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Inequality and Poverty in the United States: The Effects of Rising Male Wage Dispersion and Changing Family Behavior

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Abstract

Using semiparametric density estimation techniques, we analyze the contribution of rising dispersion of men's earnings and related changes in family behavior—rising labor force participation by women and changes in family composition—to increasing inequality in the distribution of family income in the United States during the years 1969-1998. For the period 1969 to 1989, rising dispersion of men's earnings and increases in nontraditional family structures (notably single parenthood) respectively explain up to about three-fourths and one-half of rising inequality in family income and poverty. In contrast to results from most previous research, we find that the increase in women's labor force participation offset rising inequality in family income, with its impact moving down the income distribution over time. Moreover, inequality grew at a slower rate in the 1990s than in earlier decades, largely due to slower deterioration in the relative earnings of men from low-income families.

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Introduction

Since the late 1960s, inequality in family income has risen substantially in the United States. Although some of the rising dispersion has been associated with gains in the upper portion of the income distribution (Burkhauser et al. 1999), dispersion in the bottom portion of the distribution has grown substantially. This increase in dispersion in the lower tail increased poverty rates during the 1980s, especially among families with children (Blank and Card 1993), suggesting important implications of rising inequality for social welfare and policy.

Rising dispersion in individual earnings has been offered as a leading explanation for rising inequality in family income (see for example Karoly and Burtless 1995). As the same time, however, the U.S. has witnessed extensive changes in family structure and behavior that also may have contributed to rising inequality in family income. These changes include pronounced increases in women's labor force attachment and in the share of households consisting of a single adult with children.

In this paper, we assess the effects of the changing distribution of men's earnings, rising labor force participation by women, and changes in family structure and living arrangements on family income inequality and poverty. We focus on these three factors, rather than a wider list of possible causes, due to their primacy in past research and their potential behavioral links. Our data span the years 1967-1998, and we largely abstract from business cycle effects on inequality by focusing on the business cycle peak or near-peak years of 1969, 1979, 1989, and 1998. In addition to a reexamination of the causes of rising income inequality during the 1970s and 1980s, an important contribution of this paper is to update the analyses and examine whether the deterioration in the relative status of low-income families that occurred in these decades extended into the 1990s as well.

Our methodology relies on the imposition of specific counterfactuals on the observed distribution of income. In particular, we assess the impact of changes in the distribution of men's earnings using a rank-based distributional exchange (similar to Burtless 1999), and we assess the impact of changing family behavior and underlying characteristics using the conditional density estimation method of DiNardo, Fortin, and Lemieux (1996; henceforth "DFL"). This combined approach enables us to examine the impact of the modelled factors in a relatively unified framework that accounts for inter-related impacts through conditional estimation and sequential decomposition. The technique also enables us to estimate the effects of the modelled factors on the complete distribution of family income, hence its specific features, such as dispersion in the lower half (including the poverty rate). Compared with past research in this area, our findings provide a more extensive and precise depiction of the impact of the modelled factors on the distribution of family income, and they serve to clarify conflicting results from past literature, notably in regard to the effect of rising labor force participation by women on family income inequality.

The paper proceeds by first describing our data and trends in income inequality. We describe the decomposition methodology used in Section II and present results in Section III. In Section IV, we summarize the results and discuss implications for future research.

I. Data and Trends

We use data from the March Annual Demographic Supplements to the Current Population Survey (CPS) for the years 1968-1999 (income years 1967-1998). The CPS is the monthly survey of about 60,000 U.S. households that is used for tabulations of official labor force series and related statistics. The March supplement collects detailed information on income and work experience in the previous calendar year. In the formal decomposition analysis, we focus on the years 1969, 1979, 1989, and 1998, which largely span our sample frame. These years represent either business cycle peaks or ongoing expansions, so that our analysis of changes over time will be relatively unaffected by underlying business-cycle determinants of inequality. For consistency with past work in this area, we restricted the sample to families in which the male or female head is between the ages of 16 and 64.[1]

Our measure of economic well-being starts with total pre-tax, post-transfer income for the family unit.[2] We apply the concept of equivalent personal income, which adjusts available income for family size and economies of scale in consumption. Letting Y_f denote our measure of total

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[1]We excluded families with elderly heads because the determinants and dynamics of family structure, labor supply, and family income are different for elderly individuals and couples than they are for families with working-age adults, as are the policy options for addressing income dispersion and poverty. Although the exact age range varies, this is a common restriction in papers that investigate the impact of earnings dispersion and family behavior on the distribution of family income (see for example Cancian and Reed 1998, Danziger and Gottschalk 1993); our use of this restriction enhances the comparability of our work with past work. Elderly individuals living with families headed by their working-age relatives are included in our sample.

[2]The CPS does not provide information on post-tax income. Although exclusion of tax collections may distort measurement of changes in the distribution of disposable income over time, Gottschalk and Smeeding (1997, page 670) argue that in the U.S., “changes in taxes and transfers account for only a small part of the trend in inequality during the 1980s and 1990s.” In Section III, we discuss indirect evidence on the impact of government transfers during the 1990s.

family income and *F* denote family size, our measure of family equivalent income (*Y*) for each individual in a family is:

$$Y = \frac{Y_f}{F^{\sigma}} \tag{1}$$

where $0 \leq \sigma \leq 1$ (see Buhmann et al. 1988). For the results reported in this paper, we set σ=0.5. This value lies at the midpoint of the range of assumptions regarding economies of scale in family consumption, and it has the virtue of being nearly identical to the implied equivalence scale used in the Census Bureau's official poverty thresholds (Ruggles 1990).[3] We apply this calculation for all families identified in our data except for multi-generational families that share living quarters, for whom income and family size are totalled across the sub-families.[4] We use the CPS sampling weights, which are designed to ensure that the sample is representative of the associated population. All income values are expressed in 1998 dollars, using the GDP deflator for personal consumption expenditures.

A key feature of the U.S. distribution of family income is the percentage of families or individuals who fall below the income cutoff that represents the official poverty line. We identify all individuals in a family as being in poverty if family equivalent income (defined above) falls below the official federal poverty threshold for a single-person family. Although this

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[3] The boundary values of σ—0 and 1—imply, respectively, infinite and no economies of scale in family consumption. To test the sensitivity of our results to our specific equivalence scale, we also estimated our primary model with σ set equal to 0.25 and 0.75 (see Section III).

[4] This accounts for income sharing across related families in such households, moving our measure of family income closer to a measure of total household income. Focusing entirely on household income might show less increase in inequality over our sample frame, but existing results suggest the differences are likely to be small (Burkhauser et al. 1999).

simplification causes our tabulations of poverty rates to differ slightly from the official U.S. poverty rate, our approach has the benefit of making our inequality and poverty measures comparable through the use of identical equivalence scales.

The two panels of Figure 1 display statistics for family equivalent income for the years 1967-1998, with each series normalized so that the 1967 value equals 100. For the income variables, we focus on standard dispersion measures—the coefficient of variation, the Gini and Theil coefficients, the mean logarithmic deviation, and several percentile dispersion ratios—and display the median as well.[5] All of the dispersion measures increased substantially between 1969 and 1998.[6] The trend toward increasing inequality was most evident in the bottom half of the income distribution, as indicated by the especially pronounced rise in the mean logarithmic deviation (which is relatively sensitive to dispersion toward the bottom of the distribution) and the 50-10 income ratio. Consistent with increased dispersion in the lower tail, an upward trend in poverty also was present during this period, although the cyclical variability is high. Comparing the business cycle peak or near-peak years of 1969, 1979, 1989, and 1998 indicates that the largest increases in inequality occurred during the 1980s. Pronounced increases were associated with recessionary periods; inequality subsequently increased further during the 1980s expansion

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[5] The coefficient of variation is the standard deviation divided by the mean. The Gini and Theil measures and the mean logarithmic deviation are standard parametric indices of inequality used in previous work (see Cowell 2000 for discussion of alternative inequality measures). Our percentile dispersion measures are expressed as the ratio of incomes at the indicated percentiles of the distribution of family equivalent income.

[6] The maximum recordable income levels ("topcodes") changed somewhat during our sample frame; see Appendix A for a description of our approach to mitigating their impact.

but fell a bit during the 1990s expansion (see Appendix A for the exact yearly figures for the inequality series).

Figure 2 provides a visual representation of the distribution of real family equivalent income in 1969, 1989, and 1998.[7] The figure displays kernel density estimates of the distribution of family income for these years.[8] The distribution widened substantially between 1969 and 1989, with pronounced hollowing of the middle and density shifts to both tails. Much of the density mass moved to the right, reflecting improvements in real equivalent income for many families (Burkhauser et al. 1999). However, due to the median shift, the figure somewhat masks the decline in the relative well-being of low-income families that is indicated by the changes in the percentile dispersion measures in Figure 1. In contrast to the changes between 1969 and 1989, the increase in dispersion between 1989 and 1998 was relatively modest, with discernible movement of the density mass from the lower and middle portions to the upper portion.[9]

[7] To enhance visual clarity, we do not display the distribution of income in 1979. Figure 1 shows that most of the increase in the dispersion of income between 1969 and 1989 occurred during the 1980s.

[8] Kernel density estimates serve essentially as smoothed histogram representations of the underlying distribution from which an empirical distribution is sampled. The key choices in regard to kernel density estimation are the kernel function and the bandwidth. We used the Epanechnikov kernel function and performed subjective tests to arrive at a bandwidth that provides the clearest visual representation of our data (with the number of evaluation points set to 500). Silverman (1986) discusses nonparametric density estimation in detail, and Delgado and Robinson (1992) provide a useful summary of econometric applications. To provide an adequate visual representation of the key portions of the distribution, we truncated the kernel density plots at 0 and 60,000 dollars in all figures that display family equivalent income. However, the density estimates themselves are based on the full range of data.

[9] Kolmogorov-Smirnov tests of differences between the underlying empirical distributions for these three years strongly reject the null hypothesis of no difference in each case, with p-values well below .01 (see Burkhauser et al. 1999 for discussion of this test).

To explain these changes in family income inequality, we will focus on changes in the distribution of men's earnings, women's labor force participation, family structure, and underlying family characteristics (such as age, location, etc.). We restrict ourselves to these factors because they reflect related trends at the family level in recent decades, with the possibility of causal linkages between them.[10]

For the distribution of men's earnings, we use data on yearly wage and salary earnings, for which dispersion increased substantially over our sample frame (see for example Card and DiNardo 2002). Our measure includes individuals whose yearly earnings are zero. In Figure 3, we display the median, the coefficient of variation, and the Gini and Theil coefficients (we do not display percentile dispersion measures in Figure 3, because they are not uniformly defined for our measure of yearly male earnings—for example, the 10th percentile generally is zero).[11] As of 1998, the median had not yet returned to the peak achieved in 1973, although growth in the 1990s was been rapid. The three measures of dispersion all exhibit nearly monotonic increases between 1967 and 1993, with the net increase ranging from 29 percent for the Gini coefficient to 55 percent for Theil's coefficient. However, each declined a bit between 1993 and 1998, indicating growth in yearly male earnings that was more evenly distributed than it had been in past decades (see Appendix A for a listing of the exact yearly figures).

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[10] Exploring the causal links is beyond the scope of this paper. In recent related work, Dickens and Ellwood (2004 forthcoming) found that demography, wage changes, and social policy were key determinants of changing poverty rates in the U.S. and Great Britain between 1979 and 1999. Work on British inequality and poverty has largely focused on the impact of government tax and transfer policies and changes in work status; see for example Jenkins (1995, 1996) in regard to inequality and Piachaud and Sutherland (2001, 2003) in regard to poverty.

[11] See Appendix A for a discussion of how we handled changing topcodes for the male earnings variable over our sample frame.

Concurrent with the increased dispersion in male earnings were substantial changes in female labor force participation and family structure. Application of our conditional density estimation technique to these factors requires that they be defined in categorical terms. To measure trends in female participation, we defined a variable that is coded as 1 if the family includes a woman who worked positive hours in the previous year and 0 otherwise. Although this measure ignores hours variation other than the participation decision (such as changes in part-time work), our tabulations indicated that nearly all of the trend increase in women's labor force participation and hours worked during our sample frame was associated with a shift from non-participation to full-time work. As measured by this variable, women's participation rose steadily during most of our sample frame, with the most rapid increases evident in the late 1970s and only a small increase evident in the 1990s (see Figure 4, top panel, and the exact figures listed in Appendix A).

In regard to family structure, the trend in the U.S. during the past several decades has been away from traditional husband-wife families with children, with corresponding increases in families consisting of single (mostly female) heads with children and single individuals without children. To capture these trends, we characterize families as falling into five categories: married with children, married without children, never married individuals with children, other single individuals with children, and single individuals without children.[12] As displayed in Figure 4

[12] In most cases households contain one family. For households that contain multiple families related by blood or marriage, we treat sub-families as separate families for the identification of family structure. This accounts for important changes in living arrangements such as the rise in multi-generational families living in the same household — for example, single mothers living with their own parents. As noted by London (1998), prior to 1984 the CPS surveys did not properly account for the household relationships of children living in multi-generational households, producing an undercount of the number of single mothers due to misidentification of those who

(bottom panel), the share of married couples with children declined substantially, along with pronounced percentage increases in the share of never married individuals with children and singles without children.[13] For example, the share of families consisting of a never-married single head with children increased by a factor of nearly six between 1969 and 1998, rising from 0.010 to 0.059 (see the exact time series tabulations in Appendix A). Although this still represents a small share of all families, the increase of about five percentage points is large relative to the share of families likely to have low income or be in poverty; for example, the poverty rate during our sample frame is mostly in the range of 0.10 to 0.15.

II. Decomposition Methodology and Counterfactuals

Family Behavior and Characteristics

Our formal analysis of the changing distribution of family equivalent income involves the imposition of specific counterfactuals on the observed distribution of family equivalent income. We treat the changing dispersion of men's earning as exogenous and model this factor in an unconditional framework, through a rank-based distributional exchange. The other factors — women's labor force participation, family structure, and underlying family characteristics — are treated as interrelated, through application of the conditional density approach of DiNardo, Fortin, and Lemieux (1996; henceforth DFL). Most previous work that investigates the

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live with their parents. To reduce the impact of undercounting this key family type in the early portion of our sample, we applied London's correction to the pre-1984 data.

[13]The incidence of cohabitation without legal marriage increased during our sample frame (Casper et al. 1999), which may to some degree offset the impact of declining marriage rates on equivalent income. Direct measures of cohabitation only became available in the CPS beginning in 1996; see Section III for an assessment of the sensitivity of our results to this issue.

determinants of rising dispersion of family income is based on decompositions of family income into its components.[14] These approaches in general are unconditional, in that they do not account for changes in related variables that each affect the distribution of family income. By contrast, our approach to female labor supply, family structure, and related characteristics accounts for the inter-relationships between these factors, thereby providing a more clear understanding of their independent effects. We describe these procedures heuristically in the text; for the formal derivations, see Appendix B.

We begin by defining an individual observation as a vector (Y, Z, t) , consisting of family equivalent income Y, a vector of characteristics Z, and a year t (this discussion closely parallels Section 3 of DFL). This distribution may also depend on other factors, such as the structure of men's earnings, denoted m_t. Using notation similar to DFL, we can express the density of family equivalent income in a particular year, $f_t(Y)$, as:

$$f_t(Y) \equiv f(Y;\ t_Y = t,\ t_Z = t,\ m_t) \qquad \textbf{(2)}$$

[14] Numerous papers examine the effects of female earnings and labor supply on family income inequality in the U.S. Among them, Shorrocks (1983), Lerman and Yitzhaki (1985), and Karoly and Burtless (1995) decomposed inequality indices by income source and concluded that wives' earnings magnify family income inequality. In contrast, Cancian, Danziger and Gottschalk (1993) found that wives' earnings decrease income inequality in married couple families, and Blackburn and Bloom (1989) found little overall impact of women's labor supply on inequality in family income. As noted by Cancian and Reed (1998), the impact of wives' earnings and labor supply on inequality depends on the specific counterfactual assumption that is made. Several papers have found a substantial effect of changing family structure on inequality in family income poverty; these include Gottschalk and Danziger (1993) and Lerman (1996). In a different approach, Bradbury (1996) applied regression analyses to data aggregated at the regional level and found that changing family structure increased inequality while increases in women's labor force participation reduced it.

Our estimation of counterfactual densities involves the combination of different years, or "datings," and equation (2) introduces notation that accounts for these datings. For example, while the expression f(Y; t_Y=89, t_Z=89, m_{89}) represents the observed density of family equivalent income in 1989, the expression f(Y; t_Y=89, t_Z=69, m_{89}) represents the density of family equivalent income that would have prevailed in 1989 had the distribution of characteristics, and their relationship to family equivalent income, remained as they were in 1969.

In our application, the vector of characteristics Z has three components: women's labor force participation *L*, family structure *S*, and a vector of other family characteristics *X*. Then following DFL, and applying the multiplicative properties of conditional distributions to equation (2), the density of family equivalent income in 1989 can be expressed as:

$$f_{89}(Y) \equiv f(Y;\ t_Y = 89,\ t_{L|S,X} = 89,\ t_{S|X} = 89,\ t_X = 89,\ m_{89}) \qquad \textbf{(3)}$$

For the labor force participation, family structure, and family characteristics components, estimation of counterfactual densities based on alternative datings proceeds through the estimation and application of conditioning weights. Consider first the application to women's labor force participation. Assuming that women's participation rose between 1969 and 1989, the simplest way to impose the 1969 distribution of women's participation on the 1989 distribution of family equivalent income is to downweight families that include a working woman by a factor that is equal to the percentage increase in the share of families in which women work (and, similarly, upweight families in which no working woman is present).

This simple adjustment, however, ignores any changes in the relationship between women's labor force participation, family structure, and family characteristics. For example, suppose the increase in women's labor force participation between 1969 and 1989 was concentrated among divorced women who were attempting to replace income lost upon marital separation. Then the rise in women's labor force participation largely is associated with changes in family structure rather than changing labor supply behavior *per se*. More generally, the likelihood of changes in the joint distribution of *L*, *S*, and *X* necessitates estimation of the 1989 distribution of family income with women's labor force participation, *and its relationship to S and X*, held to its earlier period level.

In terms of the notation in (3), we are interested in:

$$f(Y;\ t_Y = 89,\ t_{L|S,X} = 69,\ t_{S|X} = 89,\ t_X = 89,\ m_{89}) \qquad \textbf{(4)}$$

This expression represents the density that would be observed if the probability of women's labor force participation retained its 1969 level and structure (conditional on family structure *S* and family characteristics *X*), but family equivalent income otherwise was determined by the same distributional characteristics prevailing in 1989. As derived in Appendix B, this distribution can be expressed as the original unconditional distribution of family income in 1989, with individual observations reweighted by the function $\hat{\psi}_{L|S,X}(L,S,X)$. This function represents the change between 1969 and 1989 in the conditional probability that the female head works in a family defined by characteristics $\{S,X\}$. The conditional probabilities can be estimated as the fitted values obtained from standard binary dependent variable models, in which female participation is regressed on family structure and family characteristics. In the empirical analysis, we used the logit model to estimate these conditional probabilities.

The process by which we account for the impact of changes in family structure is similar to that for women's labor force participation, although we must account for the added complexity of family structure as an outcome variable. Our measure of family structure is a variable with five mutually exclusive categories: married with children, married without children, never married with children, other single individuals with children, and single individuals with no children. To form the appropriate conditioning weights, $\hat{\psi}_{S|X}(S,X)$, we estimated multinomial logit models for each year of data being compared; the outcome variable *S* is defined by the five family structure categories, and the covariates are defined by our *X* vector of family characteristics. We then use the fitted coefficients from these models to form conditioning weights similar to those described for the labor force participation step. Estimating the separate effects of changes in other family characteristics (*X*) proceeds in a similar fashion, through estimation of conditioning weights $\hat{\psi}_X(X)$ based on bivariate logit models.[15]

This procedure for assessing the impact of changes in family behavior and characteristics relies on simple assumptions that are embodied in the counterfactuals. For women's labor force participation, we are assuming that if the distribution of family structure and characteristics were as observed in 1969, then the distribution of women's labor force participation also would be as observed in 1969. A similar assumption is made for the distribution of family structure. This approach ignores the possibility that unmeasured influences may have changed the conditional relationships between these factors. Rather than attempting to incorporate such influences, we set

[15]See Appendix B for the exact derivation of these weights. We extend the original DFL technique by examining a three-way (rather than two-way) breakdown of the conditioning characteristics Z and by including a conditioning variable that takes on more than two outcome categories (see also Biewen 2001, Hyslop and Maré 2001).

a more modest goal that is consistent with the usual ceteris paribus assumptions employed in standard decomposition analyses.

Men's earnings

Whereas the estimated effects of changes in the various family characteristics are obtained through conditioning weights, the impact of changes in the distribution of men's earnings is obtained through a direct adjustment to family equivalent income. This is based on a rank-preserving exchange of the men's earnings distributions between two periods (similar to Burtless 1999). Focusing again on the years 1969 and 1989 as examples, assume that men's earnings are ranked from lowest to highest in each year, and consider a man ranked at a particular point. Our underlying conceptual goal is to replace the earnings of each man in the 1989 data with the earnings of the man ranked at the same position in the 1969 data. This is trivial in the case of identically sized samples without weights; we simply rank men's earnings for each year, do a one-to-one merge, and exchange the earnings values. In the case of samples of different sizes, and where it is important to incorporate sampling weights, the procedure is based on the calculation and matching of median earnings by percentile group (quantile) for each distribution. This procedure preserves the rank of each man in the 1989 distribution but imposes the location and dispersion characteristics of the 1969 distribution of men's earnings on the 1989 distribution.[16]

[16] This is very similar to the procedure used by Burtless (1999). However, if we applied Burtless' exact procedure, we would hold constant the sum of men's earnings in 1989, thereby accounting for the change in dispersion only. In contrast, our approach also accounts for the change in the midpoint (median and mean) of the distribution of men's earnings.

To apply this adjustment to family income, we subtracted from 1989 total family income the wage and salary earnings from the median of the 1989 quantile to which the male head in that family belongs, then we added back the wage and salary earnings from the median of the 1969 quantile that is at the same rank as the 1989 quantile to which that male head belonged.[17] After again applying our equivalence adjustment, the resulting income variable Y_m replaces our original equivalent income variable Y in the analysis. It is important note that we included men with earnings equal to zero in these distributional calculations, to account for the possibility that declining labor force participation by low-wage men (Juhn 1992) has contributed to rising inequality in family income. By contrast, Burtless (1999) restricted the calculation to positive earners only. See Appendix C for a formal description of our procedure.[18]

Our approach to the changing distribution of male earnings takes its underlying determinants as given, and as such represents an unconditional approach to assessing its effect, rather than a conditional approach (as for our other factors).[19] We take this approach because accounting for the causes of rising dispersion of men's earnings has been the focus of other work

[17] We used wage and salary earnings for this calculation and excluded self-employment earnings, because the literature on rising dispersion in earnings has focused on the former. Although much of that literature has focused on hourly earnings, we used yearly earnings for consistency with our yearly measure of family income.

[18] In contrast to our unconditional ranking procedure, Fortin and Lemieux (1998) developed a ranking procedure that accounts for covariates and applied it to the analysis of the gender wage gap.

[19] Our approach to changes in the distribution of men's earnings could be extended to women's earnings as well. However, this would supplant our analysis of the change in women's labor force participation, which is one of the primary behavioral changes that we wish to investigate. Moreover, Burtless (1999) found that rising dispersion of female earnings had a much smaller impact on the distribution of family income than did rising dispersion of men's earnings.

(see Katz and Autor 1999 for an overview), and it would detract from our focus on changes occurring at the family level. Beyond that, our approach to the contribution of men's earnings relies on simple assumptions that are directly related to the counterfactual used. In particular, we assume that the decision to take a job in the analysis period (for example, 1989) is based on the labor supply and demand conditions of the counterfactual period (for example, 1969). In other words, we assume that if a man at the 10th percentile of the earnings distribution worked in 1969, a man at that percentile of the earnings distribution also would work in 1989 (even if the observed 10th percentile in 1989 had zero earnings). No attempt is made to account for general equilibrium effects of the changing distribution of men's earnings on labor supply and demand decisions over our analysis period; we leave estimation of a more general model of this type to future work.

Sequential Decomposition

For the complete decomposition, we consider all four factors discussed above and assess their contributions to rising inequality and poverty. The factors first are considered in the following primary sequence: the distribution of men's earnings, women's labor force participation, family structure, and underlying family characteristics. We assess the contribution of men's earnings first because leading explanations of changes in the distribution of men's earnings—which include skill-biased technological change and institutional factors such as unionization and the minimum wage—imply that shifts in the distribution of men's earnings are likely to be exogenous in our setting. In regard to the additional steps of the decomposition, conditioning female participation on family structure and characteristics seems most sensible from a causal perspective.

One concern regarding our approach, however, is that their estimated effects may depend on the ordering of the decomposition, due to inter-relationships between the explanatory factors. We assess the extent of this dependence by reversing the order of the decomposition, considering in sequence the effects of changing family characteristics, changing family structure, changes in women's labor force participation, and changes in the distribution of men's earnings. The derivation of the reverse-order conditioning weights, and the necessary calculations, are substantively different than for the primary-order weights; the exact procedure is described in Appendix B.

Table 1 lists the various distributions of family equivalent income that we analyse in the primary-order and reverse-order decompositions, for the comparison between 1969 and 1989. Each distribution is defined by the income measure used and a weight. The first row in both panels lists the observed distribution for 1989, which is defined by the unadjusted family weight θ and the 1989 time dimension for all relationships. The next four rows in both panels identify the adjusted (counterfactual) distributions of equivalent family income in the 1989 data. The first step of the primary-order decomposition adjusts family equivalent income (Y) for the change in the distribution of men's earnings, which produces the adjusted equivalent income measure Y_m. We then hold the relationships between L, S, and X to their earlier period structure, using the estimated weighting functions ($\hat{\psi}_{L|S,X}$, $\hat{\psi}_{S|X}$, and $\hat{\psi}_{X}$). The conditionally weighted density estimates are obtained by using weights that are the product of the family weights (θ) and the estimated conditioning weights. The effects of the explanatory factors are obtained by comparison of the distributions in sequence, while the impact of non-modelled "residual" factors are obtained through comparison of the final adjusted 1989 distribution and the unadjusted 1969 distribution. Comparison of the distributions can be made in visual terms, using kernel density estimates, or in quantitative terms, using standard dispersion measures.

III. Results

Changes Between 1969 and 1989 (Primary Order Decomposition)

We now apply our complete conditional reweighting and decomposition procedure. We first apply our analysis to the years 1969 and 1989, which correspond roughly to the period of rising inequality that has been the focus of previous work. Our vector of family characteristic variables *X* includes some characteristics, such as location, that are shared among family members, and others, such as educational attainment, that differ across members. The variables are age, years of education, the interaction of age and education, whether black, an SMSA residence indicator, and nine census geographic division indicators. In the women's labor force participation equations used to form the conditioning weights, we include the age, education, and race variables separately for male and female heads if both were present. As discussed above, we also estimated the labor force participation equations as a function of the family structure categories, which we use separate dummy variables to identify. Finally, in estimation of the family structure weights, the multinomial logit equations include controls for the average values of age and education for male and female heads if both are present, and for the sole head values in other families.

Figure 5 displays the impact of the estimated factors on the distribution of real equivalent family income. We display the effects in sequence; each panel holds an additional modelled factor to its 1969 structure and examines the impact of this adjustment compared with the prior distribution. Panel A shows the effect of the changing distribution of men's earnings. Had the distribution of men's earnings remained constant between 1969 and 1989, the distribution of family equivalent income would have been substantially narrower in 1989. The adjusted distribution contains more mass in the middle and less in the lower tail, with only a limited difference in the upper tail. Panel B displays the sequential effect of women's labor force participation. The effect of the changing conditional distribution of women's labor force participation appears to be a relatively uniform shift to the right in the distribution of family income. In Panel C, the changing conditional distribution of family structure caused a shift in density mass from the middle of the distribution to the left tail; the impact appears similar to that of the changing distribution of men's earnings, but smaller in magnitude. Finally, in panel D, the changing distribution of family characteristics appears to have caused a relatively uniform shift to the right in the distribution of family income.

A complementary visual representation of the distribution can be obtained by calculating the differences between the densities displayed in Figure 5. Because this information is closely related to that in Figure 5, for brevity we examine only the figure corresponding to the net impact of all four explanatory factors. Figure 6 displays the difference between the unadjusted ("actual") distributions for 1969 and 1989, along with the difference between the unadjusted 1969 distribution and the 1989 distribution adjusted for all four of our explanatory factors ("fully adjusted"). A "flat line" difference between the unadjusted 1969 distribution and the fully adjusted 1989 distribution would indicate that our factors fully account for the exact observed change in the distribution of family equivalent income. Although we do not obtain a flat line, the figure shows that accounting for our four explanatory factors substantially decreases the difference in densities. In particular, the relatively greater mass at very low incomes in the 1989 distribution largely disappears, and the relatively greater mass in the lower middle in the 1969 distribution and the upper portion in the 1989 distribution is substantially reduced.

The quantitative analogue to the visual representation in Figures 5 and 6 is listed in Table 2. Due to their precision, we focus on these quantitative assessments in the remainder of the paper. The first column of the table shows the total change in the measured statistic between the two years—that is, the 1989 value minus the 1969 value. All of the dispersion measures increased substantially between these two years, although the increase in dispersion was larger in the lower half of the distribution (the 50-10 ratio) than it was in the upper half (the 90-50 ratio). The additional columns show the portion of the total change that can be attributed to changes in the explanatory factors, bootstrapped standard errors for these estimates in parentheses, and each factor's contributory share of the total change in the inequality measure listed in brackets.[20] In addition to the four explanatory factors, we also list the contribution of residual (unmeasured) factors, which is defined as the difference between the total change and the net portion accounted for by our explanatory factors.[21]

The first row of Table 2 shows that changes in the distribution of men's earnings and in family structure had a largely neutral effect on the midpoint of the distribution of family equivalent

[20] Because the estimated standard errors imply high levels of statistical significance for the estimated factor contributions discussed in the text, we do not discuss them further. The bootstrap standard errors reported in the tables are obtained using the paired approach, based on 200 random resamplings (with replacement) of the original data (see Efron and Tibshirani 1993 for discussion of the paired or XY bootstrap).

[21] In each row, the share effects across the five factor columns sum to one (except for rounding errors in some cases).

income. However, the rise in women's labor force participation and, to a lesser degree, the change in typical family characteristics both increased median family equivalent income. The large effect of rising labor force participation by women on median income indicates the importance of using inequality measures in normalized terms—that is, percentile ratios and parametric measures such as the coefficient of variation rather than the standard deviation. Given the large effect of female participation on the location (median) of the income distribution, non-normalized dispersion measures such as the standard deviation would mistakenly suggest that rising female participation either had a neutral effect on or increased inequality in family income.

The remaining rows of Table 2 list the total changes and the effects of the explanatory factors for our key dispersion measures. The estimates listed in the second column indicate that relative to the other factors considered, the changing distribution of men's earnings had the largest and most uniform effect on family income inequality. Excluding changes in the standard deviation, changes in the distribution of men's earnings explain from just over one-half to about four-fifths of rising dispersion in family equivalent income. Although the changing distribution of men's earnings explains a greater share of increased inequality in the upper half of the income distribution, in absolute terms it had a greater impact on the lower half, where the distribution widened by a larger amount.[22]

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[22]We find a larger effect of changes in the distribution of men's earnings than did Burtless (1999), for two reasons: we account for the rising incidence of zero earnings among men, whereas' Burtless did not; and the increase in the dispersion of men's earnings slowed in the 1990s, which reduced the contribution of this factor to rising inequality in family income during Burtless' sample period (1979 to 1996). See our Table 4 and discussion below for results from the 1990s.

The third column of Table 2 shows that rising labor force participation by women (conditional on family structure and family characteristics) tended to reduce dispersion in family equivalent income between 1969 and 1989. Although increased female participation increased the standard deviation of family equivalent income, it increased the mean by a larger amount and therefore reduced the coefficient of variation by an amount equal to nearly one-quarter of the actual increase in that statistic. Rising female participation also reduced the other dispersion measures listed, with the size of its impact ranging from 5 to 32 percent of the increase in dispersion. In share terms, the largest impact of female participation was on dispersion in the upper half of the distribution (the 90-50 ratio); relative to the total change in the statistic, rising female participation explains much less of the increase in the bottom half (the 50-10 ratio, and the mean logarithmic deviation, which is relatively sensitive to dispersion towards the bottom).

Results listed in the fourth column of Table 2 indicate that changes in family structure (conditional on the distribution of family characteristics) substantially increased inequality between 1969 and 1989. The impact of changing family structure was concentrated in the lower half of the distribution of family equivalent income; it explains 52 percent of the increase in the 50-10 ratio, 49 percent of the increase in the 95-5 ratio, and only about 9 percent of the increase in the 90-50 ratio. This suggests that increasing prevalence of family types associated with low incomes—especially never married individuals with children—was more important for rising dispersion of family equivalent income than was the increasing prevalence of family types associated with higher incomes—mainly married and single individuals without children. Changing family structure also explains about one-third of the increase in the key summary measures of overall dispersion (the Gini and Theil indices and mean logarithmic deviation).

The fifth column of Table 2 lists the effects of the changing distribution of family characteristics on the distribution of family equivalent income. Once the effects of these variables are conditioned out in the previous steps of the decomposition, they have very little independent impact on dispersion, with a mixture of small positive and negative effects across the various dispersion measures. As discussed below, however, the impact of family characteristics is somewhat larger when they are considered first in our decomposition.

The contribution of residual (unmeasured) factors, listed in the final column of Table 2, is small compared to the net impact of our modeled factors, indicating that our four explanatory factors account for the bulk of the change in our dispersion measures. Our explanatory factors account for about 60 percent of the increase in the standard deviation, about 80 percent of the increase in the Gini and Theil measures, and nearly 100 percent of the increase in several of the percentile ratio measures. Although omitted factors may have had large effects that offset each other and therefore account on net for only a small residual share, it seems clear that our explanatory factors played a key role in changing inequality between 1969 and 1989.

The final row of Table 2 lists the effects of the explanatory factors on the change in the poverty rate between 1969 and 1989. The effects of the explanatory factors on poverty are qualitatively similar to their impact on inequality, although the effects on poverty were especially large relative to the total change in the poverty rate between 1969 and 1989. Taken separately, the changing distribution of men’s earnings and changes in family structure increased the poverty rate by an amount approximately equal to the actual increase. These effects were offset somewhat by poverty-reducing impacts of rising female participation and changes in family characteristics, both of which were associated with a (counterfactual) reduction in the poverty rate of nearly a percentage point. On net, however, the explanatory factors account for 136 percent of the change in the poverty rate, indicating that unmeasured factors are associated with a (counterfactual) reduction in the poverty rate of 1 percentage point. One key omitted factor is government transfer payments; they are included in our measure of family equivalent income, but we do not directly account for receipt of government transfers in our decomposition. The population eligible for government transfer programs increased between 1969 and 1989, which may account for a substantial portion of the residual decline in poverty during that period.[23]

We investigated the sensitivity of these results to two main data issues mentioned in Section I. First, we checked the sensitivity to the specific family equivalence scale used, by setting the scaling factor equal to 0.25 and 0.75 (in addition to the value of 0.5 used for the results reported in our tables). We found very little sensitivity of our results to equivalence scale, either in regard to the rise in inequality or the contribution of our modelled factors. The increase in inequality between 1969 and 1989 is somewhat larger for higher values of the equivalence scaling factor, but the factor contributions are quite similar in share terms for each case.

We also investigated the sensitivity of our results to the treatment of cohabiting couples (unmarried partners, with or without children, who share living quarters). Indirect estimates from the U.S. Bureau of the Census (Casper et al. 1999) suggest that the number of cohabiting couples increased substantially over the past few decades, although they constitute a relatively small percentage of all families. By not identifying such couples and not accounting for their resource

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[23] Between 1969 and 1989, caseloads for the Aid to Families with Dependent Children (AFDC) program increased significantly, roughly accompanying the rise in the number of single mothers (Moffitt 2000).

sharing, we may overstate the contribution of changing living arrangements to rising income inequality.

Because the CPS only began explicit identification of cohabitants in 1996, we are unable to incorporate them directly into our analyses.[24] However, under the extreme assumption of no cohabitation in 1969, we can obtain an upper-bound estimate of the impact of cohabitation for the estimated change in inequality between 1969 and 1998. When we do so, the results change somewhat from those obtained using our primary sample definitions (ignoring cohabitants) for the same period. In particular, changing family structure makes a somewhat smaller contribution to rising income inequality in the lower half of the distribution when cohabitants are incorporated in this fashion: rather than explaining nearly one-half of the increase in the 50-10 ratio between 1969 and 1998, changing family structure (including cohabitants) explains just under 40 percent of the increased 50-10 ratio. More generally, the impact of changing family structure on overall dispersion is little affected by the incorporation of cohabitation, even in this upper-bound case.[25]

We also investigated whether the patterns identified for the entire 1969-1989 period were more pronounced in the 1970s or the 1980s, by applying our analyses to the sub-periods of 1969-1979 and 1979-1989. The basic tabulations show that increases in family income inequality were more pronounced during the 1980s than during the 1970s, especially for measures focused on the

[24] Although indirect estimates of cohabitation may be informative with respect to trends over time, they are likely to introduce substantial error into the identification of living arrangements for individual households.

[25] Our approximation also assumes that income is shared by cohabiting couples in the same way as married couples, which may not be true. Consistent with our finding of a relatively small impact, Carlson and Danziger (1999) report that accounting for rising cohabitation only slightly reduces the measured increase in child poverty in the U.S. between 1969 and 1989.

upper portion of the distribution, such as the 90-50 ratio. On the other hand, our primary explanatory factors—male earnings, female participation, and family structure—each explain a smaller share of rising dispersion in the 1980s than in the 1970s, indicating a more prominent role for residual factors in that decade. The exceptions are a larger offsetting impact of rising female participation on dispersion in the lower portion of the distribution in the 1980s, and a larger impact of changing family structure on dispersion in the upper half of the distribution in the 1980s. Table 3 displays the exact results for the sub-period 1979-1989.

Reverse-Order Decomposition

As discussed in Section II, the contribution of the explanatory factors may depend on their order in the decomposition analysis, due to the possibility of joint causation in the distribution of men's earnings, female labor supply, and family structure. To address this possibility, Table 4 presents results from the reverse-order decomposition analysis for the 1969-1989 period, for which we consider, in turn, the effects of the changing distributions of family characteristics, family structure, women's labor force participation, and men's earnings.[26] The impact of the changing distribution of family characteristics is increased in magnitude when we consider it first rather than last in the decomposition. Compared with the primary-order decomposition (Table 2), in which its effects on the dispersion of family equivalent income were small and of mixed sign, the reverse-order decomposition in Table 4 indicates that the changing distribution of family characteristics increased each of the dispersion measures. The share increase in dispersion

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[26] We do not present charts for this decomposition and the additional sub-period decompositions because the visual representation of their effects is similar to that displayed for the primary-order decomposition in Figure 5.

associated with this factor ranges from 6 percent for the coefficient of variation to 27 percent for the 95-5 ratio.

The larger impact of family characteristics in the reverse-order decomposition, however, does not substantially alter the impact of the other explanatory factors on dispersion. The effects of the changing distribution of men's earnings when it is considered last in the decomposition (Table 4) are very similar to its effects when it is considered first (Table 2), with the exception of a noticeably smaller impact on the mean logarithmic deviation in the former case. Compared with Table 2, in Table 4 the effects of family structure are somewhat smaller in the bottom portion of the distribution but a bit larger in the upper portion, and the effects on the key summary measures are only slightly smaller in the reverse-order case. Rising female participation has a larger equalizing effect on the distribution of family equivalent income in the reverse-order case than it did in the primary-order decomposition in Table 2. This larger equalizing effect is most evident in the lower portion of the distribution, with the reduction in the 50-10 ratio due to female participation increasing from 3 percent of the total change (Table 2) to about 22 percent of the total change (Table 4). In addition, the reverse-order effects of rising female participation on the coefficient of variation, the Gini and Theil coefficients, and the mean logarithmic deviation are equal to about 20-25 percent of the total change, a noticeably larger effect than in the primary-order decomposition. Overall, however, the residual portion left unexplained after accounting for our four factors is largely the same in the primary-order and reverse-order cases.

The effects of the explanatory factors on poverty are somewhat different in the reverse-order decomposition than in the primary-order decomposition. The changing distribution of family characteristics increases poverty a bit when considered first in the decomposition (Table 4), whereas it reduces poverty when considered last (Table 2). Among the other explanatory factors, the poverty-increasing effect of changing family structure is smaller in the reverse-order case, the poverty-reducing effect of female labor supply is larger, and the poverty-increasing effect of men's earnings is larger. Although the impact of the changing distribution of men's earnings is especially large in the reverse-order case, the qualitative interpretation of the effects is basically unchanged.[27]

Changes Between 1989 and 1998

We now turn our attention to a comparison of the years 1989 and 1998, which provides an updated view of changing family income inequality in the U.S., for a period which showed some diminution in the trend towards rising inequality. Table 5 lists primary-order decomposition results for changes in the distribution of family equivalent income between 1989 and 1998. The pace of rising inequality slowed in the 1990s: the total changes in our dispersion statistics during this period are positive but small compared to their increases during the preceding two decades (Table 2, first column). In assessing the factor contributions, we focus here on the actual explained change rather than the share of the total change, for purposes of direct comparison with earlier periods. Leaving aside the effects of men's earnings for the moment, the results in the third and fourth columns indicate that rising female participation reduced inequality in the 1990s and changing family structure increased it (by amounts similar to those in the 1980s; see Table 3).

[27] The residual factors account for slightly different shares of the total change in the primary-order and reverse-order cases. Although the net impact of the conditionally weighted factors does not differ in the two cases, due to equivalence of the product of the primary-order weights with the product of the reverse-order weights (see Appendix B), the full impact of the four explanatory factors is affected by the ordering of the men's earnings step.

In contrast, the impact of the changing distribution of men's earnings on rising inequality in family income was much smaller in the 1990s than it was in earlier decades. This includes a substantially reduced impact of men's earnings on the relative well-being of low-income families, as measured by the small positive effect of this factor on the change in the 50-10 ratio and an actual negative effect (albeit quite small) on the poverty rate. Two potential explanations for the effect of changes in the distribution of men's earnings on low-income families in the 1990s are changes in the share of males with zero earnings and changes in the minimum wage. Auxiliary tabulations (available on request) indicate that the share of male household heads with zero earnings remained approximately constant between 1989 and 1998, so the first explanation is not compelling. A more likely explanation is the increase in the U.S. federal minimum wage, which rose 16 percent between 1989 and 1998. Low hourly earnings for a male household head does not necessarily imply that his family's total income is low. However, the correspondence is likely to be close enough for the increase in the minimum wage to account for the reduced impact of male wage inequality on low-income families during the 1990s.

The effects of our explanatory factors on poverty during the 1990s are displayed at the bottom of Table 5. The poverty rate was essentially unchanged between 1989 and 1998. However, changes in our explanatory factors account for changes in the poverty rate ranging from three-tenths of a percentage point to just over a percentage point. Changes in family structure imply a full percentage point increase in poverty in the 1990s, which is almost as large as the family structure effect on poverty during each of the previous two decades (see Table 2, fourth column, bottom row; the total contribution of 0.026 implies a contribution of half that, or 0.013, for each of the two decades included in that table).

Among the most interesting results in Table 5 is the change in poverty attributable to residual factors (final column, bottom row). The residual effect was especially large in the 1990s, implying an increase in the poverty rate of 1.2 percentage points. This contrasts directly with a negative residual contribution between 1969 and 1989 (Tables 2 and 3), which is consistent with expansion of government transfer programs between those two years (Moffitt 2000). In contrast, the large positive residual contribution between 1989 and 1998 (Table 5) is consistent with a reduction in benefit receipt following the 1996 welfare reform law.

IV. Conclusions

Our assessment of the causes of rising inequality in U.S. family incomes reinforce some of the results from past work but also includes several important points of departure. We focussed first on the 1969 to 1989 period, which corresponds roughly to the period of rising inequality examined in much previous work. Our explanatory factors account for most of the increase in inequality and poverty during this period, with unobserved residual factors explaining an amount ranging from none to about one-third of the total change in our key dispersion measures. The changing distribution of men's earnings had the largest impact on inequality and poverty, explaining anywhere from about one-half to four-fifths of the increase. Changes in family structure—most notably the increase in families characterized by single heads with children—also explained a substantial share of the increase; these effects were concentrated among lower-income families, especially those at risk of poverty.

In contrast, we found that the increase in women's labor force participation between 1969 and 1989 tended to offset increasing inequality and poverty over this period. Its counterfactual impact ranged as high as one-third of the actual change in dispersion. Importantly, its impact spread down the income distribution over time, with a larger impact in the lower half of the distribution in the 1980s and 1990s than in the 1970s.

The equalizing effect of rising female participation might appear to conflict with previous findings of a rising correlation between husbands' and wives' earnings (see Juhn and Murphy 1997), which would tend to increase inequality. However, the impact of a rising correlation between husband and wife earnings depends on the precise income rank and related characteristics of families that generate this rising correlation. Our conditional weighting approach enables us to focus on specific portions of the distribution and to control explicitly for the changing distribution of family characteristics. The results indicate that *ceteris paribus*, the increase in women's labor force participation, and its positive impact on family income, was concentrated among lower-income families. Of note, using an approach that is similar to ours, Biewen (2001) also finds equalizing effects of women's labor force participation on the East German income distribution in the 1990s (in particular, declining female participation was associated with rising inequality in that setting).

We also examined changes in family income inequality and poverty between 1989 and 1998. The net change in inequality during this period was relatively small, largely due to changes in the distribution of men's earnings which actually served to reduce inequality somewhat in the lower portion of the distribution of family income. However, changing family structure and rising labor force participation by women continued to have substantial effects on income dispersion in the 1990s, with rising female participation largely offsetting the inequality-increasing effects of changing family structure.

Except for the poverty rate, we focused on standard measures of inequality, such as the coefficient of variation, the Gini and Theil coefficients, and percentile ratios, in order to enhance the comparability of our analyses with existing analyses. However, because our approach relies on complete density re-estimation, it could easily be adapted to provide more specific welfare analysis. For example, between 1969 and 1989, real living standards (as measured by our pre-tax, post-transfer income variable) declined for families below the 18th percentile of the distribution of U.S. equivalent family income. Our technique could be used to assess the contribution of explanatory factors to such changes in real living standards, thereby providing welfare-based information for policymakers interested in maintaining or improving living standards for specific target populations. Beyond that, extension of our approach to data from European countries might be especially interesting; the patterns of changing inequality differ substantially across these countries, as do the extent of changes in individual earnings and family behavior.

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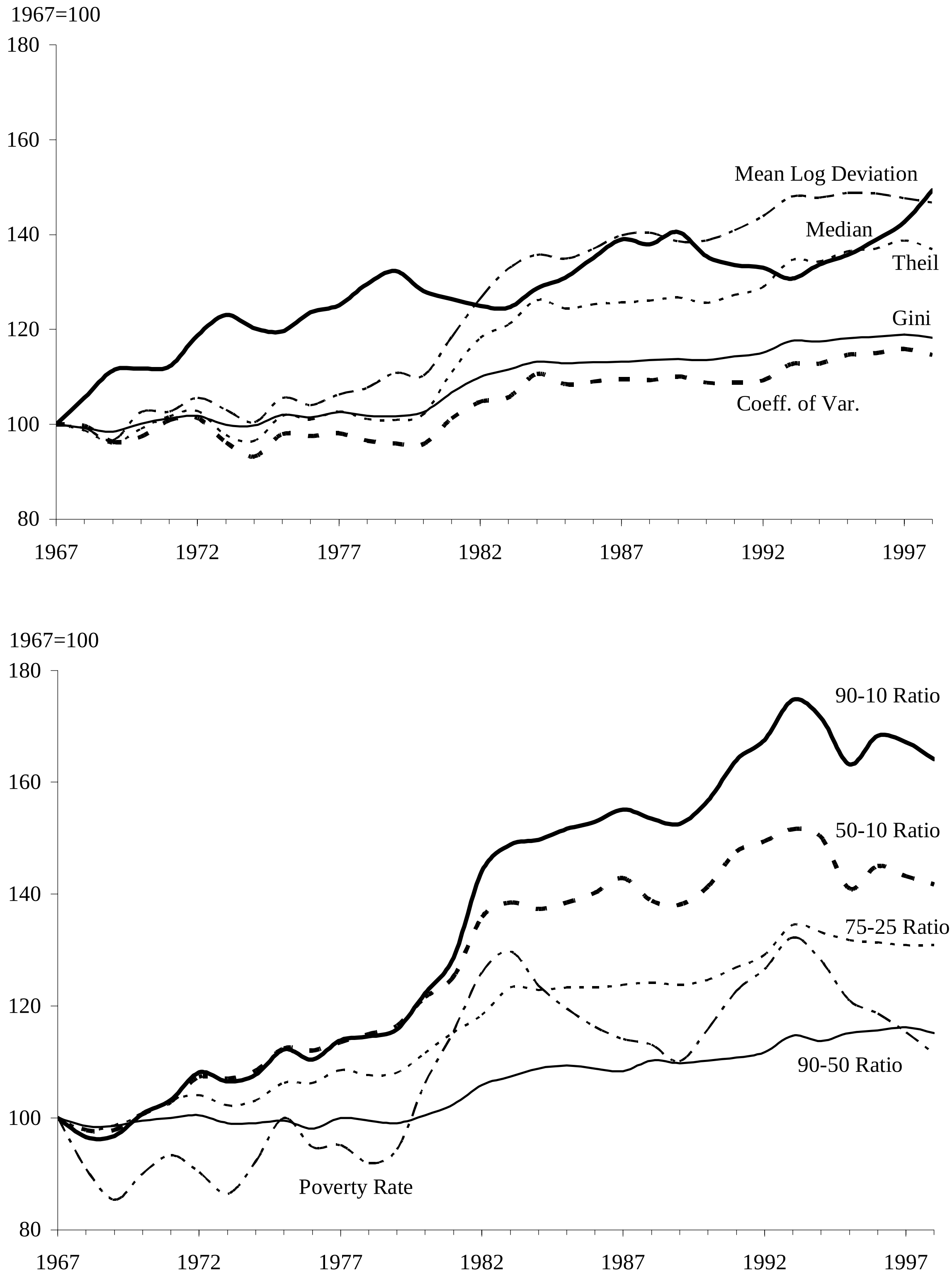
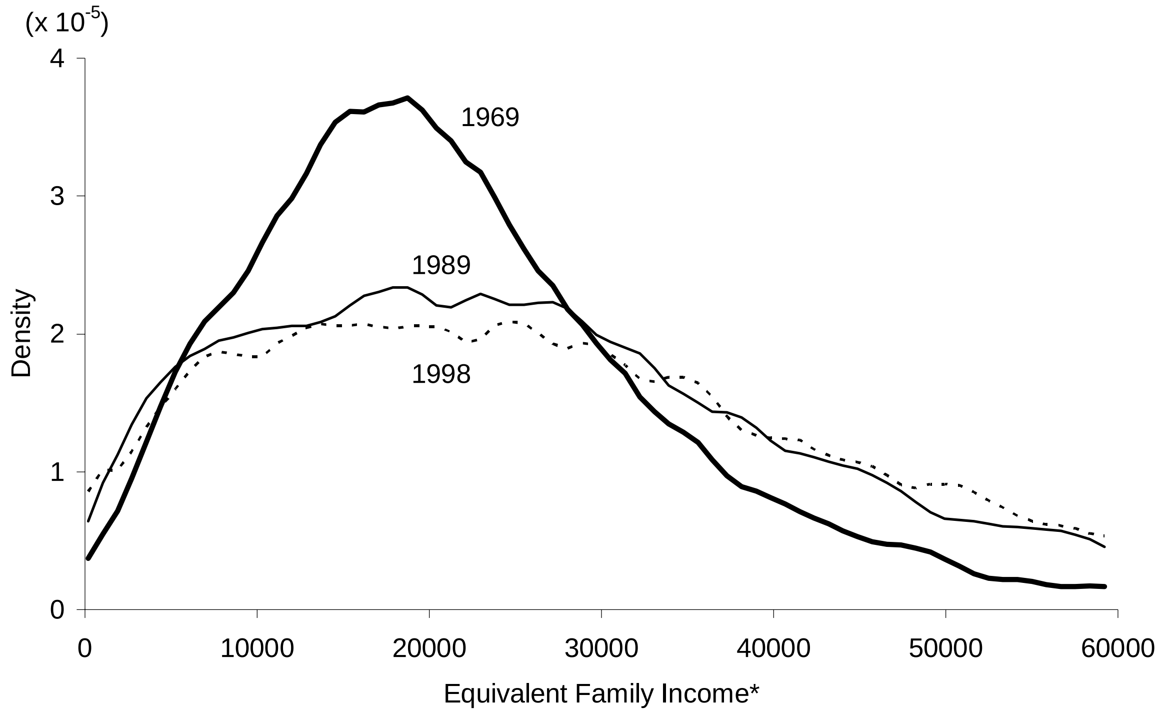
Figure 1 - Inequality in Family Equivalent Income (1967-1998)



Source: Authors' tabulations of March CPS data; figures in 1998 dollars.

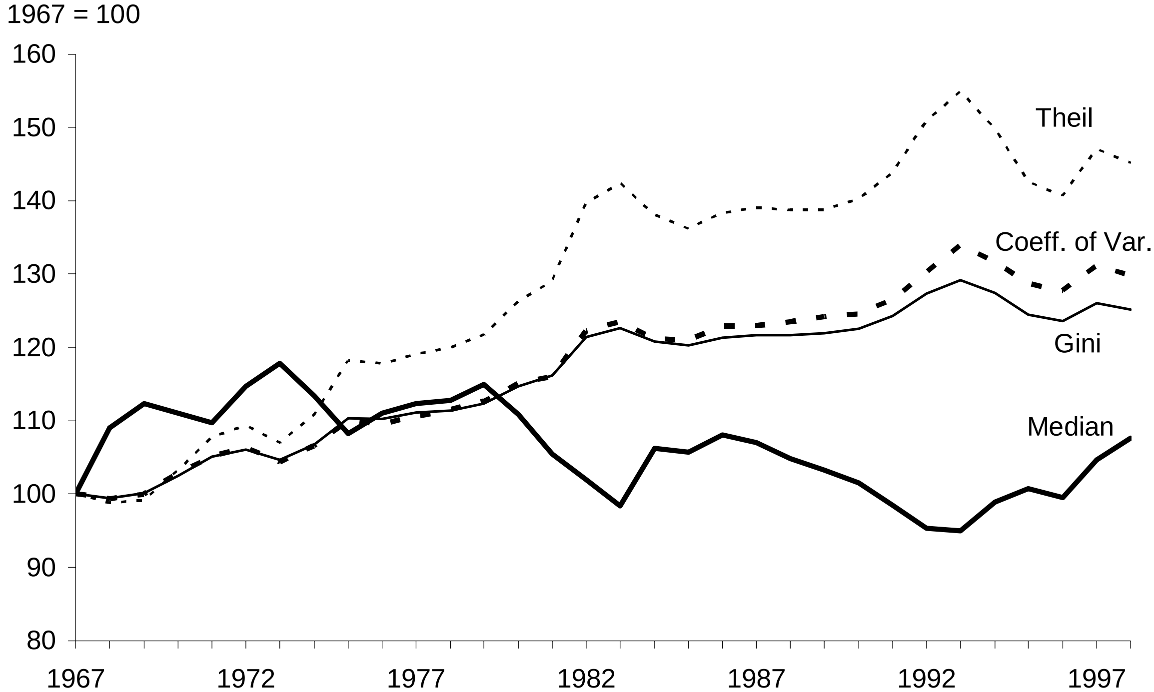
Figure 2
Density of Real Family Equivalent Income



* Distributions truncated at 0 and 60,000

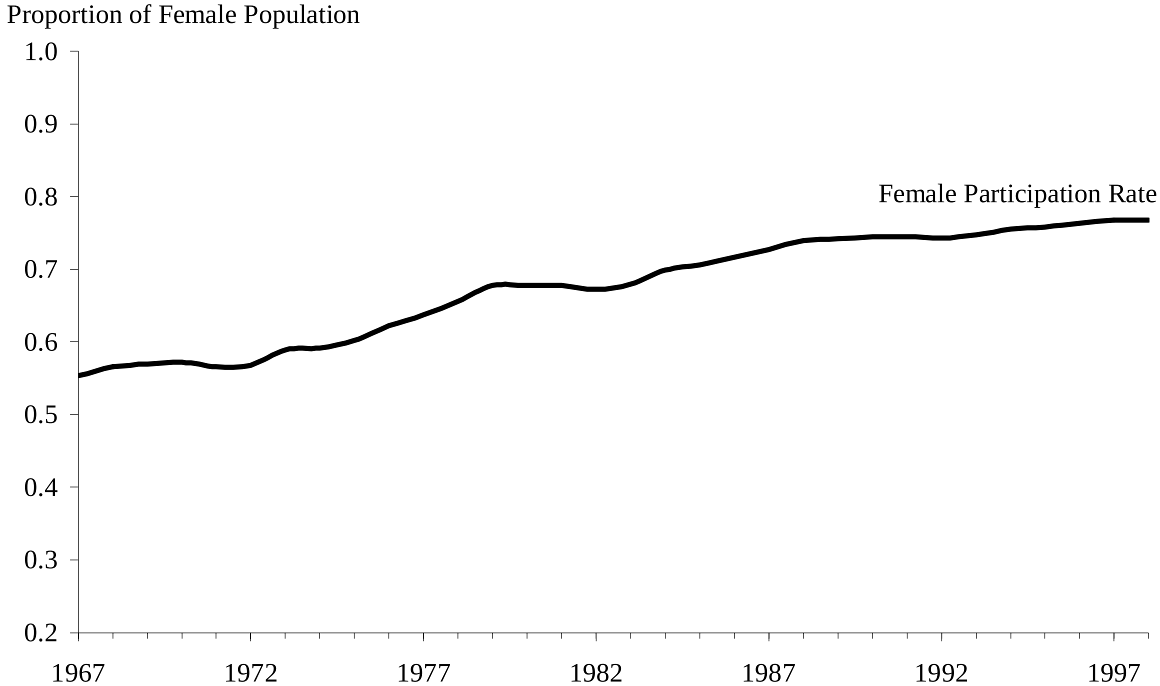
Source: Authors' tabulations of 1970, 1990, and 1999 March CPS data; figures in 1998 dollars.

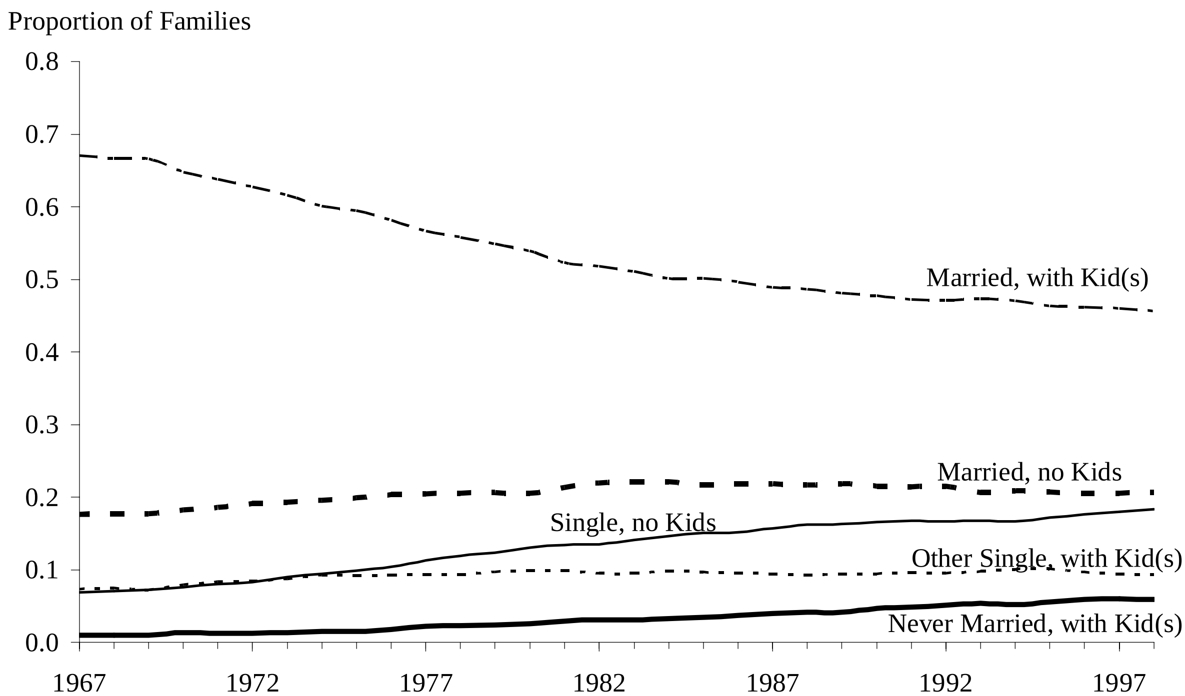
Figure 3 - Summary Measures for Male Earnings (1967-1998)



Source: Authors' tabulations of March CPS data; figures in 1998 dollars.

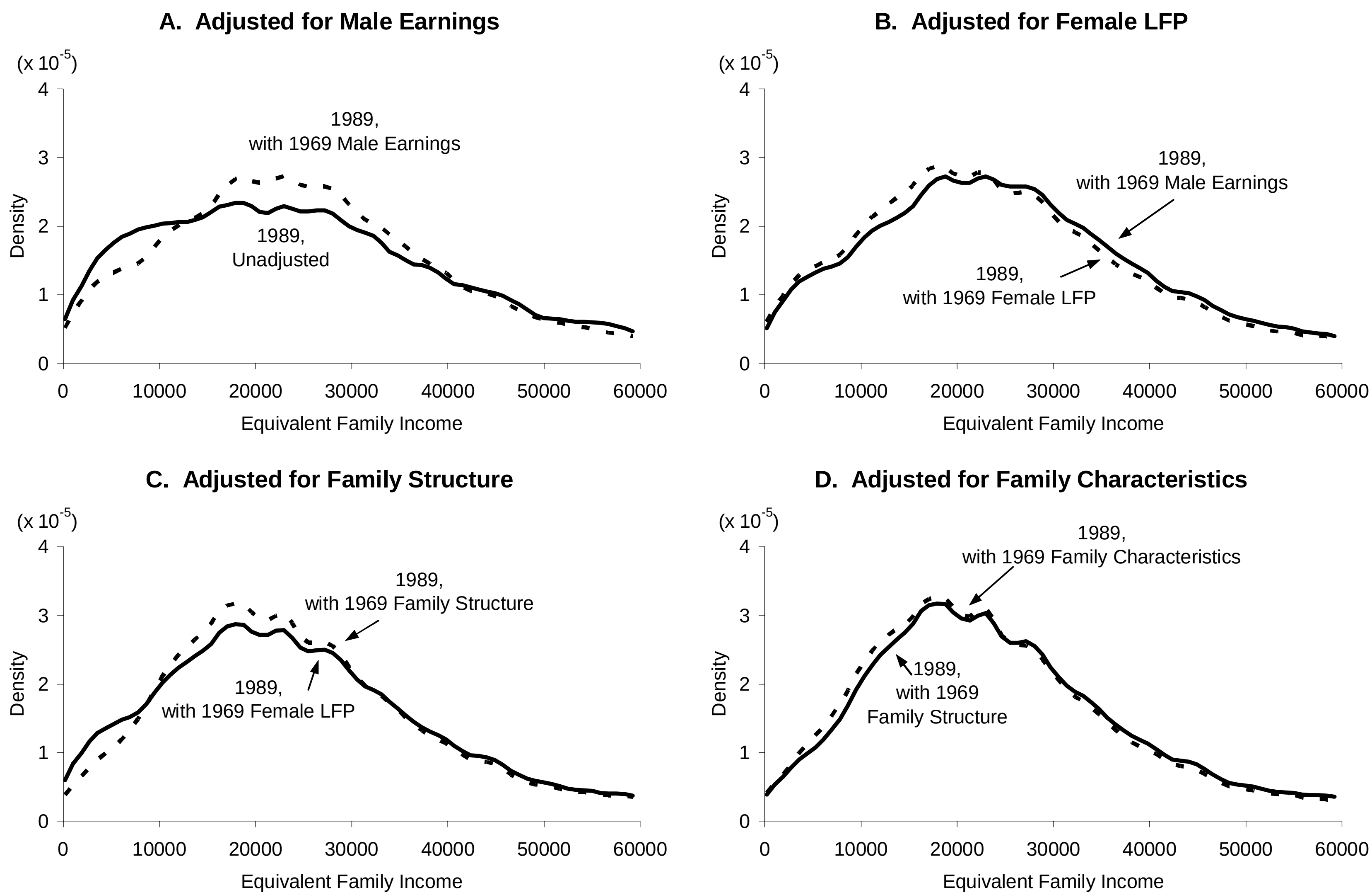
Figure 4 - Female Labor Force Participation and Family Structure (1967-1998)





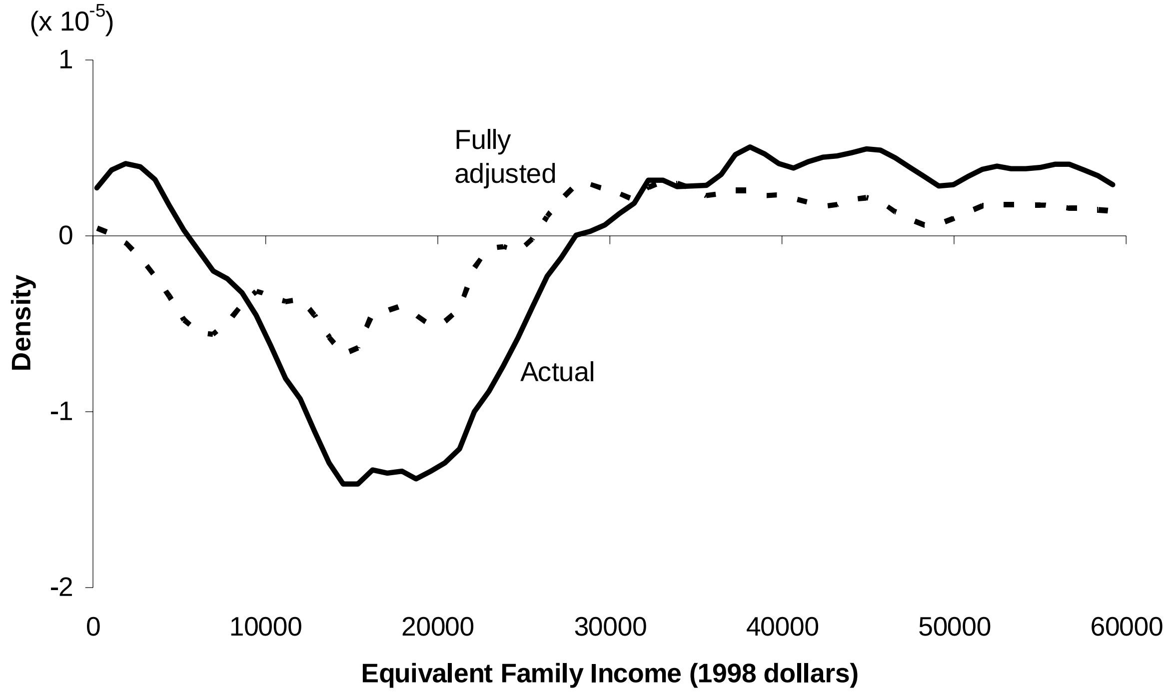
Source: Authors' tabulations of March CPS data.

Figure 5
Density of Real Family Equivalent income, 1989, Adjusted to 1969 Distribution of Indicated Factors



Note: Figures in 1998 dollars.

Figure 6
Difference between 1989 and 1969 Densities



Note: Calculated as differences between the densities in Figures 2 and 3 (panel D).

Table 1: Income Measures and Conditioning Weights Used in the Density Decomposition

Primary-order decomposition

Distribution	Income Measure	Weight
1. 1989 distribution (actual) $f(Y;\ t_Y=89,\ m_{89},\ t_{L\|S,X}=89,\ t_{S\|X}=89,\ t_X=89)$	Y	θ_{89}
2. (1) with 1969 men's earnings distribution $(Y;\ t_Y=89,\ m_{69},\ t_{L\|S,X}=89,\ t_{S\|X}=89,\ t_X=89)$	Y_m	θ_{89}
3. (2) with 1969 women's LFP $f(Y;\ t_Y=89,\ m_{69},\ t_{L\|S,X}=69,\ t_{S\|X}=89,\ t_X=89)$	Y_m	$\theta_{89}\cdot \hat{\psi}_{L\|S,X}$
4. (3) with 1969 family structure $f(Y;\ t_Y=89,\ m_{69},\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=89)$	Y_m	$\theta_{89}\cdot \hat{\psi}_{L\|S,X}\cdot \hat{\psi}_{S\|X}$
5. (4) with 1969 family characteristics $f(Y;\ t_Y=89,\ m_{69},\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=69)$	Y_m	$\theta_{89}\cdot \hat{\psi}_{L\|S,X}\cdot \hat{\psi}_{S\|X}\cdot \hat{\psi}_{X}$
6. 1969 distribution (actual) $f(Y;\ t_Y=69,\ m_{69},\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=69)$	Y	θ_{69}

Reverse-order decomposition

Distribution	Income Measure	Weight
1. 1989 distribution (actual) $f(Y;\ t_Y=89,\ t_{X\|S,L}=89,\ t_{S\|L}=89,\ t_L=89,\ m_{89})$	Y	θ_{89}
2. (1) with 1969 family characteristics $f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=89,\ t_L=89,\ m_{89})$	Y	$\theta_{89}\cdot \hat{\psi}_{X\|S,L}$
3. (2) with 1969 family structure $f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=89,\ m_{89})$	Y	$\theta_{89}\cdot \hat{\psi}_{X\|S,L}\cdot \hat{\psi}_{S\|L}$
4. (3) with 1969 women's LFP $f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m_{89})$	Y	$\theta_{89}\cdot \hat{\psi}_{X\|S,L}\cdot \hat{\psi}_{S\|L}\cdot \hat{\psi}_{L}$
5. (4) with 1969 men's earnings distribution $f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m_{69})$	Y_m	$\theta_{89}\cdot \hat{\psi}_{X\|S,L}\cdot \hat{\psi}_{S\|L}\cdot \hat{\psi}_{L}$
6. 1969 distribution (actual) $f(Y;\ t_Y=69,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m_{69})$	Y	θ_{69}

Note: Y refers to equivalent family income, θ_t is the survey sampling weight for families in year t (see Section I), Y_m is equivalent family income adjusted for the change in the distribution of men's earnings (see Section II), and the $\hat{\psi}$'s are estimated conditioning weights (see Section II). The subscripts for the adjustment variables — m, L, S, and X — refer to men's earnings, women's labor force participation, family structure, and other family characteristics, respectively.

Table 2: Primary-Order Decomposition of Changes in the
Distribution of Family Equivalent Income and Poverty, 1969-1989

Statistic	Total Change	Effect of:[a] Men's earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	5281	-40.9	1476	301	915	2,629
		(135)	(89.3)	(58.4)	(102)	(126)
		[-.008]	[0.279]	[0.057]	[0.173]	[0.498]
Standard Deviation	7,247	1990	173	941	1060	3,084
		(211)	(84.9)	(88.5)	(103)	(268)
		[0.275]	[0.024]	[0.130]	[0.146]	[0.425]
Coefficient of Variation[c]	0.095	0.053	-0.023	0.029	0.005	0.031
		(.006)	(.003)	(.003)	(.003)	(.009)
		[0.552]	[-0.240]	[0.305]	[0.055]	[0.329]
90-10[d]	3.06	2.12	-0.299	1.25	-0.073	0.060
		(.099)	(.065)	(.085)	(.060)	(.123)
		[0.693]	[-0.098]	[0.409]	[-0.024]	[0.020]
50-10	1.09	0.698	-0.034	0.565	-0.076	-0.06
		(.041)	(.029)	(.037)	(.027)	(.053)
		[0.638]	[-0.031]	[0.516]	[-0.069]	[-0.054]
90-50	0.231	0.183	-0.074	0.020	0.033	0.069
		(.012)	(.009)	(.010)	(.012)	(.019)
		[0.794]	[-0.321]	[0.085]	[0.143]	[0.298]
75-25	0.568	0.437	-0.065	0.157	-0.031	0.069
		(.017)	(.012)	(.011)	(.012)	(.023)
		[0.769]	[-0.114]	[0.276]	[-0.054]	[0.122]
95-5	8.26	4.43	-0.963	4.08	-0.168	0.883
		(.237)	(.237)	(.300)	(.161)	(0.356)
		[0.536]	[-0.117]	[0.494]	[-0.020]	[0.107]
Gini Coefficient	0.052	0.033	-0.009	0.018	0.000	0.010
		(.002)	(.001)	(.001)	(.001)	(.003)
		[0.638]	[-0.175]	[0.339]	[0.006]	[0.193]
Theil's Coefficient	0.062	0.038	-0.011	0.022	0.001	0.012
		(.002)	(.001)	(.001)	(.001)	(.003)
		[0.611]	[-0.183]	[0.355]	[0.015]	[0.202]
Mean Logarithmic Deviation	0.089	0.050	-0.009	0.030	-0.000	0.017
		(.002)	(.001)	(.002)	(.001)	(.004)
		[0.567]	[-0.099]	[0.341]	[-0.003]	[0.194]
Poverty Rate	0.028	0.029	-0.008	0.026	-0.009	-0.010
		(.002)	(.001)	(.001)	(.001)	(.004)
		[1.04]	[-0.301]	[0.933]	[-0.303]	[-0.364]

[a] Numbers in parentheses show bootstrapped standard errors for the estimated factor contributions. Numbers in brackets show the share of the explained change in the total change.
[b] See Section III in the text for a complete list of the family characteristic variables.
[c] Defined as the standard deviation divided by the mean.
[d] Ratio of the 90th and 10th percentiles of the distribution of real family equivalent income. The other percentile ratio measures are defined similarly.
Source: Authors' calculations using 1970 and 1990 March CPS data.

Table 3: Primary-Order Decomposition of Changes in the
Distribution of Family Equivalent Income and Poverty, 1979-1989

		Effect of:[a]				
Statistic	Total Change	Men's earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	1,542	-492	619	128	-1117	2404
		(138)	(60.4)	(51.9)	(102)	(132)
		[-0.319]	[0.401]	[0.083]	[-0.724]	[1.56]
Standard Deviation	4,787	1584	80.6	443	-226	2905
		(108)	(38.4)	(53.5)	(65.1)	(218)
		[0.331]	[0.017]	[0.092]	[-0.047]	[0.607]
Coefficient of Variation[c]	0.097	0.049	-0.012	0.011	0.019	0.030
		(0.003)	(0.001)	(0.002)	(0.002)	(0.007)
		[0.498]	[-0.119]	[0.113]	[0.197]	[0.311]
90-10[d]	2.00	0.877	-0.392	0.442	0.458	0.618
		(0.099)	(0.072)	(0.070)	(0.066)	(0.139)
		[0.438]	[-0.196]	[0.221]	[0.229]	[0.309]
50-10	0.589	0.177	-0.139	0.188	0.149	0.214
		(0.040)	(0.030)	(0.031)	(0.029)	(0.063)
		[0.300]	[-0.236]	[0.319]	[0.253]	[0.364]
90-50	0.216	0.134	-0.027	0.012	0.042	0.055
		(0.011)	(0.006)	(0.005)	(0.007)	(0.014)
		[.622]	[-0.123]	[0.054]	[0.195]	[0.253]
75-25	0.352	0.195	-0.061	0.055	0.101	0.061
		(0.017)	(0.012)	(0.010)	(0.012)	(0.024)
		[0.555]	[-0.172]	[0.157]	[0.286]	[0.174]
95-5	5.83	2.084	-1.106	1.26	1.25	2.349
		(0.275)	(0.207)	(0.201)	(0.216)	(0.405)
		[0.357]	[-0.190]	[0.216]	[0.214]	[0.403]
Gini Coefficient	0.041	0.020	-0.005	0.005	0.008	0.012
		(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
		[0.501]	[-0.131]	[0.131]	[0.204]	[0.295]
Theil's Coefficient	0.52	0.025	-0.007	0.007	0.010	0.016
		(0.002)	(0.001)	(0.001)	(0.001)	(0.003)
		[0.493]	[-0.133]	[0.137]	[0.198]	[0.304]
Mean Logarithmic Deviation	0.058	0.027	-0.007	0.008	0.011	0.019
		(0.002)	(0.001)	(0.001)	(0.001)	(0.004)
		[0.464]	[-0.127]	[0.143]	[0.194]	[0.26]
Poverty Rate	0.017	0.012	-0.008	0.006	0.011	-0.004
		(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
		[0.681]	[-0.448]	[0.322]	[-0.651]	[-0.207]

[a] Numbers in parentheses show bootstrapped standard errors for the estimated factor contributions. Numbers in brackets show the share of the explained change in the total change.
[b] See Section III in the text for a complete list of the family characteristic variables.
[c] Defined as the standard deviation divided by the mean.
[d] Ratio of the 90th and 10th percentiles of the distribution of real family equivalent income. The other percentile ratio measures are defined similarly.

Source: Authors' calculations using 1980 and 1990 March CPS data.

Table 4: Reverse-Order Decomposition of Changes in the
Distribution of Family Equivalent Income and Poverty, 1969-1989

Statistic	Total Change	Effect of:[a]				
		Family Characteristics[b]	Family Structure	Female LFP	Men's earnings	Residual Factors
Median	5,281	545 (153) [0.103]	-89.8 (91.5) [-0.017]	1437 (107) [0.272]	519 (131) [0.098]	2870 (134) [0.543]
Standard Deviation	7247	727 (174) [0.100]	1121 (143) [0.155]	246 (64.3) [0.034]	1693 (222) [0.271]	3189 (272) [0.440]
Coefficient of Variation[c]	0.095	0.006 (.005) [0.063]	0.038 (.004) [0.395]	-0.026 (.002) [-0.275]	0.050 (.007) [0.527]	0.028 (.009) [0.289]
90-10[d]	3.06	0.489 (.111) [0.160]	1.05 (.100) [0.344]	-0.749 (.081) [-0.245]	2.25 (.110) [0.734]	0.023 (.124) [0.008]
50-10	1.09	0.174 (.044) [0.159]	0.428 (.041) [0.391]	-0.240 (.033) [-0.220]	0.810 (.045) [0.740]	-0.078 (.054) [-0.071]
90-50	0.231	0.028 (.012) [0.120]	0.034 (.008) [0.147]	-0.067 (.014) [-0.289]	0.166 (.015) [0.721]	0.070 (.020) [0.301]
75-25	0.568	0.043 (.019) [0.075]	0.174 (.016) [0.307]	-0.142 (.019) [-0.249]	0.439 (.021) [0.772]	0.054 (.024) [0.096]
95-5	8.26	2.22 (.349) [0.268]	2.99 (.310) [0.362]	-1.60 (.182) [-0.194]	4.04 (.265) [0.489]	0.618 (.335) [0.075]
Gini Coefficient	0.052	0.006 (.001) [0.110]	0.016 (.001) [0.302]	-0.013 (.001) [-0.248]	0.035 (.002) [0.635]	0.008 (.003) [0.161]
Theil's Coefficient	0.062	0.007 (.002) [0.114]	0.022 (.001) [0.352]	-0.016 (.001) [-0.263]	0.039 (.003) [0.629]	0.010 (.004) [0.169]
Mean Logarithmic Deviation	0.089	0.011 (.002) [0.121]	0.027 (.002) [0.300]	-0.018 (.002) [-0.207]	0.030 (.005) [0.338]	0.040 (.006) [0.448]
Poverty Rate	0.028	0.004 (.002) [0.137]	0.018 (.002) [0.646]	-0.022 (.001) [-0.773]	0.040 (.003) [1.43]	-0.0132 (.004) [-0.439]

[a] Numbers in parentheses show bootstrapped standard errors for the estimated factor contributions. Numbers in brackets show the share of the explained change in the total change.
[b] See Section III in the text for a complete list of the family characteristic variables.
[c] Defined as the standard deviation divided by the mean.
[d] Ratio of the 90th and 10th percentiles of the distribution of real family equivalent income. The other percentile ratio measures are defined similarly.

Source: Authors' calculations using 1970 and 1990 March CPS data.

Table 5: Primary-Order Decomposition of Changes in the Distribution of Family Equivalent Income and Poverty, 1989-1998

Statistic	Total Change	Effect of:[a] Men's earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	1,565	721	180	-543	1702	-495
		(107)	(49.4)	(61.9)	(107)	(175)
		[0.460]	[0.115]	[-0.347]	[1.09]	[-0.317]
Standard Deviation	1,852	1547	56.4	-118	1200	-833
		(122)	(39.2)	(45.4)	(83.5)	(220)
		[0.835]	[0.030]	[-0.064]	[0.648]	[-0.450]
Coefficient of Variation[c]	0.004	0.017	-0.003	0.010	-0.008	-0.012
		(.003)	(.001)	(.001)	(.002)	(.006)
		[4.45]	[-0.821]	[2.56]	[-2.01]	[-3.19]
90-10[d]	0.496	0.275	-0.250	0.488	-0.060	0.044
		(.094)	(.058)	(.067)	(.081)	(.175)
		[0.554]	[-0.504]	[0.982]	[-0.121]	[0.088]
50-10	0.121	0.080	-0.095	0.179	-0.012	0.041
		(.035)	(.025)	(.028)	(.033)	(.072)
		[0.065]	[-0.785]	[1.481]	[-0.100]	[0.339]
90-50	0.058	0.066	-0.010	0.023	-0.009	-0.013
		(.011)	(.004)	(.006)	(.010)	(.021)
		[1.141]	[-0.165]	[0.391]	[-0.150]	[-0.217]
75-25	0.155	0.066	-0.028	0.067	-0.037	0.087
		(.015)	(.009)	(.012)	(.015)	(.032)
		[0.428]	[-0.179]	[0.431]	[-0.239]	[0.559]
95-5	1.91	1.074	-1.382	1.750	-0.369	0.838
		(.366)	(.274)	(.269)	(.323)	(.649)
		[0.562]	[-0.723]	[0.916]	[-0.193]	[0.438]
Gini Coefficient	0.008	0.007	-0.002	0.005	-0.003	0.001
		(.001)	(.000)	(.001)	(.001)	(.002)
		[0.877]	[-0.247]	[0.621]	[-0.322]	[0.071]
Theil's Coefficient	0.01	0.009	-0.003	0.007	-0.004	0.000
		(.002)	(.001)	(.001)	(.001)	(.003)
		[0.981]	[-0.294]	[0.737]	[-0.387]	[0.037]
Mean Logarithmic Deviation	0.009	0.007	-0.004	0.008	-0.001	0.000
		(.002)	(.001)	(.001)	(.001)	(.005)
		[0.725]	[-0.456]	[0.875]	[-0.164]	[0.021]
Poverty Rate	0.002	-0.003	-0.004	0.010	-0.012	0.012
		(.001)	(.001)	(.001)	(.001)	(.003)
		[-1.65]	[-2.09]	[5.03]	[-6.25]	[5.96]

[a] Numbers in parentheses show bootstrapped standard errors for the estimated factor contributions. Numbers in brackets show the share of the explained change in the total change.
[b] See Section III in the text for a complete list of the family characteristic variables.
[c] Defined as the standard deviation divided by the mean.
[d] Ratio of the 90^{th} and 10^{th} percentiles of the distribution of real family equivalent income. The other percentile ratio measures are defined similarly.

Source: Authors' calculations using 1990 and 1999 March CPS data.

Appendix A — Data Changes and Tabulations Over Time

As described in the text (Section I), we use data from the March Annual Demographic Supplements to the Current Population Survey (CPS) for the years 1968-1999 (income years 1967-1998). This survey is conducted by the U.S. Bureau of the Census for the U.S. Bureau of Labor Statistics (BLS). In this Appendix, we describe changing treatment of high income values over time in the CPS, and we display the exact time-series tabulations underlying Figures 1, 3, and 4 in the text.

To preserve respondent confidentiality, the Census Bureau truncates recorded income values at an upper limit ("topcode"). For the results reported in this paper, we did not adjust total family income for changes in the topcode prior to the 1996 survey, because past work and our examination of the data suggested that year-to-year changes and the trend in inequality during this period are largely unaffected by changes in the nominal topcodes. However, beginning with the 1996 survey (income year 1995), the Census Bureau recorded values for several topcoded variables at the group means of the actual topcoded incomes rather than at the topcode itself. For consistency with previous years of data, we recoded these variables to equal the topcode value and adjusted total family income accordingly in income years 1995-1998. If this adjustment is not applied, inequality measures that are sensitive to outliers, such as the coefficient of variation and the Gini and Theil indices, indicate a substantial increase in inequality between income years 1994 and 1995 that is maintained in later years as well. Appendix Table A1 lists the exact tabulations for the yearly median and inequality series (which are displayed in Figure 1 in the text).

The change in topcoding procedure in the 1996 survey (income year 1995) also was applied to yearly earnings; we use these data for men in our sample. This change only affects the top few percentiles of the distribution of men's earnings, but it raises the coefficient of variation and the Gini and Theil measures substantially. As previous authors have done, we topcoded men's earnings in each year at a fixed percentile (the 97.5 percentile) in order to smooth out the impact of changing topcodes on the dispersion measures (see for example Burtless 1999). The yearly median and inequality series for men's earnings are displayed in Figure 3 in the text, and the exact underlying values are listed in Appendix Table A2. Our data on men's earnings includes individuals whose yearly earnings are zero. As a result, we did not calculate percentile dispersion measures for these data, because these measures are not uniformly defined (the 10th percentile generally is zero).

The change in the topcode procedure for income year 1995, along with our approach of imposing a uniform percentile topcode, may have held down our estimated growth in male earnings inequality in the last few years of our sample frame (1967-1998). However, the decline in male earnings inequality discussed in the text began in income year 1994, prior to use of the new topcode procedure. In addition, measures of dispersion that are relatively insensitive to the topcode, such as the 75-25 ratio, also exhibit a decline between 1993 and 1998.

Our analysis also focuses on measures of female labor force participation and family structure. Figure 4 provides a time-series depiction of the variables that we use, and the exact underlying are listed in Appendix Table A3.

Appendix Table A1 – Summary Statistics for the
Distribution of Family Equivalent Income, 1967-1998

Income Year	Median	Coefficient of Variation	Gini	Theil	90-10 Ratio	50-10 Ratio	90-50 Ratio	75-25 Ratio	% in Poverty
1967	17,968	0.691	0.341	0.201	5.55	2.73	2.04	2.23	0.114
1968	18,966	0.688	0.338	0.199	5.36	2.67	2.01	2.18	0.104
1969	20,012	0.665	0.336	0.193	5.37	2.67	2.01	2.20	0.098
1970	20,076	0.673	0.341	0.199	5.59	2.75	2.03	2.24	0.103
1971	20,166	0.696	0.345	0.205	5.73	2.81	2.04	2.30	0.107
1972	21,332	0.700	0.347	0.207	6.00	2.93	2.05	2.32	0.103
1973	22,109	0.664	0.341	0.197	5.91	2.92	2.02	2.28	0.099
1974	21,607	0.644	0.340	0.194	5.98	2.96	2.02	2.30	0.106
1975	21,494	0.676	0.348	0.205	6.23	3.07	2.03	2.37	0.114
1976	22,209	0.674	0.346	0.203	6.13	3.06	2.00	2.37	0.108
1977	22,485	0.678	0.350	0.207	6.32	3.10	2.04	2.42	0.109
1978	23,283	0.667	0.347	0.203	6.36	3.14	2.03	2.40	0.105
1979	23,770	0.663	0.347	0.203	6.43	3.18	2.02	2.41	0.108
1980	23,024	0.662	0.350	0.206	6.79	3.32	2.05	2.49	0.122
1981	22,727	0.699	0.364	0.224	7.14	3.42	2.09	2.57	0.132
1982	22,467	0.724	0.375	0.238	8.00	3.71	2.16	2.64	0.144
1983	22,396	0.730	0.380	0.244	8.26	3.78	2.19	2.75	0.149
1984	23,121	0.764	0.386	0.254	8.31	3.75	2.22	2.74	0.142
1985	23,524	0.749	0.385	0.251	8.42	3.78	2.23	2.75	0.137
1986	24,273	0.753	0.386	0.252	8.49	3.83	2.22	2.75	0.133
1987	24,978	0.757	0.386	0.253	8.61	3.90	2.21	2.76	0.131
1988	24,793	0.755	0.387	0.254	8.52	3.79	2.25	2.77	0.129
1989	25,260	0.761	0.388	0.255	8.47	3.77	2.24	2.76	0.126
1990	24,347	0.751	0.387	0.253	8.70	3.86	2.25	2.78	0.133
1991	24,011	0.752	0.390	0.256	9.10	4.03	2.26	2.83	0.140
1992	23,908	0.754	0.392	0.259	9.30	4.08	2.28	2.88	0.145
1993	23,480	0.779	0.401	0.271	9.70	4.14	2.34	3.00	0.151
1994	24,025	0.779	0.400	0.270	9.52	4.10	2.32	2.97	0.147
1995	24,390	0.792	0.403	0.275	9.06	3.85	2.35	2.94	0.139
1996	24,959	0.795	0.404	0.276	9.34	3.96	2.36	2.93	0.136
1997	25,624	0.800	0.405	0.279	9.28	3.91	2.37	2.92	0.132
1998	26,842	0.792	0.403	0.276	9.11	3.87	2.35	2.92	0.128

Source: Authors' tabulations of March CPS data (weighted); figures in 1998 dollars.

Appendix Table A2 – Summary Statistics for the
Distribution of Male Wage and Salary Earnings, 1967-1998

Income Year	Median	Coefficient of Variation	Gini	Theil
1967	26,010	0.674	0.378	0.279
1968	28,345	0.670	0.375	0.276
1969	29,212	0.674	0.378	0.277
1970	28,865	0.694	0.387	0.288
1971	28,534	0.709	0.397	0.301
1972	29,830	0.717	0.400	0.305
1973	30,630	0.704	0.395	0.299
1974	29,489	0.719	0.403	0.309
1975	28,145	0.742	0.417	0.330
1976	28,871	0.739	0.416	0.329
1977	29,221	0.746	0.419	0.333
1978	29,336	0.752	0.420	0.335
1979	29,908	0.760	0.424	0.340
1980	28,825	0.776	0.433	0.353
1981	27,414	0.783	0.438	0.361
1982	26,522	0.824	0.458	0.390
1983	25,592	0.833	0.463	0.398
1984	27,618	0.817	0.456	0.386
1985	27,493	0.816	0.454	0.380
1986	28,114	0.829	0.458	0.386
1987	27,840	0.829	0.459	0.388
1988	27,265	0.833	0.459	0.388
1989	26,851	0.837	0.460	0.388
1990	26,406	0.840	0.463	0.392
1991	25,608	0.853	0.469	0.402
1992	24,794	0.879	0.481	0.421
1993	24,701	0.903	0.488	0.433
1994	25,723	0.888	0.481	0.419
1995	26,194	0.868	0.470	0.398
1996	25,878	0.862	0.467	0.393
1997	27,215	0.884	0.476	0.411
1998	28,000	0.875	0.472	0.405

Source: Authors' tabulations of March CPS data (weighted); figures in 1998 dollars.

Appendix Table A3 – Means for Female Labor Force Participation and
Family Structure Variables, 1967-1998

Income Year	Female Participation Rate	Married, no Kids	Married, with Kid(s)	Never Married, with Kid(s)	Other Single with Kid(s)	Single, no Kids
1967	0.554	0.177	0.670	0.010	0.074	0.069
1968	0.566	0.178	0.667	0.010	0.075	0.071
1969	0.570	0.178	0.666	0.010	0.073	0.073
1970	0.572	0.182	0.648	0.014	0.080	0.077
1971	0.566	0.186	0.638	0.012	0.084	0.080
1972	0.568	0.191	0.628	0.013	0.085	0.084
1973	0.589	0.193	0.616	0.014	0.088	0.090
1974	0.592	0.196	0.601	0.015	0.093	0.095
1975	0.602	0.199	0.595	0.015	0.092	0.099
1976	0.622	0.203	0.582	0.018	0.093	0.104
1977	0.637	0.205	0.566	0.022	0.094	0.113
1978	0.656	0.206	0.558	0.023	0.094	0.119
1979	0.677	0.206	0.549	0.024	0.097	0.124
1980	0.678	0.206	0.539	0.026	0.099	0.131
1981	0.677	0.214	0.524	0.029	0.099	0.134
1982	0.672	0.220	0.518	0.031	0.095	0.135
1983	0.679	0.221	0.511	0.031	0.095	0.141
1984	0.699	0.221	0.501	0.033	0.098	0.146
1985	0.706	0.217	0.501	0.034	0.097	0.151
1986	0.717	0.219	0.497	0.038	0.096	0.152
1987	0.727	0.219	0.489	0.040	0.095	0.158
1988	0.739	0.217	0.487	0.042	0.092	0.163
1989	0.742	0.219	0.481	0.042	0.095	0.163
1990	0.745	0.216	0.477	0.047	0.095	0.166
1991	0.745	0.215	0.473	0.049	0.096	0.168
1992	0.743	0.215	0.471	0.051	0.096	0.167
1993	0.747	0.207	0.473	0.054	0.098	0.168
1994	0.755	0.209	0.471	0.052	0.101	0.167
1995	0.758	0.207	0.464	0.056	0.101	0.172
1996	0.763	0.206	0.461	0.059	0.097	0.177
1997	0.767	0.206	0.460	0.060	0.094	0.180
1998	0.767	0.207	0.457	0.059	0.094	0.184

Source: Authors’ tabulations of March CPS data (weighted).

Appendix B — Derivation of the Conditioning Weights

This appendix provides the derivation of the conditioning weights $\psi_{L|S,X}$, $\psi_{S|X}$, and ψ_X (and the corresponding reverse-order weights), described heuristically in Section II. This discussion largely follows that in DiNardo, Fortin, and Lemieux (1996; DFL), with modification to our specific setting.

Our formal derivation picks up from equation (3) (Section II) in the text, generalizing to an arbitrary year "t"and excluding the term representing the distribution of men's earnings (m_{89}). The following equation expresses the distribution of family income Y in year t, conditional on female labor force participation L, family structure S, and other family and individual characteristics X:

$$f_t(Y) \equiv f(Y;\ t_Y = t,\ t_{L|S,X} = t,\ t_{S|X} = t,\ t_X = t) \tag{B1}$$

A distribution such as (B1) can be expressed as:

$$f_t(Y) = \int\int\int f(Y|L,S,X,t_Y = t)\, dF(L|S,X,t_{L|S,X} = t)\, dF(S|X,t_{S|X}=t)\, dF(X|t_X=t) \tag{B2}$$

In this equation, $f_t(Y)$ is the density of Y and can be seen as the integral of the density of income Y conditional on a set of individual and family attributes and on a date t, f(Y | L, S, X, t_y=t), integrated over the distribution of individual attributes at date t, F(L, S, X; t).

We are interested (for example) in the distribution of Y in 1989 if the distribution of L conditional on S and X is held to its 1969 structure:

$$f(Y; t_Y = 89,\ t_{L|S,X} = 69,\ t_{S|X} = 89,\ t_X = 89) \tag{B3}$$

Using (B2), this distribution can be expressed as:

$$\begin{aligned} f_t(Y; t_Y{=}89, t_{L|S,X}{=}69, t_{S|X}{=}89, t_X{=}89) &= \int\int\int f(Y|L,S,X,t_Y{=}89)\, dF(L|S,X,t_{L|S,X}{=}69)\cdot \\ &\qquad dF(S|X,t_{S|X}{=}89)\, dF(X|t_X{=}89) \\ &= \int\int\int f(Y|L,S,X,t_Y{=}89)\, \psi_{L|S,X}\, dF(L|S,X,t_{L|S,X}{=}89)\cdot \\ &\qquad dF(S|X,t_{S|X}{=}89)\, dF(X|t_X{=}89) \end{aligned} \tag{B4}$$

where $\psi_{L|S,X}$ is a reweighting function to be defined momentarily. Note that except for $\psi_{L|S,X}$, the bottom line of (B4) is identical to (B2) with t=89—i.e., the distribution we want to estimate is equal to the unconditional distribution of income in 1989 with observations reweighted by the function $\psi_{L|S,X}$. If we can estimate $\psi_{L|S,X}$, it is straightforward to incorporate it and obtain the counterfactual distribution expressed in (B4) by using the observed univariate, unconditional distribution of income in 1989.

The reweighting function for female labor force participation is defined (identically) as:

$$\begin{aligned} \psi_{L|S,X}(L,S,X) &\equiv \frac{dF(L\,|\,S,X,t_{L|S,X}=69)}{dF(L\,|\,S,X,t_{L|S,X}=89)} \\ &= L\cdot\left(\frac{Pr(L=1\,|\,S,X,t_{L|S,X}=69)}{Pr(L=1\,|\,S,X,t_{L|S,X}=89)}\right) + (1-L)\cdot\left(\frac{Pr(L=0\,|\,S,X,t_{L|S,X}=69)}{Pr(L=0\,|\,S,X,t_{L|S,X}=89)}\right) \end{aligned} \tag{B5}$$

The first line identity in (B5) is obtained by substituting the expression on the right side into (B4) and canceling-out the denominator. Regarding the second equality in (B5), note first that L only takes the values 0 or 1, so that:

$$dF(L|S,X,t_{L|S,X}=t) \equiv L\cdot Pr(L=1|S,X,t_{L|S,X}=t) + (1-L)\cdot Pr(L=0|S,X,t_{L|S,X}=t) \tag{B6}$$

The second equality in (B5) follows from the recognition that one term on the right-hand side of (B6) will always equal zero.

The weight $\psi_{L|S,X}$ represents the change in the probability between 1969 and 1989 that a family defined by characteristics (S,X) is observed to have a female head who works. The probabilities in (B5) are easily recognized as expressions from standard binary dependent variable models. These conditional probabilities can be obtained by estimating a model such as a probit or logit and then using the fitted values. We used the logit equation:[28]

$$\begin{aligned} Pr(L=1|S,X,t_{L|S,X}=t) &= pr(\mu > -H(S,X)\beta) = 1 - G(-H(S,X)\beta) \\ &= \frac{\exp(-H(S,X)\beta)}{1 + \exp(-H(S,X)\beta)} \end{aligned} \qquad \text{(B7)}$$

Using data from year t, this estimated equation provides the structure of (L|S,X) in year t, where the cumulative distribution of μ is a logistic function denoted by G. In (B7), H(X) is a vector function of X designed to capture the conditional relationship being modeled, and β is a vector of estimated coefficients (in the simplest case, H is purely linear). The regressions are weighted using the March supplement (individual) weights attached to female heads in the sample. This equation is estimated for both the 1969 and 1989 samples, and the coefficients are retained. We used the results to fit the probabilities in (B5) using the values of (S,X) from the 1989 sample, combined with the 1969 coefficients for the numerator and the 1989 coefficients for the

[1] DFL used probit equations. We use logits, for consistency with our multinomial logit estimation of the family structure conditioning model, and because the underlying distribution function for the logit model has a closed-form representation that may be useful in other settings. Use of probits rather than logits for our labor supply and characteristics (X) conditioning steps does not alter the results noticeably.

denominator. For families in which no female head is present, the reweighting function $\psi_{L|S,X}$ is set to 1.

Formation of the family structure weight is similar but requires extension to the case of multiple outcome categories. In this case, we are interested in the distribution of family income if female labor force participation and family structure are both held to the levels and relationship with X prevailing in 1969 :

$$\begin{aligned} f_t(Y;t_Y{=}89,t_{L|S,X}{=}69,t_{S|X}{=}69,t_X{=}89) &= \int\int\int f(Y|L,S,X,t_Y{=}89)\,dF(L|S,X,t_{L|S,X}{=}69)\cdot \\ &\quad dF(S|X,t_{S|X}{=}69)\,dF(X|t_X{=}89) \\ &= \int\int\int f(Y|L,S,X,t_Y{=}89)dF(L|S,X,t_{L|S,X}{=}89)\,\psi_{L|S,X}\cdot\psi_{S|X}\cdot \\ &\quad dF(S|X,t_{S|X}{=}89)\,dF(X|t_X{=}89) \end{aligned} \tag{B8}$$

The second line of this equation is identical to (B4) except for inclusion of the conditioning weight $\psi_{S|X.}$ Thus, to estimate the distributional impact of holding family structure to its 1969 levels and relationship with X, we can use the 1989 distribution adjusted for female labor supply, with a further reweighting specified by the family structure weight, $\psi_{S|X}$.

To derive the appropriate weight, assume that family structure S is characterized by C mutually exclusive, exhaustive categories, so that we can represent S as a single variable with C possible values (in our specific application, C=5). Then the family structure conditioning weight, $\psi_{S|X}$, is defined as:

$$\begin{aligned} \psi_{S|X}(S,X) &\equiv \frac{dF(S\,|\,X,t_{S|X}=69)}{dF(S\,|\,X,t_{S|X}=89)} \\ &= \sum_{c=1}^{C} I_c\cdot\frac{Pr(S{=}c\,|\,X,t_{S|X}=69)}{Pr(S{=}c\,|\,X,t_{S|X}=89)} \end{aligned} \tag{B9}$$

In this expression, I_c is an indicator variable that takes on the value 1 if S=c and 0 otherwise. Note that the second line of (B9) is a generalization of the second line of (B5), with S|X replacing L|S,X; in (B5), C is set equal to two. For each observation in the data, I_c=1 for only a single value of c, so the second line of (B9) is generated by the same property as illustrated in (B6). The expression to the right of the summation sign in the bottom line of (B8) is the relative conditional probability that a family with characteristics X falls into family type c in 1969 and 1989; each family in the 1989 data is upweighted or downweighted depending on whether it was more or less likely to be observed as such in 1969 than in 1989. These conditional probabilities can be estimated using a model designed to handle unordered polychotomous dependent variables. We use the multinomial logit model for these estimates, weighted by the family weight.

The final conditioning weight adjusts for the change in the underlying distribution of the family characteristics X. Using a derivation similar to that for the labor supply and family structure weights, this weight is defined as:

$$\begin{aligned} \psi_X(X) &\equiv \frac{dF(X|t_X=69)}{dF(X|t_X=89)} \\ &= \frac{Pr(t_X=89)}{Pr(t_X=69)} \cdot \frac{Pr(t_X=69|X)}{Pr(t_X=89|X)} \end{aligned} \tag{B10}$$

The second equality is derived through a simple rearrangement of the conditional probabilities based on Bayes' Law. This weighting function represents the relative probability of observing a family with characteristics X in the 1969 versus the 1989 sample, normalized by the unconditional probabilities of being in either sample.

The function ψ_X is estimated by pooling the 1969 and 1989 data sets, and then estimating a binary dependent variable model for a dummy variable indicating the sample from which the observation is obtained. The conditional probabilities $Pr(t_X=t|X)$ are obtained by forming fitted probabilities for families in the 1989 sample, based on their X values. As with the labor force participation weights, we estimate logit equations to form the conditional probabilities. The unconditional probabilities, $Pr(t_X=t)$, are the weighted shares of the 1969 and 1989 samples in the pooled sample.

As described in Section II in the paper, we combined the estimated weights in sequence to produce the decomposition results, after adjusting the data for changes in the distribution of male earnings (see Appendix C). Due to the possibility of general equilibrium or endogenous relationships between male earnings, family structure, and female labor force participation, we also reversed the order of the decomposition. The representation of the underlying general distribution in reverse conditioning sequence is:

$$f_t(Y) = \int\int\int f(Y|X,S,L,t_Y=t)\,dF(X|S,L,t_{X|S,L}=t)\,dF(S|L,t_{S|L}=t)\,dF(L|t_L=t) \tag{B11}$$

We consider the associated conditioning weights from right to left in this equation. First, using steps similar to those used to derive ψ_X in (B8) above, we have:

$$\begin{aligned}\psi_L(L) &\equiv \frac{dF(L|t_L=69)}{dF(L|t_L=89)}\\ &= \frac{Pr(t_L=89)}{Pr(t_L=69)}\cdot\frac{Pr(t_L=69|L)}{Pr(t_L=89|L)}\\ &= \frac{Pr(t_L=69|L)/Pr(t_L=69)}{Pr(t_L=89|L)/Pr(t_L=89)}\end{aligned} \tag{B12}$$

Then the estimated weight $\psi_L(L)$ is a simple function of the female labor force participation rate (estimated using the individual sampling weights) in both years. In particular, observations in the 1989 data in which the female head participates are downweighted by the proportional increase between 1969 and 1989 in the percentage of female heads who participate; observations in the 1989 data in which the female head does not participate are upweighted by the proportional decrease in the percentage of female heads who do not participate. For families in which no female head is present, the reweighting function ψ_L is set to 1.

In the reverse-order decomposition, the family structure weight is conditioned on female participation. Using a derivation similar to (B9), we have:

$$\begin{aligned}\psi_{S|L}(S,L) &\equiv \frac{dF(S|L, t_{S|L}=69)}{dF(S|L, t_{S|L}=89)} \\ &= \sum_{c=1}^{C} I_c \cdot \frac{Pr(S{=}c|L, t_{S|L}=69)}{Pr(S{=}c|L, t_{S|L}=89)}\end{aligned} \tag{B13}$$

The probabilities in (B13) can be estimated through simple cross-tabulation of the family structure and female labor force participation outcomes. For both years, we calculated the percentage of observations that fall into each of the 10 categories defined by the five family structure categories and two possible values (0 and 1) for the labor force participation variable. Observations falling into a particular cell in the 1989 data are upweighted or downweighted by the proportional change in the percentage share of that cell between 1969 and 1989. For consistency between the primary-order and reverse-order decompositions in the treatment of observations, we included families in which no female head is present with families in which the female head did not participate.

Finally, we estimated the X weight, $\psi_{X|S,L}(X,S,L)$, using our prior estimates of the primary-order weights and the two reverse-order weights described above. In particular, because $F(X,S,L)=F(X|S,L)\cdot F(S|L)\cdot F(L)=F(L|S,X)\cdot F(S|X)\cdot F(X)$, the product of the complete set of reverse-order weights is equal to the product of the complete set of primary-order weights, and we can rearrange to obtain $\psi_{X|S,L}$:

$$\psi_{X|S,L}\cdot\psi_{S|L}\cdot\psi_{L}=\psi_{L|S,X}\cdot\psi_{S|X}\cdot\psi_{X} \quad \Rightarrow$$
$$\psi_{X|S,L}=\frac{\psi_{L|S,X}\cdot\psi_{S|X}\cdot\psi_{X}}{\psi_{S|L}\cdot\psi_{L}} \tag{B14}$$

We used this equality to estimate $\psi_{X|S,L}$ for our sample. One further implication of this equality is equivalence of the net effect of the three conditioning factors in the primary-order and reverse-order cases. However, inclusion of the male earnings step in our exact setting produces slight differences across the two cases in the net impact of the explanatory factors.

Appendix C— Adjusting for the Distribution of Men's Earnings

In this appendix, we provide a detailed description of our adjustment for the changing distribution of men's earnings. This procedure is described heuristically in the text. We specify the adjustment more precisely here, using 1969 and 1989 as our example years. The basic idea is to adjust 1989 family income by subtracting out each male head's 1989 yearly earnings and adding back the yearly earnings from the median of the 1969 percentile group (quantile) that is at the same rank as the 1989 quantile to which that male belonged.

Let M_j^t denote (wage and salary) earnings of male family head j in year t. Assume that the observations in each year of data are rank ordered from lowest to highest earnings for men. We divide the men's earnings data into Q equally sized quantiles, with the size of each quantile defined by the sum of the individual sampling weights for males in the quantile.[1] Let $M_{j(q)}^t$ denote the median earnings in the quantile (q) to which male j belonged in year t. To implement the adjustment, we calculated median earnings by quantile for the observation year and for an earlier year that serves as the men's earnings counterfactual.

Consider adjustment of the 1989 distribution of total family income (T^{89}) for families with male heads present based on substitution of the quantile median from the 1989 distribution of men's earnings with the appropriate quantile median from the 1969 distribution of men's

[1] We sorted the earnings observations by descending values of the individual weight. This is equivalent to basing the procedure on writing a new data set with the sampling weights treated as frequency weights, so that each observation on men's earnings is expanded to the number of observations represented by its sampling weight. The quantile cut points exhibit minor variation depending on how the weights are treated in the sorting process.

earnings.[2] In this case, our adjusted measure of total family income (with individual identifier j suppressed) is:

$$T_m^{89} = T^{89} - M_q^{89} + M_q^{69} \qquad \text{(C1)}$$

For families in which no male head is present, no adjustment is applied. For the application in this paper, we set the number of quantiles Q to 100, which results in a distribution of earnings with midpoint and dispersion measures that are virtually identical to the original distribution. We verified that we obtain nearly identical results when we set Q equal to 500.

Applying the same equivalence adjustment as applied to unadjusted total family income (equation 1 in the text) produces our measure of equivalent family income adjusted for the changing dispersion of men's earnings:

$$Y_m = \frac{T_m}{F^{\sigma}} \qquad \text{(C2)}$$

where F=family size and σ=0.5.

[2]We use the quantile median from the 1989 male earnings distribution, rather than the actual 1989 value of male earnings, to avoid an artificial reduction in family income inequality that might occur because inequality calculated at quantile midpoints is smaller than in the full distribution. We thank an anonymous referee for suggesting this approach.