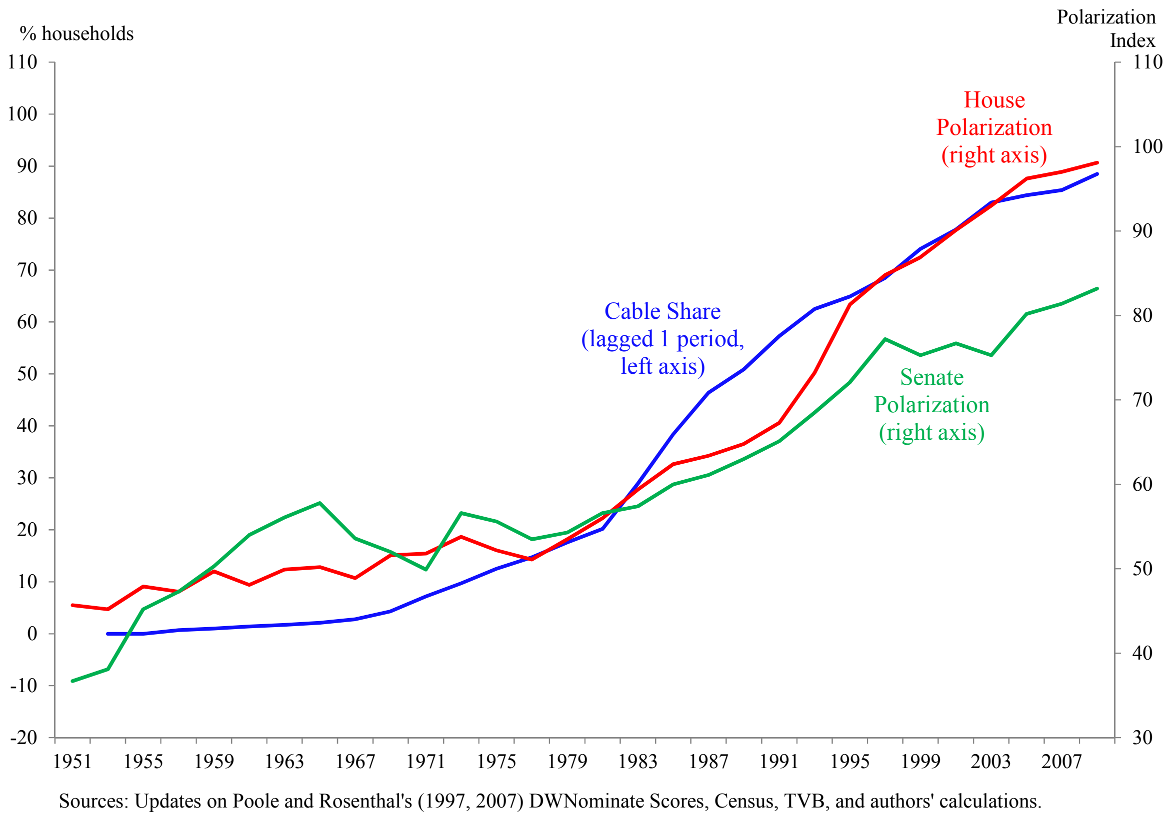


Figure 1: Cable TV Share, Polarization in the U.S. House and Senate

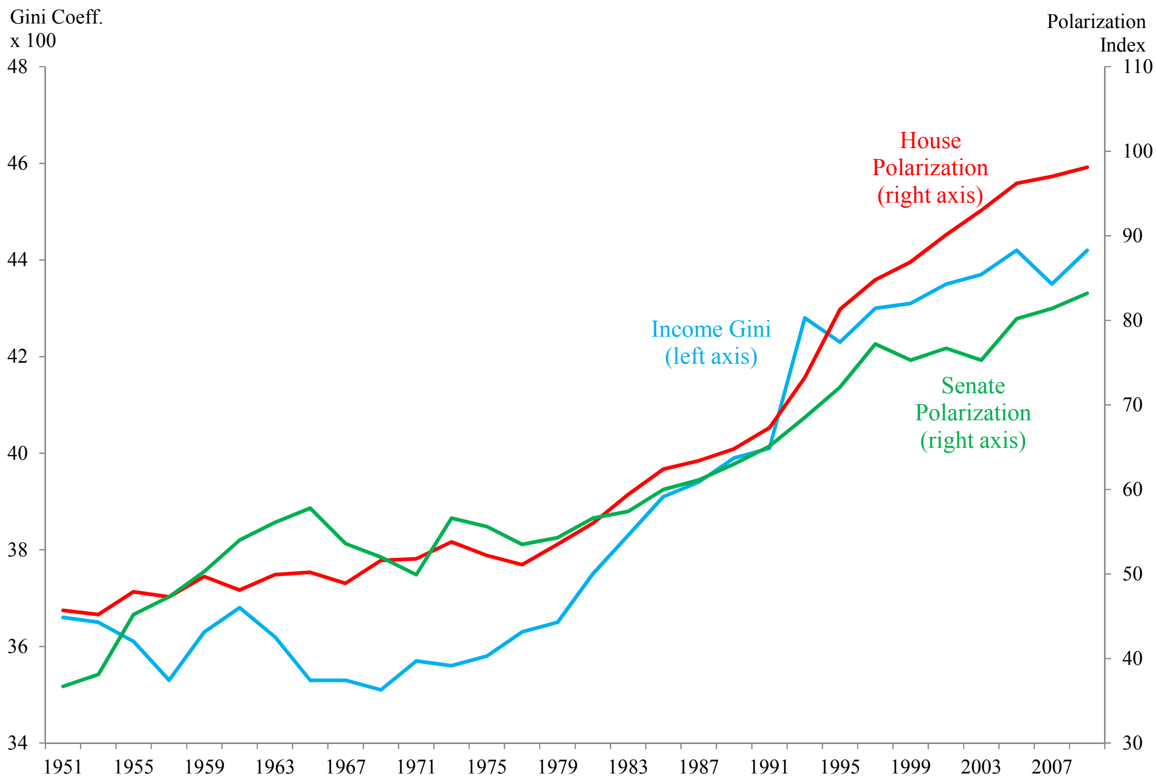
Past studies have used cross-sectional testing to link both of these factors to polarization – income inequality as discussed in McCarty, Poole, and Rosenthal, 2006 and forthcoming, and media fragmentation as done by Prior (2005, 2007). But this approach cannot rule out that cross-sectional factors may shift over time. Although these studies provide partial or suggestive evidence (as in Figure 2) relating higher inequality or cable TV usage to greater partisanship, these linkages have not been rigorously tested using appropriate time series techniques for nonstationary variables, i.e., those whose averages shift over time.

Moreover, there are crucial causal questions to be considered. If media fragmentation is a mere side effect of income inequality, then a study that looks solely at media fragmentation and polarization runs the risk of erroneously attributing increased polarization to media fragmentation when, in fact, income inequality is the cause. On the other hand, if Prior and others are correct that media fragmentation reduces perceived common interests and thereby fosters both income inequality and polarization, then a study that omits fragmentation runs the risk of erroneously attributing the rise of polarization to inequality.

The present study adopts such techniques to assess whether increased polarization can statistically be linked to shifting trends in inequality or media fragmentation, as measured by cable TV use. These time series tests are best viewed in conjunction with cross section evidence, as neither time series nor cross section evidence alone is sufficient to definitively test competing hypotheses about polarization.

This paper assesses whether the rise of media fragmentation or income inequality is linked to the House and Senate measures of polarization from Poole and Rosenthal (2007), using indexes based on the primary dimension of differences: the role of the government in the economy in the modern sense of the terms liberal-moderate-conservative. Using time series techniques, we test whether the rise of political polarization in Congress can be statistically attributed to a measure of media fragmentation–the share of households with cable or pay TV—

or income inequality, as tracked by the Gini coefficient for the dispersion of income across American families.[1] Using the cointegration methods of Engle and Granger (1987), Johansen and Juselius (1990), and Johansen (1991), we find that periods when the share of households with cable or pay TV has risen are followed by upward shifts in Congressional polarization (Figure 2), with less convincing evidence of a link between income inequality and polarization.[2]



Sources: Updates on Poole and Rosenthal's (1997, 2007) DWNominate Scores, Census, and authors' calculations.

Figure 2: Income Inequality Misses Pre-1981 Polarization Trends

A favorable feature of the long-run variables is that cable TV share is likely exogenous to the polarization index for two reasons. First on intuitive grounds, the time series pattern of the cable/pay TV share variable is likely exogenous to shifts in the degree to which Congressional voting is polarized reflecting the impact of technology on TV delivery systems. Second, formal statistical tests indicate that cable TV share is exogenous to polarization, but polarization is not

[1]We follow Baum and Kernell (1999) in using the share of families with cable or pay TV as an independent variable.
[2]These findings are consistent with two non-mutually exclusive hypotheses. One is that the shift from network to cable TV reinforced the political views of viewers who self-select into watching stations that reinforce prior beliefs (Jamieson and Cappella, 2008; Iyengar and Hahn, 2009; and Sunstein, 2007). The other, "entertainment" hypothesis is that the rise of cable TV reduced learning about political issues as a by-product of watching network TV news coverage (Downs, 1957; Baum and Kernel, 1999; and Prior, 2007).

exogenous to the cable TV share.[3] Overall, the time series results complement the media fragmentation findings and theories of Baum and Kernel (1999) and Prior (2005, 2007), and to some degree cut against the income inequality findings of Rosenthal (2011), though we cannot entirely rule out the income-inequality channel as a secondary source for political polarization.

To establish these findings, our study is organized as follows. Section II tests for long-run relationships between polarization and either the cable/pay TV share of households or the income Gini coefficients for American families. Section III assesses whether the cable share, as reflected in this long-run relationship, explains short-run changes in polarization in the presence of any significant short-run controls. The fourth section addresses whether the cable share or Gini coefficients are temporally exogenous to political polarization, while the fifth section discusses the potential role of Internet use. The last section interprets the statistical findings.

II. Are Cable/Pay TV or Income Inequality Related to Polarization in the Long-Run?

IIA. The Long-Run Variables Tracking Polarization, Income Inequality, and Cable/payTV

The four long-run variables are the Poole and Rosenthal indexes of polarization in the House and Senate (*PolarH* and *PolarS*, respectively), the U.S. Census's Gini coefficient for family income (*Gini*), and the share of American households with cable or satellite (pay) TV (*Cable*). We use *Gini* to track income inequality because in other estimation models not shown to conserve space, we found that the Gini coefficient had a much stronger statistical relationship with polarization than the income share of the top 1 percent of families and the inverted Pareto-Lorenz coefficient for the top 10 percent of families (favored by Atkinson, et al., 2011).[4] Because the polarization indexes are biennial measures of Congressional polarization, *Gini* and *Cable* are biennial readings equal to two-year averages of values for the two years covered by the

[3] In the absence of the cable share, income inequality is exogenous to polarization while the opposite is not true.

[4] Results are available upon request. The income share of the top 1 percent are updates of Piketty and Saez's (2006) data on income with or excluding all capital gains (http://emlab.berkeley.edu/users/saez/piketty-saezOUP04US.pdf). We find that *Gini* outperformed both top 1% shares. See Atkinson, Piketty, and Saez (2011) for more on inequality.

Congressional sessions. *Cable* is constructed by splicing data on the share of Americans with cable TV using pre 1996 data from the Census and post-1995 Nielsen-based estimates from TVB on cable TV and the rising use of satellite TV. The data account for the emergence of satellite TV since 1996 to reflect the use of both wired and satellite sources of pay TV.

IIB. Empirical Approach to Testing for Long-Run Relationships

Cointegration techniques are used because of strong evidence that our four main variables are nonstationary, meaning that they have trends that can complicate statistical analysis. In particular, *PolarH*, *PolarS*, and *Gini* is integrated of order 1 according to Dickey-Fuller GLS tests, meaning that the first differences of these variables are stationary (Table 1). This test does not reject that the first difference of *Cable* is stationary at a biannual frequency. However, unit root tests have low power, especially when there are few observations. Indeed, when *Cable* is assessed at an annual frequency (recall the biannual data are from averages of annual data) *Cable* is integrated of order 1, reflecting the low power of unit root tests applied to a sample with half as many observations owing to the biannual frequency. We use the Johansen-Juselius approach rather than the DOLS approach to estimate the long-run relationships, as the latter's use of future changes in variables seems implausible given that expectations of getting cable in the future would induce Congressional behavior in advance.

Cointegration analysis is amenable to testing whether right-hand side variables are exogenous to the dependent variable, providing evidence on whether media structure or income inequality drives political polarization 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:

$$\ln(Y_1) = \alpha_0 + \alpha_1 \ln(Y_2)$$

$$\Delta\ln(Y_1) = \beta_1[\ln(Y_1) - \alpha_0 + \alpha_1\ln(Y_2)]_{t-1} + \Sigma_{i=1}\gamma_i\Delta\ln(Y_1)_{t-i} + \Sigma\delta_i\Delta\ln(Y_2)_{t-i} + \lambda_1 X_t + \varepsilon_{1t}$$

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

where the lags of first difference endogenous variables minimize the AIC, X is a vector of exogenous factors, ε_{it} are residuals, and the λ_i, γ_i, and $\delta_{i\,v}$ are row vectors of coefficients.

The estimation of long-run and any short-run relationships is joint, and depends on the sets of exogenous, short-run factors included (the vector X). We tried several variables tracking possible effects of war deaths—with or without a dummy for the draft era, midterm congressional elections, the election of a new or reelection of an incumbent president (*Pres2nd*), and the era when broadcast media using the public airwaves were subject to the Fairness Doctrine, which encouraged balanced news reporting. None of these was consistently statistically significant, with the exception of *Pres2nd* (=1 for the first Congress following a President's re-election), which was associated with higher polarization. This last finding may reflect the frequent transformation of re-election campaigns into referendums on incumbents and lingering hard feelings from its aftermath, helping generate the traditional "second term blues" that follow successful re-elections (Zacher, 1996). 0-1 variables for controversial leaders, such as Nixon or the Gingrich-led House, were also significant but are omitted from the tables given their more arbitrary nature and that their inclusion did not affect the other qualitative results.

IIC. Tests of Whether the Cable TV and House Polarization Are Related in the Long-Run

The upper panel of Table 1 reports cointegration tests for polarization in the House using different sets of long-run variables with or without the one significant short-run factor in VECMs for data over 1951-2011. Since researchers may differ over which short-run variables to include, we estimated a number of models with alternative exogenous variables, and then used a model selection procedure that progressively dropped the most statistically insignificant of these short-run variables. In the end, only one of these variables remained, the dummy for the first Congress following the re-election of a president (*Pres2nd*). The upper portion of Table 1 presents results about what drives polarization in the long-run, while the lower portion provides results pertaining to modeling short-term movements in polarization, as reflected in the first difference of *polarization*. The first model only includes cable TV share as a long-run determinant. The second model lags cable TV share by one period, based on the tendency for cable trends to precede polarization trends displayed in Figure 1. To model 2, model 3 adds *Pres2nd* as a short-term variable. Model 4 is equivalent to model 1 except that it replaces cable share with income inequality as a long-run determinant of polarization. Model 5 adds the t-1 lag of cable share to Model 4 to test the relative significance of cable share and income inequality, while Model 6 adds the dummy for the congressional term that follows the reelection of a president.

In every model including the cable share (all models except Model 4), test statistics indicate the existence of only one significant cointegrating vector, as the eigenvalue and trace statistics reject the null hypothesis of no significant long-run relationship exists between cable share and polarization in the U.S. House. The lag lengths used to estimate these vectors were chosen based on the minimum lag length needed to obtain a unique, significant cointegrating variable and, if possible, also yielded clean model residuals using the VECLM statistics on lags t-1 through t-6. 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 periods, most of which start in 1959. In each case, the cable TV share variable has a significant and negative long-run coefficient (in the upper part of Table 1) ranging from -0.60 to -0.75, a finding that indicates a robust long-run effect. This tight range implies that increases in the share of households with cable/pay TV are consistently associated with increases in the DW-Nominate index of polarization in the House of Representatives. The magnitude of the estimated long-run coefficients on *Cable* imply that the 90 percentage point rise in the share of households with cable/pay TV since the 1950s is associated with about a 55 point rise in the DW-Nominate index of polarization in the House.

While the income inequality measure is statistically significant in the absence of the cable share in model 4, there is only weak evidence at the 90% confidence level of a unique, long-run relationship. Furthermore, *Gini* is statistically insignificant in the presence of the cable share variable in Models 5 and 6. These results point to using models without the Gini coefficient. Comparing the fit of models 1 and 2 and noting the insignificant t-1 lag on the first difference of *Cable* in Model 1, reveals that lagging the cable share as in Model 2 yields a more parsimonious model with fewer variables (2 lags of first differences instead of 3). Adding *Pres2nd* to Model 2 as in Model 3 hardly changes the long-run coefficents, but improves the model fit when examining the fit of models tracking the change (the first difference) in polarization. The baseline (Model 2) and its preferred variant (Model 3) imply similar long-run relationships:

$$PolarH = 41.599 - 0.638*Cable, \qquad \text{(Model 2, Table 1) and} \qquad (2)$$

$$PolarH = 42.905 - 0.606*Cable, \qquad \text{(Model 3, Table 1),} \qquad (3)$$

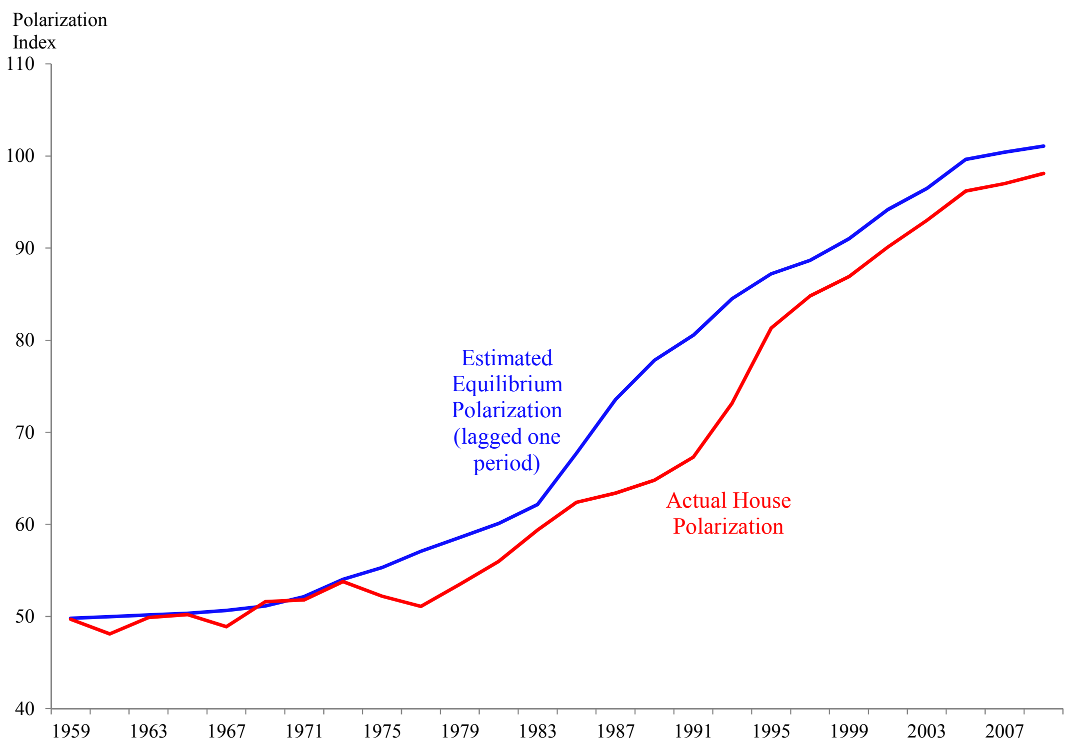
where *Cable* is in percentage points. The magnitude of the estimated long-run coefficients on *Cable* imply that the 90 percentage point rise in the share of households with cable/pay TV since the 1950s is associated with about a 55 point rise in the DW-Nominate index of polarization in the House of Representatives. As shown in Figure 3, the long-run equilibrium values from Model 3 (with an adjustment for the constant in the first difference equation) track the long-run movements in the index of House polarization over the estimation period 1959-2009.[5]

Table 2 presents results from modeling polarization in the Senate using models that correspond to those in Table 1 in terms of which long-run variables (*Cable*, *Cable*(t-1), or *Gini*) are included and whether the short-run variable for terms following the reelection of a president (*Pres2nd*) is included. As before, a statistically significant and unique cointegrating vector can be identified when *Cable* is included. And similar to the model 3 results for the House, the long-run equilibrium values from the cable share model (which includes a short-run control for presidential reelections) track the long-run movements in the index of Senate polarization over the estimation period 1959-2009 as shown in Figure 4.

[5]The adjustment adds the constant in the lower panel divided by minus the coefficient on the error-correction term, and is also similarly done in Figure 4.

IID. Tests of Whether the Cable TV and Senate Polarization Are Related in the Long-Run

Nevertheless, the relationships for the models that omit *Gini* are significant at only the 95 percent confidence level for Senate Models 1-3 in Table 2, rather than at the 99 percent confidence level for corresponding House Models 1-3 in Table 1. Another difference is that a unique cointegrating vector can be identified in the Senate model that only includes the income Gini coefficient, and this relationship is significant with 99 percent confidence. However, as before, the long-run coefficient on *Gini* becomes insignificant in the presence of the cable share in Models 5 and 6. Overall, the long-run evidence clearly favors a stronger statistical relationship between cable share and polarization than between income inequality and polarization for the behavior of the House of Representatives, with the evidence showing a less compelling case for a stronger statistical relationship using the cable share rather than income inequality for explaining the long-run polarization trends in the Senate.



Sources: Updates on Poole and Rosenthal's (1997, 2007) DWNominate Scores and authors' estimates.

Figure 3: Cable-Implied Equilibrium Estimates Track House Polarization Trends Well

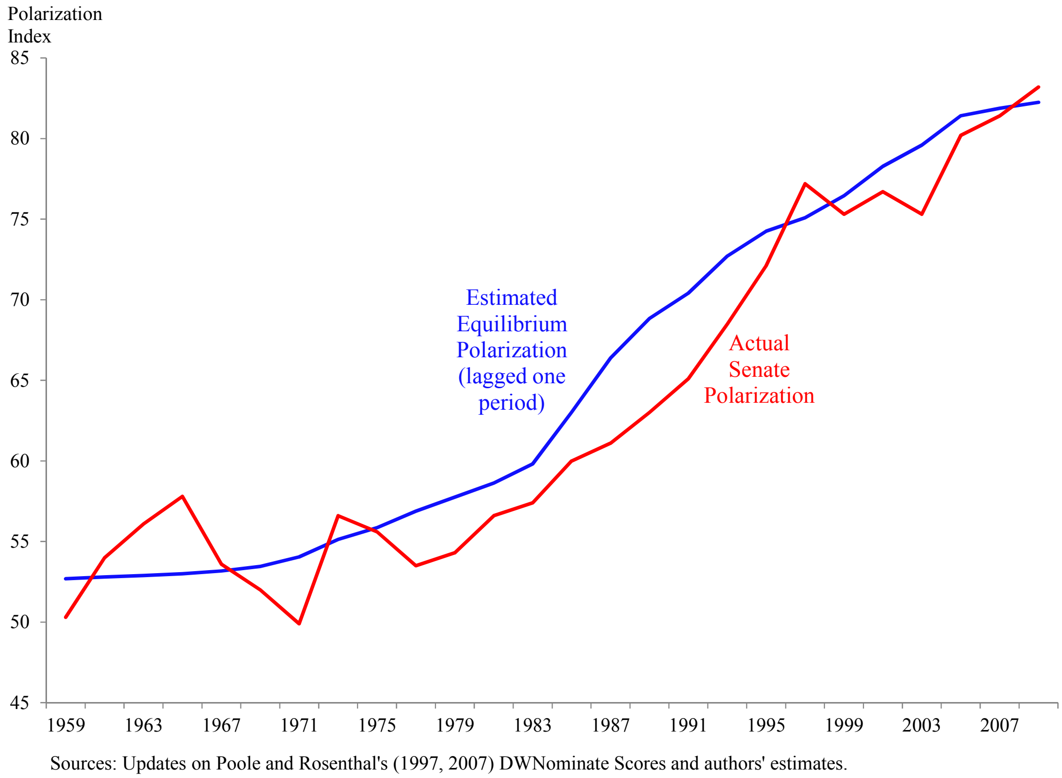


Figure 4: Cable-Implied Equilibrium Estimates Track Senate Polarization Trends Well

Because Senate terms span different congresses, the House polarization evidence is arguably stronger. The reason is that polarization effects from either the rise of cable share or shifts in income inequality may owe to changes in both the behavior of reelected members and shifts in the composition of members from elections, which are slower in the Senate. Indeed, polarization has changed more markedly in the House than in the Senate, both in terms of a larger long-run rise over the 1959-2009 estimation period in the House than in the Senate (about 50 versus 33 points, respectively) and a larger standard deviation in the House versus Senate polarization indexes (17.8 versus 10.8, respectively). Still, the impact of media fragmentation on polarization is evident in our Senate results as well as our House results.

III. Are Cable TV and Political Polarization Related in the Short-Run?

To see if long-run relationships help explain short-run movements in the polarization measures, we examine short-run changes in the polarization indexes from using the same set of VECMs used to estimate long-run relationships.

IIIA. Tests of Whether the Cable TV and House Polarization Are Related in the Short-Run

The lower panel of Table 1 reports the results from six models of the change in the House polarization index. In the presence of cointegration, proper specification requires including an error-correction term (*EC)* that essentially implies that long-run deviations of actual from equilibrium levels of no opinion add information about short-run movements in the 'no opinion' share. Accordingly, models 1 through 6 include error-correction terms testing for the impact of cable share and/or income inequality on short-run changes in polarization. In each case, EC_{t-1}, equals the gap between the actual polarization reading and the equilibrium share implied by *Cable* and/or *Gini* from the corresponding cointegrating vector in the upper-panel of Table 1. When the actual share exceeds the equilibrium share in the prior quarter, long-run equilibrium implies a tendency for polarization share to fall in the short-run, thus implying a negatively signed coefficient on the variable EC_{t-1}.

In each model, the error-correction term is negative and highly statistically significant, implying that the polarization tends to fall when actual exceeds the equilibrium share in the prior Congress, where the latter is increasing in the share of households with cable TV or in income inequality. The size of the error-correction coefficients suggests that roughly 30 to 80 percent of the gap between actual and equilibrium polarization is closed over the next two-year period. Another nice feature of the models is that their fits are reasonably high, ranging between 47 and 78 percent. Given the weak long-run evidence of income inequality, it is interesting that the preferred cable models 2 and 3 account for between 66 and 74 percent of the variation in the change of polarization in the House. The estimated and statistically significant coefficient on

Pres2nd implies that political polarization index rises about 2 points in the Congress following the reelection of a president.

IIIB. Tests of Whether the Cable TV and Senate Polarization Are Related in the Short-Run

Results from modeling the change in the Senate polarization index are reported in the lower panel of Table 2, where the models correspond to the House models in Table 1. In each model, the error-correction term is negative and highly statistically significant, implying that the polarization tends to fall when actual exceeds the equilibrium share in the prior Congress, where the latter is increasing in the share of households with cable TV or in income inequality.

There are some differences between the House and Senate short-run results. First, the adjusted R squares of the House models tend to notably exceed those of corresponding Senate models. This may reflect the tendency for House behavior to vary more in response to shifting long-run patterns, as reflected in the higher standard deviation of the House versus Senate polarization index over the sample period, and for the greater tendency for lagged first difference terms to be significant in the House versus Senate models. Second, the error-correction speeds tend to be faster in the corresponding Senate models. Coupled with the relative less importance of first difference terms in corresponding Senate models, this suggests that Senate behavior is more stable in the sense of being driven more by long-run factors (the error-correction term) than by short-run changes (the lagged first difference terms). This may reflect that the composition of the Senate tends to shift more slowly than that of the House. Finally, the model with income inequality by itself (Model 4) has a better fit than the corresponding cable share model (Model 2) or if Pres2nd were added to it compared to Model 3. This suggests that income inequality may be more important than the cable share in a statistical sense for modeling changes in polarization in the Senate. However, the insignificance of the inequality measure in the presence of the cable share in the long-run relationships (the upper panels of Tables 1 and 2) favor using the cable share rather than the income inequality Gini coefficient.

IV. Is the Cable Share Exogenous to Polarization?

One attractive feature of the *Cable* variable is that the polarization indexes are likely exogenous to the cable share for two reasons. First, on intuitive grounds, the cable share is plausibly driven by the impact of technology on TV delivery systems rather than by polarization. Nevertheless, it is hypothetically conceivable that greater polarization in congressional voting might result in legislation that enhances the relative use of cable or that a common factor—the populace becoming inherently more divided in their preferences—drives both Americans to increasingly purchase cable or satellite TV to accommodate their less uniform tastes and their elected representatives to vote in more divisive ways.

To address this latter possibility, the models presented earlier were estimated as a vector error-correction model that contained separate equations for changes in polarization and cable TV shares (and income inequality), which were regressed on the same error-correction term, the same lags of changes in the long-run variables, and the same sets of short-run variables. If the error-correction term is significant in the model of polarization but is insignificant in the model of cable share, then formal econometric evidence would indicate that polarization is 'weakly exogenous' to the cable share as discussed in Urbain (1992) and that polarization index is caused, in a long-run sense, by the cable share according to Granger and Lin (1995). As reported in upper and middle panels of Table 3, respectively, this is indeed the case in every model shown in Table 1 and Table 2. This finding is also evident in Figure 1, where the cable share shifts about one or two biennial periods ahead of the polarization indexes.

Interestingly, in model 4 where the cable variable is omitted, *Gini* is exogenous to polarization (see the lower panel, Table 3) while polarization is not exogenous to *Gini*. However, in the presence of the cable share, polarization and income inequality are not always exogenous to each other. This implies cleaner evidence of causality running from *Cable* to polarization than from *Gini* to polarization.[6]

V. Is it Cable TV Share, Income Inequality, or Internet Usage that Drives Polarization?

One possible objection to the findings concerns whether something other than cable TV use or inequality is inducing more polarization. Among the more plausible alternative factors is the rising use of the Internet, through which many access information.[7] As shown in Figure 5, Internet use (*InternetUse*) among adults (U.S. Census data) rose rapidly in the 1990s, but does not reflect the increased trends in polarization that occurred in the 1980s and early 1990s, in contrast to the cable TV share and income inequality variables.

To more technically assess the possible role of the Internet, biennial models from Table 1 were re-estimated adding either the time t or t-1 lag of *InternetUse*. At most lag lengths a significant, unique cointegrating vector could not be identified, and in the remainder where one could be identified, *Cable* remained significant with the expected sign, but the error-correction term had the wrong sign. This implies that the changes in House polarization did not tend to eliminate past deviations of actual from the estimated equilibrium values. The same was also true in models where Internet use was the only long-run variable, except that in cases where a unique cointegrating vector was identified, the error-correction term was significant with the wrong sign. These patterns reflect that the information in Internet use does not have a reliable relationship with polarization, an interpretation consistent with both a short history of Internet use and that its first difference is not stationary at biennial and annual frequencies.

[6] Nevertheless, there are reasons why bidirectional long-run causality might be expected between inequality and polarization (see Duca and Saving, 2012b; and McCarty, 2012).

[7] The linkage of Internet usage rates to accessing the media is more direct than using personal computer ownership rates, which rose in the 1980s well before the Internet became available. The adult Internet usage rates are from the infrastructure section of the World Bank's World Development Indicators database. Tewksbury (2005) provides evidence that Internet news viewers self-select into different silos that tend to reinforce viewer's prior beliefs.

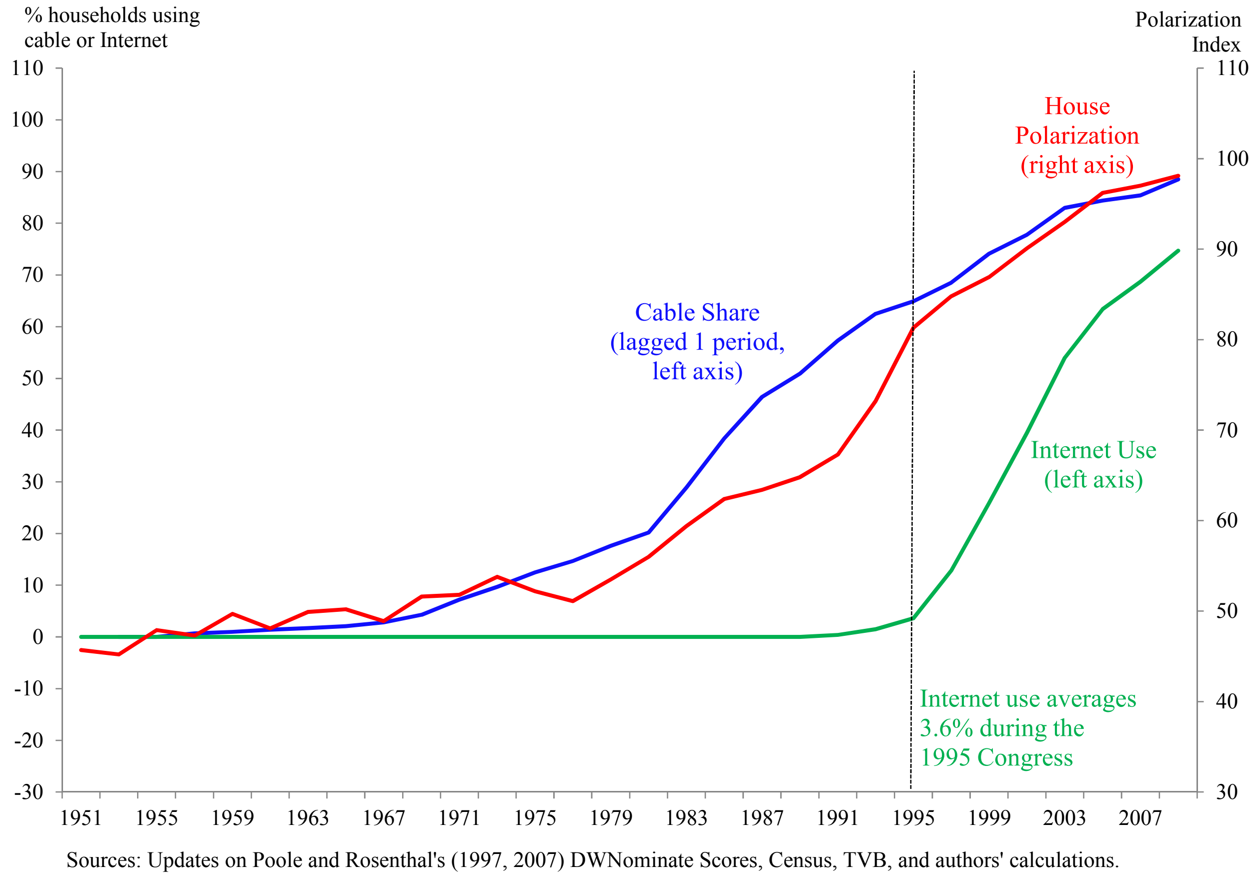


Figure 5: Polarization Trends More in Line With Cable Share Than With Internet Use, Especially Before the Mid-1990s

VI. Conclusion

This paper examines the upward trend over the last six decades in the Congressional polarization indexes of Poole and Rosenthal (1997, 2007) and asks whether the increasingly polarized policymaking environment in the United States stems primarily from rising income inequality or increased media fragmentation. Using time series techniques, we find that this uptrend largely has a stronger statistical link to media fragmentation than to income inequality.[8] The analysis also finds some evidence consistent with the view that Congressional voting temporarily becomes more polarized at the beginning of a President's second term, hampering efforts to reach common ground at those particular points in time. In these two dimensions, the

[8]Because cable share and inequality are measured on a national basis over time, tests on regional polarization are not feasible. Duca and Saving (2012b) find bidirectional causality between polarization and inequality over 1913-2009.

paper contributes to the literature that empirically assesses the links between media fragmentation and public opinion.

It is conceivable that gradual policy shifts over time toward less regulation and less tax progressivity might have induced both (1) a less regulated communications industry that gave rise to a more fragmented structure of TV, and (2) a less equal distribution of income. While we demonstrate that cable share statistically outperforms an income Gini coefficient series in tracking polarization in both the short- and long-runs, this does not necessarily invalidate income inequality as a useful proxy for polarization or even a secondary channel through which polarization could be induced. From this point of view, our findings are not necessarily inconsistent with the hypothesis of McCarty, Poole, and Rosenthal (2006, forthcoming) that inequality has contributed to political polarization. On the other hand, the observed effects of cable share on polarization may have emanated more from changes in communications technology than from shifts in regulation. Because technological and regulatory shifts have often occurred in sync, it is hard to distinguish between these plausible, underlying causal factors, underlining the difficulty in sorting out the ultimate roles of technology, inequality, and regulation in driving political polarization over the past half century.

With this caveat in mind, our findings strongly suggest that greater media fragmentation has contributed to increased political polarization. This may occur as individuals seek out self-reinforcing viewpoints rather than be exposed to a common "nightly news" broadcast, or alternatively, may occur as individuals opt out of news entirely in favor of entertainment, thereby reducing incidental or by-product learning about politics (Downs, 1957; Duca and Saving, 2012a; Prior, 2005, 2007). Within our empirical framework, these respective silo and entertainment hypotheses are observationally equivalent. Other trends such as declines in public viewership of presidential addresses and debates (Baum and Kernel, 1999), along with downtrends in voter participation and knowledge of political issues (Prior, 2005, 2007) suggest both may be occurring, in contrast to hopes expressed in some quarters that media fragmentation might on net reduce polarization by producing a more well-informed citizenry (Habermas, 1998). In these ways, our time series findings are broadly consistent with the conditional political learning framework of Prior (2007), complement the earlier media fragmentation findings of Baum and Kernel (1999) and Prior (2007), and support some of the insights of Downs (1957) and Popkin (1991).

The results also suggest that until other structural changes (e.g., media or political reforms) shift the legislative landscape, an elevated degree of polarization is likely to persist. If so, this could have far-reaching implications for the nation's political economy, as the U.S. struggles to address not only the short-run threats posed by the 2013 'fiscal cliff', but also its long-run fiscal challenges, a concern that Standard & Poors (2011) cited when it downgraded the credit rating of U.S. government debt in summer 2011. We leave it to future research to address these and other possible economic ramifications of the political polarization that has been fostered by media fragmentation, and whether the rise of new, more interactive media may alter these trends.

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Table 1: Biennial Models of Political Polarization in the U.S. House of Representatives

(Congresses spanning votes over 1957-2010 in *Cable* Models, 1959-2010 in *Gini* Models)

Equilibrium Long-Run Relationship: $PolarH_t = \lambda_0 + \lambda_1 Cable_t$ or $= \lambda_0 + \lambda_1 Gini_t$

	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	40.069	41.559	42.905	-194.607	94.921	37.363
$Cable_{t/t-1}$	0.625^{**}	0.638^{**}	0.606^{**}		0.752^{**}	0.614^{**}
	(20.17)	(36.91)	(26.70)		(7.67)	(9.06)
$Gini_t$				667.294^{**}	147.38	13.047
				(12.36)	(1.52)	(0.20)
Eig. (1 v.)	0.565	0.712	0.618	0.409	0.730	0.758
Eig. (2 v.)	0.022	0.033	0.0141	0.001	0.272	0.295
Trace (1 v.)	22.23^{**}	31.95^{**}	25.39^{**}	13.68^{+}	42.36^{**}	44.24^{**}
Trace (2 v.)	0.57	0.83	0.38	0.03	8.27	8.79
Max-Eig (1v)	21.66^{**}	31.12^{**}	25.01^{**}	13.66^{+}	34.09^{**}	35.45^{**}
Max-Eig (2v)	0.57	0.83	0.38	0.03	8.27	8.75
Cointegration?	Yes^{**}	Yes^{**}	Yes^{**}	Weak($^{+}$,$^{+}$)	Yes^{**}	Yes^{**}

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

Sample	1957-2009	1957-2009	1957-2009	1959-2009	1959-2009	1959-2009
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	4.720^{**}	6.240^{**}	3.422^{**}	2.718^{**}	3.191^{**}	6.035^{**}
	(4.77)	(6.15)	(5.01)	(3.88)	(4.31)	(5.35)
EC_{t-1}	-0.499^{**}	-0.803^{**}	-0.534^{**}	-0.297^{**}	-0.439^{**}	-0.828^{**}
	(-4.83)	(-6.48)	(-5.53)	(-3.42)	(-4.37)	(-5.59)
$Pres2nd_t$			1.892^{*}			1.681^{*}
			(2.33)			(2.00)
$\Delta PolarH_{t-1}$	0.289^{+}	0.233^{+}		0.089	0.220	0.133
	(1.90)	(0.32)		(0.46)	(1.30)	(1.00)
$\Delta PolarH_{-2}$	0.048	0.129	0.221	0.019	0.157	0.156
	(0.32)	(1.02)	(1.92)	(1.62)	(0.10)	(1.17)
$\Delta PolarH_{t-3}$	-0.158	-0.175	-0.062	-0.333^{+}		-0.051
	(-1.09)	(-1.52)	(-0.49)	(-1.97)		(-0.45)
$\Delta Cable_{t-1}$	-0.363^{*}					-0.387^{*}
	(-2.11)					(-2.32)
$\Delta Cable_{t-2}$	0.028	-0.204	-0.222		-0.140	-0.362^{*}
	(0.14)	(-1.39)	(-1.32)		(-0.80)	(-2.15)

$\Delta Cable_{t-3}$	-0.595^{**}	-0.475^{**}	-0.470^{*}		-0.614^{**}	-0.675^{**}
	(-2.90)	(-2.91)	(-2.52)		(-3.09)	(-2.76)
$\Delta Cable_{t-4}$		-0.694^{**}				
		(-6.15)				
$\Delta Gini_{t-1}$				-12.676	134.81	-4.985
				(-0.26)	(2.79)	(-0.10)
$\Delta Gini_{t-2}$				-36.967	51.12	33.553
				(-0.60)	(-0.75)	(-0.50)
$\Delta Gini_{t-3}$				-79.666		-24.171^{**}
				(-1.43)		(-0.59)
Adjusted R^2	.560	.741	.658	.472	.576	.782
S.E.	1.445	1.130	1.274	1.584	1.419	1.037
VECLM(1)	8.97	8.89	5.24	3.29	3.86	5.41
VECLM(6)	3.27	2.70	3.17	1.53	6.03	9.11

Dickey-Fuller GLS Unit Root Tests Modified Schwartz Information Criterion
(1951-2009 Congresses, covering 1951-2010)

	Level (SIC lag)			***Level (SIC lag)***	
PolarH	-2.3911	(5)	Δ *PolarH*	-4.2121^{**}	(0)
PolarS	-1.8722	(0)	Δ *PolarS*	-5.1059^{**}	(0)
Gini	-1.8125	(0)	Δ *Gini*	-6.4524^{**}	(0)
Cable (biennial)	-1.9931	(1)	Δ *Cable* (biennial)	-2.0726	(0)
Cable (annual)	-1.3389	(1)	Δ *Cable* (annual)	-3.3739^{*}	(0)

Notes: "v." denotes vector, while $^{+}$, *and ** denote 90% , 95%, and 99% significance levels, respectively. t-statistics are in parentheses. Lag lengths of 3, 3, 2, 3, 2, and 3 for models 1-3 and 4-5, respectively, yielded unique, significant vectors and clean residuals, while a lag length of 3 gave the strongest evidence of cointegration for model 4. The significance level of VECLM statistics accounts for size of the vector. Lag lengths for unit root tests are based on the Schwartz Information. Note that *Cable* is in percentage point units while *Gini* is in decimals. To make them comparable, multiply the *Gini* coefficients by 0.01.

Table 2: Biennial Models of Political Polarization in the U.S. Senate

(Congresses spanning votes over 1957-2010 in *Cable* Models, 1959-2010 in *Gini* Models)

Equilibrium Long-Run Relationship: $PolarS_t = \lambda_0 + \lambda_1 Cable_t$ or $= \lambda_0 + \lambda_1 Gini_t$

	Model 1	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	47.761	49.051	49.352	-67.876	68.139	80.590
$Cable_{t/t-1}$	0.359**	0.358**	0.350**		0.410**	0.449**
	(18.27)	(19.47)	(17.95)		(5.93)	(6.17)
$Gini_t$				334.026**	-53.981	89.632
				(22.40)	(-0.82)	(1.29)
Eig. (1 v.)	0.497	0.512	0.481	0.593	0.789	0.786
Eig. (2 v.)	0.030	0.001	0.008	0.001	0.301	0.390
Trace (1 v.)	18.67*	18.67*	17.26*	23.368**	49.75**	52.96**
Trace (2 v.)	0.79	0.03	0.21	0.02	9.31	12.91
Max-Eig (1v)	17.88*	18.67*	17.05*	23.347**	40.44**	40.04**
Max-Eig (2v)	0.79	0.03	0.21	0.02	9.31	12.84
Cointegration?	Yes*	Yes*	Yes*	Yes**	Yes**	Yes**

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

Sample	1959-2009	1959-2009	1959-2009	1959-2009	1959-2009	1959-2009
Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
Constant	4.232**	2.882**	2.110**	2.057**	2.941**	2.294**
	(3.97)	(3.23)	(2.66)	(3.84)	(3.68)	(3.04)
EC_{t-1}	-0.768**	-0.813**	-0.691**	-0.735**	-0.703**	-0.590**
	(-4.19)	(-4.49)	(-4.15)	(-4.73)	(-4.36)	(-4.13)
$Pres2nd_t$			3.048**			3.049*
			(2.72)			(2.57)
$\Delta PolarS_{t-1}$	0.258	0.237	0.242	0.057	0.096	0.151
	(1.48)	(1.38)	(1.60)	(0.39)	(0.58)	(1.00)
$\Delta PolarS_{t-2}$	0.064	0.096	0.159	0.007	0.132	0.162
	(0.44)	(0.67)	(1.26)	(0.05)	(1.00)	(1.37)
$\Delta PolarS_{t-3}$	-0.118			-0.281*		
	(-0.78)			(-2.29)		
$\Delta Cable_{t-1}$	-0.543*					
	(-2.43)					
$\Delta Cable_{t-2}$	-0.027	-0.219	-0.448*		-0.361	-0.533*
	(-0.09)	(-0.92)	(-2.00)		(-1.63)	(-2.53)
$\Delta Cable_{t-3}$	-0.360	-0.379	-0.089		-0.469*	-0.202
	(-1.40)	(-1.51)	(-0.38)		(-2.05)	(-0.87)

$\Delta Gini_{t-1}$				-107.745	112.577^{+}	47.071
				(-1.56)	(1.90)	(0.76)
$\Delta Gini_{t-2}$				33.384	178.745**	156.168
				(062)	(3.01)	(2.87)
$\Delta Gini_{t-3}$				-52.168		
				(-1.01)		
Adjusted R^2	.416	.414	.552	.602	.525	.617
S.E.	1.961	1.964	1.718	1.618	1.768	1.588
VECLM(1)	5.53	6.60	6.95	4.39	12.11	11.20
VECLM(6)	1.17	0.63	1.76	5.39	10.17	4.15

Dickey-Fuller GLS Unit Root Tests Modified Schwartz Information Criterion (1951-2009 Congresses, covering 1951-2010)

	Level (SIC lag)			Level (SIC lag)	
PolarH	-2.3911	(5)	Δ *PolarH*	-4.2121**	(0)
PolarS	-1.8722	(0)	Δ *PolarS*	-5.1059**	(0)
Gini	-1.8125	(0)	Δ *Gini*	-6.4524**	(0)
Cable (biennial)	-1.9931	(1)	Δ *Cable* (biennial)	-2.0726	(0)
Cable (annual)	-1.3389	(1)	Δ *Cable* (annual)	-3.3739*	(0)

Notes: "v." denotes vector, while $^{+}$,*and ** denote 90% , 95%, and 99% significance levels, respectively. t-statistics are in parentheses. Lag lengths of 3, 2, 2, 3, 2, and 2 for models 1-3 and 4-5, respectively, yielded unique, significant vectors and clean residuals, while a lag length of 3 gave the strongest evidence of cointegration for model 4. The significance level of VECLM statistics accounts for size of the vector. Lag lengths for unit root tests are based on the Schwartz Information. Note that *Cable* is in percentage point units while *Gini* is in decimals. To make them comparable, multiply the *Gini* coefficients by 0.01.

Table 3: Weak Exogeneity Tests

A. Testing Whether *Polarization* is Weakly Exogenous to *Cable* or *Gini*

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

Test whether α_1 is equal to zero: resoundingly rejected in Models 1-66. Note that the EC term in Models 5-6 includes information from the statistically insignificant long-run *Gini*.

Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
House Polarization						
EC_{t-1}	-0.499**	-0.803**	-0.534**	-0.297**	-0.439**	-0.828**
	(-4.83)	(-6.48)	(-5.53)	(-3.42)	(-4.37)	(-5.59)
Senate Polarization						
EC_{t-1}	-0.768**	-0.813**	-0.691**	-0.735**	-0.703**	-0.590**
	(-4.19)	(-4.49)	(-4.15)	(-4.73)	(-4.36)	(-4.13)

B. Testing Whether *Cable* is Weakly Exogenous to *Polarization*

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

Test whether α_1 is equal to zero: resoundingly NOT rejected in Models 1-3 and 5-6. Note that the EC term in Models 5-6 includes information from the statistically insignificant long-run *Gini*.

Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
House Polarization						
EC_{t-1}	0.014	-0.021	0.012	n.a.	0.038	-0.111
	(0.15)	(-0.14)	(0.13)		(0.33)	(-0.58)
Senate Polarization						
EC_{t-1}	-0.024	-0.024	0.000	n.a.	-0.024	0.005
	(-0.03)	(-0.36)	(0.00)		(0.31)	(0.06)

C. **Testing Whether *Gini* is Weakly Exogenous to *Polarization***

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

Test whether α_1 is equal to zero: resoundingly NOT rejected in Models 4-6. Note that the EC term in Models 5-6 includes information from the statistically insignificant long-run *Gini*.

Variable	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**	**Model 6**
House Polarization						
EC_{t-1}	n.a.	n.a.	n.a.	0.017	0.014**	0.015
				(0.40)	(3.85)	(0.03)
Senate Polarization						
EC_{t-1}	n.a.	n.a.	n.a.	0.010	0.019**	0.021**
				(1.32)	(3.92)	(5.92)

*and ** denote 95% and 99% significance levels, respectively. t-statistics are in parentheses.

Appendix: Constructing Consistent Annual Data on Households with Cable or Satellite TV

For the years 1951-2000 the source data are annual readings from the Census on the share of U.S. households with cable TV. Satellite TV (e.g., Dish TV rather than hard-wired cable) essentially began in 1996, for which data are available from TVB. Adding the shares of households with satellite and those with cable for between 1996 and 2004 provide an estimate of households with either cable or satellite TV that may double count some household having both services. This double-counting is implied by TVB estimates of the share of households with either type of service, which are a little lower than the sum of Census cable estimates and TVB estimates of satellite TV for the years 2002-2004. However, in 2001, the TVB estimate of 81.0 percent of households slightly exceeded the sum (80.4) of cable share from the Census and satellite (wireless cable) TV from TVB. To provide a consistent series, we spliced the TVB estimates since 2001 with the product of the sum-based series before 2000 and the multiplicative break adjustment ratio of (81.0/80.4) based on the 2000 ratio of the comprehensive TVB estimate and the summed series for that year.