Political Polarization and Income Inequality

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Abstract

Since the early 1970s, American society has undergone two important parallel transformations, one political and one economic. Following a period with mild partisan divisions, post-1970s politics is increasingly characterized by an ideologically polarized party system. Similarly, the 1970s mark an end to several decades of increasing economic equality and the beginning of a trend towards greater inequality of wealth and income. 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. We explore the relationship between voter partisan self-identification and income from 1956 to 1996. We find that over this period of time 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 much less a reflection of increased economic inequality. The partisanship results replicate for presidential vote choice.

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, Brown, Princeton, and Salerno. We thank participants for their comments. Correspondence to: Nolan McCarty, Woodrow Wilson School, Princeton University, Princeton, NJ. 08544, nmccarty@Princeton.edu

1. Introduction

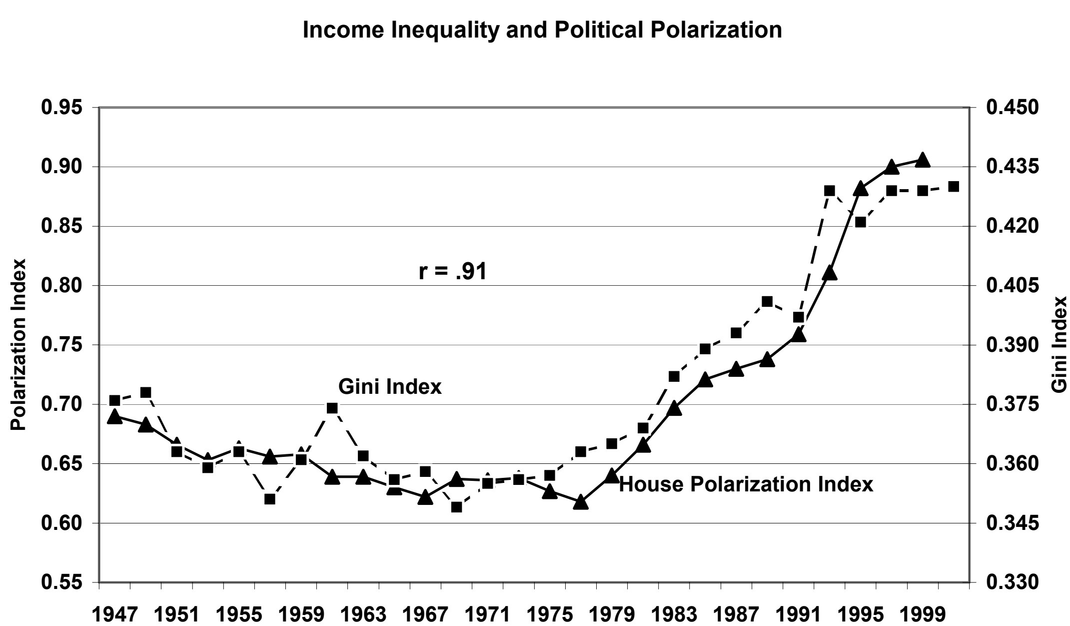
The decade of the 1970s marked many fundamental changes in the structure of American society. In particular, America witnessed almost parallel transformations of both its economic structure and the nature of its political conflict.

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, 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 has given 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 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



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.[5] 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. 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.

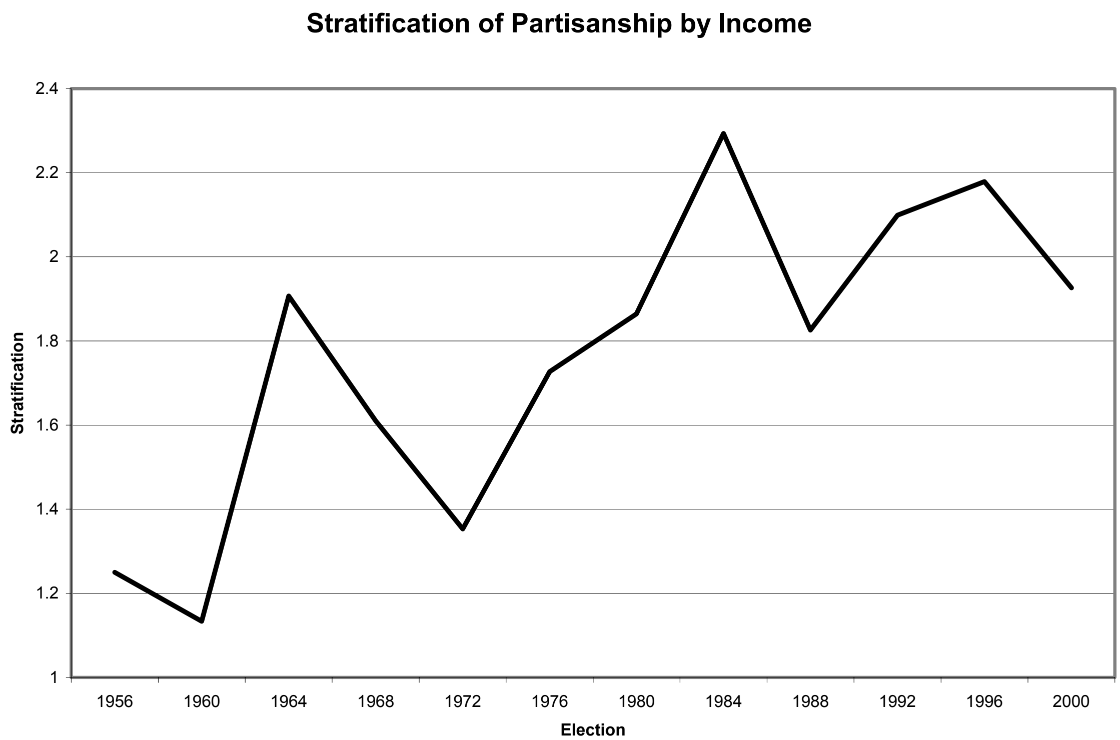
Partisanship (as measured by the National Election Study) was, in fact, only weakly related to income in the period following World War II. In 1956, a respondent from the highest income quintile was only 25% more likely to identify as a Republican than was a respondent from the lowest economic quintile. In 1960, that number was only 13%. In contrast, throughout the 1990s, a respondent was more than twice (100%) as likely to identify as a Republican if she is in the highest quintile than if she is in the lowest.

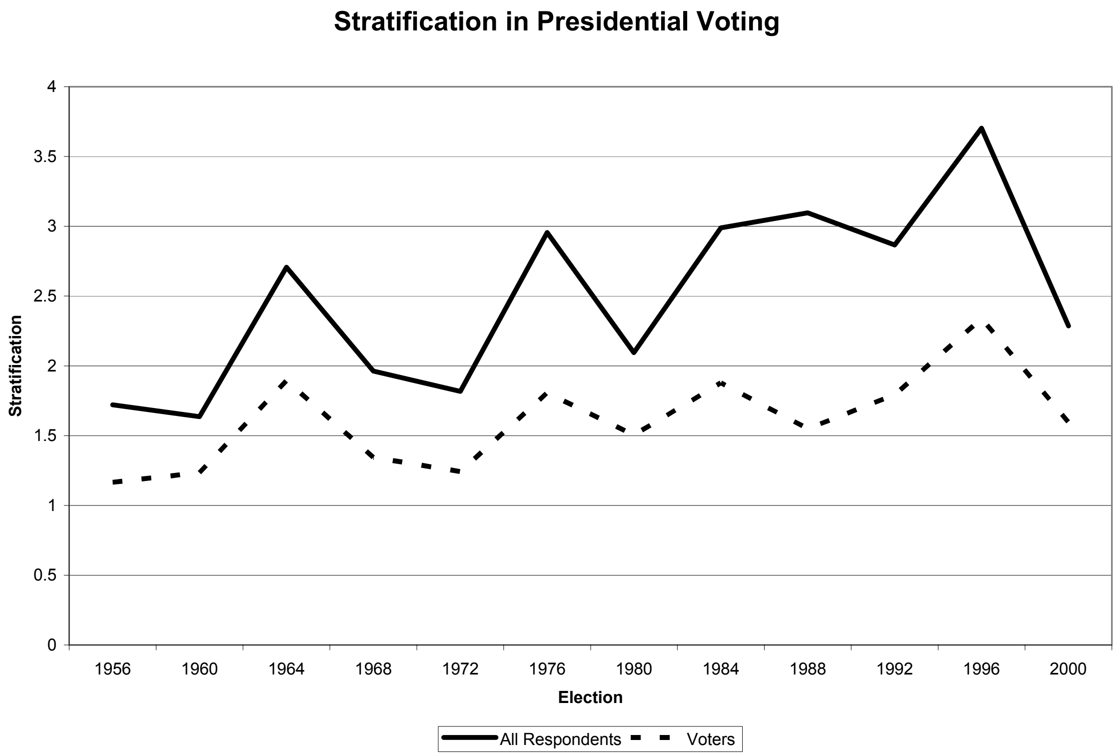
An alternative way to summarize how partisanship has acquired an income basis is through an index of party-income stratification. The index we use is simply the proportion of Republican identifiers (strong and weak) in the top income quintile divided by the proportion of Republican identifiers in the bottom quintile.[6] As seen in figure 2, the stratification of partisanship by income has steadily increased over the past 40 years, leading to an increasing class 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, although somewhat less smooth than for partisan identification, is also evident for the presidential vote. Note, however, that there is a rather sharp downturn in presidential stratification for the 2000 election.[7] We return to this observation later.

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





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 1956 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, the effect of income on partisanship may have increased. Below we argue that this is consistent with party polarization on economic policy issues.

Second, increased inequality might have made low income groups poorer and high-income groups richer so that with even a small income effect stratification would increase. Third, increased stratification might be a result of a change in the joint distribution of other characteristics and income. For example, pro-Democratic groups may have gotten poorer while pro-Republican groups got richer. Finally, groups with high incomes may have moved toward the Republicans while poorer groups moved toward the Democrats.

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. 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.[8] Assuming that tax schedules are either proportional or progressive, individuals with higher incomes prefer lower tax rates since they pay large sums of money for a given tax rate. Alternatively, when aggregate income is larger, higher tax rates produce more money for redistribution and public goods. Thus, *ceterus 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 $t(y_i/\overline{y}) \equiv t(r_i)$. The ideal rate is decreasing in relative income, so $t' < 0$.[9].

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 $u(t_R \mid r_i) > u(t_D \mid r_i)$. 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 income of a voter who would have ideal tax rate t_R. To facilitate estimation, we assume quadratic utility: $u(t_R \mid r_i) = -(r_i - r_R)^2$.[10] 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

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

$$= \alpha + \beta\left(r_D^2 - r_R^2\right) + 2\beta\left(r_R - r_D\right) r_i + \theta \boldsymbol{x}_i + \varepsilon_i$$

$$= \tilde{\alpha} + \tilde{\beta} r_i + \theta \boldsymbol{x}_i + \varepsilon_i \qquad [1]$$

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: 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: 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 data National Election Studies from 1956 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 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 numerous other studies have found to be related to partisanship. These include whether the respondent is African-American, female, or a southerner. 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 12 National Election Study presidential year survey from 1956 through 2000.

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 $\tilde{\beta}$ in various ways. We report four sets of results corresponding to a constant income effect, a distinct income effect in each election, an effect with a linear trend, and an income effect which is a step function of different eras. We also allow the effects of other variables to change over time. We report only estimates where these coefficients move 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.[11]

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 0.137 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. Column (2) shows the model that allows a separate income coefficient for each election. We find significant variance over time in the effect of income. The coefficient ranges from a low of 0.065 in 1960 to a high of 0.227 in 1996, more than a three-fold increase. While the coefficient for 1996 is still not huge, the increased importance of income since the 1950s is substantial. The constant coefficient model is easily rejected by model (2). As we discussed above, this is strong evidence for party polarization on economic issues.

While the income coefficients bounce around a bit, there is a definite trend over the entire period. Model (3) simplifies matters by assuming that the income effect changes only linearly. This model produces a statistically significant growth rate in the income coefficient of. 0026 per year. Thus, over the 44 years spanned by our data, we estimate the income effect has increased by 0.1144, representing a more than doubling of the base line value of 0.080. Model (4) is a specification in which a separate income effect for the periods of 1956-1960, 1964-1972, 1976-1984, 1988-1992, and 1996-2000 is allowed. Each subsequent period has a higher estimated income effect. While the estimated income effects are not particularly large, they have grown substantially over time. The effect for the final period is nearly triple that of the initial period.

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.

Note, however, that the constant in the formal model, $\tilde{\alpha} = \alpha + \beta\left(r_D^2 - r_R^2\right)$, should be falling in time, as captured by the (unreported) year fixed-effect variables. Note that the second term can be expanded to $\beta(r_D - r_R)(r_D + r_R)$. 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. Therefore, the second term 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. This suggests, given our strong priors about the second term, that the behavioral parameter α is growing sharply in time. 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. For example, in the column (3) model of table 1, the year fixed-effects are near zero in 1956, 1960, and 1968 but move to over 0.5 ($p<0.000$) in 1996 and 2000. Our analysis suggests an even faster growth in the Republican "bias" term α, given that the estimated year-fixed effect should also embed a growing, negative term from polarization.

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 any casual observer. African-Americans and females have moved away from the Republican Party, just as 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 is in part reflected by the fact that average levels of education have increased.

The effect of income is very important compared to that of the demographics. Consider the estimates from column (4) of table 1. In 2000, moving an individual from quite poor, at half average income, to upper middle class, at twice average income, has a net effect of 0.796, much greater than shifting the individual from female to male (0.145) from north to south (0.151), or giving a college degree to someone who had never attended college (0.052). The income effect is nearly as large as shifting from black to non-black (.920).

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

	(1) Constant Income Effect	(2) Unrestricted Effect	(3) Trended Income Effect	(4) Step Income Effect
Relative Income	0.137 (0.011)		0.080 (0.021)	
Relative Income x (Year-1956)			0.0026 (0.0008)	
Relative Income x 1956		0.078 (0.036)		
Relative Income x 1960		0.065 (0.040)		
Relative Income x 1964		0.144 (0.034)		
Relative Income x 1968		0.113 (0.036)		
Relative Income x 1972		0.082 (0.027)		
Relative Income x 1976		0.166 (0.031)		
Relative Income x 1980		0.151 (0.037)		
Relative Income x 1984		0.132 (0.032)		
Relative Income x 1988		0.174 (0.036)		
Relative Income x 1992		0.169 (0.030)		
Relative Income x 1996		0.227 (0.036)		
Relative Income x 2000		0.154 (0.044)		
Relative Income x (1956+1960)				0.070 (0.027)
Relative Income x (1964+1968+1972)				0.107 (0.019)
Relative Income x (1976+1980+1984)				0.150 (0.019)
Relative Income x (1988+1992)				0.172 (0.023)
Relative Income x (1996+2000)				0.199 (0.029)

	(1) Constant Income Effect	(2) Unrestricted Income Effect	(3) Trended Income Effect	(4) Step Income Effect
African-American	-0.535 (0.049)	-0.551 (0.050)	-0.555 (0.050)	-0.557 (0.050)
African-American x (Year-1956)	-0.009 (0.002)	-0.008 (0.002)	-0.008 (0.002)	-0.008 (0.002)
Female	0.143 (0.028)	0.137 (0.029)	0.137 (0.029)	0.137 (0.029)
Female x (Year-1956)	-0.007 (0.001)	-0.006 (0.001)	-0.006 (0.001)	-0.006 (0.001)
South	-0.458 (0.033)	-0.466 (0.033)	-0.467 (0.033)	-0.468 (0.033)
South x (Year-1956)	0.014 (0.001)	0.014 (0.001)	0.014 (0.001)	0.014 (0.001)
Some College	0.277 (0.043)	0.294 (0.044	0.294 (0.044)	0.297 (0.044)
Some College x (Year-1956)	-0.004 (0.002)	-0.005 (0.002)	-0.005 (0.002)	-0.005 (0.002)
College Degree	0.352 (0.047)	0.397 (0.049)	0.396 (0.049)	0.402 (0.048)
College Degree x (Year-1956)	-0.006 (0.002)	-0.008 (0.002)	-0.008 (0.002)	-0.008 (0.002)
Age	0.066 (0.009)	0.062 (0.009)	0.062 (0.009)	0.062 (0.009)
Age x (Year-1956)	-0.003 (0.000)	-0.003 (0.000)	-0.003 (0.000)	-0.003 (0.000)
μ_1	-0.520 (0.056)	-0.603 (0.070)	-0.600 (0.062)	-0.612 (0.065)
μ_2	0.181 (0.056)	0.099 (0.070)	0.101 (0.062)	0.089 (0.065)
μ_3	0.481 (0.056)	0.398 (0.070)	0.401 (0.062)	0.388 (0.065)
μ_4	0.781 (0.056)	0.699 (0.070)	0.701 (0.062)	0.689 (0.065)
μ_5	1.100 (0.056)	1.018 (0.070)	1.102 (0.062)	1.008 (0.065)
μ_6	1.664 (0.057)	1.582 (0.070)	1.584 (0.062)	1.572 (0.065)
Log-likelihood	-37950.2	-37941.0.	-37945.2	-37943.2
Likelihood Ratio p-value (H0 = Constant Effect)		0.073	0.002	0.007
Number of Observations	20550	20550	20550	20550

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 1956 and 1996 under different scenarios using the results of the "stepwise" specification in column 4 of Table 1.

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 2 below gives the profiles of the lowest and highest income quintiles for the 1956 and 1996 surveys.

Table 2: Characteristics of Income Quintiles, 1956 and 1996

Variable	Top Quintile 1996	Bottom Quintile 1996	Ratio 1996	Top Quintile 1956	Bottom Quintile 1956	Ratio 1956
Average Relative Income	2.063	0.169	12.207	2.338	0.314	7.446
% African-American	4.5	24.6	0.183	0.6	17.1	0.035
% Female	42.1	69.6	0.605	46.3	62.1	0.746
% Southern	30.5	47.6	0.641	20.2	42.5	0.475
% Some College	15.8	16.8	0.940	18.0	2.8	6.429
% College Degree	64.3	14.6	4.404	24.8	1.2	20.667
Average Age	44	51	0.863	43	54	0.796

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. Beyond the striking change in the distribution of income, we find large changes in placement of other 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 1956 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 1956.

However, changes in the income distribution of the other demographic categories clearly work to increase stratification. Females compose a notably larger share of the lowest quintile and a lower share of the top quintile. Since they have moved steadily towards the Democratic Party, the effects on party-income stratification are quite apparent.[12] 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. The increase at the top is relatively larger than the increase at the bottom, which would tend to increase stratification. However, since blacks remain substantially overrepresented at the bottom and under represented at the top, the fact that they have become more Democratic increases stratification. This second effect dominates the first.

To quantify the magnitude of some these effects, we compute stratification scores given the typical respondent profiles from 1956 and 1996 using the results of the step function model. We then manipulate the model and the profiles in order to assess which factors most contributed to the increased stratification. These results are given in Table 3. The first two rows of Table 3 reflect the estimated stratification for each year using the actual model and profiles for that year. These results are benchmarks for comparison with other counterfactuals. In row 3, we estimate the stratification that would have occurred using the coefficients from 1996 and the profiles from 1956. The result is a stratification score of 2.006, which is only slightly smaller than the actual estimated score of 2.074. Alternatively, row 4 shows the estimated stratification using the 1956 model with the 1996 profiles to capture the effects of the demographic shifts. The resulting stratification of 1.777 is only slightly larger than that for 1956 -- only about ¼ of the difference between 1956 and 1996. These two results imply that the changes in the relationship between partisanship and the demographic variables accounts for much more of the 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 examples, 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 income profile of the other year. These results show that the aggregate distribution of income has relatively little effect on stratification. In both cases, the counterfactual stratification indices are almost identical to the actual ones. This suggests that the changes in the aggregate distribution of income accounts for very little of the change.

Finally, we turn to the effects of the increased impact of relative income on partisanship. In row 7, we find that the 1996 stratification with the 1956 income coefficient is indistinguishable from the actual 1956 stratification. In row 8, we get a similar result after reversing the roles of 1956 and 1996. 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 3: Determinants of Party/Income Stratification

Scenario	**Republican Proportion of Lowest Quintile**	**Republican Proportion of Highest Quintile**	**Party/Income Stratification**
1956 Profiles and Coefficients	0.178	0.296	1.666
1996 Profiles and Coefficients	0.205	0.424	2.074
1996 Coefficients with 1956 Profiles	0.224	0.449	2.006
1996 Profiles with 1956 Coefficients	0.182	0.323	1.777
1956 Profiles and Coefficients except 1996 Income Profiles	0.175	0.289	1.653
1996 Profiles and Coefficients except 1956 Income Profiles	0.213	0.445	2.093
1996 Profiles and Coefficients except 1956 Income Coeffi.	0.199	0.331	1.665
1956 Profiles and Coefficients except 1996 Income Coeff.	0.188	0.398	2.124

Note: Based on coefficient estimates from the model with step income effects.

6. 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 two major party candidates in the presidential year.

Declared abstentions, votes for minor parties, and non-responses resulted in our having only about 2/3s 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 4, 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 4. The linear trend effect remains significant in column (3), but it is less precisely estimated than in the case of partisan identification. The drop in precision is shown both in the estimated standard error for the trend coefficient and in the likelihood-ratio test for column (3). The p-value for vote choice is 0.039 for vote choice, substantially larger than the 0.002 for identification. Similarly, the estimated relative income effects on a year-by-year basis, shown in column (2) are more varied than for partisan identification. In particular, the three coefficients for the Clinton-Gore candidacies are below those for the elections of 1976 through 1988. The 2000 election coefficient is below all other coefficients later than 1972.

To some degree these more varied results may just reflect worse data. We have fewer observations with vote choice and the dichotomy may contain much less information than the six categories of partisan identification. More importantly, perhaps, presidential vote choice may just be noisier than partisan identification. Vote choice may reflect variations in candidate competence or charisma and in short-term issue salience. In principle, our year fixed-effects control for some of this noise. This is indeed the case. The R^2 between our 12-year fixed effects and the Republican presidential vote share is 0.66 (p=.0015). Republican landslides are associated with positive fixed effects.

Since our year fixed effects appear to be effective controls, it is possible that the presidential vote choice results may indicate a fundamental shift in American politics, or at least presidential politics. Brady (2001) and many others have noted that Gore won the rich states along the two oceans and Bush won the poorer 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 2, in contrast to that of table 4, 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 4:
Effects of Relative Income on Presidential Vote Choice
Probit
(s.e. in parentheses)

	(1)	(2)	(3)	(4)
Relative Income	0.190 (0.016)		0.136 (0.031)	
Relative Income x (Year-1956)			0.0025 (0.0012)	
Relative Income x 1956		0.115 (0.050)		
Relative Income x 1960		0.124 (0.056)		
Relative Income x 1964		0.178 (0.050)		
Relative Income x 1968		0.077 (0.058)		
Relative Income x 1972		0.123 (0.046)		
Relative Income x 1976		0.302 (0.049)		
Relative Income x 1980		0.242 (0.064)		
Relative Income x 1984		0.294 (0.053)		
Relative Income x 1988		0.244 (0.055)		
Relative Income x 1992		0.181 (0.048)		
Relative Income x 1996		0.224 (0.054)		
Relative Income x 2000		0.158 (0.065)		
Relative Income x (1956+1960)				0.118 (0.038)
Relative Income x (1964+1968+1972)				0.130 (0.030)
Relative Income x (1976+1980+1984)				0.285 (0.032)
Relative Income x (1988+1992)				0.209 (0.036)
Relative Income x (1996+2000)				0.198 (0.042)
Log-likelihood	-7946.3	-7935.8	-7944.2	-7937.7
Likelihood Ratio p-value		0.033	0.039	0.003
Number of Observations	12926	12926	12926	12926

7. Conclusion

In this paper, we have attempted to lay some of the groundwork for an expanded study of the links between economic inequality and political polarization in the United States. Specifically, we attempted to explain the increasingly strong relationship between income and partisan identification. We have also shown a strong relationship between income and presidential voting choice.

Given our interest in both inequality and polarization, we were somewhat surprised to find that polarization, but not inequality, seemed to be the primary factor behind the increased party-income stratification. This leaves us with an important puzzle: why did the parties polarize given the lack of a large direct effect of increased inequality on the composition of the parties?

Appendix

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,

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

$$EI_{kt} = \begin{cases} \int_0^{x_{1t}} x dF(x \mid \hat{\theta}_t) & k = 1 \\ \int_{x_{k-1,t}}^{x_{kt}} x dF(x \mid \hat{\theta}_t) & otherwise \end{cases}$$

In this paper, we assume that $F(\cdot)$ is log-normal and that $\theta_t = \{\mu_t, \sigma_t\}$. 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 $\hat{\theta}_t$ for each presidential election year. These results underscore the extent to which the income distribution has become more unequal.

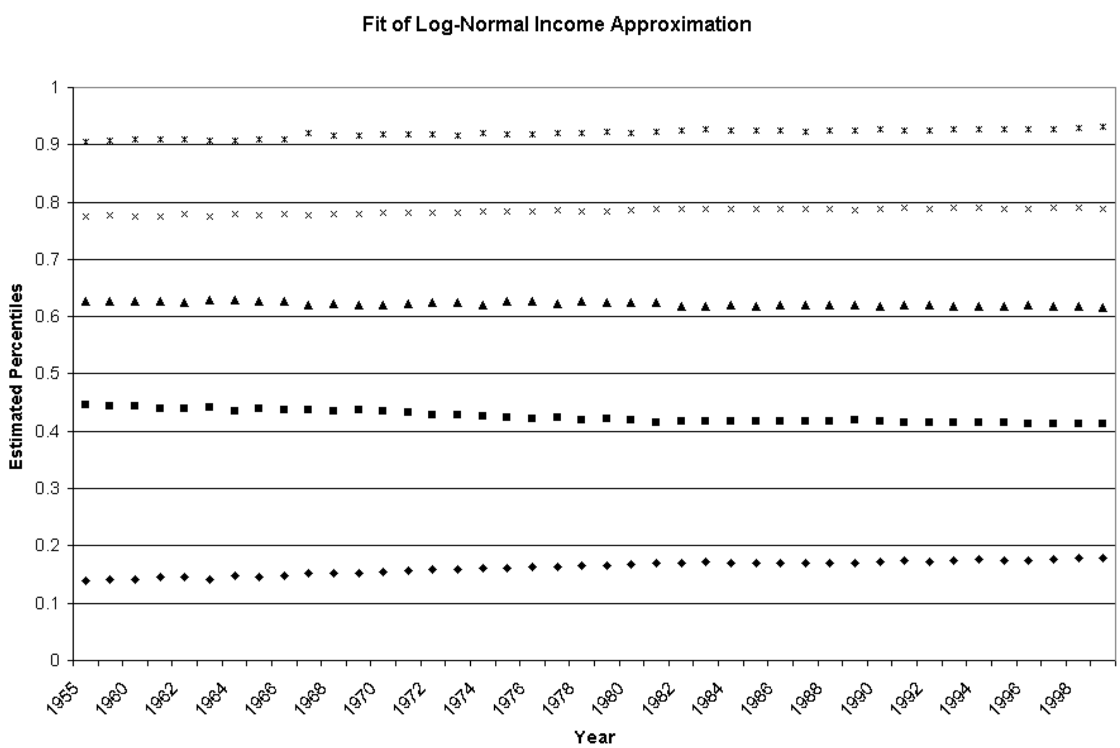
Table A1

Election	μ_t	σ_t
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

Figure A1 plots $F\left(\boldsymbol{y}_t \mid \hat{\theta}_t\right)$ against $\boldsymbol{z}_t$ and shows how well the log-normal approximates the distribution of income. While the approximation is generally very good, the log-normal is a poor approximation of incomes at lower levels since the true distribution of income has positive mass at incomes of zero. The effect is that EI_{kt} has a slight positive bias for low values of k.[13]

Figure A1



References

Acemoglu, Daron and James A. Robinson, forthcoming, “A Theory of Political Transitions”, *American Economic Review*.

Alesina, Alberto and Roberto Perotti. 1995. “Income Distribution, Political Instability, and Investment,” *European Economic Review* 40:1203-1228.

Alesina, Alberto and Dani Rodrick. 1993. “Income Distribution and Economic Growth: A Simple Theory and Some Empirical Evidence,” In Alex Cukierman, Zvi Herscovitz, and Leonardo Leiderman, eds. *The Political Economy of Business Cycles and Growth*. Cambridge: MIT Press.

Atkinson, A.B. 1997. “Bringing Income Distribution in from the Cold,” *The Economic Journal* 107:297-321.

Benabou, Roland. 2000. “Unequal Societies: Income Distribution and the Social Contract,” *American Economic Review* 90: 96-129.

Brady, Henry E. 2001. “Trust the People: Party Coalitions and the 2000 Elections”, Working paper, University of California, Berkeley.

Edlund, Lena and Rohini Pande. 2002. “Why Have Women Become Left-Wing? The Political Gender Gap and the Decline of Marriage,” *Quarterly Journal of Economics* 117: 917-962..

Kuznets, Simon. 1956. “Economic Growth and Income Inequality,” *American Economic Review*, 45:1-28.

Londregan, John and Keith Poole. 1990. “Poverty, the Coup Trap, and the Seizure of Executive Power,” *World Politics* 62:151-183.

McCarty, Nolan, Keith T. Poole, and Howard Rosenthal. 1997. *Income Redistribution and the Realignment of American Politics*. Washington D.C: AEI Press.

McKelvey, Richard R. and William Zavoina. 1975. “A statistical model for the analysis of ordinal level dependent variables,” *Journal of Mathematical Sociology*, 4:103-120.

Meltzer, Allan H. and Scott F. Richard. 1981. “A Rational Theory of the Size of Government,” *Journal of Political Economy* 89:914-27.

Perotti, Roberto. 1996. “Political Equilibrium, Income Distribution, and Growth,” *Review of Economic Studies.*

Persson, Torsten and Guido Tabellini. 1994. “Is Inequality Harmful for Growth? Theory and Evidence,” *American Economic Review*.

Phillips, Kevin. 1990. *The Politics of Rich and Poor: Wealth and the American Electorate in the Reagan Aftermath.* New York: Harper Collins.

Poole, Keith T. and Howard Rosenthal. 1984. "The Polarization of American Politics," *Journal of Politics* 46:1061-79.

Poole, Keith T. and Howard Rosenthal. 1997. *Congress: A Political-Economic History of Roll Call Voting.* New York: Oxford University Press.

Roberts, Kevin W.S. 1977. "Voting over Income Tax Schedules," *Journal of Public Economics*. 8:329-340.

Roemer, John. 1999. "The Democratic Political Economy of Progressive Income Taxation, *Econometrica*. 67:1-19.

Romer, Thomas. 1975. "Individual Welfare, Majority Voting, and the Properties of a Linear Income Tax," *Journal of Public Economics*. 14:163-185.

U.S. Census Bureau; Current Population Survey. Income Limits for Each Fifth and Top 5 Percent of Households (All Races): 1967 to 1999.

U.S. Census Bureau (1975). *Historical Statistics of the United States from Colonial Times to 1970*.

U. S. Census Bureau (2002); "Table F-4. Gini Ratios for Families, by Race and Hispanic Origin of Householder: 1947 to 2001 (Families as of March of the following year)"; Published 30 September; <http://148.129.75.3/hhes/income/histinc/f04.html>

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] An important exception (though by a non-political scientist) 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).

[6] 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 all three Republican categories. In our ordered probit analysis we use all seven categories.

7 We believe that this downturn is largely a consequence of the NES severely under sampling lower income citizens in 2000. We will update in a revision to the manuscript.

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

9 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.

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

11 We also estimated the model on each year separately which allows all the coefficients and thresholds to vary over time. The results were substantively identical.

12 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).

13 When we have more income distribution data, we should be able to estimate a truncated lognormal to account for this effect.