Political Polarization and Income Inequality

Nolan McCarty
Princeton University

Keith T. Poole
University of Houston

Howard Rosenthal
Princeton University and Russell Sage Foundation

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Abstract

Politics in the United States can now be characterized as an ideologically polarized two-party system. The economy features increased income inequality. While the literature on comparative political economy has focused on the links between economic inequality and political conflict, the relationship between these trends in the United States remains essentially unexplored. Using National Election Study data from 1952 to 2000, we explore the relationship between income and voter partisan self-identification. We find that partisanship has become more stratified by income. We argue that this trend is largely the consequence of polarization of the parties on economic issues and the development of a two-party system in the South. The trend is much less a reflection of increased economic inequality. The partisanship results replicate for presidential vote choice. We also find that the two-party system has adjusted to remain competitive in spite of the large increases in real income on the last half of the twentieth century. If voters in 2000 were voting as if real average income were only that of 1960, partisanship would have swung strongly in the Republicans favor.

For Richard McKelvey, in memoriam. This research was conducted as part of the Russell Sage Foundation's Inequality Project at Princeton. Previous versions were presented in seminars or conferences at Harvard, NYU, Princeton, Russell Sage Foundation, Salerno, Stanford and Washington University. We thank participants for their comments. Correspondence to: Nolan McCarty, Woodrow Wilson School, Princeton University, Princeton, NJ. 08544, nmccarty@Princeton.edu

1. Introduction

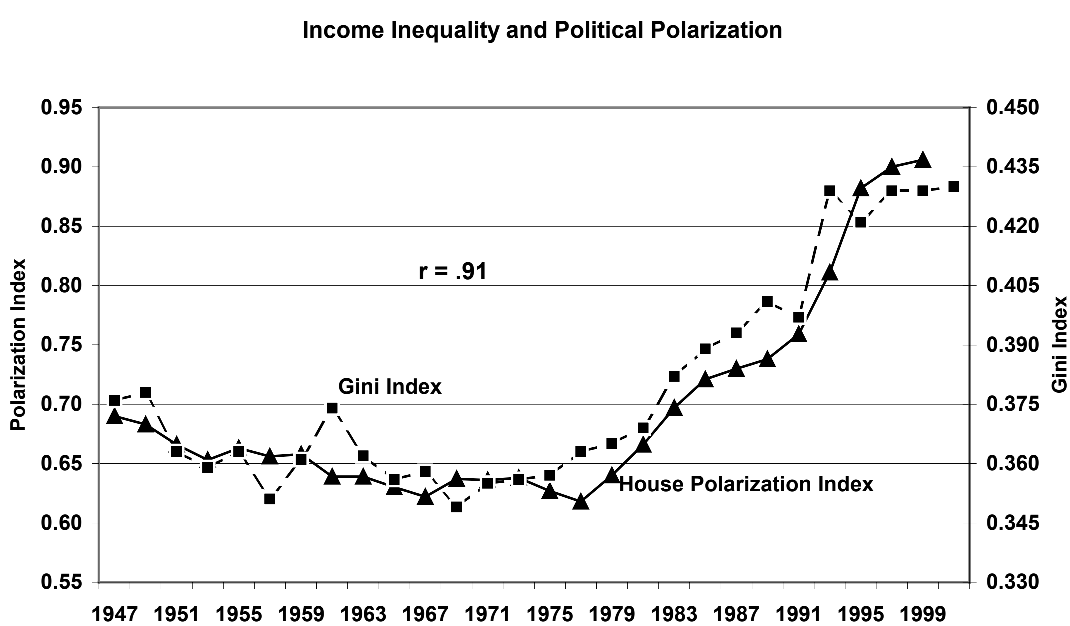
The decade of the 1970s marked many fundamental changes in the structure of American society. In particular, America witnessed almost parallel transformations of both its economic structure and the nature of its political conflict.

The fundamental economic transformation has led to greater economic inequality with incomes at the lowest levels stagnant or declining while individuals at the top have prospered. The Gini coefficient of family income, a standard measure of inequality, has risen by more than 20% since its low point in 1968 (U.S. Census Bureau, 2002).[1] A remarkable fact about this trend is that it began after a long period of increasing equality.[2] Economists and sociologists have allocated tremendous effort into discovering the root causes of this transformation. Numerous hypotheses have been put forward including greater trade liberalization, increased levels of immigration, declining rates of trade unionization, the fall in the real minimum wage, technological change increasing the returns to education, and the increased rates of family dissolution and female headed households.[3]

Within the political realm, the 1970s were also transformative. The decade witnessed both a partisan realignment in the Southern states and increased polarization in the policy positions of Democrats and Republicans. As we, together and separately, have documented in previous work, the bipartisan consensus among elites (Congress in particular) about economic issues that characterized the 1960s gave way to the deep ideological divisions of the 1990s (Poole and Rosenthal, 1984, Poole and Rosenthal 1997, McCarty, Poole, and Rosenthal 1997). Furthermore, we have found that previously orthogonal conflicts have disappeared or been incorporated into the conflicts over economic liberalism and conservatism. Most importantly, issues linked to race are now largely expressed as part of the main ideological division over redistribution.

Remarkably, the trends of economic inequality and elite political polarization have moved almost in tandem for the past half-century. Figure 1 plots the levels of inequality as measured by the Gini coefficient along with a measure of political polarization which is the average distance between Democratic and Republican members of Congress in DW-NOMINATE scores.[4] The polarization measure reflects the average difference between the parties on a liberal-conservative scale. The proximity of these trends is uncanny. In fact, inequality and polarization start increasing at approximately the same time.

Figure 1



There is similar evidence that partisan affiliation in the mass public is increasingly polarized in terms of liberal-conservative views. Green, Palmquist, and Schickler (2002) report that the difference between the "percentage of Republicans who call themselves conservatives" and the "percentage of Democrats who call themselves conservatives" has doubled between 1972 and 1996, moving from 25% to 50%.[5]

While it makes intuitive sense that economic inequality may breed political conflict (or even the converse), almost no work has been done to explain such a conjunction within the context of American politics.[6] Perhaps one reason for this dearth of interest is that traditionally income or wealth has not been seen as a reliable predictor of political beliefs and partisanship in the mass public, especially in comparison to other cleavages such as race and region or in comparison to other democracies.[7] If political conflict does not have an income basis, it makes little sense that changes in economic inequality would disturb existing patterns of political conflict.

However, the fact that American politics has not always been organized as a contest of the haves and have-nots does not mean that it will always be that way. If income and wealth are distributed in a fairly equitable way, little is to be gained for politicians to organize politics around non-existent conflicts. In this context, it is interesting that much of our empirical knowledge about the nature of American political attitudes and partisanship is drawn from surveys conducted during an era of relatively equal economic outcomes.

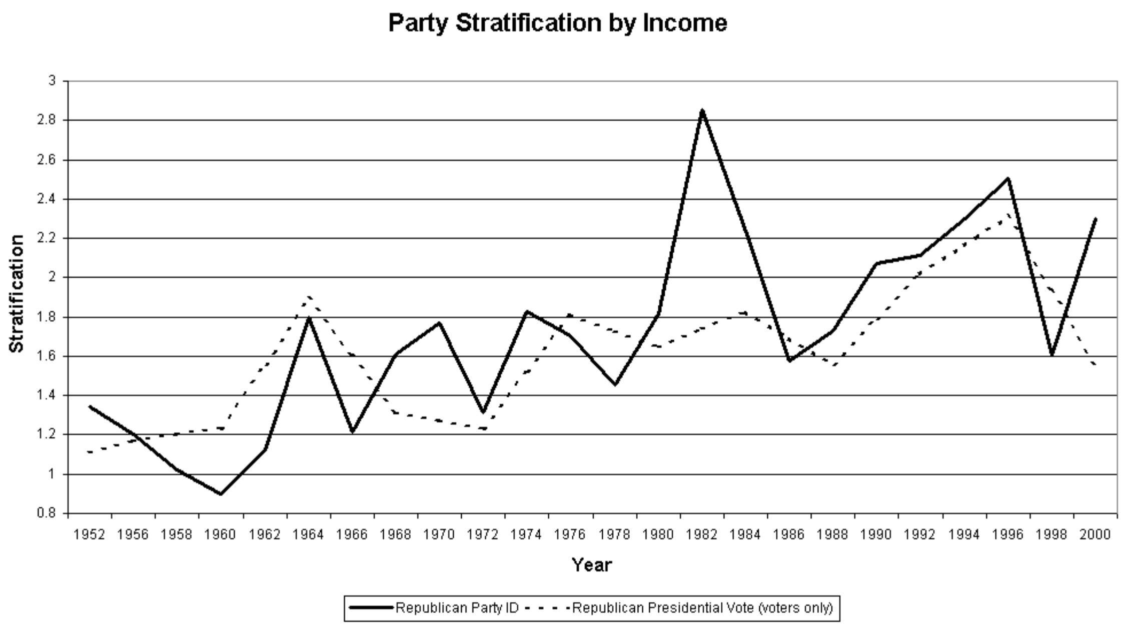
Partisanship (as measured by the National Election Study) was, in fact, only weakly related to income in the period following World War II. In the presidential election years of 1956 and 1960, respondents from the highest income quintile were hardly more likely to identify as a Republican than were respondents from the lowest quintile. In contrast, in the two presidential election years of the 1990s, respondents in the highest quintile were more than twice as likely to identify as a Republican than were those in the lowest.

We summarize how partisanship has acquired an income basis through an index of party-income stratification. Our index is simply the proportion of Republican identifiers (strong and weak) in the top income quintile divided by the proportion of Republican identifiers in the bottom quintile.[8] As seen in figure 2, the stratification of partisanship by income has steadily increased over the past 40 years, leading to an increasing rich-poor cleavage between the parties.

In figure 2, we have also plotted stratification for the presidential vote. Here we compute the ratio of the fraction of Republican voters among voters for the two major parties in the top quintile to the same fraction in the bottom quintile. The upward trend, in stratification is also evident for the presidential vote.

Of course, the simple bivariate relationship between stratification and income does not show that the party system is increasingly organized along income lines. These results could be due to changing income characteristics of party constituencies based on other cleavages. While not denying this claim (in fact, we present some evidence for it below), we insist that regardless of the mechanism that created the stratification of partisanship by income, the mere fact that there are substantial income differences across the constituencies of the two parties has important implications for political conflict. As parties are generally presumed to represent the interests of their base constituencies, the income stratification should contribute to the parties pursuing very different economic policies. Moreover, public policy may be shifting away from policies that are based on self-identified racial, ethnic, or gender characteristics. The shift could be to policies, such as preferred access to higher education for children from poor homes or earned income tax credits, which are income or wealth based. Such a shift would reinforce interest in studying the income stratification of partisan identification.

Figure 2



Notes: Stratification of partisanship is calculated in each NES survey from 1952 to 2000. Stratification of presidential votes is calculated only in presidential election year surveys.

To explore the relationship between the economic and political transformations that we have discussed, the rest of the paper attempts to provide some explanations for the causes of the increased party-income stratification. We focus on partisan identification because it is one item from the National Election Study that is present in every study from 1952 to 2000. Moreover, unlike presidential vote intention or choice, it is less influenced by election-specific factors, such as the perceived extremity of the candidates or their "charisma". We are able, however, to show that our results for partisanship largely replicate for presidential vote choice, although, as is to be expected, the results are somewhat noisier.

Logically, there are four non-mutually exclusive reasons why stratification might have increased. First, there could be a response effect. There can be a temporal increase in the coefficient for income in our model of partisanship. Below we argue that this is consistent with party polarization on economic policy issues. Second, there may be an *inequality* effect. Increased inequality might have made low income groups relatively poorer and high-income groups richer so that with even a small, constant *response* effect stratification would increase. Third, increased stratification might be a result of a change in the joint distribution of other demographic characteristics and income. Pro-Democratic groups may have gotten poorer while pro-Republican groups got richer. For example, African-American partisanship may have remained unchanged but the relative poverty of African-Americans may have increased. Finally, groups with high incomes may have moved toward the Republicans while poorer groups moved toward the Democrats. For example, the relative poverty of African-Americans may have remained unchanged but their propensity for Democratic identification increased.

To quantify each of these effects, we estimate a model of party identification and its relationship to income and other characteristics. We then use the estimates of this model as well as data about the changing distribution of income to calculate the level of party-income stratification under many different counterfactual scenarios. The results show that almost all of the increase can be attributed to an increased effect of income on partisanship and changes in party allegiances of certain groups. This increased income effect largely reflects a greatly increased income effect in the South.[9] Changes in the incomes of different groups and the widening income distribution do not play as large a role.

2. A Simple Model of the Relationship between Income and Partisanship

To motivate our empirical analysis, we begin with the canonical prediction of political-economic models of voter preferences over tax rates and the size of government. These models predict that a voter's preferred tax rate is a function of both her own income and the aggregate income of society.[10] Assuming that tax schedules are either proportional or progressive, individuals with higher incomes prefer lower tax rates since they pay a large share of taxes but receive only an equal share of public expenditure. Alternatively, when aggregate income is larger, higher tax rates produce more money for redistribution and public goods. Thus, *ceteris paribus* individuals prefer higher tax rates as aggregate income increases. To capture the intuition of these models, we assume that voter i's ideal tax rate is a function not of his income y_i but of relative income $r_i = y_i / \overline{y}$, where $\overline{y}$ is the average income of all taxpayers. The ideal tax rate is then $t\left(y_i/\overline{y}\right) \equiv t\left(r_i\right)$. The ideal rate is decreasing in relative income, so $t' < 0$.[11] .

We assume that each of the parties support different tax rates and sizes of government. Let $t_D > t_R$ be the tax platforms of the Democratic and Republican parties. Voter i then supports the Republicans on economic issues when the utility for the Republican platform is greater than that of the Democratic platform, or $u\left(t_R \mid r_i\right) > u\left(t_D \mid r_i\right)$. Unfortunately, these platforms are not observable. In order to specify an estimable model, we invert each platform into a relative income so that

$r_R = t^{-1}(t_R)$ and $r_D = t^{-1}(t_D)$. That is, r_R is the relative income of a voter who would have ideal tax rate t_R. To facilitate estimation, we assume quadratic utility: $u(t_R \mid r_i) = -(r_i - r_R)^2$.[12] Since a voter's party identification may depend on factors other than relative income, let $\boldsymbol{x}_i$ be a vector of other factors that determine support for the Republican party and ε_i be individually idiosyncratic factors.

Our model of Republican Party ID is therefore

$$\begin{aligned}
\text{Republican ID} &= \alpha + \beta\left[-(r_i - r_R)^2 + (r_i - r_D)^2\right] + \theta \boldsymbol{x}_i + \varepsilon_i \\
&= \alpha + \beta\left(r_D^2 - r_R^2\right) + 2\beta\left(r_R - r_D\right) r_i + \theta \boldsymbol{x}_i + \varepsilon_i \\
&= \tilde{\alpha} + \tilde{\beta} r_i + \theta \boldsymbol{x}_i + \varepsilon_i \qquad [1]
\end{aligned}$$

where $\tilde{\alpha} = \alpha + \beta\left(r_D^2 - r_R^2\right)$ and $\tilde{\beta} = 2\beta\left(r_R - r_D\right)$.

Given this model, we can identify several factors that in principle could account for the increased stratification of partisanship by income.

H1: *Inequality*. Increases in economic inequality may have led to more extreme values of the r_i. A standard measure of economic inequality is the ratio of the income of the top quintile to that of the bottom quintile. Thus, increased inequality would raise the mean value of *r* for the upper quintile and/or reduce the mean value of *r* for the lowest quintile.

H2: *Response polarization.* Party polarization on economic issues as reflected by $r_R - r_D$ has increased. From equation [1], this increases $\tilde{\beta}$.

H3: Other determinants of party identification such as race, gender, region, education, and age have become more related to income. Therefore, income stratification may be a by-product of the differential economic success of the demographic groups that compose each party.

H4: Poorer demographic and social groups have moved towards the Democrats while wealthier groups have identified more with the Republicans.

Before assessing these different possibilities, we turn to some important data and estimation issues.

Data

We employ the National Election Studies from 1952 to 2000 to estimate equation [1]. Our dependent variable is the seven-point scale of partisanship that ranges from Strong Democrat to Strong Republican. Unfortunately, NES data poses a number of problems specific to the estimation of our model. Perhaps the biggest problem is that the NES does not report actual incomes, but allows respondents to place themselves into various income categories. We use Census data on the distribution of household income to estimate the expected income within each category. These estimates provide an income measure that preserves cardinality and comparability over time. The details of our procedure are in the Appendix.

In addition to the constructed income variable, we include a number of control variables that other studies have found to be related to partisanship. These include race, region, gender, age, and education. We combine race and region to create two categorical variables, African-Americans and southern non-African-Americans. The residual category is northern non-African-Americans. We measure education by distinguishing between those respondents who have “Some College” or a “College Degree” from those who have only a high school diploma or less. We also include the age of the respondent. In fact, race, gender, age, and education are the only demographics that are available on all of the 24 National Election Study presidential and midterm year surveys from 1952 through 2000. (There was no midterm study in 1954.)

It is important to note that these additional variables are not only statistical controls, but they are also variables that are not distributed randomly across income levels. Thus, both changes in the joint distribution of these variables with income and changes in their relationship to partisanship may have effects on the extent to which partisanship is stratified by income.

Finally, to control for election-specific effects on partisanship, we include election fixed effects (year dummies) in the estimation.

3. Estimation

Given the fact that our dependent variable, partisan identification, is multichotomous and distributed bimodally, ordinary least squares is a highly inappropriate way of estimating our model. As is standard, we assume that the partisanship variable is a set of ordered categories and estimate an ordered probit model (McKelvey and Zavoina, 1975). To capture changes in the relationship between income and other variables to partisanship, we assume that the coefficients of equation [1] can change over time. For our relative income variable, we estimate several different specifications that restrict the movement of β^t in various ways. We report four sets of results corresponding to a constant income effect, an effect with a linear trend, an effect with cubic trend, and an income effect “dummied” for each of the five decades represented in our dataset. We also allow the effects of other variables to change over time with linear trends. Finally, we assume that the category thresholds estimated by the ordered probit are constant over time. Thus, the distribution of responses across categories changes only with respect to changes in the substantive coefficients and the distribution of the independent variables.[13]

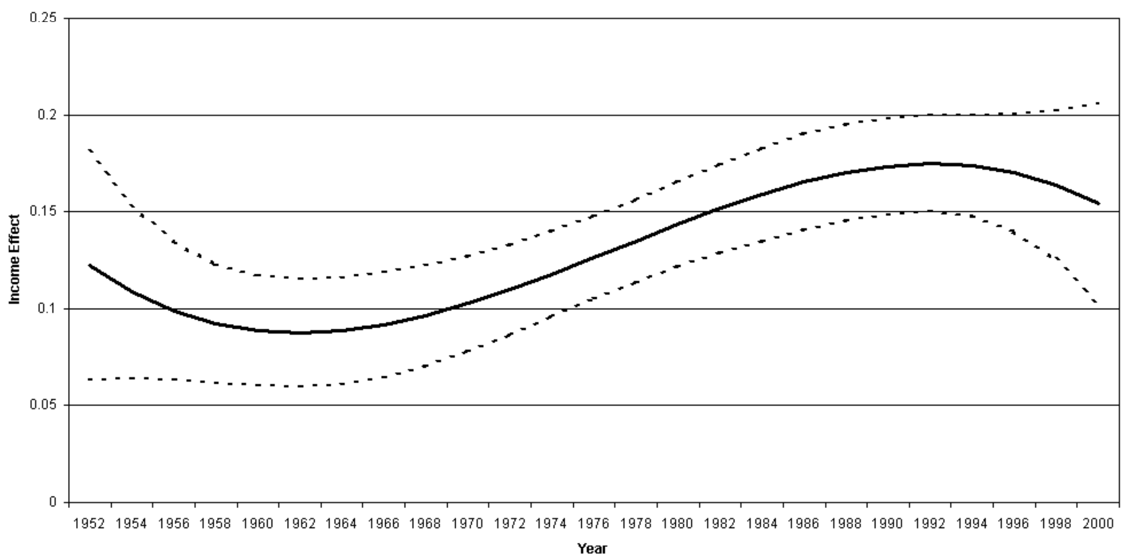
4. Results

Table 1 presents the estimates of our model for the four specifications of the income effect. Not surprisingly, across all four specifications, relative income is a statistically significant factor in the level of Republican partisanship. Column (1) presents the model with a constant income effect. While statistically significant, the estimate of the constant effect is rather small. An individual with twice the average income ($r_i = 2$) has latent partisanship measure that is only 0.139 larger than an individual with an average income. Given that the distance between the category thresholds averages more than 0.3, this effect is less than one position on the partisanship scale.

The small average effect of income masks a definite trend over the entire period. Model (2) simplifies matters by assuming that the income effect changes only linearly. This model produces a statistically significant growth rate in the income coefficient of .0021 per year. From 1952 to 2000, the income effect is estimated to have doubled, rising from 0.100 to 0.208. There results are echoed by model (4), a specification with a separate income effect for each decade. Each subsequent decade has a higher estimated income effect. The estimated income effects have grown substantially over time.

Our results on elite polarization show a decline in polarization after World War II followed by a subsequent rise. In addition, figure 1 suggests that there may have been a fall in income stratification at the end of the 90s. To allow for both of these effects, we estimated model (3) where income effects follow a cubic trend. The results are displayed in figure 3.

Figure 3
Income Effect in Party Identification
Cubic Polynomial Estimates



A first observation from figure 3 is that the confidence interval is always well above 0—income matters. The results, moreover, through the mid 90s, roughly match our earlier observations about elite polarization and income inequality. There is an initial (albeit imprecisely estimated) decline followed by an increase. The turning point, however, precedes the turning point in figure 1 by about a decade. At the end of the 90s, the income effect appears to decline, echoing popular claims that politics has now turned to abortion and other social issues. The decline, however, is imprecisely estimated—the upper bound of the confidence interval is still slightly increasing. Caution is in order as to the import of recent changes.

The effect of income in the formal theory represented in equation (1) calls for the effect of income to be expressed as the product of an underlying behavioral parameter (β) and the difference in the party platforms. The theory predicts that if we include relative income and the product of the difference in party platforms and income, only the interacted variable should be significant. To do an explicit test, we interacted relative income with the party polarization measures shown in Figure 1. Although the specification did not fit as well as the linear trend specification, the results are encouraging. The estimated coefficient of relative income, -0.099, was less in magnitude that its standard error, 0.101. In contrast, the estimated coefficient on the interacted variable was 0.393 and was significant at p=0.02.

We now turn to the results for coefficients other than income.

The constant in the formal model is contained in the (unreported) year fixed-effects. Were the term, α, constant in time, the measured fixed-effects should, theoretically, be decreasing in time. Observe that the second term in $\tilde{\alpha} = \alpha + \beta\left(r_D^2 - r_R^2\right)$ can be expanded to $\beta\left(r_D - r_R\right)\left(r_D + r_R\right)$. The coefficient β should be positive, as should the sum of the relative incomes that represent the platforms. In fact, the sum should be growing given that the Republicans have moved to the right and the Democrats have tread water (Poole and Rosenthal, 2001). The difference in platforms should be negative and growing, given the increase in polarization. The second term, therefore, should be negative. If the behavioral parameter α were constant, we should see a decreasing sequence of year fixed-effects.

In fact, the reverse occurs. The result is illustrated by the step model where the year fixed-effects are less than 0.20 before 1964 and greater than 0.65 after 1992. The regression of the year fixed-effects on time shows an R^2 of 0.92 and a t-statistic of 15.5. These results suggest, given our strong priors about the second term, that the behavioral parameter α is growing even more sharply in time than the fixed effects. That is, there are trends favoring Republican identification that are not picked up in the effects of income and demographics, including trends in these effects.

Turning to the effects of the other demographic variables, we find that the effect of each has changed dramatically over the period of our study. These changes should not be surprising to casual observers. African-Americans and females have moved away from the Republican Party, just as non-black southerners have flocked towards it. While older voters supported the Republicans in the mid-twentieth century, their allegiance deteriorated by the twenty-first. The effects of education have diminished in size, but this in part reflects the fact that college attendance and graduation have sharply increased.

The effect of income is very important compared to that of the demographics. Consider the estimates from column (4) of table 1. In table 2, we show, for the step model in 2000, how much the income of a respondent at half average income would need to increase to match the change of the other variables. Only in the case of race, would an extreme income change be needed to match the effect of the other demographic.

Table 1:
Effects of Relative Income on Republican Partisanship
Ordered Probit
(s.e. in parentheses)

	(1) Constant Income Effect	(3) Trended Income Effect	(3) Cubic Income Effect	(4) Step Income Effect
Relative Income	0.134 (0.008)	0.079 (0.017)	0.130 (0.034)	
Relative Income x [(Year-1951)/10]		0.021 (0.006)	-0.086 (0.054)	
Relative Income x $[(\text{Year-1951})/10]^2$			0.049 (0.024)	
Relative Income x $[(\text{Year-1951})/10]^3$			-0.006 (0.003)	
Relative Income x (1952-1960)				0.089 (0.020)
Relative Income x (1962-1970)				0.112 (0.018)
Relative Income x (1972-1980)				0.117 (0.015)
Relative Income x (1982-1990)				0.160 (0.016)
Relative Income x (1992-2000)				0.175 (0.016)

Table 1 (continued)

	(1) Constant Effect	(3) Trended Income Effect	(3) Cubic Income Effect	(4) Step Income Effect
African-American	-0.663	-0.685	-0.686	-0.686
	(0.042)	(0.042)	(0.042)	(0.042)
African-Amer. x (Year-1951)/10	-0.051	-0.043	-0.043	-0.043
	(0.014)	(0.014)	(0.014)	(0.014)
Female	0.145	0.139	0.140	0.139
	(0.024)	(0.024)	(0.024)	(0.024)
Female x (Year-1951)/10	-0.056	-0.054	-0.054	-0.054
	(0.008)	(0.008)	(0.008)	(0.008)
Southern Non-Black	-0.603	-0.610	-0.610	-0.610
	(0.029)	(0.029)	(0.029)	(0.029)
South Non-Black x (Year-1951)/10	0.146	0.149	0.149	0.149
	(0.009)	(0.009)	(0.009)	(0.009)
Some College	0.293	0.310	0.310	0.311
	(0.036)	(0.036)	(0.036)	(0.036)
Some College x (Year-1951)/10	-0.039	-0.045	-0.046	-0.046
	(0.011)	(0.011)	(0.011)	(0.011)
College Degree	0.410	0.450	0.451	0.453
	(0.038)	(0.040)	(0.040)	(0.040)
College Degree x (Year-1951)/10	-0.072	-0.087	-0.087	-0.088
	(0.012)	(0.012)	(0.012)	(0.012)
Age/10	0.081	0.078	0.077	0.078
	(0.008)	(0.008)	(0.008)	(0.008)
Age/10 x (Year-1951)/10	-0.029	-0.029	-0.028	-0.028
	(0.002)	(0.002)	(0.002)	(0.002)
μ_1	-0.418	-0.489	-0.447	-0.481
	(0.046)	(0.050)	(0.057)	(0.051)
μ_2	0.283	0.213	0.255	0.221
	(0.046)	(0.050)	(0.057)	(0.051)
μ_3	0.587	0.516	0.558	0.524
	(0.046)	(0.050)	(0.057)	(0.051)
μ_4	0.883	0.812	0.855	0.821
	(0.046)	(0.050)	(0.057)	(0.051)
μ_5	1.184	1.114	1.156	1.122
	(0.047)	(0.051)	(0.057)	(0.051)
μ_6	1.769	1.698	1.741	1.706
	(0.047)	(0.051)	(0.057)	(0.052)
Log-likelihood	-71849.2	-71842.8	-71840.8	-71841.3
Likelihood Ratio p-value (H0 = Constant Effect)		0.000	0.001	0.003
Number of Observations	38949	38949	38949	38949

Table 2.

Demographics and Income Shifts Compared

Demographic Change	Pro-Republican Party Identification Shift	Equivalent Relative Income Shift is from ½ ave. income to:
Black to non-black northern	0.896	5.62 ave.
85 to 25 years old	0.355	2.52 ave.
Female to male	0.126	1.72 ave.
Non-black northern to non-black southern	0.120	1.19 ave.
No college to college grad	0.021	0.62 ave.

5. What Caused the Increase in Party-Income Stratification?

In this section, we attempt to assess the relative importance of H1-H3 in increasing income/party stratification. We will use our estimates of equation [1] to compute implied levels of stratification under various scenarios. Consistent with testing H1-H3, we can manipulate the coefficients of the model, the distribution of r_i, and the joint distribution of r_i and the other demographic variables. To assess the relative importance of each of these changes, we compute the levels of party-income stratification in 1960 and 1996 under different scenarios using the results of the "cubic" specification in column 3 of Table 1.

Table 3: Characteristics of Income Quintiles, 1960 and 1996

Variable	Top Quintile 1996	Bottom Quintile 1996	Ratio 1996	Top Quintile 1960	Bottom Quintile 1960	Ratio 1960
Average Relative Income	2.423	0.183	13.250	2.098	0.210	9.972
% African-American	2.9	25.0	0.118	1.7	17.3	0.097
% Female	44.6	69.0	0.647	50.0	63.0	0.794
% Southern	33.3	48.0	0.695	24.2	45.7	0.529
% Some College	13.7	17.2	0.796	18.5	4.3	4.271
% College Degree	72.1	14.4	5.015	24.2	3.1	7.828
Average Age	45	51	0.874	45	63	0.711

Before turning to the question of what accounts for the change in party-income stratification, we first consider the types of demographic changes that have occurred over this period. Table 3 gives the profiles of the lowest and highest income quintiles for the 1956 and 1996 surveys. A respondent is in the lowest quintile if his or her family income or single income is below the 20th percentile point of the March CPS household income distribution and in the highest quintile if above the 80th percentile point.

A comparison of the quintile ratio columns shows the magnitude by which the income distribution and the joint distribution of income and other attributes have changed over the past 40 years. The top-bottom quintile ratio for average relative income has increased from under 10 to over 13. Beyond this striking change in the distribution of income, we find large changes in the placement of groups within the distribution.

Some changes have worked against the increased stratification of partisanship on income. This is true of education. Both measures of education are distributed more equitably in 1996 than 1960 while their correlation with Republican partisanship has diminished substantially. The changing distribution of age and its relation to partisanship also works against the increased overrepresentation of Republican identifiers in the top quintile. This reflects the fact that the bottom quintile is relatively younger in 1996 while age is negatively correlated with Republican identification 1996 whereas it was positively correlated in 1960.

However, changes in the income distribution of the other demographic categories clearly work to increase stratification. With the increase in single females from 1960 to 1996, females have become a notably larger share of the lowest quintile respondents and a lower share of the top quintile. Since females have moved steadily towards the Democratic Party, the effects on party-income stratification are quite apparent.[14] Alternatively, southerners have become better represented in the top quintile as they moved into the Republican Party. This also contributes to stratification.

The changes with respect to race are more ambiguous. Income inequality *among* African-Americans has increased dramatically so that blacks now compose a greater fraction of *both* of the extreme income quintiles. African-Americans as a group are largely Democratic identifiers. If the propensity to choose a Democratic identification were independent of income, the black increase at the top quintile would decrease stratification and the increase at the bottom would increase it. However, controlling for income, we find that the propensity of African-Americans to identify with the Democrats has increased. Since blacks remain substantially over represented at the bottom and under represented at the top, the fact that they have become more Democratic increases stratification. This effect of increased African-American identification with the

Democrats dominates the effects arising from the changes in the income distribution of blacks.

To quantify the magnitude of some these effects, we simulate stratification scores for 1960 and 1996 using the results of the cubic income effect model. We manipulate the model and the profiles in order to assess which factors most contributed to the increased stratification. These results are given in Table 4.

The first two rows of Table 4 reflect the estimated stratification for each year using the actual model. That is, for each respondent in the top quintile, we use the estimated coefficients to compute the probability that the respondent is a Republican (strong and weak) identifier. We then sum these probabilities to estimate the fraction of the top quintile that are Republican identifiers. We do the same for the bottom quintile. The ratio of the two fractions is our stratification measure. These results are benchmarks for comparison with other counterfactuals. They are somewhat greater for 1960 and somewhat less for 1996 than the actual stratifications reported in Figure 2.

Our first exercise untangles whether the change in stratification is driven by changes in estimated model effects or by changes in demographics. In row 3, we estimate stratification using the estimated coefficients for 1996 applied to the 1960 sample respondents. The result is a stratification score of 1.950, which is only slightly smaller than the actual 1996 estimated score of 2.036. Alternatively, row 4 shows the estimated stratification applying the 1960 coefficients to the 1996 respondents to capture the effects of the demographic shifts. The resulting stratification of 1.804 is substantially further in the direction of the estimated stratification for 1960 (row 1). These two results imply that the changes in the relationship between partisanship and the demographic variables accounts for much more of the stratification change than the demographic shifts. That is, while there have been important changes in the distribution of our demographic variables in the last half of the twentieth century—for example, more income inequality, higher levels of education, and a greater share of the population in the South—the effect of these changes on aggregate partisan identification have been largely offsetting. In contrast, how demographic characteristics relate to identification—for examples, the increasing effect of income, the flip in the gender gap—have had important net effects.

The remaining rows of table 3 deal specifically with the direct effects of relative income. Rows 5 and 6 correspond to counterfactual estimates of stratification in each year using the degree of income inequality in the other year. For row 5, we use the results in table 2 to multiply top 1960 quintile incomes by 2.201/1.900 (see table 3) and bottom quintile incomes by .210/.279. Otherwise, we use the 1960 sample and coefficients. For row 6, we reverse the process to simulate 1996 stratification with the 1960 income distribution. These results show that the aggregate distribution of income has barely any effect on stratification. In both cases, the counterfactual stratification indices are almost identical to the actual ones. This suggests that increasing income inequality accounts for very little of the change in stratification. The change is largely one of increased "pocketbook" partisanship.

To see this, we turn, finally, to the effects of the increased impact of relative income on partisanship. In row 7, we estimate 1996 stratification using the 1996 sample and all coefficients except for that on relative income, where we substitute the 1960 coefficient. In row 8, we reverse the roles of 1960 and 1996. The two resulting stratifications are about equal. That is, holding demographics and other effects constant, the change in the income effect substantially increases stratification in 1960 and decreases it in 1996. The 1996 stratification in row 8, 1.790, is nearly identical to that in row 4, 1.804, where all the coefficients were changed. That is, just as the changes in demographic profiles offset, the changes in demographic coefficients offset, except for the increased effect of income.

These results suggest that the driving force behind the increased stratification was the increased correlation between income and partisanship. Thus, if we interpret this increase as party polarization, these findings suggest that the changes in the bivariate relationship can be best accounted for by the actions of the party elites and not the voters.

Table 4: Determinants of Party/Income Stratification

Scenario	Average Republican Probability of Lowest Quintile	Average Republican Probability of Highest Quintile	Party/Income Stratification
1960	0.205	0.311	1.518
1996	0.195	0.397	2.036
1996 with 1960 Sample	0.190	0.370	1.950
1990 with 1996 Sample	0.214	0.393	1.804
1960 with 1996 Income	0.204	0.321	1.574
1996 with 1960 Income	0.196	0.376	1.914
1996 with 1960 Income Effect	0.192	0.332	1.737
1960 with 1996 Income Effect	0.210	0.376	1.790

Note: Probabilities are the estimated probabilities of weak or strong Republican identification from the model with a cubic specification of income effects.

6. Political Competition in a Richer Society

In the period of our study, the American political system has remained remarkably competitive. The Republicans have won seven of the presidential elections, just one more than an even split. The NES sample percentages of Republican partisans from 1956 to 1998 have fluctuated, with no apparent trend, from 21% to 30% (Green, Schickler, and Palmquist, 2002, p. 15). (There has been a decline for the Democrats, to the benefit of Independents.)

Should this balance have been maintained? Real median income doubled between 1952 and 1996. Average income increased even more sharply. Should not this change have benefited the Republicans?

If respondents computed their relative income based not on average income in the year of the survey but on average income in 1960, this would clearly have been the case. Table 5 shows the actual fraction of Republican identifiers in the sample, the estimated fraction using the model coefficients and the estimated fraction replacing average real income in the year with average real income in 1960 in computing relative income, r_i.

The results show that from 1988 onward that the increase in real incomes would have generated a gain of over 3 percent in Republican identification had respondents compared their current incomes to 1960 average incomes. Given that the actual system is very competitive, a gain of 3 percent would likely have swung many offices to Republicans. Arguably, the real incomes represented by the parties have increased in a way that preserves a competitive two-party system. The richer voters represented by both parties are, again to speculate, less likely to favor redistribution and social insurance than were the counterparts of these voters a half-century earlier.

Table 5. Republican Identification and the Change in Real Income.

Year	Actual	Estimated from Model	Estimated Using 1960 Mean Income in Computing Relative Income
1956	0.299	0.260	0.258
1960	0.297	0.245	0.245
1964	0.219	0.173	0.177
1968	0.229	0.214	0.224
1972	0.240	0.262	0.276
1976	0.238	0.259	0.279
1980	0.228	0.245	0.266
1984	0.270	0.279	0.300
1988	0.285	0.295	0.325
1992	0.253	0.270	0.298
1996	0.276	0.269	0.305
2000	0.263	0.285	0.341

7. Did the South Do It?

In the period of our study, the last half of the twentieth century, the American South transited from a one-party system to a two-party system. In the results we have just presented, we have seen how the Southern non-blacks switched from being substantially more Democratic than Northern non-blacks to now being substantially more Republican. This change in partisan identification has been examined previously by Green, Schickler, and Palmquist (2002). Our contribution is to indicate that pocketbook voting is an important part of the story of the dramatic switch of partisan allegiances in the South.

What happened in the South with respect to income is vividly illustrated by figures 4 and 5. Figure 4 shows results for the cubic polynomial estimation when the model is estimated with only non-black respondents in the South. Figure 5 is the comparable figure for northern non-blacks.

Figure 4
Income Effect in Party Identification
Southern non-blacks
Cubic Polynomial Estimates

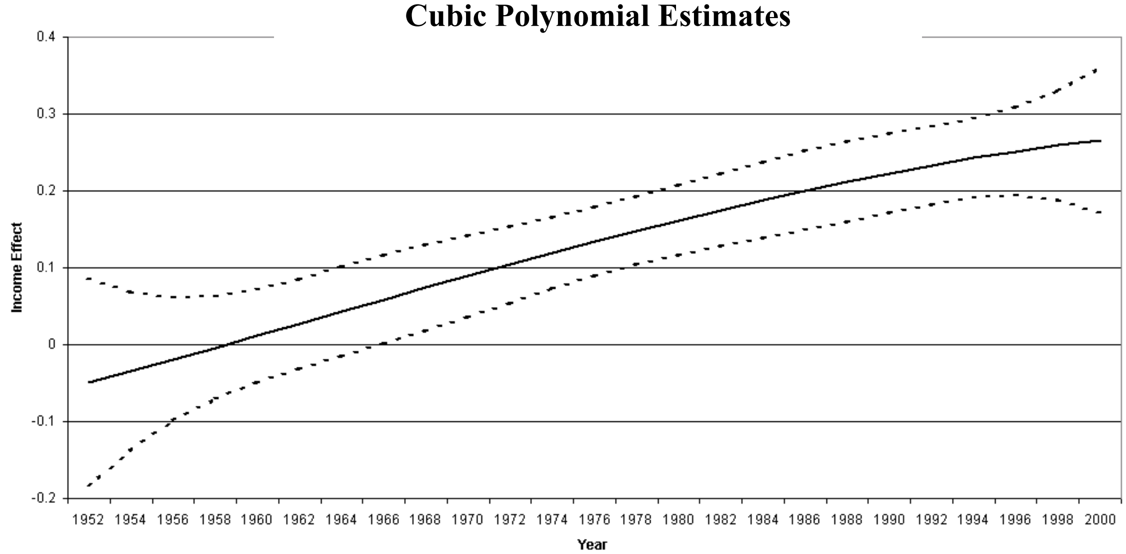
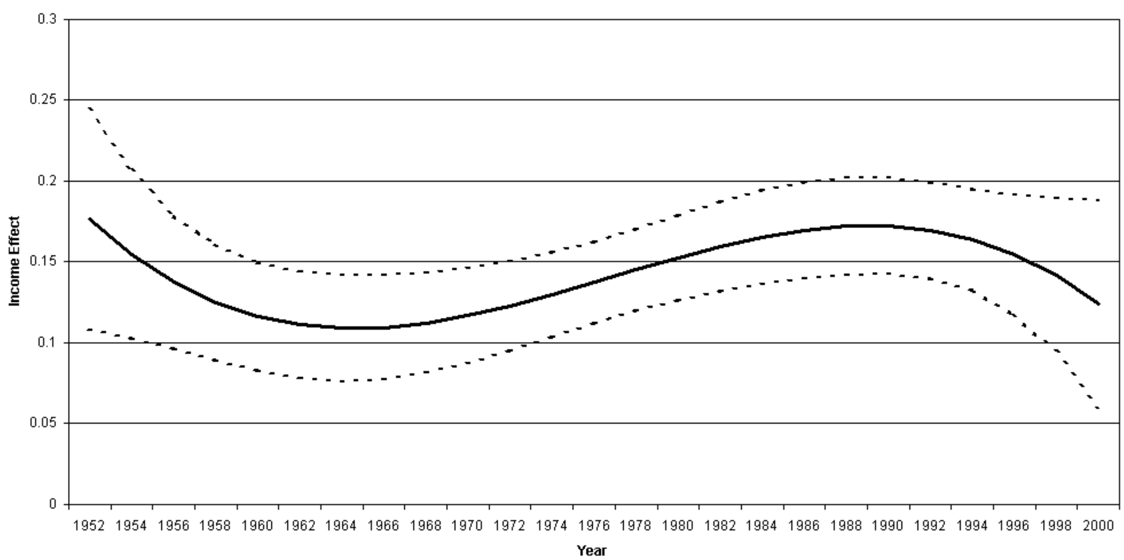


Figure 5
Income Effect in Party Identification
Northern non-blacks
Cubic Polynomial Estimates



The South shows a sharply increasing income effect. Income had essentially no effect on southern partisanship in the 1950s. The confidence interval shown in figure 4 includes 0 until 1966. A likelihood ratio test of the linear effect rejects the null hypothesis of a constant income effect. As suggested by the figure, testing a cubic model against the linear does not lead to rejection of the null hypothesis that the increase is linear.

The results for the North form a stark contrast to the South. While the confidence interval always lies above 0, the constant income effect model is not rejected for the linear model but is rejected for the cubic. The North went through a declining income effect in the 1950s when elite polarization was also decreasing and an increasing effect in the 70s and 80s when elite polarization was increasing. But, intriguingly, the income effect in the North declined in the 90s to be no higher than in the 60s. The overall increase in the income effect shown in figure 3 is largely the result of the transformation of southern politics and the increased demographic weight of the South.

8. Presidential Voting: A Replication

Our results for partisan identification replicate nicely in a dichotomous probit analysis of presidential vote choice. The dependent variable is coded "1" for Republican and "0" for Democrat. Our sample here is defined only by those individuals who expressed a choice for one of the major party candidates in the presidential year. Declared abstentions, votes for minor party candidates, and non-responses resulted in our having only about two-thirds as many observations as for the partisan identification analysis. We use the same independent variable specifications as in our analysis of partisan identification. In table 6, we show the results for the income variables. The results for the demographics are available on request from the authors.

For presidential voting, relative income continues to have a significant effect as shown in column (1) of table 6. The result for the linear trend model in column (2) is very similar to that for partisan identification, particularly after considering the effects of sample size on precision. Similarly, the pattern of the time polynomial coefficients for the cubic model is quite similar in comparing models (3) from table 1 and table 6.

The major distinction between the partisan identification and vote choice estimates lies in column (4). Whereas we saw a steady increase in income based voting for partisan identification, the vote choice coefficient for the last decade is smaller than that for the previous one.

This reversal is largely driven, from inspection of year-by-year estimates, by a sharp drop in the income effect for the 2000 election. The 2000 election presidential vote choice results may indicate a fundamental shift in American politics. Brady (2001) and many others have noted that Gore won the rich, red states along the two oceans and Bush won the poorer, blue ones in the interior. (Of course, within each state, high income and Republican voting may still go together.) Brady attributes the geographic shift to a rise in the importance of "moral" as against economic issues.

Whether "moral" will come to dominate "economic" politics is an interesting, but open question. We note that, while 2000 showed a drop in the income coefficient for partisan identification as well as for vote choice, the variation is well within other election-to-election shifts. Column (4) of table 1, in contrast to that of table 6, shows a steady increase in the effect of income. When one looks at partisan identification as well as vote choice, it is hard to see a fundamental shift in the political system.

Table 6:
Effects of Relative Income on Presidential Vote Choice
Probit (s.e. in parentheses)

	(1) Constant Income Effect	(3) Trended Income Effect	(3) Cubic Income Effect	(4) Step Income Effect
Relative Income	0.173	0.114	0.128	
	(0.014)	(0.027)	(0.042)	
Relative Income x [(Year-1951)/10]		0.025	-0.074	
		(0.010)	(0.077)	
Relative Income x $[(\text{Year-1951})/10]^2$			0.068	
			(0.039)	
Relative Income x $[(\text{Year-1951})/10]^3$			-0.011	
			(0.005)	
Relative Income x (1952+1956)				0.118
				(0.034)
Relative Income x (1960+1964+1968)				0.116
				(0.029)
Relative Income x (1972+1976)				0.192
				(0.031)
Relative Income x (1980+1984+1988)				0.245
				(0.031)
Relative Income x (1992+1996+2000)				0.188
				(0.030)
Log-likelihood	-8853.672	-8842.253	-8850.115	-8844.121
Likelihood Ratio p-value (H0 = Constant Effect)		0.010	0.000	0.016
Number of Observations	14328	14328	14328	14328

9. Conclusion

High income Americans have consistently, over the second half of the twentieth century, been more prone to identify with and vote for the Republican party than have low income Americans, who have sided with the Democrats. The impact of income persists when controlling for other demographics, and the impact's magnitude is important. Moreover, there has been a rather substantial transformation in the economic basis of the American party system. In the 1990s, income was far more important than it had been in the 1950s. While American politics is certainly far from purely class based, the divergence in partisan identifications and voting between high and low income individuals has been striking. Certainly, this trend helps to explain the conflicts over taxation of estates and dividends in an era generally presumed to be dominated by "hot button" social issues like abortion and guns.

In our simple theoretical model, we posited that relative, not absolute, income was important to voting behavior. As average incomes rose in the last half of the twentieth century, voters and political parties, we believe, made adjustments that maintained an extremely competitive, most strikingly in the 2000 presidential race, two-party system. Indeed a simulation suggested that the Republicans would have a more than three percent additional advantage in partisan identification were voters comparing their current incomes to 1960 average income.

There are, of course, multiple sources to the increased political divergence between high and low income voters. But our evidence shows changes in both overall income inequality and the incomes of various demographic groups have only marginally contributed to increased partisan stratification on income; the most important contributions seem to come from partisan polarization and the southern realignment. As our model would suggest, the coefficient of relative income roughly tracks patterns of elite polarization derived from congressional voting studies. Indeed, as our simulation results show, the increase in this coefficient seems to be primary responsible for the increased connection between income and partisanship.

It is not terribly surprising that the southern realignment also plays an important role in our findings because it is the most important change in the American party system during the 20th century. However, our results about the effects of changes in southern politics differ substantially from arguments stressing the role of race and social issues. While not denying the importance of these factors, we find that the political attachments of the contemporary South are driven by income and economic status to an extent even greater than the rest of the country.

It is probably too early to tell whether recent declines in income-based partisanship and voting in the north are anything more than the effects of fat wallets produced by the economic boom of the 1990s. However, even if this decline proves to be fundamental and enduring, the accelerating role of income in southern politics and the South's increasing share of the national electorate will likely prevent any significant de-polarization of American politics in the near future.

Appendix 1: Approximating Incomes for NES Categories

Given categorical income data, there are two typical approaches to comparing income responses at different points in time. Let $\boldsymbol{x}_t = \{x_{1t}, \ldots, x_{kt} = \infty\}$ be the vector of upper bounds for the NES income categories at time *t*. The first approach is to use the categories ordinally by converting them to income percentiles for each time period. However, this approach throws away potentially useful cardinal information about income. Further, as it is unlikely that income categories will always coincide with a particular set of income percentiles, some respondents will have to be assigned *ad hoc* to percentile categories. A second approach is to assume that the true income is a weighted average of the income bounds. Formally, one might assume that the true income for response *k* at time *t* is $\alpha x_{k-1,t} + (1-\alpha) x_{kt}$ for some $\alpha \in [0,1]$. However, the true weight will depend on the exact shape of the income distribution. When the income density is increasing in $[x_{k-1,t}, x_{kt}]$, the weight on x_{kt} should be higher than when the density is decreasing over the interval. Thus, the same weights cannot be used for each category at a particular point in time, or even the same category over time.

Since neither of these two approaches can be used to generate the appropriate data, we use Census data on the distribution of income to estimate the expected income within each category. These estimates provide an income measure that preserves cardinality and comparability over time.

To outline our procedure, let $\boldsymbol{y}_t = \{y_{1t}, \ldots, y_{mt}\}$ be the income levels reported by the census corresponding to a vector of percentiles $\boldsymbol{z}_t = \{z_{1t}, \ldots, z_{mt}\}$. We use family income quintiles and the top 5%. Therefore, for 1996,

$y_{1996} = \{\$18485, \$33830, \$52565, \$81199, \$146500\}$ and $z_{1996} = \{.2, .4, .6, .8, .95\}$. We assume that the true distribution of income has a distribution function $F(\cdot \mid \mathbf{\Omega}_t)$ where $\mathbf{\Omega}_t$ is a vector of time specific parameters. Therefore, $F(\mathbf{y}_t \mid \mathbf{\Omega}_t) = \mathbf{z}_t$. In order to generate estimates $\hat{\mathbf{\Omega}}_t$, let $w(\hat{\mathbf{\Omega}}_t) = F(\mathbf{y}_t \mid \hat{\mathbf{\Omega}}_t) - \mathbf{z}_t$. We then choose $\hat{\mathbf{\Omega}}_t$ to minimize $w(\hat{\mathbf{\Omega}}_t)' \, w(\hat{\mathbf{\Omega}}_t)$. Given an estimate of $\hat{\mathbf{\Omega}}_t$, we can compute the expected income within each NES category as

$$EI_{kt} = \begin{cases} \left[F\left(x_{1t} \mid \hat{\mathbf{\Omega}}_t\right)\right]^{-1} \int_0^{x_{1t}} x dF\left(x \mid \hat{\mathbf{\Omega}}_t\right) & k = 1 \\ \left[F\left(x_{kt} \mid \hat{\mathbf{\Omega}}_t\right) - F\left(x_{k-1,t} \mid \hat{\mathbf{\Omega}}_t\right)\right]^{-1} \int_{x_{k-1,t}}^{x_{kt}} x dF\left(x \mid \hat{\mathbf{\Omega}}_t\right) & otherwise \end{cases}$$

We assume $F(\cdot)$ log-normal with $\mathbf{\Omega}_t = \{\mu_t, \sigma_t\}$. These parameters have very straightforward interpretations. The median income at time *t* is simply e^{μ_t} while σ_t^2 is the variance of log income that is a commonly used measure of inequality. Table A1 gives the estimates of $\hat{\mathbf{\Omega}}_t$ for each presidential election year. These results underscore the extent to which the income distribution has become more unequal.

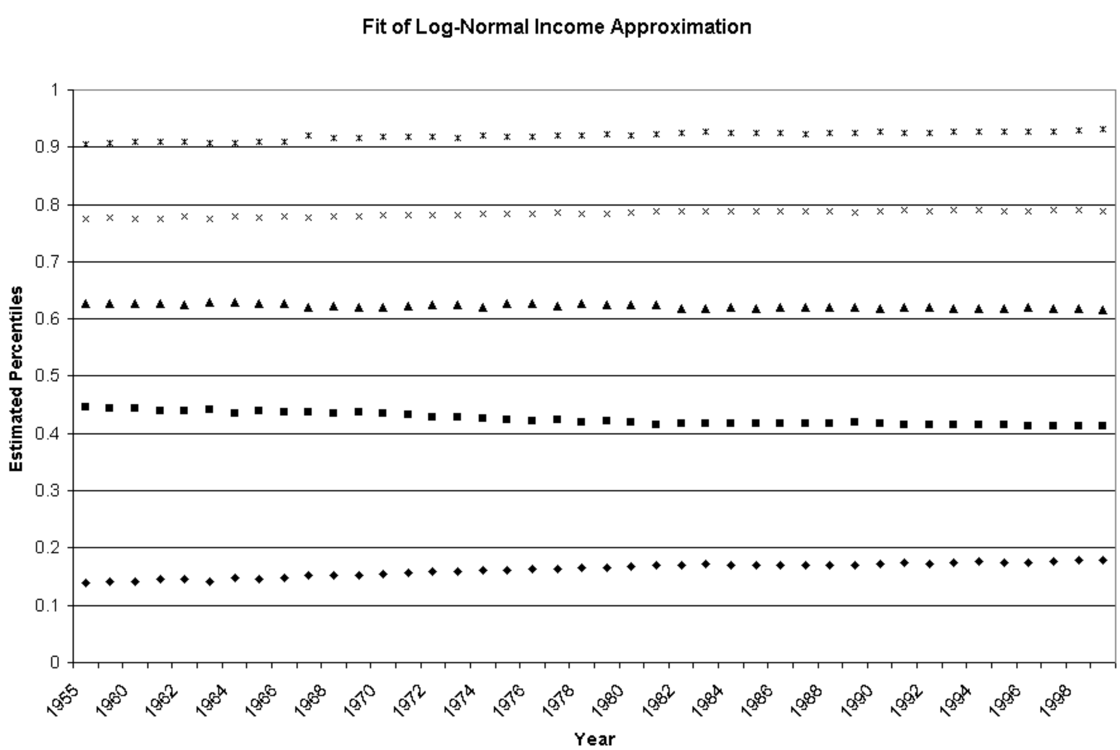
Table A1

Election	μ_t	σ_t
1952	7.894	0.817
1956	8.172	0.804
1960	8.401	0.795
1964	8.544	0.811
1968	8.834	0.738
1972	9.076	0.746
1976	9.357	0.765
1980	9.698	0.776
1984	9.959	0.794

1988	10.167	0.812
1992	10.314	0.824
1996	10.437	0.843
2000	10.617	0.857

Figure A1 plots $F\left(\mathbf{y}_t \mid \hat{\mathbf{\Omega}}_t\right)$ against z_t and shows how well the log-normal approximates the distribution of income -- if it were a perfect fit, the lines would track .2, 4, .6, .8, and .95 exactly. While the approximation is generally very good, the log-normal is a poor approximation of incomes at lower levels as the lowest line is generally below .2. This is because the true distribution of income has a larger mass near zero and a larger tail than the log-normal. The effect is that EI_{kt} has a slight positive bias for low incomes and a slight negative bias for large incomes.

Figure A1



Appendix 2: Measuring Income-Party Stratification

Three complications arise in using the NES to measure the partisanship of the top and bottom income quintiles:

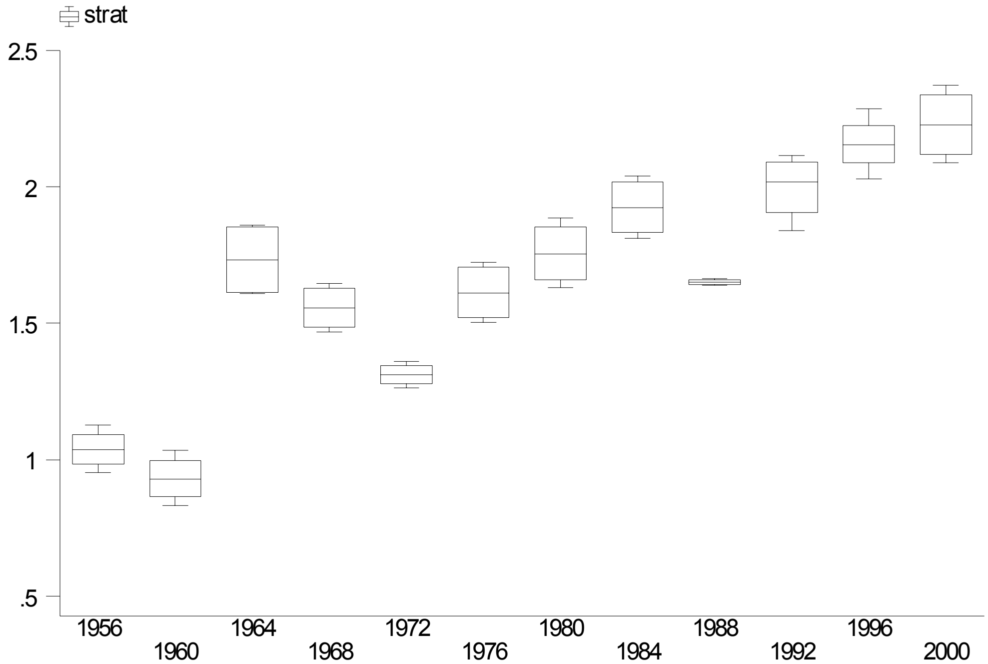
1. **Unrepresentative NES samples:** There are several years in which the distribution of respondent's income is very unrepresentative of the income distribution reported by the Census Bureau.
2. **NES sample matches neither the "Family" nor "Household" samples for which the Census Bureau reports income quintiles.** The NES asks respondents for the income of their family for the previous year. For single voters, the NES asks their individual income. Thus, the NES sample includes families and single person households. However, the Census family sample does not include single persons living alone, and the household sample aggregates multiple families living at the same household, but does include single householders. Thus, neither Census sample matches the NES.
3. **Income quintile measures will often fall within NES income categories.** When a quintile measure falls within an income category, the issue arises as to how to allocate the respondents in that category into the adjoining quintiles.

It is very difficult to solve all three of these problems for the entire period from 1952 to 2000. Problem 1 necessitates matching the NES sample with the income distribution from the Current Population Survey, but problem 2 necessitates recomputing that distribution for units more closely resembling those of the NES. However, even with appropriate measures of the income quintiles, problem 3 has no obvious solution. In figure 2, we use our log-normal approximation of the distribution of household income to compute expected income for each NES category. We use these estimates to classify respondents into income quintiles based on the Census Bureau's reported income limits for household income quintiles.

Given the limitations of these choices, we did a number of other calculations to see if the results in figure 2 are robust. To deal with problem 1, we recomputed the stratification measures using both household and family income distributions to classify respondents. We also use samples from the November Current Population Survey's from 1972 and 1996 which include single individuals and families so as to approximate the NES population and minimize problem 2.[15] Since the November CPS data is categorical, we use both linear and exponential extrapolation to compute the 20^{th} and 80^{th} percentiles. Thus, combining all of these data sources, we have four quintiles estimates for 1972 to 1996 and two for each the other years.

To deal with problem 3, we experiment with various ways of allocating respondents into quintiles. We do four computations for each quintile measure by including or excluding the relevant NES category in the top and bottom quintiles. Thus, we have 16 total stratification measure for 1972-1996 and 8 for the other years. Figure A2 is a "box and whiskers" plot showing the variation across the different measures in each year. Fortunately, the variation tends to be quite small and the central pattern is close to that of Figure 2.

Figure A2



Box Plot of Stratification Measures

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Endnotes

[1] The Gini coefficient is the average squared deviation of the income shares of different percentile groups from proportionality. Other measures of inequality such as the variance of log income, the proportion of the income going to the top percentiles, and the ratio of the income of the top quintile to the bottom quintile show essentially the same pattern.

[2] This prior trend was so pronounced that it gave Kuznets (1956) the confidence to argue that increasing equality was a central feature of developed capitalist economies.

[3] The literature on the reasons for increased inequality is voluminous, but see Atkinson (1997) for a good review.

[4] The Gini coefficients are taken from U.S. Census (2002). The DW-NOMINATE scores are based on a scaling of Congresses 1-106. They can be downloaded at http://voteview.uh.edu/dwnomin.htm. See McCarty, Poole, and Rosenthal (1997) for an exposition of the derivation of these scores.

[5] Computed from Green, Palmquist, and Schickler (2002), Table 2.3, p. 31. The percentage differences for presidential and midterm election years running from 1972 to 1996 are 25, 30, 32, 36, 34, 35, 36, 34, 36, 29, 38, 48, 50.

[6] An important exception (though by a non-academic) is Phillips (1990). This lack of interest is not true, however, of recent work in comparative political economy that has sought to link inequality to political conflict and back to economic policy. See Acemoglu and Robinson (forthcoming), Alesina and Perotti (1995), Alesina and Rodrick (1993), Benabou (2000), Londregan and Poole (1990), Perotti (1996), and Persson and Tabelini (1994).

[7] For example, a major recent work on partisan identification, Green, Schickler and Palmquist (2002) makes little or no use of the demographic variables employed in this paper with the exception of a chapter on partisan realignment in the South. They focus on the stability of individual partisan self-identification. Our focus is on important changes in how demographics relate to partisan identification. Our main concern is income but we also find, in addition to the South, an important shift with regard to gender.

[8] Party identification is measured on a seven-point scale in which the categories are "Strong Democrat, Weak Democrat, Lean Democrat, Independent, Lean Republican, Weak Republican, Strong Republican". This is constructed from several questions. Respondents are first asked to choose between Democrat, Independent, and Republicans. "Democrats" are then asked if they are Strong or Weak. Ditto for Republicans. "Independents" are asked if they "lean" to one of the parties. In our analysis of stratification in Figure 1, we combine the strong and weak Republican categories. In our ordered probit analysis we use all seven categories. We divide the respondents into income quintiles using the Census Bureau's series on the distribution of household income. The details of the computation of our stratification measure are relegated to the appendix.

[9] Throughout we defined the South as the eleven Confederate states plus Kentucky and Oklahoma.

[10] See Romer (1975), Roberts (1977), Meltzer and Richard (1978), Perotti (1996), and Roemer (1999).

[11] As an example, Meltzer and Richard (1978) argue that the optimal linear income tax rate for voter i is $t(r_i) = \frac{r_i(1+\eta_1)+1}{r_i(1+\eta_1)+(1+\eta_2)}$ where the η's are tax elasticities that are assumed to be less than 0. Since the elasticities are negative, it is easy to show that t is decreasing in r_i.

[12] While this quadratic functional form is difficult to derive from economic fundamentals, it should be a reasonable approximation.

[13] We also estimated the model both with income effects "dummied" for each year and with each year estimated separately. The separate estimations allow all the coefficients and thresholds to vary over time. The results were substantively identical. As would be expected from inspection of figure 2, 1982 and 1998 are least consistent with the general pattern of the results.

[14] For a study that links changes in the income distribution across genders to increased divorce rates and changes in the partisanship of women see Edlund and Pande (2000). Since the NES income variable for female respondents records family income for a respondent from a family and individual income for a respondent from a single-person household, the fall in female income undoubtedly reflects the increased number of females now living in single households.

[15] We thank Christine Eibner for sharing this data with us.