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

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Abstract

The trend toward increasing inequality in family income in the United States since the late 1960s is well documented. Among key possible explanations for this increase are rising dispersion in individual earnings, changes in female labor supply decisions, and changes in family composition and living arrangements. We analyze the contribution of these factors to changes in family income inequality and poverty during the years 1969-1998, focusing on labor supply and family structure as behavioral changes but accounting also for changes in the distribution of male earnings. Our analyses rely on conditionally weighted density estimation, a semiparametric decomposition technique recently developed by DiNardo, Fortin, and Lemieux (1996). We also use a relatively novel rank-based distributional exchange to assess the effects of changes in the distribution of male earnings.

In our empirical work, we first analyze changes between 1969 and 1989, which corresponds roughly to the period of rising inequality that has been the focus of previous work. Our results indicate that rising dispersion of male earnings and the decline of traditional forms of family structure respectively explain up to about three-fourths and about one-half of rising inequality in family income during this period. The impact of changing family structure was most pronounced in the lower half of the distribution. In contrast, the increase in female labor force participation offset rising inequality to some degree, mainly in the upper half of the distribution, although its impact has moved down the distribution over time. In extending the analyses to the 1990s, we find that the rate at which inequality grew slowed after 1989, but the explanatory factors continued to have substantial effects. In each decade, the effects of the explanatory factors on poverty were especially large and followed a pattern similar to that for inequality.

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I. Introduction

During the past three decades, inequality in family income has risen substantially in the United States. A commonly cited official measure of inequality in household and family income, the Gini coefficient tabulated by the U.S. Census Bureau, exhibited a relatively consistent upward trend between the late 1960s and at least 1994 (Weinberg 1996). 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).

Rising inequality in family income and rising poverty have been accompanied by a corresponding increase in the dispersion of male earnings. The general consensus among researchers is that inequality in male earnings began to rise in the early to mid-1970s, with an acceleration evident in the 1980s (Karoly 1993, Levy and Murnane 1992). This increase appears to have played a key role in rising inequality of family income. For example, Karoly and Burtless (1995) found that 40 percent of the total increase in family income inequality between 1969 and 1989 can be attributed to changes in the dispersion of individual earnings. Burtless (1999) found a similar result for the years 1979-96, with virtually all of the increase attributable to rising dispersion of male earnings and very little to rising dispersion of female earnings.

Perhaps due in part to changes in the distribution of male earnings, family composition and living arrangements and the extent of female labor force participation also have changed. For example, between 1969 and 1989, the proportion of persons living in single-head families with children and the proportion of individuals living in single-person families rose substantially (Karoly and Burtless 1995). During the same period, the labor force participation rate of married women climbed steadily, thereby increasing the percentage of dual-earner families. This increase in labor force participation has been relatively uniform across all racial and ethnic groups (Cancian et al. 1993), within fertility classes, and across low-wage and high-wage households (Juhn and Murphy 1997).

It is likely that these changes in family structure and labor supply have contributed to the increase in the dispersion of family income. Income tends to be low in single-parent families, so growth in this family type is likely to increase the prevalence of low-income families. Existing work has found evidence consistent with the claim that changes in family structure have contributed to rising inequality in family income and poverty (Blank and Card 1993, Burtless 1999, Bradbury 1996, Gottschalk and Danziger 1993, Karoly and Burtless 1995, Lerman 1996).

Similarly, depending on the income class of families in which women's labor force participation has risen, rising female participation may increase the income gaps between high-income and low-income families. Evidence on changing female labor supply suggests that increases in labor force participation rates and growth in wages have been largest for women whose husbands have relatively high earnings, which will tend to increase inequality in family income (Blackburn and Bloom 1995, Cancian, Danziger, and Gottschalk 1993, Juhn and Murphy 1997). However, investigations of the effect that increased earnings and labor supply by wives have had on family income inequality have produced mixed results. Moreover, the methodologies used in previous work generally have not accounted for related covariates, such as demographic variables, that may affect family structure, labor force participation, and family income.

In this paper, we assess the effects of the changing distribution of male earnings, rising female labor force participation, and changes in family structure and living arrangements on family income inequality and poverty. We use data from the March Annual Demographic Supplement to the Current Population Surveys (March CPS) for the years 1968-1999, focusing primarily on the business cycle peak or near-peak years of 1969, 1979, 1989, and 1998. In addition to a reexamination of the causes of 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 the 1980s extended into the 1990s.

The key contribution of this paper, however, is its application of a relatively new methodology—the conditional density estimation technique recently developed by DiNardo, Fortin, and Lemieux (1996). This technique enables estimation of the conditional contribution of a set of factors on the entire distribution of an outcome variable. In applying this technique, we improve on the existing literature in two ways. First, we produce counterfactual income distributions that incorporate changes in the conditional relationship between the explanatory factors and a set of related individual and family characteristics; our results therefore have greater behavioral content than results from previous analyses. Second, we estimate the effects of the modeled changes on the entire distribution of family income rather than a narrow set of summary measures. This technique is especially useful for estimation of specific features of a distribution, such as dispersion in the lower half or cumulative percentage cutoffs—for example, the poverty rate. We extend the original DFL technique by applying it to a conditioning factor (family structure) that takes on more than two outcome categories. Moreover, we use a separate technique, based on a relatively novel rank-based distributional exchange, to account for the changing distribution of male earnings.

The paper proceeds by first describing our data source and variable definitions and presenting tabulations that illustrate basic trends in the data. We describe the decomposition methodology used in Section III and present results in Section IV. In Section V, we summarize the results and discuss implications for future research.

II. Data and Trends

We use data from the March CPS for the years 1968-1999, as administered by the U.S. Census Bureau for the U.S. Bureau of Labor Statistics. These files provide information on income and related variables for the calendar year prior to the survey date; thus, our analyses apply to income years 1967-1998. In the formal decomposition analysis, we focus on the years 1969, 1979, 1989, and 1998. These years largely span our sample frame, and they 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.

Given our focus on family structure and living arrangements, our definition of families is key for the analysis. In most cases, households contain one family. For multi-family households, we treat sub-families as separate families for purposes of identifying family structure (although we pool income for our income measure for multi-family households, as discussed below). 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 live with their parents. To reduce the impact of undercounting this key family type in the early portion of our sample, we applied her correction to the pre-1984 data. We also restricted the sample to families in which at least one head is between the ages of 16 and 64. We excluded families with elderly heads because the determinants and dynamics of family structure, labor supply, and family income are different for the elderly than they are for families with working-age adults. As a matter of terminology, we refer to the husband and wife in married couples as the male head and female head, respectively; families headed by a single adult have either a male head or a female head.

In all tabulations and analyses of family income, we set our basic sampling weight equal to the sum of the individual weights for all persons in the family unit. Thus, although our analyses are conducted at the family level, the results should be interpreted as characterizing the experience of individuals who constitute the associated population.

Measurement of income is key for our analysis. We began with the survey information on total pre-tax, post-transfer income for the family unit.[1] Because we are interested in the distribution of economic well-being, we analyze the distribution of total family income divided by a function of family size, using the definition of families given above for most cases. For

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[1] 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 the upper tail of the distribution, 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.

households that contain multiple families related by blood or marriage (including multi-generational families), we treat sub-families as separate families for the analysis of family structure, but we pool the income of all primary family and sub-family members to form total family income, under the assumption that related families sharing living quarters share income in the same manner as nuclear family units.[2]

In general, the well being of family members depends on income per member. However, given a particular level of total family income, well being per member may not decline by the same amount for each additional family member added, due to economies of scale in consumption. Letting T denote our measure of total family income and F denote family size, our measure of family equivalent income is:

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

We set σ=0.5 in the present work. 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 related families that share living quarters, for which income and family size are totaled across the sub-families.[4]

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[2] We treat unrelated sub-families that share living quarters as separate families for the calculation of family structure and total family income.

[3] The boundary values of θ—0 and 1—imply, respectively, infinite and no economies of scale in family consumption.

[4] Our equivalent income measure for related families that share living quarters therefore is constructed under the assumption that income and consumption are shared *across* sub-families in these households in the same way that they are shared *within* other families in our sample.

We also analyze yearly wage and salary earnings for males as a separate component of our decomposition (weighted by the individual March supplement weight). All of our income measures are expressed in 1998 dollars, using the GDP deflator for personal consumption expenditure as our adjustment factor.

A key feature of the distribution of family income is the percentage of families or individuals who fall below the income cutoff that represents the official poverty rate. The official poverty scales are based on the concept of subsistence income, and the official poverty rate is a useful indicator of levels of deprivation that are likely to engender policy remedies. We define families as being in poverty if their equivalent income falls below the relevant official federal poverty threshold.

A precise definition of family structure also is important for our analysis. The trend in family structure 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.

Our final key analysis variable is female labor force participation. We coded labor force participation as 1 if the respondent reported positive hours worked in the previous year, and 0 otherwise. This measure is relatively broad and produces higher participation rates than would a

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However, we still treat sub-families in multi-generational households as separate families for purposes of identifying changes in family structure.

standard point-in-time measure. We chose to use a participation variable that corresponds most closely to our yearly income measure. Moreover, our preliminary tabulations indicated that most of the trend increase in female labor force participation during the past several decades was associated with increases in full-time work.

Figures 1 through 5 illustrate the trends in our data for our measure of equivalent family income and for our key explanatory variables. For the income variables, we focus on standard dispersion measures—the coefficient of variation, the Gini and Theil coefficients, and several percentile dispersion ratios—and display the median as well.[5] The complete set of values that underlie these figures are listed in Appendix Tables 1-3.

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. All of the dispersion measures increased substantially between 1969 and 1998. 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 50-10 income ratio. These patterns are consistent with those found in other analyses of income inequality (see for example Karoly 1993). 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. In the 1980s and 1990s, the increases were discontinuous, with pronounced increases associated with recessionary periods (especially 1982-1983).

Figure 1 also indicates that income inequality leveled out after 1993. This finding is

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[5] The coefficient of variation is the standard deviation divided by the mean. The Gini and Theil measures are standard parametric indices of inequality used by a variety of previous authors (see Karoly 1992 for a useful discussion of alternative inequality measures). Our percentile dispersion measures are expressed as the ratio of income at the indicated percentiles of the distribution of family equivalent income.

consistent with the latest Census Bureau tabulations for family income unadjusted for family size (U.S. Bureau of the Census 1999). Between 1993 and 1998, the Gini index in our data changed only slightly, rising from 0.401 to 0.403, the 90-50 ratio remained virtually constant, and the 50-10 ratio fell from 4.14 to 3.87. Moreover, the median value of real equivalent family income rose by about 14 percent ($3400) during this expansionary period, after falling noticeably between 1989 and 1993. By comparison, the same measure grew 12.8 percent during the slightly longer measured expansionary period between 1983 and 1989. Similarly, the 50-10 ratio stayed virtually constant during the expansionary years of 1983 to 1989, and the 90-50 ratio rose only slightly, suggesting that the expansionary periods of the 1980s and 1990s produced gains in family income that were similarly distributed. Thus, the 1980s were an unusual decade not due to the uniqueness of the expansion but instead because the recession early in the decade had an extensive and durable effect on inequality in family income.

Figure 2 provides a visual representation of the distribution of real family equivalent income in 1969, 1989, and 1998.[6] The figure displays kernel density estimates of the distribution of family income for these years, weighted by our measure of the family sampling weight.[7] The

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[6] To preserve visual clarity, we do not display the distribution of income in 1979. As listed in Appendix Table 1, most of the increase in the dispersion of income between 1969 and 1989 occurred during the 1980s.

[7] 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 kernel

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 masks somewhat 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 with the changes between 1969 and 1989, the increase in dispersion between 1989 and 1998 was relatively modest, with apparent movement of the density mass from the lower and middle portions to the upper portion.

Figure 3 displays the poverty rate. Poverty exhibits a pronounced cyclical pattern, with comparable peaks associated with the early 1980s and early 1990s recessions. A slight upward trend since the late 1960s is evident as well, consistent with rising dispersion in the lower portion of the income distribution. Due to substantial income gains in the bottom of the distribution between 1993 and 1998, the poverty rate fell from 15.1 percent in 1993 (the historical peak in our data) to 12.8 percent in 1998. However, this latest figure is slightly higher than the 12.6 percent figure recorded at the end of the 1980s expansion and well above the levels from the late 1960s and 1970s.

Figure 4 displays trends in the distribution of yearly wage and salary earnings for men. Our measure includes individuals whose yearly earnings are zero. We display the median, the coefficient of variation, and the Gini and Theil coefficients; see Appendix Table 2 for a listing of

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density estimates themselves are based on the full range of data.

the complete time series for each.[8] The median has not yet returned to the peak achieved in 1973, although growth in recent years has 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.[9]

Figure 5 displays trends in our other key explanatory variables, female labor force participation and family structure (see Appendix Table 3 for a complete listing). Female labor force participation rose steadily during our sample frame, with the most rapid increases evident in the late 1970s. In regard to family structure, a long-run decline in the share of married couples with children is evident, along with pronounced percentage increases in the share of never married individuals with children and singles without children. For example, the share of families consisting of a never-married single head with children increased by nearly a factor of six between 1969 and 1998, rising from 0.010 to 0.059. Although this still represents a small share of all

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[8] The change in the topcode procedure beginning with the 1996 survey (see footnote 2) only affects the top few percentiles of this distribution, but it raises the coefficient of variation and the Gini and Theil measures substantially. As previous authors have done, we topcoded male 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). In addition, we do not display percentile dispersion measures in Figure 4, because they are not uniformly defined for our measure of yearly male earnings (the 10th percentile generally is zero).

[9] The change in the topcode procedure for income year 1995, along with our response of imposing a uniform percentile topcode, may have held down our estimated growth in male earnings inequality in the last few years. However, the decline in male earnings inequality 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.

families, the increase of about five percentage 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 mostly is in the range of 0.10 to 0.15. The net impact of these changes in family structure was to reduce average family size in our data from about 3.5 in 1969 to about 2.6 in 1998.

III. Decomposition Methodology

A. Background

The most common approach to the analysis of income inequality involves calculation of a standard set of dispersion measures. These include summary measures such as the coefficient of variation and the Gini and Theil indices, percentile dispersion measures, and breakdowns of income shares by population sub-groups. Most previous work that investigates the determinants of rising dispersion of family income is based on decompositions of family income into its components, rather than model-based approaches that account for changes in family behavioral outcomes and underlying characteristics.[10]

An appropriate model-based methodology for evaluating changes in the distribution of earnings and income recently was developed and applied by DiNardo, Fortin, and Lemieux (1996; henceforth DFL). Their general technique enables estimation of how a set of factors have

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[10] Among papers that examine the role of female earnings and labor supply, 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 family income inequality, and Blackburn and Bloom (1989) found little overall impact of female labor supply on inequality in family income. These differences in results are due in large part to differences in sample, time period, and investigation of female labor supply versus earnings. Moreover, Cancian and Reed (1995) report that the impact of wives' earnings on inequality depends on the specific counterfactual assumption that is made.

affected a distribution of interest, through estimation and application of conditioning weights. The estimated conditioning weights are combined with sample survey weights to produce an adjusted distribution. This is a flexible procedure that provides semiparametric estimates of the entire distribution of an outcome variable under counterfactual assumptions.

The conditional reweighting procedure is related to standard regression-based decompositions of variance. Regressions typically are used, however, to estimate the mean of a distribution. The advantage of the weighted density procedure is that it estimates the entire conditional distribution, as opposed to analyzing distributional characteristics one-by-one, and therefore provides a more flexible method than regression techniques for investigating distributional changes. DFL used their technique to estimate the amount by which inequality in individual earnings would have risen between 1979 and 1989 if the real minimum wage and union membership density in the U.S. had remained at their 1979 levels and structure. Modification and expansion of this technique is required for application in our setting, which we describe in the next section.[11]

B. Conditionally Weighted Density Estimation

In this section, we describe how simple estimated reweighting functions can be obtained and applied to the estimation of income distributions that embody counterfactual assumptions regarding conditional outcomes. We describe these procedures heuristically in the text, focusing on 1969 and 1989 as our comparison years. The exact derivation is provided in Appendix A.

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[11] Gottschalk and Danziger (1993) applied a related regression-based approach to estimate the contribution of changing family structure to changes in child poverty.

Consider the distribution of family equivalent income Y in year t, conditional on female labor force participation L, family structure S, and family characteristics X:

$$f_t(Y) \equiv f(Y; t_Y{=}t, t_{L|S,X}{=}t, t_{S|X}{=}t, t_X{=}t) \qquad \textbf{(2)}$$

This identity is notational; it shows that the distribution of Y is defined in year t, conditional on the distributions of L, S, and X in the same year. Expression of the joint distribution in terms of conditional distributions relies on the simple recursion property that is characteristic of joint distributions. In the empirical work, t_Y defines the analysis sample, and it is in general set equal to various years that fall in the later portion of our sample frame.

The essence of the test is to investigate the effects of holding $t_{L|S,X}$ and $t_{S|X}$ at earlier year levels—i.e., to estimate what the distribution of family income would be if the distributions of female labor force participation (conditional on S and X) and family structure (conditional on X) had remained the same as in the earlier period. Consider the application to female labor force participation. Assuming that female labor force participation has risen over time, the simplest way to impose the earlier distribution of female labor force participation on the current family income distribution is to downweight families in which the female head works by a factor that is equal to the percentage increase in the share of families in which female heads work (and similarly upweight families in which the female head does not work).[12]

This simple test, however, ignores any changes in the relationship between female labor force participation, family structure, and family characteristics. For example, suppose the increase

[12] This is essentially a nonparametric, individual-data version of the grouped-data procedure ("shift-share analysis") that is commonly used to reweight mean values across groups.

in female 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 female labor force participation largely is attributable to 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 female labor force participation, *and its relationship to S and X*, held to its earlier period level.

In terms of the notation in (2), we are interested in:

$$f(Y; t_Y\!=\!89, t_{L|S,X}\!=\!69, t_{S|X}\!=\!89, t_X\!=\!89) \qquad \textbf{(3)}$$

This expression represents the density that would be observed if the probability of female labor force participation retained its 1969 level and structure (conditional on family structure S and family characteristics X), but family income otherwise was determined by the distributional characteristics prevailing in 1989. As shown in Appendix A, this distribution can be expressed as the original unconditional distribution of family income in 1989, with individual observations reweighted by the function $\psi_{L|S,X}(L,S,X)$. 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 female labor force participation, although we must account for the added complexity of family structure as an outcome variable. To capture the important changes in family structure and living arrangements during our sample frame, we divided families into the five groups listed earlier and in Figure 5: married with children, married without children, never married with children, other single individuals with children, and single with no children. We want to estimate:

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

To form the appropriate conditioning weights, $\psi_{S|X}(S,X)$, 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. Using the fitted coefficients from this model, we estimated the conditional probability that a family defined by characteristics X is observed to be in its actual family structure category. For the conditionally weighted estimates, we upweighted or downweighted families in the 1989 sample by the proportional difference in the probability of observing a particular family falling into the same family structure category in 1989 as in 1969. Appendix A contains the exact derivation of the estimated conditioning weight. This derivation requires extension of the original DFL technique to the case of a conditioning variable that takes on multiple outcome categories.

We also estimated the separate effects of changes in the distribution of the X vector of family characteristics.[13] The distributional effect of holding t_x=69 can be modeled by again estimating weights and applying them to the 1989 earnings distribution. The estimated weighting function—$\psi_X(X)$—is equal to the relative probability of observing a family with characteristics X

[13] In the decomposition analysis, the characteristics included in X serve as demographic controls in the labor supply and family structure models and as underlying determinants of family income.

in the 1969 versus the 1989 sample, normalized by the unconditional probabilities of being in either sample. If the distribution of the X's changed between the two years, the weights $\psi_X(X)$ will alter the unconditional distribution. Families in the 1989 sample with characteristics that make them relatively more likely to be observed in the 1969 sample will be upweighted in the conditional density estimation. Appendix A describes the details of this derivation, and the list of X variables is provided in the results section (Section IV).

C. Accounting for the Dispersion of Male Earnings

In addition to changes in family structure and female labor supply, existing research has identified the widening distribution of individual earnings as a key factor underlying the widening distribution of family income. Leading explanations for the increase in earnings dispersion include skill-biased technological change, changing patterns of international trade and immigration, and institutional factors such as changes in unionization and the real value of the minimum wage (see Katz and Autor 1999 for a comprehensive overview of this literature). Because our analysis focuses primarily on family behavior and income, we take the underlying determinants of rising earnings dispersion as given. In addition, because Burtless (1999) found that rising dispersion of female earnings had very little impact, we adjust for rising dispersion of male earnings only.

We implement a straightforward procedure for holding constant the characteristics of the male earnings distribution, based on a rank-preserving exchange of the male earnings distributions between two periods. Focusing again on the years 1969 and 1989 as examples, assume that we have ranked the male earnings data in each year from lowest to highest, and consider a man ranked at a particular point. Our underlying conceptual goal is to replace the earnings of each male head in the 1989 data with the earnings of the male head ranked at the same position in the 1969 data. This is trivial in the case of identically sized samples without weights; we simply rank the male earnings data 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 male in the 1989 distribution but imposes the location and dispersion characteristics of the 1969 distribution on the 1989 distribution.[14]

To adjust family equivalent income for the changing distribution of male earnings, we subtracted from 1989 total family income the male head's actual 1989 wage and salary earnings, 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 belonged.[15] After again applying our equivalence adjustment, the resulting income variable (Y_m) replaces our original equivalent income variable (Y) in the analysis. See Appendix B for a formal description of this procedure. This adjustment step requires that we add a factor to our decomposition notation. In particular, letting m=t denote the distribution of male earnings prevailing in year t, estimation of the 1989

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[14] 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 male 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 male earnings distribution.

[15] 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 and comparability with our family income measure. In addition, we included men with earnings equal to zero in the distribution, due to the possibility of declining labor force participation by low-wage men (Juhn 1992).

distribution of family income with the male earnings distribution held to its 1969 structure is indicated by:

$$f(Y; t_Y\!=\!89, m\!=\!69, t_{L|S,X}\!=\!89, t_{S|X}\!=\!89, t_X\!=\!89) \tag{5}$$

D. Decomposition of Factor Contributions

For the complete decomposition, we consider all four factors discussed above and estimate their contributions to rising inequality and poverty. The factors first are considered in the following primary sequence: the distribution of male earnings, female labor force participation, family structure, and underlying family characteristics. We assess the contribution of male earnings first, because leading explanations of changes in the distribution of male earnings—such as skill-biased technological change and institutional factors such as unionization and the minimum wage—imply that shifts in the male earnings distribution 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. However, as discussed below, we also reversed the order of the decomposition, which implies a different causal ordering.

Table 1 lists the various distributions that we analyze, which are defined by the income measure used and the weight associated with each observation. The separate panels identify the distributions estimated in the primary-order and reverse-order decompositions. The first two rows in both panels list the population-weighted distributions for 1969 and 1989; these are defined by the family weight θ and consistent time dimensions for all relationships. The next four rows in both panels identify the adjusted (counterfactual) distributions of 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 male 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. This results in the four adjusted distributions of 1989 family equivalent income that are listed in the table. The effects of the explanatory factors are obtained by comparison of the distributions in sequence, with the adjusted income measure and the product of the sampling weight and the conditioning weights incorporated as listed in Table 1. Comparison of the original and modified distributions can be made in visual terms, using kernel density estimates, or in quantitative terms, using any well-defined dispersion measure.

One concern regarding our approach is the possibility of endogenous or general equilibrium relationships between the distribution of male earnings, female labor supply, and family structure. As a result, the contribution of the explanatory factors may depend on their order in the decomposition analysis. 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 female labor force participation, and changes in the distribution of male earnings. Comparison of the results using our primary-order and reverse-order decompositions enable us to assess the sensitivity of the decomposition to alternative assumptions about the causal relationships between the explanatory factors. The income variable and weights used in the reverse-order decomposition steps are listed in the second panel of Table

1. The exact derivation and estimation of the reverse-order weights is described in Appendix A.

IV. Results

A. 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, SMSA residence, and nine census geographic division dummies. In the female labor force participation equations used to form the conditioning weights, we included the age, education, and race variables separately for male and female heads if both were present. As discussed above, we also estimated the female labor force participation equations as a function of the family structure categories, for which we use separate dummy variables. 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 value in other families.

Figure 6 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 modeled factor to its 1969 structure and examines the impact of this adjustment compared to the distribution adjusted for the immediately preceding factor (or the unadjusted distribution). Panel A shows the effect of the changing distribution of male earnings. Had the distribution of male 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 female labor force participation. The effect of the changing conditional distribution of female labor force participation appears to be a relatively uniform shift to the right in the distribution of family income; the distribution held to the 1969 structure of female labor force participation is to the left of the comparison distribution. 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 male 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.

The quantitative analogue to the visual representation in Figure 6 is listed in Table 2. 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, with the corresponding share of the total change listed in parentheses. 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.[16]

The first row of Table 2 shows that changes in the distribution of male earnings and in family structure had a largely neutral effect on the midpoint of the distribution of family equivalent income. However, the rise in female 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 female labor force participation 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 (midpoint) of the income distribution, non-normalized dispersion measures such as the standard deviation would mistakenly suggest that rising female participation had a neutral effect 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 male earnings had the largest and most uniform effect on family income inequality. Excluding changes in the standard deviation, changes in the distribution of male earnings explain from just over one-half to about four-fifths of rising dispersion in family equivalent income. Although the changing distribution of male 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

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

widened by a larger amount.[17]

The third column of Table 2 shows that rising female labor force participation (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 31 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).

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, 48 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 income—mainly married and single individuals without children. Changing

[17] For the period 1969-1989, we find a larger effect of changes in the distribution of male earnings than did Burtless (1999) for the period 1979-1996, largely because the impact of the male earnings distribution was smaller in the 1990s than earlier (see Table 5).

family structure also explains about one-third of the increase in the key summary measures of overall dispersion (the coefficient of variation and the Gini and Theil indices).

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 relatively small in general, indicating that our four explanatory factors account for a large share 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 all 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 explanatory factors' effects 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 male 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 134 percent of the change in the poverty rate, indicating that unmeasured factors are associated with a (counterfactual) reduction in the poverty rate of about 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.[18]

Reverse-Order Decomposition

As discussed in section III.D, 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 male earnings, female labor supply, and family structure. To address this possibility, Table 3 presents results from the reverse-order decomposition analysis, for which we consider in turn the effects of the changing distributions of family characteristics, family structure, female labor force participation, and male earnings.[19] The impact of the changing distribution of family characteristics is increased in magnitude when we consider it first rather than last in the

[18] 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).

[19] 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 6.

decomposition. Compared to the primary-order decomposition (Table 2), in which its effects on dispersion in family equivalent income were small and of mixed sign, the reverse-order decomposition in Table 3 indicates that the changing distribution of family characteristics increased each of the dispersion measures. The share increase in dispersion 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 male earnings when it is considered last in the decomposition (Table 3) are very similar to its effects when it is considered first (Table 2). Compared to Table 2, in Table 3 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 similar across the two tables. Rising female participation has a somewhat larger equalizing effect on the distribution of family equivalent income in Table 3 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 5 percent of the total change to about 22 percent of the total change. In addition, the reverse-order effects of rising female participation on the coefficient of variation, the Gini coefficient, and Theil's coefficient are equal to about one-fourth of the total change, a noticeably larger effect than in the primary-order decomposition.

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 3), 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 male earnings is larger. Although the impact of the changing distribution of male earnings is especially large in the reverse-order case, the qualitative interpretation of the effects is basically unchanged.[20]

B. Alternative Periods

In the preceding analysis, we examined changes in the distribution of family income between 1969 and 1989, which corresponds roughly to the period of rising inequality that has been the focus of past work. In this section, we examine whether the patterns identified for this period also hold for the key 1980s sub-period. We also examine the role played by our explanatory factors during the 1990s, when the net increase in inequality was relatively modest.

Table 4 lists primary-order decomposition results for changes in the distribution of family equivalent income between 1979 and 1989. Comparing the list of total changes from the first column of this table to those for the period 1969 to 1989 (Tables 2 and 3), we see that increases in family income inequality were more pronounced during the 1980s than during the 1970s. In particular, the total changes between 1979 and 1989 generally account for well above half of the total changes between 1969 and 1989. As extreme examples, the full increase in the coefficient of

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[20] 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 and reverse-order weights (see Appendix A), the full impact of the four explanatory factors is affected by the ordering of the male earnings step.

variation between 1969 and 1989 occurred in the 1980s, and the change in the 90-50 ratio was almost entirely restricted to the 1980s.

Although the increase in inequality was concentrated in the 1980s, comparison of the final columns of Tables 2 and 4 indicates that the net share effect of our explanatory factors was smaller in the 1980s than in the 1970s and 1980s combined. 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 full two-decade period. One key exception is a larger negative effect of rising female participation on dispersion in the lower portion of the distribution of family income during the 1980s: rising female participation accounts for a relatively large (counterfactual) reduction in the 50-10 ratio and the poverty rate during the 1980s. Another noteworthy difference between the results in Tables 2 and 4 is a shift in the family structure effect to the upper half of the distribution in the 1980s, as indicated by the large share of this factor in explaining the increase in the 90-50 ratio.

Our final test involves examining changes in the family income distribution between 1989 and 1998. The key difference between this analysis and the analysis for earlier periods lies in our treatment of the male earnings data. The change in the topcode procedure in 1995 has a substantial effect on dispersion of male earnings and a substantial impact on our estimates of the contribution of male earnings to changing inequality in family income between 1989 and 1998. Therefore, as we did for the display of this variable's distributional characteristics in Figure 4, we topcoded the values of this variable at the 97.5 percentile (in both years) for the formal decomposition analysis.[21]

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 decade (Table 4). 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 male earnings for the moment, the results in the third column indicate that rising female participation decreased inequality in the 1990s by amounts similar to those in the 1980s. Changing family structure increased dispersion in the 1990s by amounts comparable to or slightly larger than those in the 1980s, thereby explaining a larger share of the total change in the 1990s.

In contrast, the impact of the changing distribution of male earnings on rising inequality in family income was much smaller in the 1990s than it was in earlier decades. Moreover, the change in the distribution of male earnings actually improved the relative well-being of low-income families, as indicated by the associated reductions in the 50-10 and 95-5 ratios and the poverty rate. Two key potential explanations for the effect of changes in the male earnings distribution on low-income families in the 1990s are changes in the share of males with zero earnings and changes in the minimum wage. Our auxiliary tabulations (available on request) indicated that the share of male household heads with zero earnings remained approximately

[21] We did not do so for the previous time periods because the changing topcodes had very little impact on the distribution of this variable prior to income year 1995. We verified in auxiliary runs (available on request) that the results listed in Tables 2-4 are not substantially affected by the male earnings topcode.

constant between 1989 and 1998 in our data. However, the real value of the U.S. federal minimum wage 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 during the 1990s to account for much of the relative increase in the income of low-income families.

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 nearly five-tenths of a percentage point to just over a percentage point. Changes in family structure accounted for a full percentage point increase in poverty in the 1990s, which is substantially larger than the family structure effect on poverty in the 1980s.

Among the most interesting results in Table 5 is the change in poverty attributable to residual factors. The residual effect was especially large in the 1990s, implying an increase in the poverty rate of 1.5 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. In contrast, the large positive residual contribution between 1989 and 1998 is consistent with a reduction in benefit receipt associated with recent welfare reform.

V. Conclusions

In this paper, we applied relatively novel semiparametric techniques to assess the contributions of the changing distribution of male earnings, rising female labor force participation, and changing family structure to increasing inequality in family equivalent income in the United States between 1969 and 1998. We estimated the effects of rising female participation and changing family structure in a conditional framework. Our analyses focused on the years 1969, 1979, 1989, and 1998, which represent business cycle peaks or near-peaks, so that our results are relatively unaffected by business-cycle determinants of inequality.

We focused 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 during this period, with unobserved residual factors explaining an amount ranging from none to about one-fourth of the total change in our key dispersion measures. The changing distribution of male earnings had the largest impact on rising dispersion of family equivalent income during this period, explaining anywhere from about one-half to four-fifths of the increase. Changes in family structure also substantially increased inequality between 1969 and 1989. This factor explained up to one-half of the increase in inequality, with its effects concentrated in the lower half of the distribution of family income.

In contrast, we found that the increase in female labor force participation between 1969 and 1989 had an equalizing effect on the distribution of family equivalent income. Its counterfactual impact ranged as high as one-third of the actual change in dispersion and was largest in the upper half of the distribution of family equivalent income. Moreover, the impact of rising female participation shifted to the lower half of the distribution in the 1980s and 1990s, suggesting that rising female participation as a response to growing income gaps spread down the income distribution over time.

The equalizing effect of rising female participation might appear to conflict with previous findings of a rising correlation between husband and wife 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 income and related characteristics of families that generate this rising correlation. Our approach and results explicitly indicate that rising female earnings, in particular those associated with rising female participation rates, occurred among families with relatively lower income in most portions of the distribution.

The effects of our explanatory factors on changes in poverty between 1969 and 1989 followed a pattern similar to their effects on dispersion, especially dispersion in the lower half of the distribution. Because the peak-to-peak changes in poverty were relatively small, our explanatory factors predict very large changes in poverty relative to the actual observed change. The changing distribution of male earnings had the largest impact on poverty, although the impact of changing family structure also was large. Thus, the declining earnings position of low-skilled men played an important role in rising poverty in recent decades, but this was substantially exacerbated by changing family structure, most notably the increase in families characterized by single heads with children.

Within our full sample period, we also focused on changes in family income inequality between 1979 and 1989 and between 1989 and 1998. The sharpest increase in family income inequality occurred during the 1980s, and the role of our explanatory factors was similar to that for the longer period that includes the 1970s as well. In contrast, the net change in inequality between 1989 and 1998 was relatively small, due in part to flat or slightly declining inequality between 1993 and 1998. The changing distribution of male earnings continued to make a substantial contribution to rising inequality during the 1990s, although its absolute impact was smaller than in the 1980s. Also, in contrast to earlier decades, in the 1990s changing family structure increased inequality to a greater degree and rising female participation had a larger offsetting impact, with the effects of both being especially pronounced in the lower half of the distribution.

Each of our explanatory factors also had substantial impacts on poverty in the 1990s. Changing family structure had an especially large effect, causing an increase of 1.2 percentage points. This was more than offset by changes in the underlying characteristics of low-income families, however, and on net our explanatory factors predict a decline in the poverty rate during the 1990s. The large residual increase in poverty (1.5 percentage points) between 1989 and 1998 may be attributable to changes in government transfer policies that reduced the incomes of low-income families.

Except for the poverty rate, we focused in general 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 available income) declined for families below the eighteenth percentile of the distribution of equivalent family income in our data. 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 policy-makers interested in maintaining or improving living standards for specific target populations.

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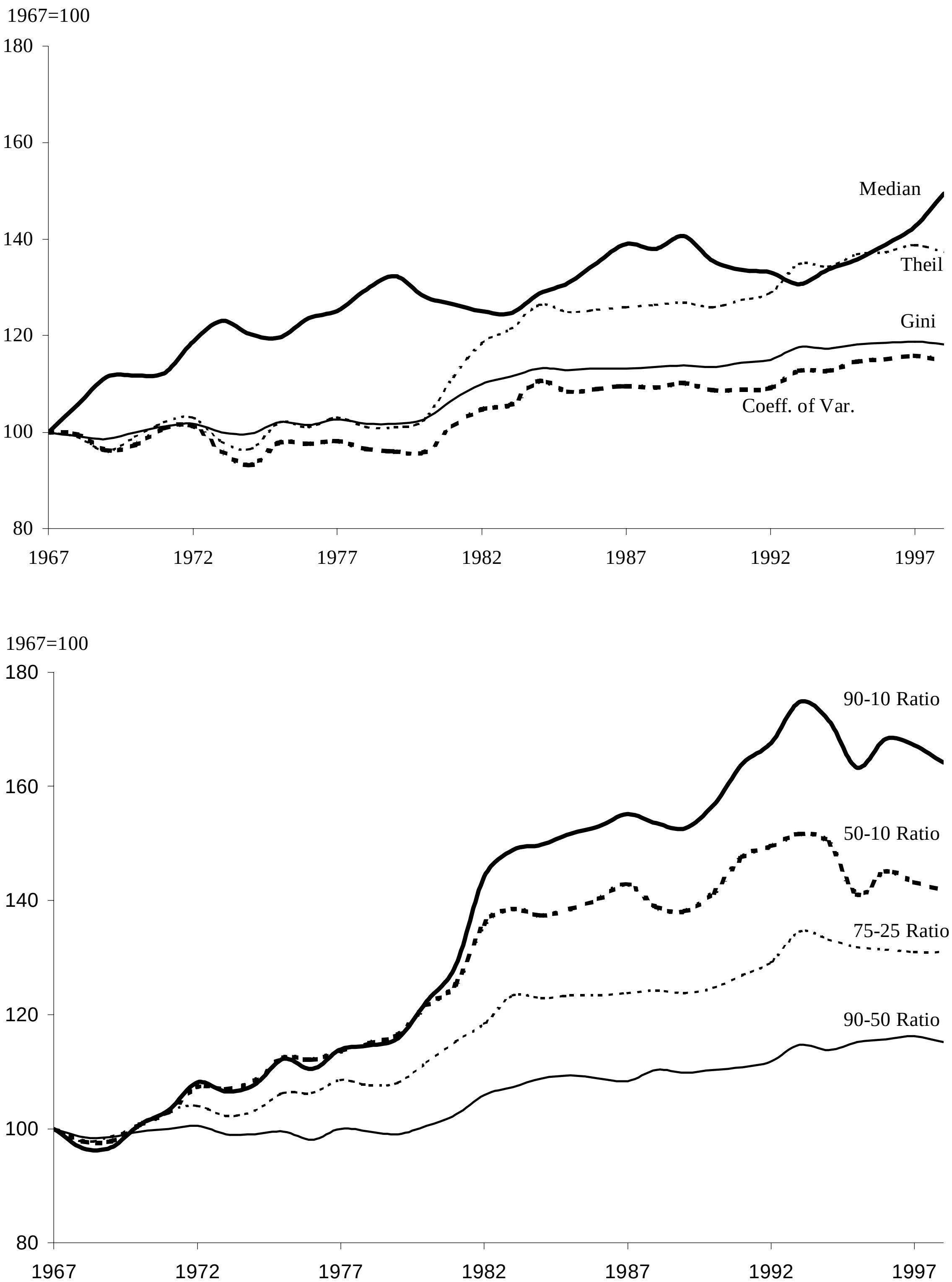
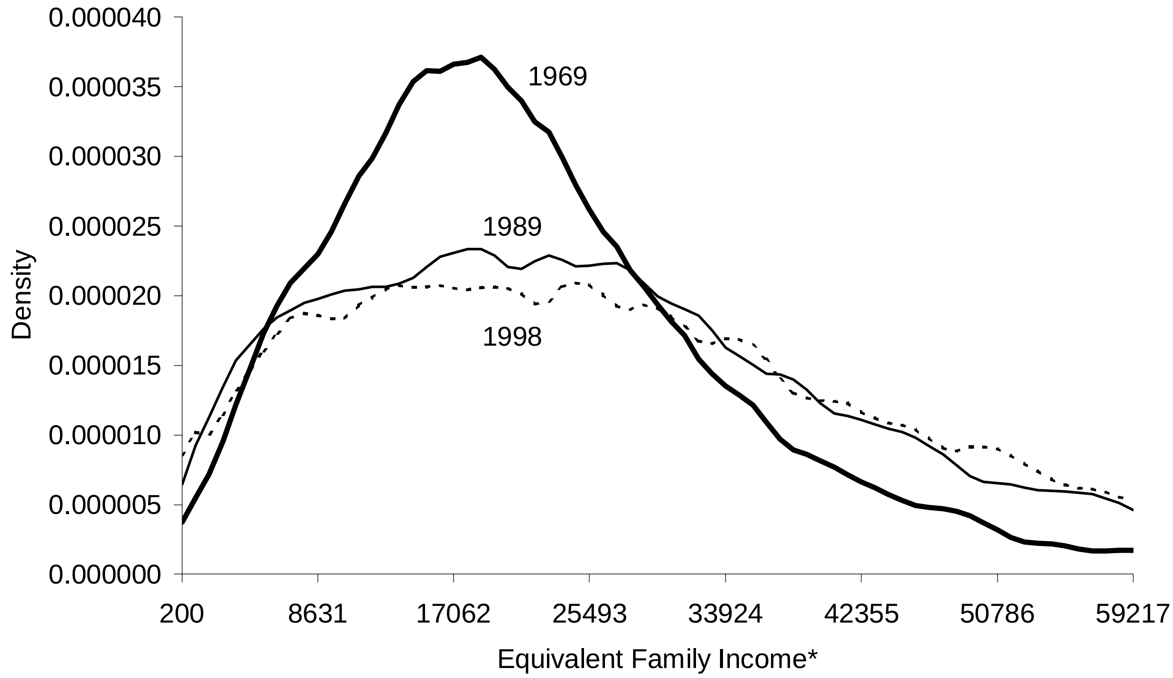
Figure 1 - Summary Measures for Family Equivalent Income (1967-1998)



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

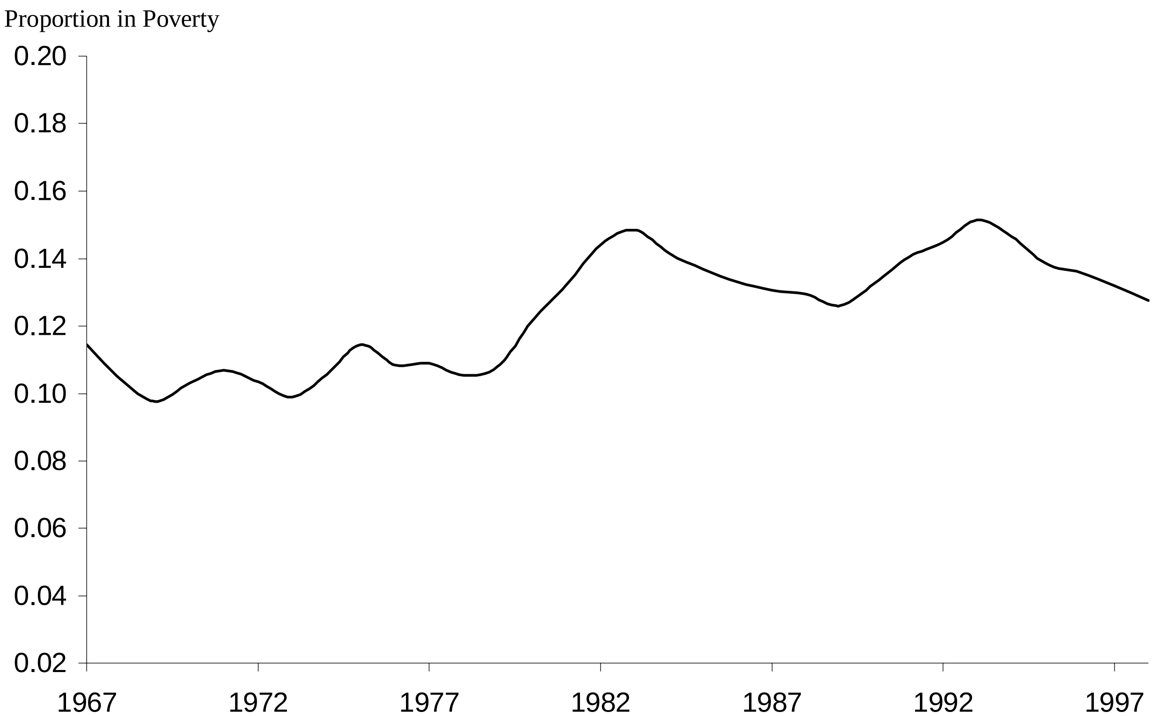
Figure 2
Density Plots for Equivalent Family Income (Unadjusted)



* Distributions truncated at 0 and 60,000

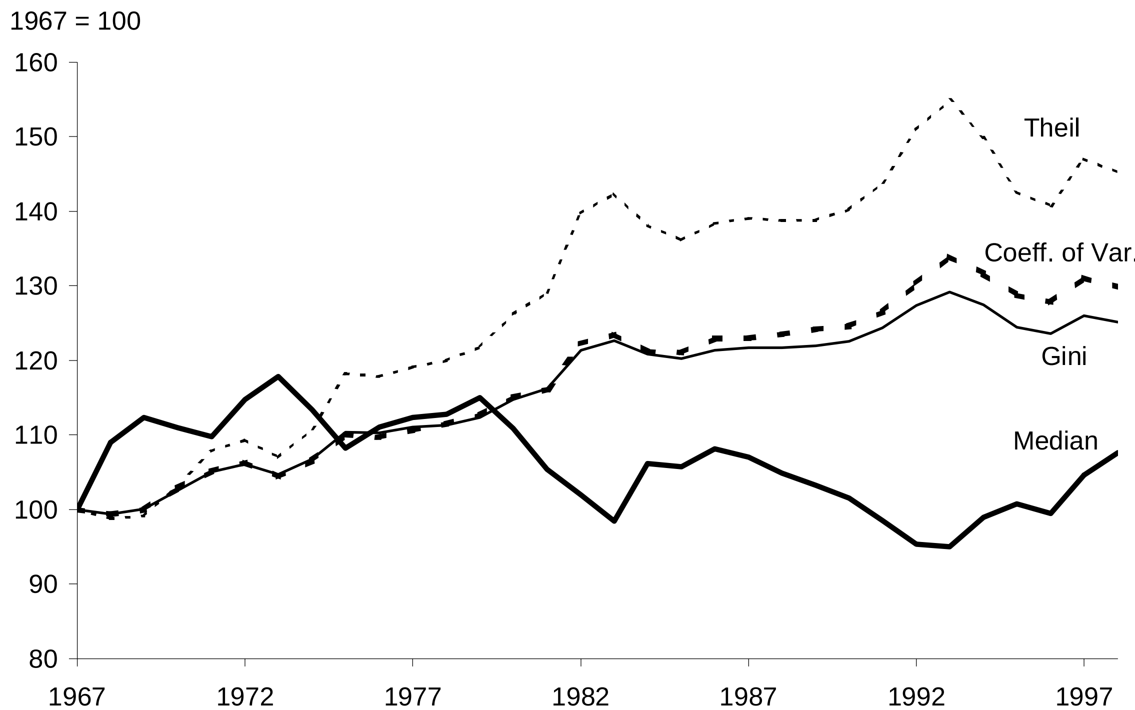
Source: Authors' tabulations of 1970, 1990, and 1999 March CPS data; figures in 1998 dollars.

Figure 3 - Poverty Rate (1967-1998)



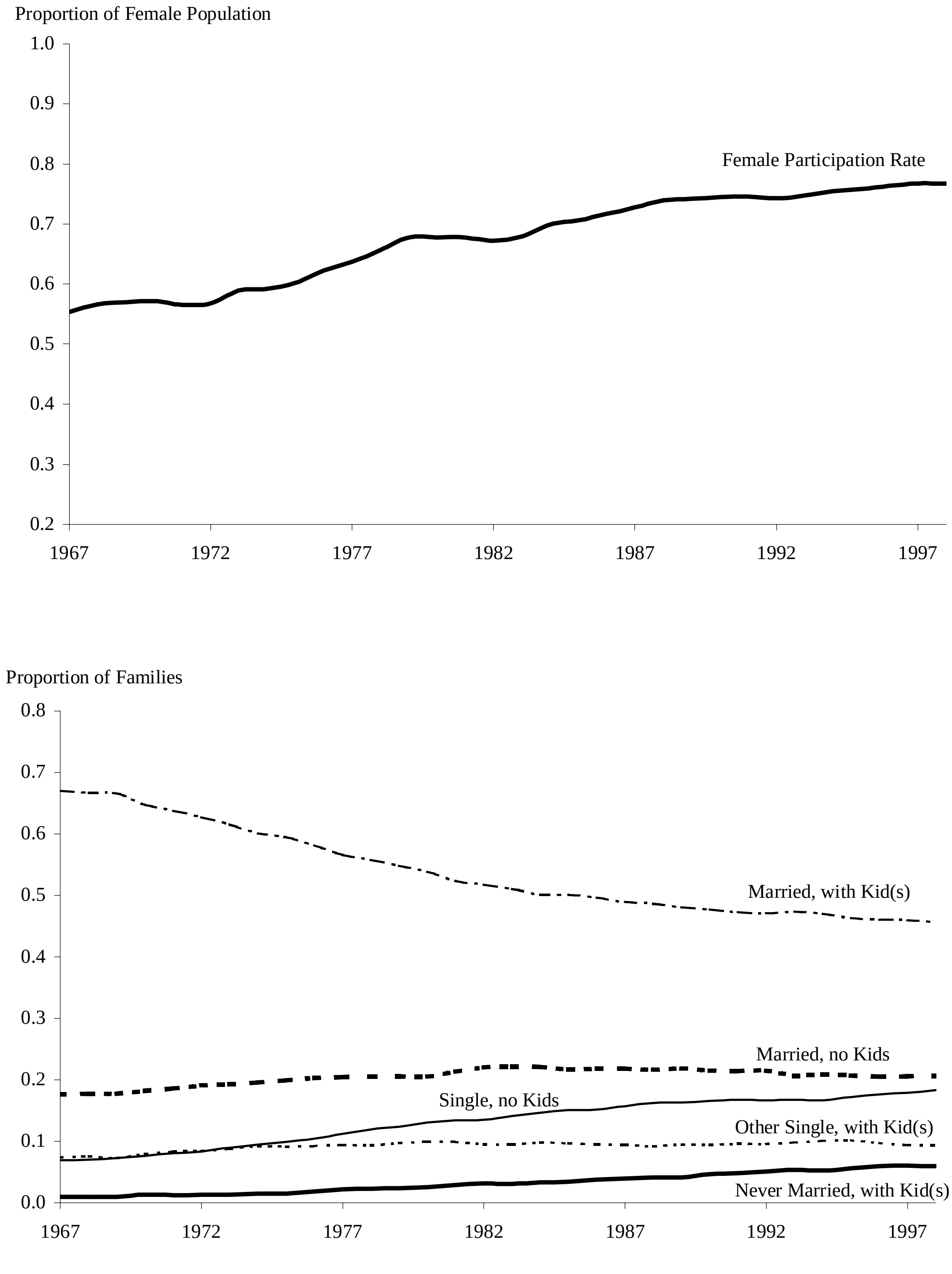
Source: Authors' tabulations of March CPS data.

Figure 4 - Summary Measures for Male Earnings (1967-1998)



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

Figure 5 - Female Labor Force Participation and Family Structure (1967-1998)



Source: Authors' tabulations of March CPS data.

Figure 6
Primary-Order Decomposition, 1989 (1969 base)

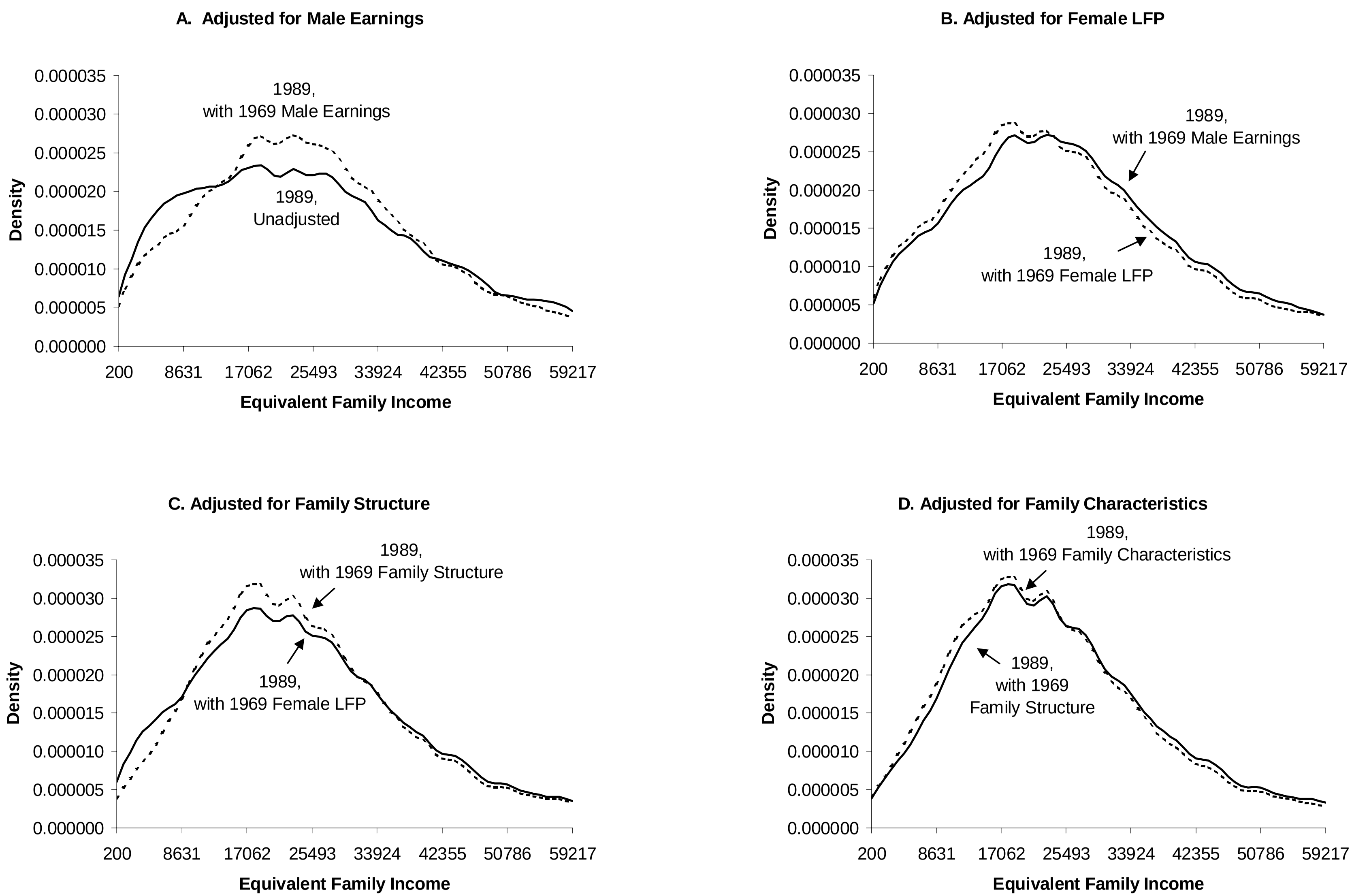


Table 1: Income Measures and Conditioning Weights Used in the Density Decompositions

Primary-Order Decomposition

Distribution	**Income Measure**	**Weight**
$f(Y;\ t_Y=69,\ m=69,\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=69)$	Y	θ_{69}
$f(Y;\ t_Y=89,\ m=89,\ t_{L\|S,X}=89,\ t_{S\|X}=89,\ t_X=89)$	Y	θ_{89}
$f(Y;\ t_Y=89,\ m=69,\ t_{L\|S,X}=89,\ t_{S\|X}=89,\ t_X=89)$	Y_m	θ_{89}
$f(Y;\ t_Y=89,\ m=69,\ t_{L\|S,X}=69,\ t_{S\|X}=89,\ t_X=89)$	Y_m	$\theta_{89}\cdot\psi_{L\|S,X}$
$f(Y;\ t_Y=89,\ m=69,\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=89)$	Y_m	$\theta_{89}\cdot\psi_{L\|S,X}\cdot\psi_{S\|X}$
$f(Y;\ t_Y=89,\ m=69,\ t_{L\|S,X}=69,\ t_{S\|X}=69,\ t_X=69)$	Y_m	$\theta_{89}\cdot\psi_{L\|S,X}\cdot\psi_{S\|X}\cdot\psi_X$

Reverse-Order Decomposition

Distribution	**Income Measure**	**Weight**
$f(Y;\ t_Y=69,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m=69)$	Y	θ_{69}
$f(Y;\ t_Y=89,\ t_{X\|S,L}=89,\ t_{S\|L}=89,\ t_L=89,\ m=89)$	Y	θ_{89}
$f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=89,\ t_L=89,\ m=89)$	Y	$\theta_{89}\cdot\psi_{X\|S,L}$
$f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=89,\ m=89)$	Y	$\theta_{89}\cdot\psi_{X\|S,L}\cdot\psi_{S\|L}$
$f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m=89)$	Y	$\theta_{89}\cdot\psi_{X\|S,L}\cdot\psi_{S\|L}\cdot\psi_L$
$f(Y;\ t_Y=89,\ t_{X\|S,L}=69,\ t_{S\|L}=69,\ t_L=69,\ m=69)$	Y_m	$\theta_{89}\cdot\psi_{X\|S,L}\cdot\psi_{S\|L}\cdot\psi_L$

Note: Y refers to equivalent family income, θ_t is the survey sampling weight for families in year t (see Section II), Y_m is equivalent family income adjusted for the change in the distribution of male earnings (see Section III), and the ψ's are estimated conditioning weights (see Section III). The subscripts for the adjustment variables—m, L, S, and X—refer to male earnings, female labor force participation, family structure, and 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] Male earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	5,281	-27.0 (-.005)	1449 (0.274)	279 (0.053)	897 (0.170)	2,683 (0.508)
Standard Deviation	7,247	2079 (0.287)	208 (0.029)	960 (0.132)	1015 (0.140)	2,985 (0.412)
Coefficient of Variation[c]	0.095	0.055 (0.576)	-0.022 (-0.230)	0.030 (0.311)	0.005 (0.053)	0.027 (0.284)
90-10[d]	3.06	2.14 (0.699)	-0.334 (-0.109)	1.26 (0.411)	-0.058 (-0.019)	0.050 (0.017)
50-10	1.09	0.700 (0.640)	-0.055 (-0.050)	0.568 (0.519)	-0.069 (-0.063)	-0.054 (-0.050)
90-50	0.231	0.188 (0.816)	-0.071 (-0.308)	0.021 (0.091)	0.033 (0.143)	0.060 (0.260)
75-25	0.568	0.438 (0.771)	-0.065 (-0.114)	0.159 (0.279)	-0.028 (-0.049)	0.064 (0.113)
95-5	8.26	4.55 (0.551)	-0.773 (-0.094)	3.97 (0.480)	-0.136 (-0.016)	0.649 (0.079)
Gini Coefficient	0.052	0.034 (0.649)	-0.009 (-0.169)	0.018 (0.342)	0.000 (0.008)	0.009 (0.173)
Theil's Coefficient	0.062	0.039 (0.623)	-0.011 (-0.175)	0.022 (0.358)	0.001 (0.017)	0.011 (0.177)
Poverty Rate	0.028	0.029 (1.02)	-0.009 (-0.309)	0.026 (0.920)	-0.008 (-0.294)	-0.010 (-0.337)

[a] Numbers in parentheses show the share of the explained change in the total change.
[b] See Section IV-A 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: 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	Male Earnings	Residual Factors
Median	5,281	545 (0.103)	-89.8 (-0.017)	1437 (0.272)	485 (0.092)	2904 (0.550)
Standard Deviation	7,247	727 (0.100)	1121 (0.155)	246 (0.034)	2057 (0.284)	3096 (0.427)
Coefficient of Variation[c]	0.095	0.006 (0.063)	0.038 (0.395)	-0.026 (-0.275)	0.054 (0.570)	0.023 (0.247)
90-10[d]	3.06	0.489 (0.160)	1.05 (0.344)	-0.749 (-0.245)	2.290 (0.747)	-0.020 (-0.006)
50-10	1.09	0.174 (0.159)	0.428 (0.391)	-0.240 (-0.220)	0.826 (0.755)	-0.098 (-0.085)
90-50	0.231	0.028 (0.120)	0.034 (0.147)	-0.067 (-0.289)	0.170 (0.735)	0.066 (0.287)
75-25	0.568	0.043 (0.075)	0.174 (0.307)	-0.142 (-0.249)	0.446 (0.785)	0.047 (0.082)
95-5	8.26	2.22 (0.268)	2.99 (0.362)	-1.60 (-0.194)	4.33 (0.524)	0.32 (0.040)
Gini Coefficient	0.052	0.006 (0.110)	0.016 (0.302)	-0.013 (-0.248)	0.037 (0.707)	0.006 (0.129)
Theil's Coefficient	0.062	0.007 (0.114)	0.022 (0.352)	-0.016 (-0.263)	0.041 (0.660)	0.008 (0.137)
Poverty Rate	0.028	0.004 (0.137)	0.018 (0.646)	-0.022 (-0.773)	0.041 (1.45)	-0.013 (-0.460)

[a] Numbers in parentheses show the share of the explained change in the total change.
[b] See Section IV-A 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 4: Primary-Order Decomposition of Changes in the
Distribution of Family Equivalent Income and Poverty, 1979-1989

Statistic	Total Change	Effect of:[a] Male earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	1,543	-470	603	139	-1135	2406
		(-.305)	(0.391)	(0.090)	(-0.736)	(1.56)
Standard Deviation	4,787	1590	83.1	443	-228	2899
		(0.332)	(0.017)	(0.093)	(-0.048)	(0.606)
Coefficient of Variation[c]	0.097	0.049	-0.012	0.011	0.019	0.030
		(0.499)	(-0.119)	(0.113)	(0.198)	(0.309)
90-10[d]	2.00	0.862	-0.386	0.450	0.462	0.612
		(0.430)	(-0.192)	(0.224)	(0.230)	(0.308)
50-10	0.589	0.176	-0.140	0.192	0.150	0.211
		(0.299)	(-0.238)	(0.326)	(0.254)	(0.359)
90-50	0.216	0.131	-0.024	0.115	0.043	-0.049
		(0.605)	(-0.112)	(0.534)	(0.197)	(-0.224)
75-25	0.352	0.192	-0.055	0.050	0.109	0.056
		(0.545)	(-0.155)	(0.143)	(0.310)	(0.157)
95-5	5.83	2.14	-1.11	1.21	1.24	2.35
		(0.366)	(-0.190)	(0.208)	(0.213)	(0.403)
Gini Coefficient	0.041	0.020	-0.005	0.005	0.008	0.013
		(0.501)	(-0.130)	(0.132)	(0.206)	(0.291)
Theil's Coefficient	0.052	0.025	-0.007	0.007	0.010	0.017
		(0.494)	(-0.132)	(0.137)	(0.200)	(0.301)
Poverty Rate	0.017	0.012	-0.008	0.006	0.011	-0.004
		(0.676)	(-0.447)	(0.326)	(0.657)	(-0.212)

[a] Numbers in parentheses show the share of the explained change in the total change.
[b] See Section IV-A 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 5: Primary-Order Decomposition of Changes in the
Distribution of Family Equivalent Income and Poverty, 1989-1998

Statistic	Total Change	Effect of:[a] Male earnings	Female LFP	Family Structure	Family Characteristics[b]	Residual Factors
Median	1,565	868	178	-514	1605	-572
		(0.555)	(0.114)	(-0.329)	(1.03)	(-0.370)
Standard Deviation	1,852	1749	58.4	-115	1179	-1019
		(0.944)	(0.315)	(-0.062)	(0.637)	(-0.834)
Coefficient of Variation[c]	0.004	0.018	-0.003	0.010	-0.007	-0.014
		(4.70)	(-0.849)	(2.58)	(-1.83)	(-3.60)
90-10[d]	0.496	0.155	-0.306	0.540	-0.926	1.033
		(0.312)	(-0.616)	(1.09)	(-1.87)	(2.08)
50-10	0.121	-0.051	-0.124	0.202	-0.031	0.125
		(-0.421)	(-1.02)	(1.67)	(-0.253)	(1.024)
90-50	0.058	0.069	-0.008	0.023	-0.006	-0.020
		(1.19)	(-0.130)	(0.398)	(-0.110)	(-0.348)
75-25	0.155	0.029	-0.015	0.058	-0.041	0.124
		(0.187)	(-0.100)	(0.372)	(-0.267)	(0.808)
95-5	1.91	-0.205	-0.908	1.83	0.473	0.720
		(-0.107)	(-0.475)	(0.956)	(0.247)	(0.379)
Gini Coefficient	0.008	0.007	-0.002	0.005	-0.002	0.000
		(0.793)	(-0.258)	(0.621)	(-0.283)	(0.127)
Theil's Coefficient	0.010	0.009	-0.003	0.007	-0.003	0.000
		(0.907)	(-0.307)	(0.737)	(-0.339)	(0.002)
Poverty Rate	0.002	-0.007	-0.004	0.010	-0.012	0.015
		(-3.63)	(-2.21)	(5.01)	(-5.95)	(7.78)

[a] Numbers in parentheses show the share of the explained change in the total change.
[b] See Section IV-A 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 1990 and 1999 March CPS data.

Appendix Table 1 – 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 2 – 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; figures in 1998 dollars.

Appendix Table 3 – 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 A — 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 IIIB. This discussion largely follows that in DiNardo, Fortin, and Lemieux (1996; DFL), with modification to our specific setting.

Consider 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:

$$f_t(Y) \equiv f(Y;\ t_Y = t,\ t_{L|S,X} = t,\ t_{S|X} = t,\ t_X = t) \tag{A1}$$

A distribution such as (A1) can be expressed as:

$$f_t(Y) = \int\int\int f(Y|L,S,X,t_Y = t)\, dF(L|S,X,t_{L|S,X} = t)\, dF(S|X,t_{S|X} = t)\, dF(X|t_X = t) \tag{A2}$$

In this equation, f(.) is the conditional distribution of Y and F(.) is the joint distribution of Y, L, S, and X. This right-hand side of (A2) indicates that the distribution of family income in a given year can be expressed as the product of its underlying conditional distributions and integrated over L, S, and X.

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

$$f(Y; t_Y = 89,\ t_{L|S,X} = 69,\ t_{S|X} = 89,\ t_X = 89) \tag{A3}$$

Using (A2), this distribution can be expressed as:

$$
\begin{aligned}
f_t(Y;t_Y{=}89, t_{L|S,X}{=}69, t_{S|X}{=}89, t_X{=}89) &= \int\int\int f(Y|L,S,X,t_Y{=}89)\, dF(L|S,X,t_{L|S,X}{=}69)\cdot \\
&\qquad dF(S|X,t_{S|X}{=}89)\, dF(X|t_X{=}89) \\
&= \int\int\int f(Y|L,S,X,t_Y{=}89)\, \psi_{L|S,X}\, dF(L|S,X,t_{L|S,X}{=}89)\cdot \\
&\qquad dF(S|X,t_{S|X}{=}89)\, dF(X|t_X{=}89)
\end{aligned}
\tag{A4}
$$

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 (A4) is identical to (A2) 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 (A4) by using the observed univariate, unconditional distribution of income in 1989.

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

$$
\begin{aligned}
\psi_{L|S,X}(L,S,X) &\equiv \frac{dF(L\,|\,S,X,t_{L|S,X}=69)}{dF(L\,|\,S,X,t_{L|S,X}=89)} \\
&= L\cdot\left(\frac{Pr(L{=}1\,|\,S,X,t_{L|S,X}=69)}{Pr(L{=}1\,|\,S,X,t_{L|S,X}=89)}\right) + (1-L)\cdot\left(\frac{Pr(L{=}0\,|\,S,X,t_{L|S,X}=69)}{Pr(L{=}0\,|\,S,X,t_{L|S,X}=89)}\right)
\end{aligned}
\tag{A5}
$$

The first line identity in (A5) is obtained by substituting the expression on the right side into (A4) and canceling-out the denominator. Regarding the second equality in (A5), note first that L only takes the values 0 or 1, so that:

$$dF(L|S,X,t_{L|S,X}=t) \equiv L \cdot Pr(L=1|S,X,t_{L|S,X}=t) + (1-L) \cdot Pr(L=0|S,X,t_{L|S,X}=t) \tag{A6}$$

The second equality in (A5) follows from the recognition that one term on the right-hand side of (A6) 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 (A5) 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:[1]

$$\begin{aligned} Pr(L=1|S,X,t_{L|S,X}=t) &= pr(\mu > -H(S,X)\beta) = 1 - G(-H(S,X)\beta) \\ &= \frac{\exp(-H(S,X)\beta)}{1 + \exp(-H(S,X)\beta)} \end{aligned} \tag{A7}$$

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 (A7), 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

[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.

equation is estimated for both the 1969 and 1989 samples, and the coefficients are retained. We used the results to fit the probabilities in (A5) using the values of (S,X) from the 1989 sample, combined with the 1969 coefficients for the numerator and the 1989 coefficients for the 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 :

$$\begin{aligned} f_t(Y;t_Y{=}89,t_{L|S,X}{=}69,t_{S|X}{=}69,t_X{=}89) &= \int\int\int f(Y|L,S,X,t_Y{=}89)\,dF(L|S,X,t_{L|S,X}{=}69)\cdot \\ &\qquad dF(S|X,t_{S|X}{=}69)\,dF(X|t_X{=}89) \\ &= \int\int\int f(Y|L,S,X,t_Y{=}89)dF(L|S,X,t_{L|S,X}{=}89)\,\psi_{L|S,X}\cdot\psi_{S|X}\cdot \\ &\qquad dF(S|X,t_{S|X}{=}89)\,dF(X|t_X{=}89) \end{aligned} \tag{A8}$$

The second line of this equation is identical to (A4) 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:

$$\psi_{S|X}(S,X) \equiv \frac{dF(S|X, t_{S|X}=69)}{dF(S|X, t_{S|X}=89)} = \sum_{c=1}^{C} I_c \cdot \frac{Pr(S=c|X, t_{S|X}=69)}{Pr(S=c|X, t_{S|X}=89)} \tag{A9}$$

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 (A9) is a generalization of the second line of (A5), with S|X replacing L|S,X; in (A5), 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 (A9) is generated by the same property as illustrated in (A6). The expression to the right of the summation sign in the bottom line of (A8) 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:

$$\psi_X(X) \equiv \frac{dF(X|t_X=69)}{dF(X|t_X=89)} = \frac{Pr(t_X=89)}{Pr(t_X=69)} \cdot \frac{Pr(t_X=69|X)}{Pr(t_X=89|X)} \tag{A10}$$

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 III.D 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 B). 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:

$$f_t(Y) = \int\int\int f(Y|X,S,L,t_Y = t)\, dF(X|S,L,t_{X|S,L} = t)\, dF(S|L,t_{S|L} = t)\, dF(L|t_L = t) \tag{A11}$$

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 (A8) above, we have:

$$\begin{aligned}\psi_L(L) &\equiv \frac{dF(L|t_L=69)}{dF(L|t_L=89)}\\ &= \frac{Pr(t_L=89)}{Pr(t_L=69)}\cdot\frac{Pr(t_L=69|L)}{Pr(t_L=89|L)}\\ &= \frac{Pr(t_L=69|L)/Pr(t_L=69)}{Pr(t_L=89|L)/Pr(t_L=89)}\end{aligned} \tag{A12}$$

Then the estimated weight ψ_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 (A9), we have:

$$\begin{aligned}\psi_{S|L}(S,L) &\equiv \frac{dF(S|L,t_{S|L}=69)}{dF(S|L,t_{S|L}=89)}\\ &= \sum_{c=1}^{C} I_c\cdot\frac{Pr(S{=}c|L,t_{S|L}=69)}{Pr(S{=}c|L,t_{S|L}=89)}\end{aligned} \tag{A13}$$

The probabilities in (A13) 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 $\psi_{X|S,L}$:

$$\psi_{X|S,L}\cdot\psi_{S|L}\cdot\psi_{L}=\psi_{L|S,X}\cdot\psi_{S|X}\cdot\psi_{X} \quad \Rightarrow$$
$$\psi_{X|S,L}=\frac{\psi_{L|S,X}\cdot\psi_{S|X}\cdot\psi_{X}}{\psi_{S|L}\cdot\psi_{L}} \tag{A14}$$

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 B — Adjusting for the Distribution of Male Earnings

In this appendix, we provide a detailed description of our adjustment for the changing distribution of male 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 actual 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 male earnings. We divide the male 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 male earnings counterfactual.

Consider adjustment of the 1989 distribution of total family income (T^{89}) for families with male heads present based on substitution of 1989 male earnings with the appropriate quantile

[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 male 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.

median from the 1969 distribution of male earnings. In this case, our adjusted measure of total family income (with individual identifier j suppressed) is:

$$T_m^{89} = T^{89} - M^{89} + M_q^{69} \tag{B1}$$

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.[2] Our preliminary investigations suggested that similar results are obtained 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 male earnings:

$$Y_m = \frac{T_m}{F^{\sigma}} \tag{B2}$$

where F=family size and σ=0.5.

[2] We note this equivalence between the quantile and actual distributions because one potential concern about the technique is that it may reduce dispersion per se. We also verified that the distribution of Y_m was virtually identical to the distribution of Y in terms of mid-point and dispersion when Y_m was formed by replacing actual male earnings with the appropriate quantile median from the same year.