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Welfare Reform and the Level, Composition, and Volatility of Income

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Abstract: We use data from the Current Population Survey to document changes in the level, composition, and volatility of income over the past 25 years among single mothers with dependent children. We show that there have been dramatic changes in the level and composition of income across the income distribution during the welfare reform era, with substantial substitution away from most transfers and towards labor market earnings and tax credits. Post welfare reform, fewer than half of mothers at the 10th percentile of income, which falls at about one-half the poverty level for a 3-person family, received any income from TANF, SSI, SSDI, or earnings. We identify a rise in income volatility of about 60 percent among all single mothers after 1995, and this growth in instability appears to have affected all segments of the single mother population. Quandt-Andrews tests of structural change with unknown change point indicate that the trend break in volatility is centered around 1996, which coincides with passage of PRWORA. During the same period, however, TANF income volatility declined substantially. A leading candidate for the rise in income volatility among single mothers is rising earnings volatility, though the statistical evidence is not conclusive.

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The Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996 represented a fundamental change to the delivery of cash welfare to program participants. PRWORA transformed the Aid to Families with Dependent Children (AFDC) program from an entitlement program that provided cash benefits to those households who satisfied state and federal eligibility standards into a work-based block-grant program called Temporary Assistance to Needy Families (TANF) that is almost exclusively controlled by the states. PRWORA had many goals, including ending dependence on government benefits through promotion of work and marriage, reducing and preventing out-of-wedlock pregnancies, and providing aid to needy families in order to provide home-based care. To realize these goals PRWORA established time limits on benefit receipt, work requirements for adult recipients, and work incentives such as higher earnings disregards and liquid-asset limits, among other policies. Most of the new program rules evolved out of state-level experiments conducted in the early 1990s via waivers from federal regulations granted by the U.S. Department of Health and Human Services. A significant body of research was spawned in the wake of welfare reform, but most of the analyses have relied on data and outcomes prior to 2000 (Grogger and Karoly 2005), and none has examined the volatility of income and earnings of women. In this paper we document how the level, composition, and volatility of income in single-mother families changed over the past twenty-five years inclusive of the decade following welfare reform.

The target of welfare reform was aimed squarely at low-income single mothers as this demographic group historically comprised over 90 percent of the caseload. Although the typical (non-censored) spell on AFDC for single mothers was only about 8 months (Blank and Ruggles 1996), the public perception of long-term dependence and intergenerational transmission was widespread and not altogether false from a lifetime perspective (Blank 1997). This perception

spurred policymakers first at the state level with waivers and then at the federal level with PRWORA to construct a new program that not only discourages long-term use via the 5-year federal lifetime limit on benefit receipt but also discourages entry onto the program altogether via diversion payments and work requirements (Grogger, Haider, and Klerman 2003). The TANF program, unlike AFDC, is less focused on providing cash benefits in favor of in-kind assistance. Indeed, nearly 70 percent of TANF funds are spent on in-kind transfers and 30 percent on cash, which is directly opposite that of the former AFDC program (DHHS, ACF 2006). As a consequence of the policy lens on single mothers, most welfare reform research has been directed at understanding the consequences of the legislation for single mothers.

In a widely publicized study, Primus, et al. (1999) examined changes in the earnings and income of female-headed households from 1993–1995 and 1995–1997, and found that disposable income in the lowest quintile rose from 1993–1995 during a period of rapid economic growth (and state experiments with welfare waivers), but then fell an average of $580 after passage of PRWORA. While the authors attribute more than three-fourths of the income decline to declines in cash-assistance and food stamp income, it is not possible to conclude that welfare reform per se is the reason for the income declines because the authors fail to control for other factors that might have affected earnings for this subpopulation such as the macroeconomy. Moffitt (1999) provides a more rigorous analysis of the effect of welfare reform on female-headed earnings, and concludes that in the period leading up to PRWORA (1977–1995) the state-specific welfare waivers led to an average increase in earnings of $274, although there was no significant increase in earnings of women with less than a high school degree.

Schoeni and Blank (2000) update and extend the analysis of Moffitt (1999) both to 1998 and to other outcomes such as poverty status and family structure. Unlike Moffitt, they find a

significant welfare-reform induced increase in own and family earnings for women with less than high school in the pre-PRWORA period; however, there is no additional increase after the passage of PRWORA. They do find strong evidence that welfare reform both in the waiver period and the TANF period reduced the incidence of poverty for the subpopulation of less-skilled women, which is broadly corroborated in Gundersen and Ziliak (2004). These results also appear to be broadly consistent with several welfare "leaver" studies such as Danziger, et al. (1999) and Cancian, et al. (2000), although the latter both emphasize acute post-welfare income declines among women with substantial barriers to employment such as mental health problems and drug dependencies.

Grogger (2003) used data from the 1979–1999 waves of the Current Population Survey to estimate whether and to what extent time limits affected the outcomes of female-headed families. He exploited the fact that families with young children are more likely to be affected by short time limits because of a longer eligibility horizon, and found that time limits reduced welfare use and raised employment, but had no discernable impact on earnings or income. In a break from most papers in the literature, Meyer and Rosenbaum (2001) adopt a quasi-structural approach to model the effect of the tax and transfer system on the employment of single mothers between 1984 and 1995. Although they do find some evidence that waivers encouraged employment, the striking conclusion of their analysis is that expansions in the Earned Income Tax Credit dominated all other policy reforms in the 1990s and accounted for about 60 percent of the rise in labor force participation of single mothers in the mid 1990s.

In this paper we update the previous literature on single mother families by incorporating data through the first decade of welfare reform and extend the earlier work by documenting for the first time how income and earnings volatility of single mothers has changed over the past

twenty five years. The bulk of research on income inequality and volatility has focused on men (Juhn, Murphy, and Pierce 1993; Gottschalk and Moffitt 1994, 2006; Dynarski and Gruber 1997; Blundell, Pistaferri, and Preston 2006) or families with no distinction between male and female heads (Gundersen and Ziliak 2003; Hacker 2006). Although some work on inequality has conducted analyses separately for men and women (Bound and Johnson 1992; Katz and Murphy 1992; Autor, Kearney, and Katz 2005; Lemieux 2006), to our knowledge there is no research that focuses on income volatility of single mothers in nationally representative data. This is important both because it sheds light on whether the rising inequality identified in Lemieux (2006) and Autor, Katz, and Kearney (2005) is also affecting the economically vulnerable subpopulation of single mothers and because it informs policymakers of possible financial instability in the wake of major legislative reforms.

Using data from the 1980–2005 waves of the March Annual Social and Economic Study of the Current Population Survey we first document changes in the level and composition of income across the single-mother income distribution. We then present annual estimates of volatility, and follow the literature on earnings volatility of men by decomposing total volatility in a given year into transitory and permanent components. Because our data come from repeated cross sections, rather than true panel data, we consider two alternatives to identify permanent from transitory volatility. In the first method we construct detailed birth-year by education cohorts and attribute the permanent component to cohort and cohort-year specific effects (akin to the typical person-specific fixed effect in panel data). In the second method we also net out a detailed set of covariates, along with cohort-year effects, from log income or earnings and attribute the permanent component to the observed demographic and cohort effects.

We construct our estimates of volatility separately for total income, as well as labor-market earnings and income from several transfer programs. A complication for the earnings and transfer-income volatility analysis is the fact that a sizable fraction of single mothers either do not work in the labor market in any given year, or do not participate in transfer programs; hence, our earnings and transfer-income volatility estimates are made robust to nonrandom sample selection into work and welfare using both Tobit and two-step selection correction methods. With our estimated time series of income and earnings volatilities we then conduct Quandt-Andrews type tests of structural change where the change point is treated as an unknown parameter to be estimated (Quandt 1960; Andrews 1993).

We show that there have been dramatic changes in the level and composition of income across the income distribution during the welfare reform era, with substantial substitution away from most transfers and towards labor market earnings and tax credits. We identify a rise in income volatility of about 60 percent among all single mothers after 1995, and this growth in instability appears to have affected all segments of the single mother population. Quandt-Andrews tests of structural change with unknown change point indicate that the trend break is centered around 1996, which coincides with passage of PRWORA. During the same period, however, TANF income volatility declined substantially. A leading candidate for the rise in income volatility is rising earnings volatility, though the statistical evidence is not conclusive.

II. Data

The data come from the 1980–2005 waves (1979–2004 calendar years) of the March Annual Social and Economic Study of the Current Population Survey (CPS). The unit of observation is single female family heads between the ages of 16 and 54 with dependent children present under the age of 18. Single heads includes never married women as well as those

divorced, separated, or widowed. The mothers are allocated to thirteen different five-year date of birth cohorts (starting in 1919 and ending in 1983), and within each birth cohort, three separate education groups of less than high school, high school graduate, and more than high school, yielding thirty-nine separate birth-education cohorts. The five birth cohorts from 1939 to 1963 provide complete information over the entire sample period, but the earlier and later cohorts only provide partial information for identification much like one would find in a standard unbalanced panel of families. Because the consistency of the grouping estimator described below is based in part on the number of observations per cell being large, we follow Blundell et al. (1998) and drop cohort-education cells with fewer than 50 observations.

We define earnings as total family earnings from wage and salary income, non-farm self employment, and farm self employment. Because the Census Bureau defines a family as two or more persons related by birth, marriage, or adoption, family earnings contains earnings of the mother as well as dependent children and other related adults such as a resident grandparent. It does not contain earnings of cohabiting partners or other non-family members in the household. Family income is the sum of earnings, Social Security (retirement, disability, and survivors benefits), Supplemental Security Income, Unemployment Insurance, workers' compensation, TANF and other forms of public cash welfare, veterans' payments, pension income, rent/interest/dividend income, royalties, income from estates, trusts, educational assistance, alimony, child support, assistance from outside the household, and other income sources. We append to the official Census income definition used for poverty measurement the imputed dollar value of food stamps and school lunch, along with the Earned Income Tax Credit. If the respondent refuses to supply earnings or transfer information, then the Census Bureau uses a "hotdeck" imputation method to allocate income to those with missing data. Bollinger and

Hirsch (2006) argue that including allocated data generally leads to an attenuation bias on estimated regression coefficients based on imputed data. Hence, we follow their recommendation and drop those mothers with allocated earnings or transfer income. In addition, 0.7 percent of the remaining sample has negative or zero values for total income, and we drop these observations. All income sources are deflated by the personal consumption expenditure deflator with 2005 base year. The total number of observations is 94,681 single female-headed families.

A comment on our choice of data is warranted. The bulk of the volatility literature uses data from the Panel Study of Income Dynamics, while the inequality literature uses the CPS. The distinction arises because inequality is inherently viewed as a cross-sectional issue, while volatility is generally viewed as an individual-specific issue. Thus, the advantage of panel data for volatility measurement is that it is possible to construct individual time series, which readily fosters decomposition of the variance into permanent and transitory terms. The problem with data such as PSID for our purposes is the comparatively small sample sizes of single mother families, especially when one imposes continuous single status for up to five years in order to construct time-means as in the male volatility literature (Gottschalk and Moffitt 2006; Hacker 2006). The duration of single-mother spells is relatively short, and imposing continuous single status is likely to reduce effective sample sizes (Bumpass and Raley 1995). The small sample problem is further exacerbated when splitting the sample by various demographic profiles such as education attainment, age, and marital status, each of which provides useful information regarding the heterogeneity of volatility across the population. As defined below, the estimates of total variance in a given year should be comparable whether using the PSID as Gottschalk and Moffitt (1994, 2006) or the CPS. Differences may arise, however, in the decomposition into transitory and permanent components. We discuss these distinctions in the section on volatility.

III. The Level and Composition of Income, 1979–2004

We begin with a depiction of trends in income levels and composition in Figures 1–14. Part of the observed trends described below in the figures is a mechanical artifact of the structure of transfers programs. That is, the programs are income conditioned and tax away benefits as other income sources increase. This means that as earnings rise, benefits from TANF, food stamps, and SSI necessarily fall (EITC benefits first rise then fall with earnings). The other part of the observed trends is a behavioral response to changes in economic conditions, social policies, and possible preferences towards work and welfare. At this point we do not make any claims about causal factors underlying the observed changes, and simply document facts.

In Figures 1 and 2 we show average and median income levels for all single mothers pooled together. Through the 1980s and into the mid 1990s average real income of single mothers was fairly constant at about $24,000, but beginning in 1994 there was a sizable increase in average real income, reaching a peak of $32,000 in 2004. Figure 2 demonstrates that this growth was not necessarily restricted to the far right tail of the income distribution as median income trended upward as well starting in the mid 1990s. The composition of income, however, shifted dramatically during this period. Labor market earnings at the mean and median grew substantially, and with the rise in employment and earnings, and increased EITC program generosity, income from the EITC increased as well. During this period income from SSI also expanded at the mean, which could be in part to both the *Zebley* decision—the 1990 Supreme Court ruling that liberalized child eligibility into SSI—and substitution from AFDC into SSI (Kubik 1999; Schmidt and Purvak 2004). Coincident with the large declines in welfare caseloads (Ziliak et al. 2000), mean income from TANF fell 85 percent between 1979 and 2004, and income from food stamps fell 40 percent between the peak participation year of 1994 and trough

participation year of 2001. Perhaps contrary to widespread perception of the welfare dependent single mother, Figure 2 reveals that fewer than 50 percent of mothers derived any income in any year between 1979 and 2004 from AFDC/TANF, SSI, disability, or other transfers. Most income came from labor-market earnings, supplemented with work supports such as the EITC and food stamps.

[Figures 1 and 2 here]

In Figures 3–8 we depict average and median income sources by three broad groups of education attainment—less than high school, high school graduate with no college, and more than high school. Mean income among mothers with less than high school rose 24 percent in real terms between 1993 and 2004, but this gain simply returned the average low-skilled mother to an income level comparable to the start of the period in 1979. Underlying these trends were large shifts away from TANF and food stamps and towards earnings and the EITC, along with some rise in average SSI income. The medians in Figure 4, however, put in dramatic relief the shifting composition of income among low-skilled single mothers. Prior to the mid 1990s median income was zero from most sources except AFDC and food stamps (and earnings in a few years), but after 1996 median income from TANF fell to zero and food stamps fell off dramatically. In their place arose labor market earnings and the EITC. Indeed Figure 4 highlights the interaction of the business cycle and policy reforms. Earnings fell beginning with the recession of 1980 and 1981, and then went to zero after the welfare reforms proposed by President Reagan were passed by Congress (Moffitt 1992). As the 1980s macroeconomic expansion reached deeper into the distribution median earnings became nonzero, only to fall to zero with the early 1990s recession. The large run-up in median AFDC income in the early 1990s coincided with the increase in caseloads, and then during the late 1990s expansion coupled with welfare reform median

earnings accelerated and TANF plummeted. At the end of the period, a possible cause of concern is that by 2003 the median level of the EITC was zero, which suggests that the mild recession of 2001 may have imparted longer-term negative effects among large sections of the less skilled.

[Figures 3-8 here]

In Figures 5–8 we present parallel figures for mothers with a high school degree and those with more than high school. It is clear from these figures that as one moves up the education distribution the role of transfers falls substantially even at the means, and is nearly nonexistent at the median. An important exception, especially among single mothers with high school but no college, is the EITC, and to a lesser extent food stamps. Unlike mothers with less than high school, the high school and more than high school groups had real growth of income at the means of 9 and 20 percent, respectively. This differential growth is consistent with the inequality literature that emphasizes divergence across education and labor-market experience groups (Lemieux 2006).

[Figures 9–14 here]

Because of evidence that deep poverty in the United States rose in the late 1990s and into the early 2000s (Ziliak 2006), in Figures 9–14 we take a closer look at changes in the level and composition of income at different points of the single-mother income distribution, with a particular emphasis on the bottom half of the distribution at the 10^{th}, 25^{th}, and 50^{th} percentiles. As seen in Figures 9 and 10 changes in the average and median incomes for mothers in the 10^{th} percentile of the income distribution were quite substantial. Income at the 10^{th} percentile ranged from about $6,800 to $8,200 between 1979 and 2004, and for purposes of comparison, the poverty threshold for a family with one adult and two related children in real 2005 dollars is $15,735 so that the 10^{th} percentile falls at about one-half the poverty line for a three person

family. At the 10th percentile mothers still rely quite heavily on transfers on average, but post welfare reform fewer than half of mothers even at this low level of income are receiving any cash support from TANF, SSI, Social Security and DI, or earnings except for a few years. It is important to highlight that some of these mothers may be receiving in-kind assistance from TANF such as child care subsidies or transportation assistance that is not recorded in the CPS. Indeed, because about 70 percent of TANF funds are now delivered as in-kind assistance, it is likely that the figures overstate the loss of support of TANF and may in fact reflect the change in delivery of welfare. That said, the figures underscore important changes in financial liquidity among the very poor.

Figures 11–14 depict changes in the mean and median level and composition of income at the first quartile and median of the income distribution. Real income grew about 32 percent at the mean between 1993 and 2004 for mothers at the 25th percentile, and unlike the broad group of mothers with less than high school, real income grew about 12 percent overall from 1979 to 2004 at the first quartile. This growth was driven by strong increases in earnings and the EITC. Changes at the median, however, closely resemble those of mothers with less than high school in Figure 4. Trends in mean and median income at the 50th percentile in Figures 13 and 14 are nearly coincident with those of mothers with a high school degree as seen previously in Figures 5 and 6. In other words, the typical single mother with a high school degree falls at about the median of the single mother income distribution.

IV. The Volatility of Income and Earnings

We now explore whether there have been changes in the variability of income over the past two decades. This is important both because it sheds light on whether the rising inequality identified in Lemieux (2006) and Autor, Katz, and Kearney (2005) is also affecting the

economically vulnerable subpopulation of single mothers and because it informs policymakers of possible financial instability in the wake of major legislative reforms. To estimate the volatility of income and earnings we extend the widely used error components structure of Gottschalk and Moffitt (1994, 2006) for panel data to the case of pseudo panel data based on cohorts (Deaton 1985). If we define the log of income (or earnings) for person i, $i = 1,\ldots, N$, in cohort c, $c = 1,\ldots, C$, at time t, $t = 1,\ldots, T$, as y_{ict}, then we can decompose log income as

(1) $$y_{ict} = \mu_i + \delta_c + \lambda_{ct} + \varepsilon_{ict},$$

where μ_i is the person-specific and time-invariant component, δ_c is the cohort-specific and time invariant component, λ_{ct} is the cohort-specific and time-varying component, and ε_{it} is the person-specific and time-varying component. Assuming that the four error components are orthogonal, the total (cross-sectional) variance of income in any time period t is

(2) $$V(y_{ict}) = V(\mu_i) + V(\delta_c) + V(\lambda_{ct}) + V(\varepsilon_{ict}).$$

The typical application based on panel data only has two components, μ_i and ε_{it}, to reflect "permanent" and "transitory" volatility, respectively.[1] To identify the permanent and transitory components, Gottschalk and Moffitt (2006) exploit the fact that under no autocorrelation in the transitory error terms the covariance between current income and lagged income is $Cov(y_{it}, y_{it-1}) = V(\mu_i)$. With repeated observations on the same individual over time it is straightforward to calculate the covariance, and thus the permanent variance component of volatility. Given the covariance terms, transitory volatility is found by simply differencing the covariance from the total variance, $V(\varepsilon_{it}) = V(y_{it}) - Cov(y_{it}, y_{it-1})$. Prior to constructing the variances and covariance, it is common in the literature to replace y_{it} with the corresponding

[1] The panel data model could be readily extended to the framework in (1) if there were a priori reasons to emphasize cohort differences.

residual from a pooled OLS regression of income (or earnings) on a quartic in age in order to net out life-cycle influences from the volatility measure. Because of concerns of possible autocorrelation in the transitory variance Gottschalk and Moffitt (2006) actually use the lagged residual of earnings from four years in the past to calculate the covariance terms.

With cohort data we do not have repeated observations on the same individual over time and thus must adopt an alternative identification scheme.[2] One approach is to follow the pseudo panel literature and assume that conditional on δ_c and λ_{ct} the contributions of μ_i and ε_{ict} to volatility go to zero asymptotically. In models where the focus is on the conditional mean function this is frequently a useful assumption. In practice this typically means that one calculates the cohort-specific means for each time period and treats the data akin to panel data by including cohort and year effects in the regression model. With the emphasis on the second moment we are concerned that calculating variances based on cohort averages will suppress too much idiosyncratic variation that is central to the volatility literature.

Instead, our approach is to first estimate a pooled OLS regression of income on a quartic in age and use the residual to construct the total cross-sectional variance for each time period, which yields $V(y_{ict})$ in equation (2). This is the identical procedure used in Gottschalk and Moffitt (2006) when recovering the total variance in the PSID. We next regress income on a quartic in age and a full interaction of cohort dummies and year dummies. This regression nets out both life cycle age effects as well as cohort-year effects, i.e. $y_{ict} - \delta_c - \lambda_{ct}$. The residual from this regression yields an estimate of $V(\mu_i) + V(\varepsilon_{ict})$. Provided that the contribution of μ_i is small, this term will serve as an estimate of transitory volatility. To calculate permanent volatility we simply subtract the estimate of transitory volatility from the estimated total volatility.

[2] In actuality about 70 percent of respondents in the March CPS can be matched over two consecutive years, but this is not likely sufficient to mimic analyses from the PSID that use much longer lags.

A concern with the proposed decomposition is that part of the permanent, individual-specific volatility not captured by the cohort-year effects is attributed to transitory volatility. In a bid to tighten the estimate of transitory volatility we conduct a second analysis where we estimate the total volatility as before, but now in the second stage we add to the quartic in age and cohort-year interactions a full set of person-specific covariates such as a quartic in education, dummy variables for past marital status (never married versus widowed/divorced/separated) and race (white, black, Hispanic), and interactions of these variables with age. This regression nets out life-cycle effects, cohort effects, and other person-specific factors that affect permanent income, i.e. $y_{ict} - \delta_c - \lambda_{ct} - x_{it}\beta$. Although this is a departure from the strict error components model of Gottschalk and Moffitt (1994, 2006), we believe that the residual from this model should provide an improved decomposition of permanent and transitory components and serves as a useful complement. That said, we err on the side of caution and emphasize the total volatility estimates in much of the ensuing discussion.

A. Trends in Income Volatility

In Figures 15 and 16 we depict trends in total, permanent, and transitory income volatility over the past 25 years. Income volatility was relatively stable from 1979 through 1995, but the ensuing decade has been characterized by an unprecedented growth in income volatility of over 60 percent. Although there was a clear countercyclical pattern in volatility through the mid 1990s, the cycle hovered around 0.6 with no evidence of a trend upward or downward. However, a clear break in trend volatility emerged in the later 1990s among single mother families, with the bulk of this movement being driven by transitory volatility and not permanent variation. This result coincides with Gottschalk and Moffitt's (2006) results for the earnings of white men. Many in the policy community touted the rise of employment and decline in welfare

participation in the late 1990s as barometers of success of welfare reform and EITC expansions. It appears as though this period was also characterized by historic growth in income instability. As expected, in Figure 16 we see that adding the observed demographics yields higher regression adjusted permanent volatility estimates relative to the base case with just cohort-year effects; however, the basic result of transitory factors driving rising volatility remains.

[Figures 15 and 16 here]

We next explore whether rising income volatility is being driven by certain subgroups of single mothers or whether it is an across-the-board increase in instability. In Figures 17 and 18 we depict trends in total income volatility of mothers under age 35 and those age 35 to 45. As many single mothers on welfare or at risk of entering welfare are young, it is instructive to understand whether the growth in volatility is concentrated among young mothers. Although the growth in instability is more transparent among mothers under age 35 in Figure 17 (owing in part to larger sample sizes), for both age groups volatility grew nearly 60 percent between 1995 and 2004 and thus it appears that rising instability was not concentrated among the young. In Figures 19 and 20 we depict trends in volatility for mothers with less than high school education and those with more than high school, respectively. Although low-skilled mothers are at much higher risk of entering welfare, it appears that growing instability was not restricted to this demographic group. Indeed, the magnitude of growing volatility post 1995 was U-shaped—62 percent among less than high school, 40 percent among high school (not depicted), and 106 percent among mothers with more than high school. In the 1990s some of the largest gains in employment were among never-married single mothers. In Figures 21 and 22 we depict trends in volatility among never-married single mothers and previously married (or separated) single mothers, respectively. Both figures show comparable (56 and 64 percent, respectively) growth in instability, reinforcing

the fact that growing income instability affected all segments of the single mother population, and virtually all of the growth is attributed to transitory factors.[3]

[Figures 16-22 here]

B. Volatility of Labor Market Earnings and Transfer Income

Increasing income instability could arise from any one of a number of the income sources making up total income, ranging from labor-market earnings to public and private transfers. With the decline in AFDC/TANF and food stamp participation and the rise in employment, an obvious conjecture is that volatility rose among both transfers and earnings after 1995. This may or may not have been the case as the volatility of any given income subcomponent depends on the relative role of changes in the extensive margin (in work or welfare) and intensive margin (hours of work or length of welfare spell).

To see this possible interaction between the extensive and intensive margins let I_j be income component j, j = earnings, AFDC/TANF, food stamps, SSI, SSDI, EITC, and other, and let P_j be an indicator variable equal to one if a single mother participates in a given program j and zero otherwise. The unconditional variance of I_j can thus be written as

$$(3) \qquad V(I_j) = V(E\{I_j \mid P_j\}) + E\{V(I_j \mid P_j)\}$$

where E is the expectations operator. The first term in equation (3) is the variance of the conditional mean, and noting that $E\{I_j \mid P_j = 0\} = 0$ we can write the first term as

$$(4) \qquad V(E\{I_j \mid P_j\}) = (E\{I_j \mid P_j = 1\} - E\{I_j\})^2 \times \Pr(P_j = 1) + E\{I_j\}^2 \times \Pr(P_j = 0)\,.$$

Likewise, noting that $V(I_j \mid P_j = 0) = 0$ we can write the second term in equation (3) as

[3] The near zero contribution of permanent volatility in the figures showing splits based on education attainment is purely an artifact of how we construct cohorts. Recall that a cohort is defined as the interaction of a 5-year birth window with education attainment. Thus when focusing on a specific level of education as in Figures 19 and 20, all of the permanent component is attributed to differences in birth year because education is held constant.

(5) $E\{V(I_j \mid P_j)\} = V(I_j \mid P_j = 1) \times \Pr(P_j = 1)$,

and substituting (4) and (5) into (3) yields

(6) $V(I_j) = (E\{I_j \mid P_j = 1\} - E\{I_j\})^2 \times \Pr(P_j = 1) + E\{I_j\}^2 \times \Pr(P_j = 0) + V(I_j \mid P_j = 1) \times \Pr(P_j = 1)$.

Equation (6) makes clear that the variance of earnings or transfer income depends on whether the participation rate and/or the conditional mean are rising, falling, or remaining the same. For example, suppose that the mean of earnings conditional on working is held constant but the probability of working rises. This implies that the unconditional mean $E\{I_j\}$ increases, and thus $(E\{I_j \mid P_j = 1\} - E\{I_j\})^2$ falls. Because $\Pr(P_j)$ is rising if $E\{I_j\} > E\{I_j \mid P_j = 1\} - E\{I_j\}$ (which is true if $P_j >$ ½), then a rise in P_j will have a net effect of causing the first two terms in equation (6) to fall, but the third term will be increasing and thus the variance of earnings could rise simply because more people are in the labor force.

Although we do not attempt to isolate each of the terms in equation (6), we do admit changes in the extensive and intensive margins into our estimated income component variances by employing a Tobit estimator for each of the I_j. Under the Tobit model, which is a mixture model of a discrete and continuous distribution, the unconditional mean for an observation randomly drawn from the population, whether censored at zero or not, is

(7) $E\{I_{jict}\} = \Phi\left(\frac{z'_{ict}\psi_j}{\sigma_j}\right)[z'_{ict}\psi_j + \sigma_j \lambda_{jict}]$

where Φ is the cumulative normal distribution, $\lambda_{jict} = \frac{\phi\left(z'_{jict}\psi_j / \sigma_j\right)}{\Phi\left(z'_{jict}\psi_j / \sigma_j\right)}$ is the so-called Inverse Mills Ratio, ϕ is the normal density function, and *z* is the vector of age and cohort-year variables. The Tobit residuals used in constructing the total variance of each income source using

equation (2) are then constructed by subtracting the unconditional mean from income, $I_{jict} - E\{I_{jict}\}$, after replacing the unknown ψ_j with $\hat{\psi}_j$.

We are also interested in isolating volatility at the intensive margin separate from the unconditional derived from equation (7). For the reasons given above, selection into the labor force or into a transfer program is unlikely to be random, and thus the standard OLS conditional mean ignoring zeros in the dependent variable is unlikely to be consistent. We thus construct residuals from the subpopulation of participants as $I_{jict} - E\{I_{jict} \mid P_{jict} = 1\}$, where $E\{I_{jict} \mid P_{jict} = 1\} = z'_{jict}\psi_t + \sigma_j \lambda_{jict}$ is estimated using the standard two-step procedure of Heckman (1979).

[Figures 23 and 24 here]

In Figures 23 and 24 we present Tobit-based estimates of earnings volatility of single mothers that includes both labor-market participants and non-participants and two-step selection-corrected estimates based on participants alone, respectively. In Figure 23 there is a clear counter-cyclical pattern of earnings volatility as it rises in the years surrounding the recessions of 1981, 1991, and 2001, and declines in the expansionary periods. Much like total income volatility in Figure 15, earnings volatility cycled around a stable trend in the 1980s and through the mid 1990s. However, between 1995 and 2000 earnings volatility fell substantially as total income volatility rose. This is also true at the intensive margin alone as depicted in Figure 24. The two notable distinctions between Figures 23 and 24 are the differences in variance levels and differences in trends post 2000. When labor force nonparticipants are included the estimated level of volatility is ten times larger than for participants alone. And post 2000, earnings volatility rose when nonparticipants are included, but it was comparatively stable with only

participants. Similar to total income volatility, and consistent across both figures, is the fact that the bulk of the instability in earnings is transitory and not permanent.

[Figures 25 to 36 here]

With total income volatility rising after 1995, but earnings instability falling (at least until 2000), then the growth of income instability must derive from non-labor market income sources such as cash and in-kind transfers. In Figures 25 to 36 we present Tobit and Heckman two-step estimates of volatility for each of the major income sources for participants and nonparticipants combined in the case of the Tobit estimates and for participants alone in the case to the two-step estimates. Figures 25 and 26 depict trends in volatility of EITC income. With nonparticipants included, EITC income volatility grew over 80 percent between 1986 and 1995, but was relatively flat after 1995. The growth in the late 1980s and into the mid 1990s coincided with the expanded generosity and take-up of the EITC as part of the 1986, 1990, and 1993 tax reforms. Although the total volatility of EITC income was stable post 1995, there does appear to be a marked increase in permanent volatility after 2002. Moreover, as shown in Figure 26, total EITC volatility grew about 40 percent at the intensive margin between 1995 and 2004, and this may have contributed to rising total income volatility.

Figures 27 and 28 depict TANF income volatility, and much like earnings, TANF income volatility fell by half after 1995 when both the extensive and intensive margins are incorporated. At the same time, TANF income volatility rose at the intensive margin by 36 percent between 1995 and 2002 prior to tapering off in the past two years. This suggests that overall the decline in TANF participation drove down total TANF income volatility, but among those recipients still on the program, income was less stable. On the contrary, Figures 29 and 30 demonstrate a marked rise in SSI income volatility. The Tobit-based residual variances in Figure 29 rose 40

percent after the 1990 *Zebley* decision and during the welfare waiver period, but subsequently fell after 1998. However, volatility at the intensive margin of SSI in Figure 30 rose 90 percent between 1995 and 2004.

In the remaining four figures it is clear that income from Social Security and Disability Insurance, from food stamps, and from other income sources are relatively stable compared to the previously discussed components of total income. Taken as a whole, a rather complicated picture emerges for the growing volatility of income among single mother families in the United States after the passage of PRWORA in that for many income sources there are competing forces at the extensive and intensive margins. Although the descriptive analysis presented thus far cannot identify causal channels, the figures do suggest that much of the recent widening income volatility appears to be driven by increased volatility at the intensive margin.

C. Testing for Structural Change in Volatility

The figures on volatility are striking in that for most of the income series there appears to be visual evidence of a structural break in trend volatility. In some cases, such as total income volatility, the change point in 1996 seems fairly transparent such that one may be fairly comfortable in parametrically modeling a trend break in 1996. However, in other cases the actual change point may be a bit more subtle. For example, for earnings volatility with both participants and nonparticipants in Figure 23 it is not clear whether the change point is in 1993, 1994, or 1995, and possibly a second change point after 2000, and thus one could easily misspecify the model by incorrectly choosing the break point parametrically.

In lieu of arbitrarily choosing break points in the volatility series, we turn to the recent time series literature on testing for structural change with an unknown change point (Andrews 1993; Bai 1997; Hansen 2001). The new tests build off an idea due to Quandt (1960), who

proposed splitting the sample at every possible breakdate, estimating the model parameters on the remaining data, and constructing the associated Chow test statistic with each sample split. The estimated breakdate is that sample split with the largest value of the Chow test statistic. If the breakdate is known a priori then one can appeal to the usual chi-square tables for critical values. However, in many cases the breakdate is not known and the chi-squared critical values are not valid. Andrews (1993) developed the asymptotic theory for the case of unknown change point and provided tables of critical values, and consequently the new tests are generally known as the Quandt-Andrews statistic. As noted by Hansen (2001), this method of least squares testing for structural change is valid for the linear regression model with homoskedastic variances, and Bai (1997) proposed a straightforward method of constructing confidence intervals around the estimated breakdate.

We are interested in estimating whether there is a statistically identifiable trend break in income and earnings volatility, the date of the possible break, and the associated confidence interval around the estimated breakdate. To implement the tests we take our estimated time series of total volatilities displayed in several of the figures above and run the following regression:

$$\begin{aligned} \widehat{V(y_t)} &= \sigma_1^2 + \theta_1 t + \zeta_1, t = 1, ..., \tau \\ \widehat{V(y_t)} &= \sigma_2^2 + \theta_2 t + \zeta_2, t = \tau + 1, ..., T \end{aligned} \tag{8}$$

where $\widehat{V(y_t)}$ is the estimated volatility, σ^2 is a constant term (reflecting the constant variance), θ is the coefficient on the linear trend *t*, ζ is an iid random error term, τ is the unknown breakdate, and $T = 26$ for the years 1979–2004. For each possible breakdate τ we conduct the joint test of the null hypothesis of constant variance ($\sigma_1^2 = \sigma_2^2$ *and* $\theta_1 = \theta_2$) by constructing the following Wald test statistic

(9) $$W = (SSE_{pooled} - (SSE_1 + SSE_2)) / \left((SSE_1 + SSE_2)/(T - 2*k)\right)$$

where SSE_{pooled} is the sum of squared errors for the pooled regression with no break, SSE_1 and SSE_2 are the sum of squared errors for the pre- and post-break periods, respectively, and *k* is the number of parameters in each subsample (= 2 is this case). The estimated breakdate is the $\hat{\tau}$ with the maximum test statistic $\hat{W}$, i.e. the supWald statistic. The associated Bai (1997) confidence interval for $\hat{\tau}$ with trending regressors is $\left[\hat{\tau} \pm (\frac{c}{\hat{L}} + 1)\right]$. The term *c* is the critical value for a test of size α (*c*=7 when $\alpha = 0.1$ and *c*=11 when $\alpha = 0.05$) and $\hat{L}$ is the outer product of the fitted valued of the regression standardized by the estimated error variance, $\hat{L} = \frac{\hat{\kappa}' Z Z' \hat{\kappa}'}{\hat{\sigma}^2}$, with $\hat{\kappa} = [\hat{\rho}, \hat{\theta}]$ and $Z = [1, t]$ and where *t* is set at the estimated breakdate $\hat{\tau}$ (See Bai (1997 p. 555) for additional details).

[Table 1 here]

In Table 1 we report the Quandt-Andrews supWald test statistics, along with the estimated breakdate, the associated 95 percent confidence interval, and the signs on the trend coefficients in equation (8) for the sample split prior to the breakdate and after the breakdate, i.e. $\hat{\theta}_1\Big|_{t=1}^{\hat{\tau}}$ and $\hat{\theta}_2\Big|_{t=\hat{\tau}+1}^{T}$. We conduct our tests for total income volatility, earnings, EITC, TANF, SSI, SSDI, food stamps, and other income. Except for total income, all tests are based on the Tobit model that includes both participants and nonparticipants (all) as well as the Heckman two-step selection-correction model (participants).

With a critical value of 11 at the 5 percent level the supWald statistic of 80.9 clearly rejects the null of no structural change for total income volatility. As one might suspect from Figure 15 the estimated breakdate occurs in 1996, but highlighting the importance of confidence

intervals, this break occurred statistically sometime between 1994 and 1998 with 95 percent confidence. Because the tests of structural change in Table 1 are not based on a multivariate model we are cautious about attributing the break and subsequent increasing volatility to welfare reform, but we do note that this result persists even when controlling for the lagged level of volatility, the unemployment rate, and the growth rate of real GDP. Indeed, in this expanded model the confidence interval around 1996 is a tighter +/– 1 year, which is as tight as is feasible.

A number of common themes emerge from the remaining rows of Table 1. First, many of the estimated breakdates are associated with policy reforms pertinent to the program in question, especially for results based on the pooled sample of participants and non-participants. The change point for earnings is 1997 plus or minus one year, which coincides with welfare reform. The breakdates of 1986 and 1990 for the EITC are associated with the first major expansions in the EITC with the 1986 and 1990 tax reforms. The change point for TANF is a very tightly estimated one year around passage of PRWORA in 1996. The breakdate of 1992 for SSI came on the heals of the Supreme Court's 1990 *Zebley* decision expanding child eligibility for the program. The change point for SSDI in 1991 is not tied directly to legislative changes (which occurred in 1984), but rather the interaction of the recession, growing inequality, and a rising earnings replacement rate (Autor and Duggan 2006). The food stamp breakdate of 1983 is subsequent to the budget acts of 1981 and 1982 that restricted program eligibility (both directly and indirectly via reduced eligibility for AFDC) and benefit generosity. And, the 1995 breakdate for other income coincides with welfare reform.

A second theme in the table is that the breakdate among participants frequently differs and lies outside the confidence interval from the date associated with the pooled sample of participants and nonparticipants. This underscores the importance of entry and exit into the labor

market and transfer programs. A third theme is that the estimated confidence intervals for participants are typically much wider than for the pooled sample, which again is a likely reflection of entry and exit.

Finally, although we find a strong increase in total income volatility after the 1996 breakdate, with the exception of EITC income and other income among participants, in the last column of Table 1 there is little statistical evidence of a rise in volatility among the other income sources at the same time and thus it is not transparent as to which income source is driving the rise in total income volatility after welfare reform. Figure 23 is suggestive that rising earnings volatility in the last five years of the sample is a major contributor to rising income volatility. Indeed, the trend coefficient for the post-breakdate pooled earnings sample is a positive 0.2 (0.18), which suggests that rising earnings volatility in the early 2000s contributed to rising income volatility, but the post-break trend is not statistically different from 0 or from the pre-break trend of –0.075 (0.042). This lack of statistical significance may simply be the result of small sample sizes post breakdate. At the same, it is important to recognize that total income volatility reflects both the variance of income sources, as well as the covariances among those sources, and that the rise in income volatility may also be driven by the covariance terms. For example, earnings and TANF negatively covary, and given that TANF income volatility fell, the covariance between TANF and earnings likely rose. A similar story exists for food stamps and earnings. As the EITC and earnings covary positively at low earnings levels, and then covary negatively at higher levels, the rise in total income volatility could also be due in part to a rise in labor force participation of single mothers at low earnings levels. In order to isolate these linkages a full model of covariance structures is needed, and is left for future work.

V. Conclusion

We document dramatic changes in the level and composition of income across the distribution during the welfare reform era. At the 10th percentile mothers still rely quite heavily on transfers on average, but post welfare reform fewer than half of mothers even at this low level of income are receiving any cash support from TANF, SSI, Social Security and DI, or earnings. At the same time, average real income at the 25th percentile grew about 32 percent between 1993 and 2004, and grew about 12 percent overall from 1979 to 2004 at the first quartile. This growth was driven by strong increases in earnings and the EITC.

We also estimate the volatility of income among single mother families, and through formal tests of structural change, we identify a large trend rise in income volatility beginning in 1996. This increase is manifested primarily through a rise of transitory volatility. During the same period, TANF income volatility declined substantially. At the same time, SSI and EITC volatility rose, especially at the intensive margin, thus underscoring the importance of accounting for entry and exit into the labor market and transfer programs. A leading candidate for the rise in income volatility among single mothers is rising earnings volatility, though the statistical evidence is not conclusive. A common result is that the Quandt-Andrews tests identify trend breaks in the volatility of total income, earnings, TANF, and other income after passage of welfare reform, suggesting that major legislative reforms may be financially destabilizing for single mother families.

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Figure 1: Mean Income by Source for Single Mothers, Ages 16 to 54, All Education Levels

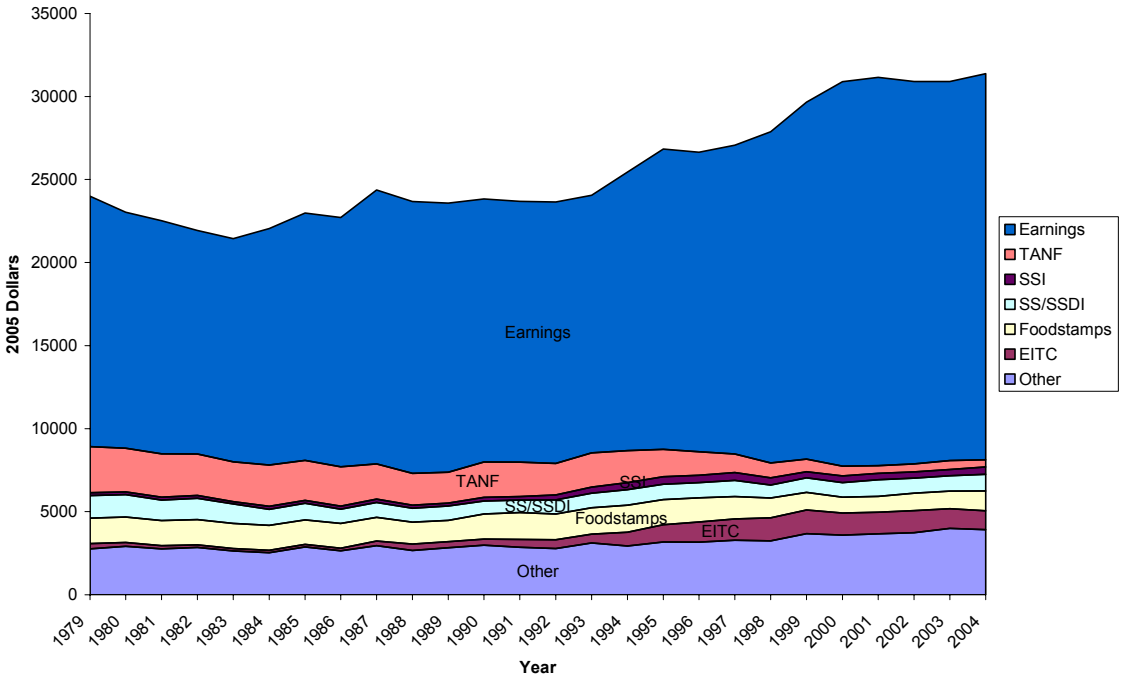

Figure 2: Median Income by Source for Single Mothers, Ages 16 to 54, All Education Levels

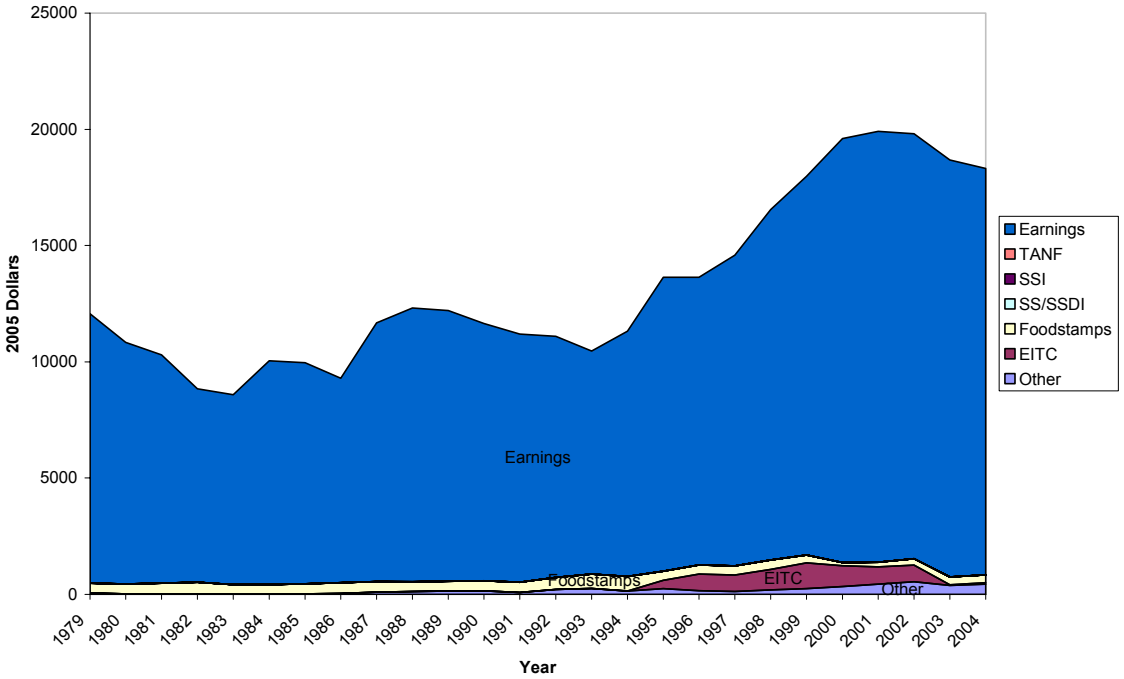

Figure 3: Mean Income by Source for Single Mothers, Ages 16 to 54, Less than High School Education

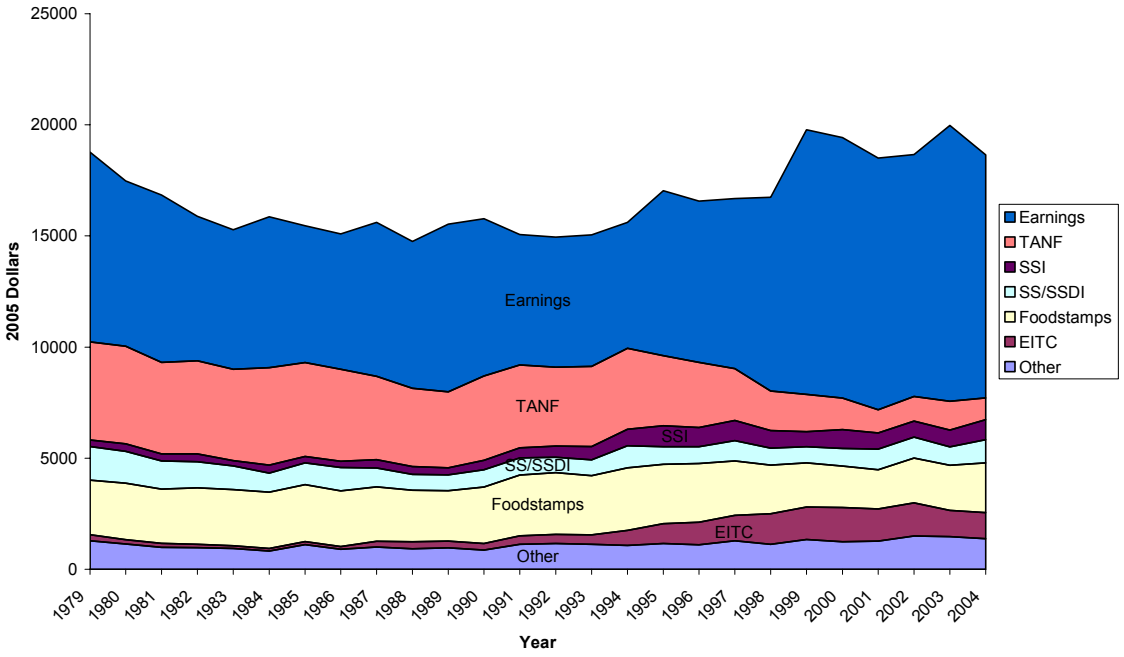

Figure 4: Median Income by Source for Single Mothers, Ages 16 to 54, Less than High School Education

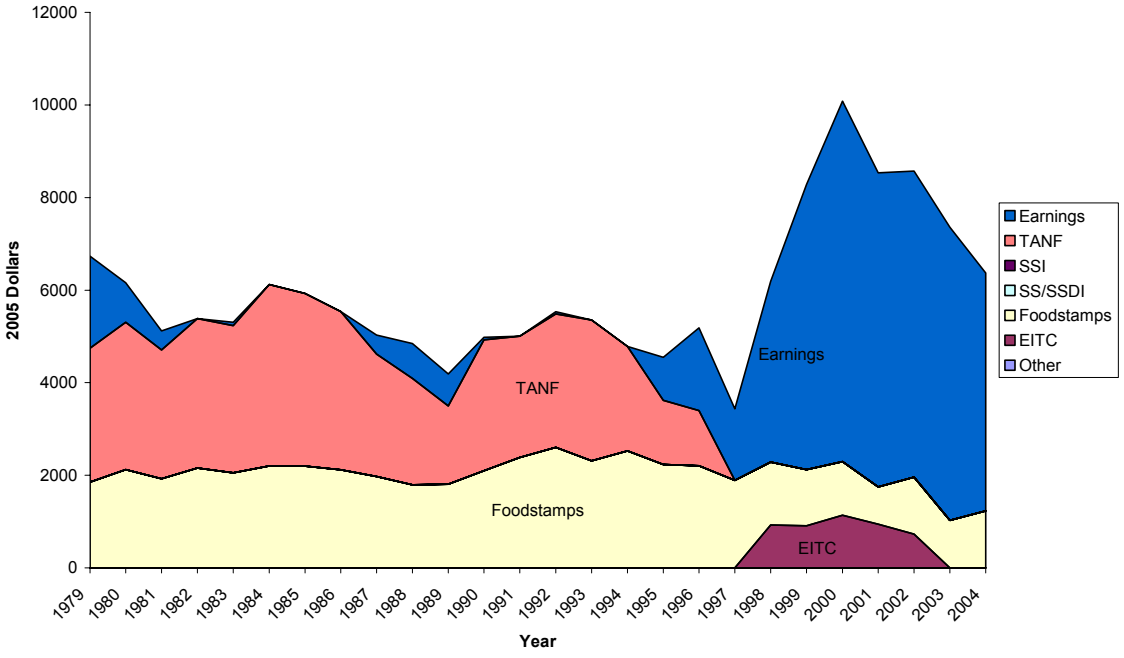

Figure 5: Mean Income by Source for Single Mothers, Ages 16 to 54, High School Education

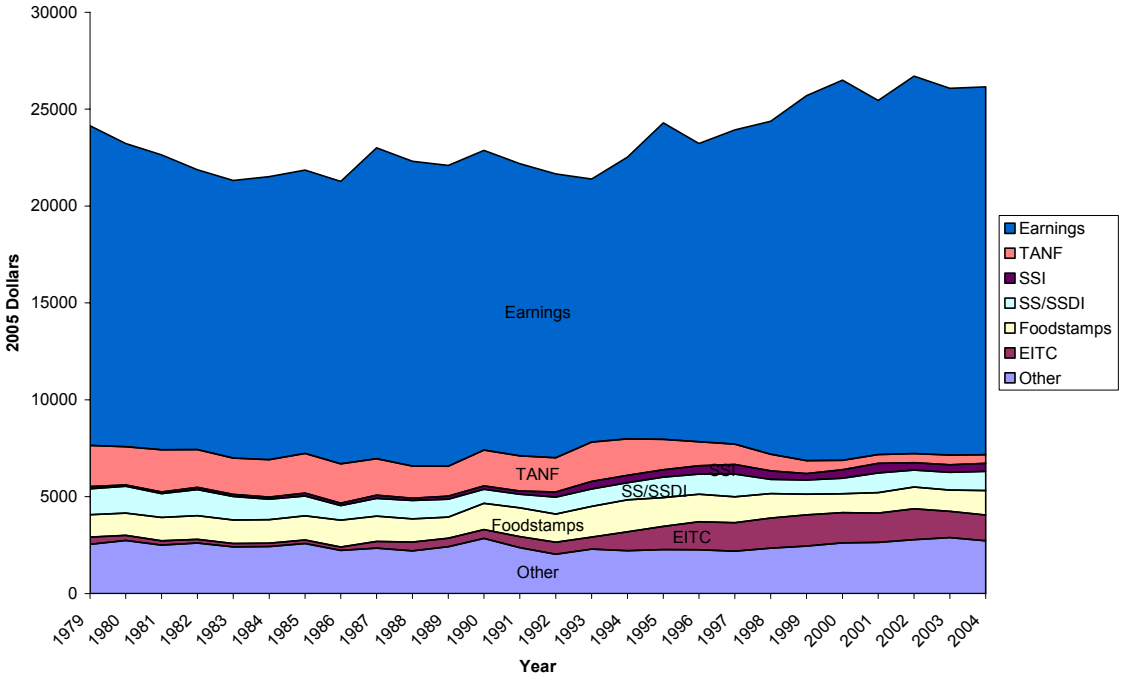

Figure 6: Median Income by Source for Single Mothers, Ages 16 to 54, High School Education

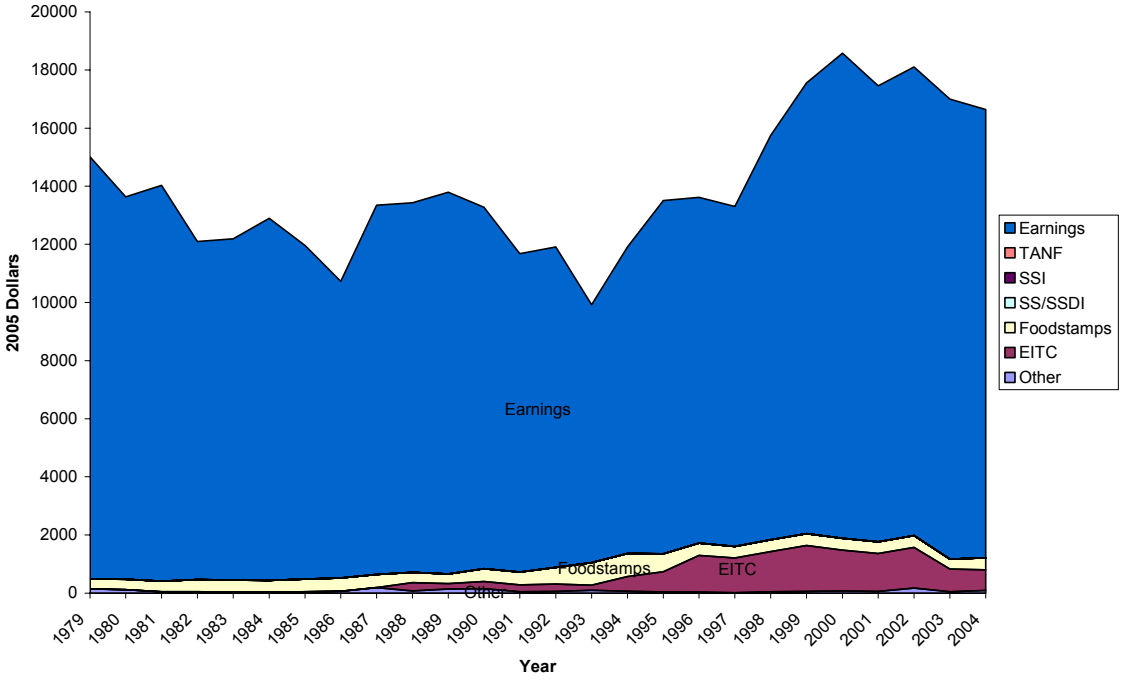

Figure 7: Mean Income by Source for Single Mothers, Ages 16 to 54, More than High School Education

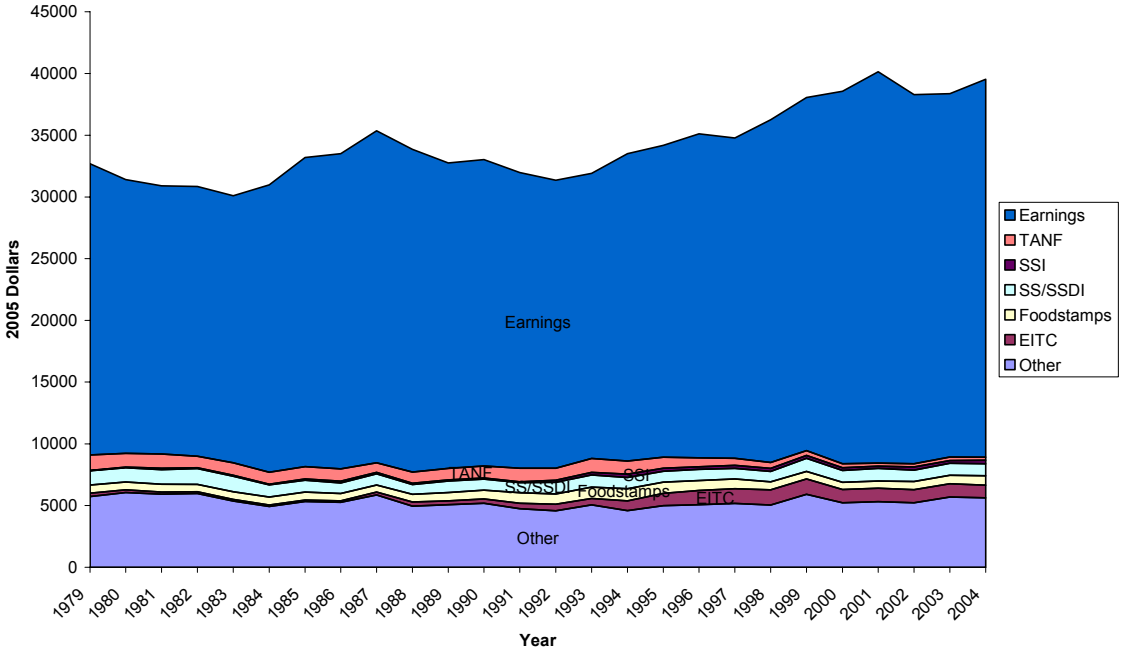

Figure 8: Median Income by Source for Single Mothers, Ages 16 to 54, More than High School Education

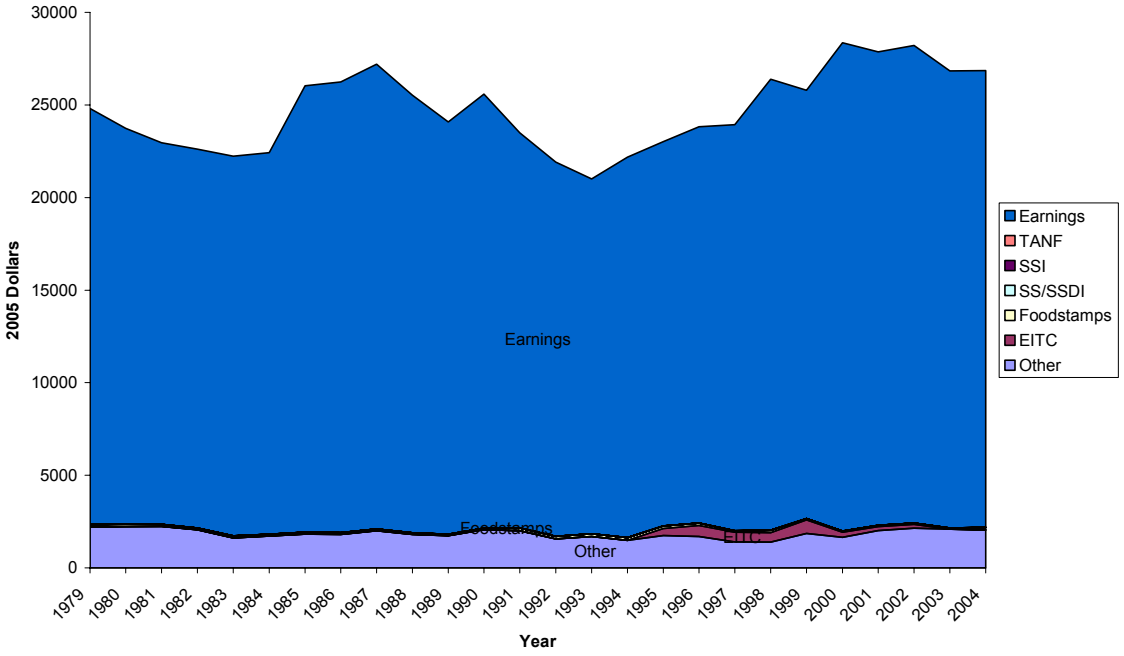

Figure 9: Mean Income by Source for Single Mothers, Ages 16-54, 10th Percentile of Income Distribution

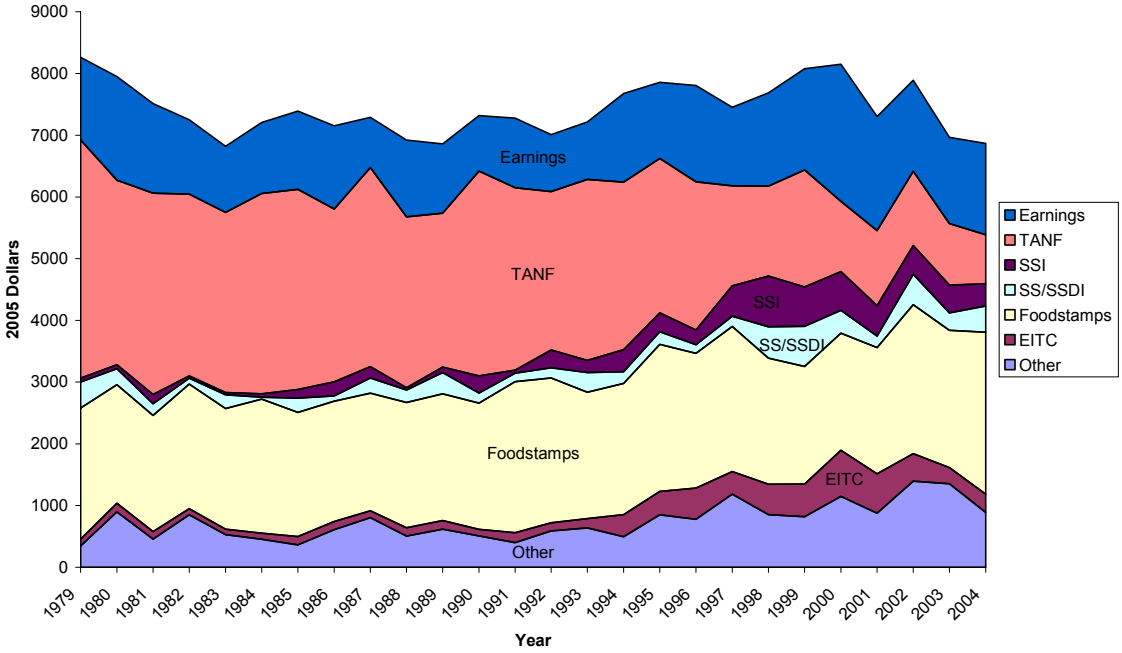

Figure 10: Median Income by Source for Single Mothers, Ages 16-54, 10th Percentile of Income Distribution

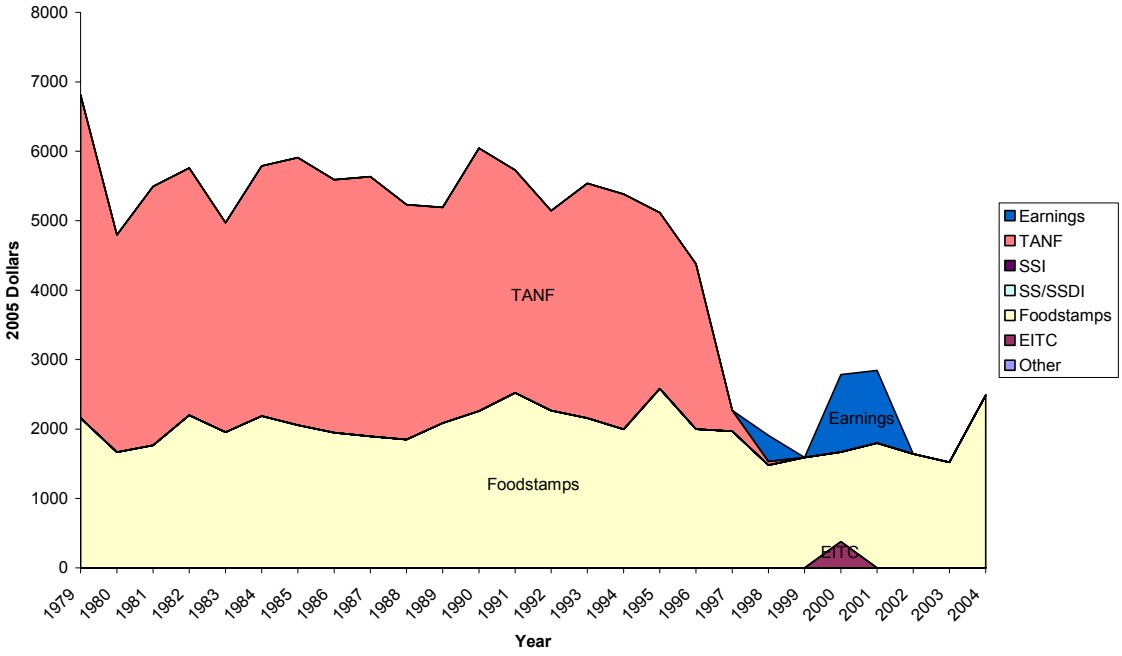

Figure 11: Mean Income by Source for Single Mothers, Ages 16-54, 25th Percentile of Income Distribution

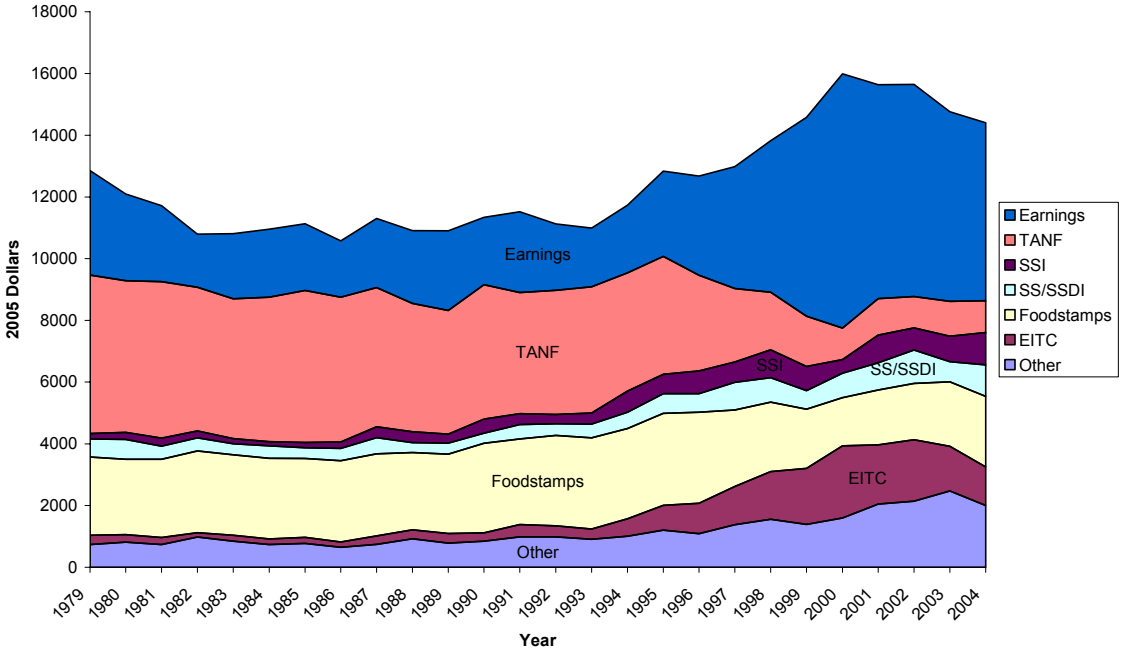

Figure 12: Median Income by Source for Single Mothers, Ages 16-54, 25th Percentile of Income Distribution

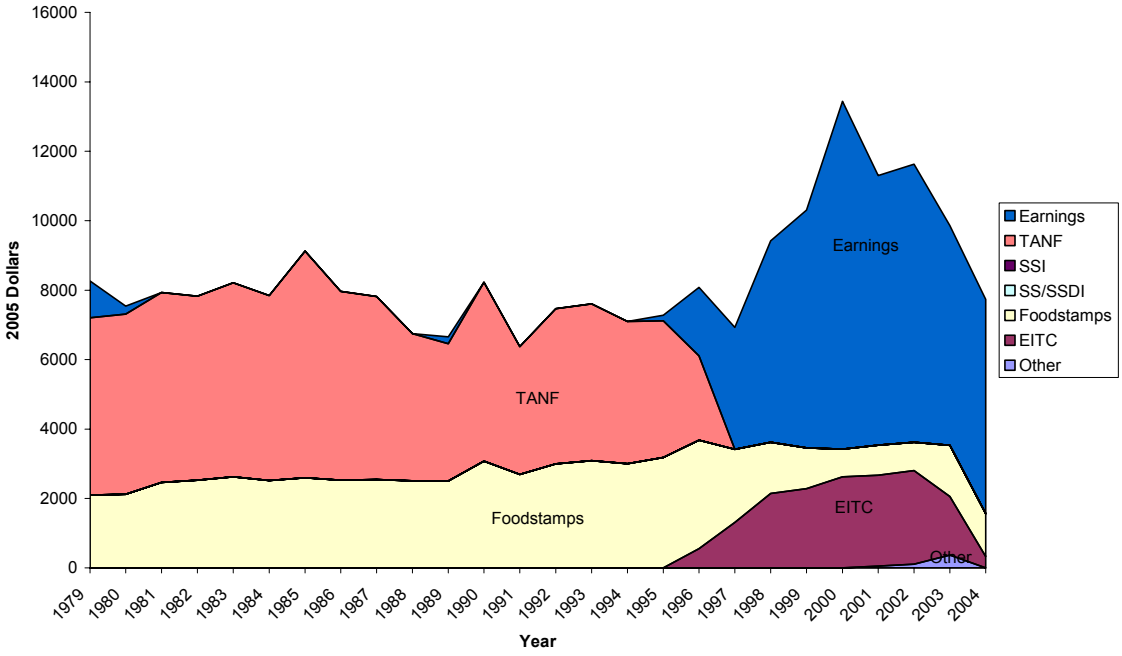

Figure 13: Mean Income by Source for Single Mothers, Ages 16-54, Median of Income Distribution

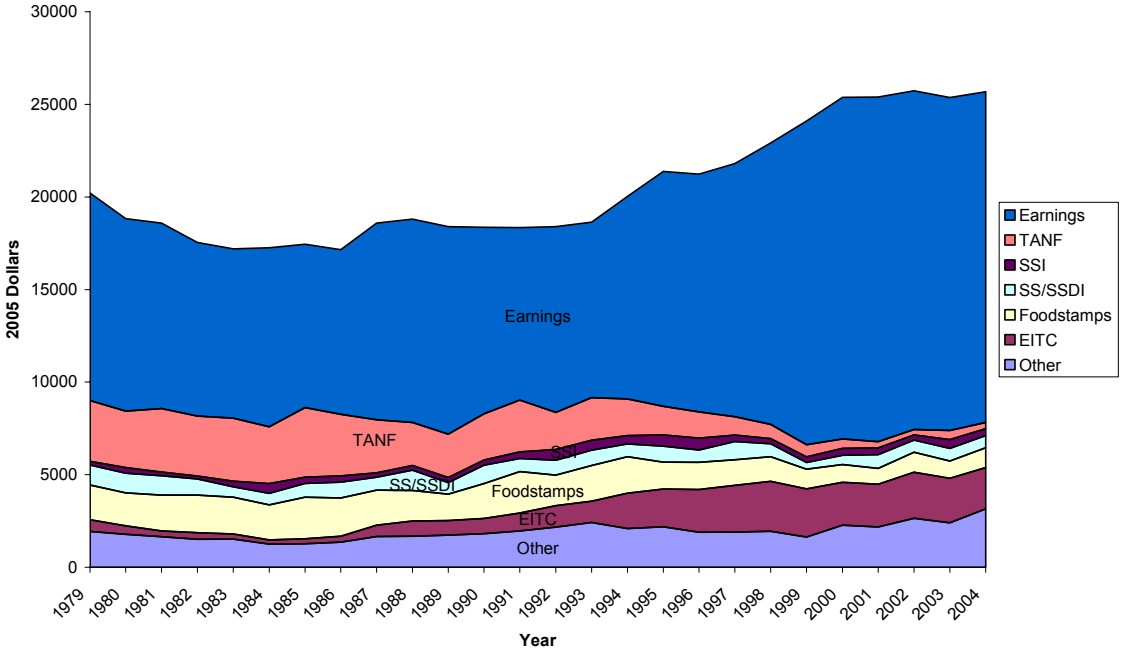

Figure 14: Median Income by Source for Single Mothers, Ages 16-54, Median of Income Distribution

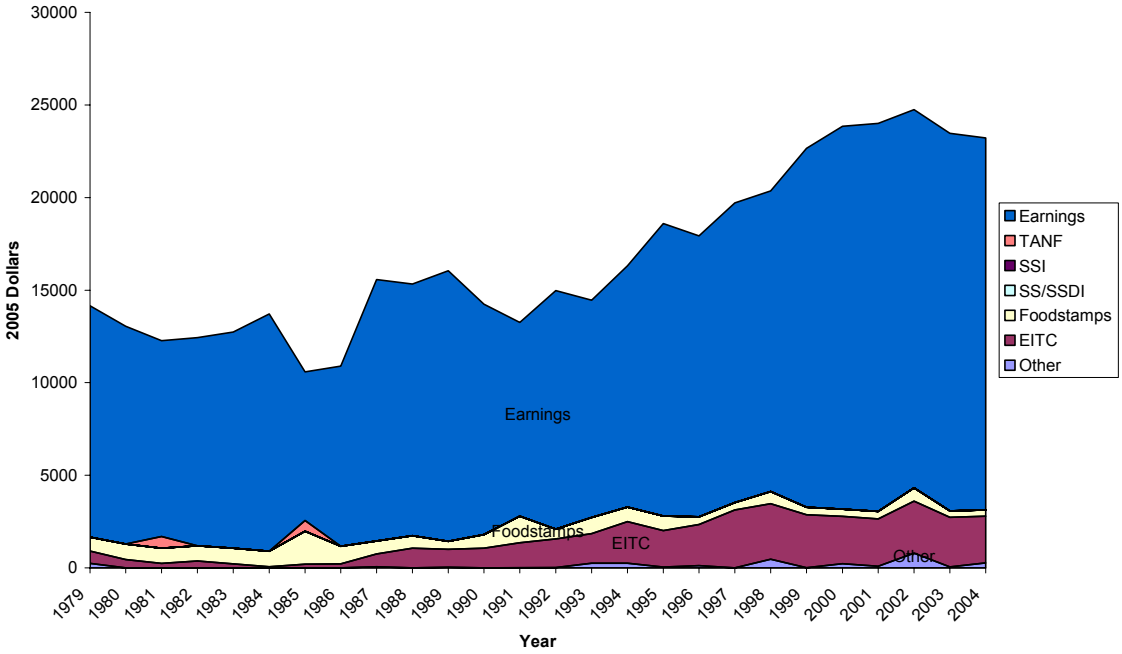

Figure 15: Total Income Volatility of Single Mothers, Ages 16 to 54

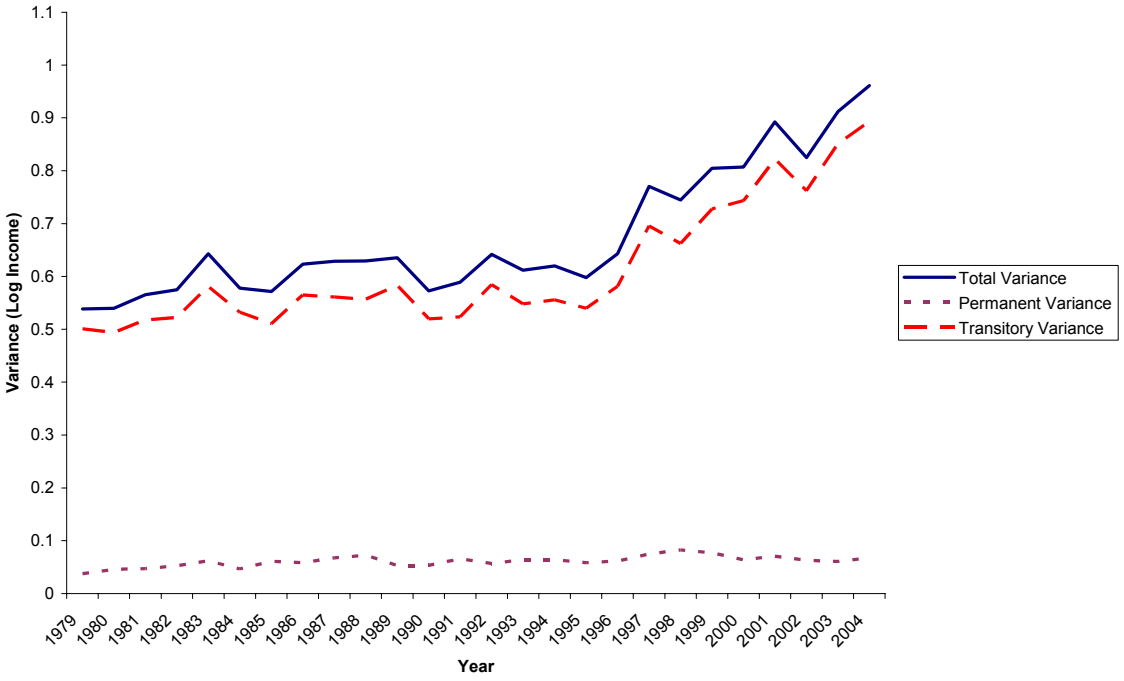

Figure 16: Total Income Volatility of Single Mothers, Ages 16 to 54, Regression Adjusted

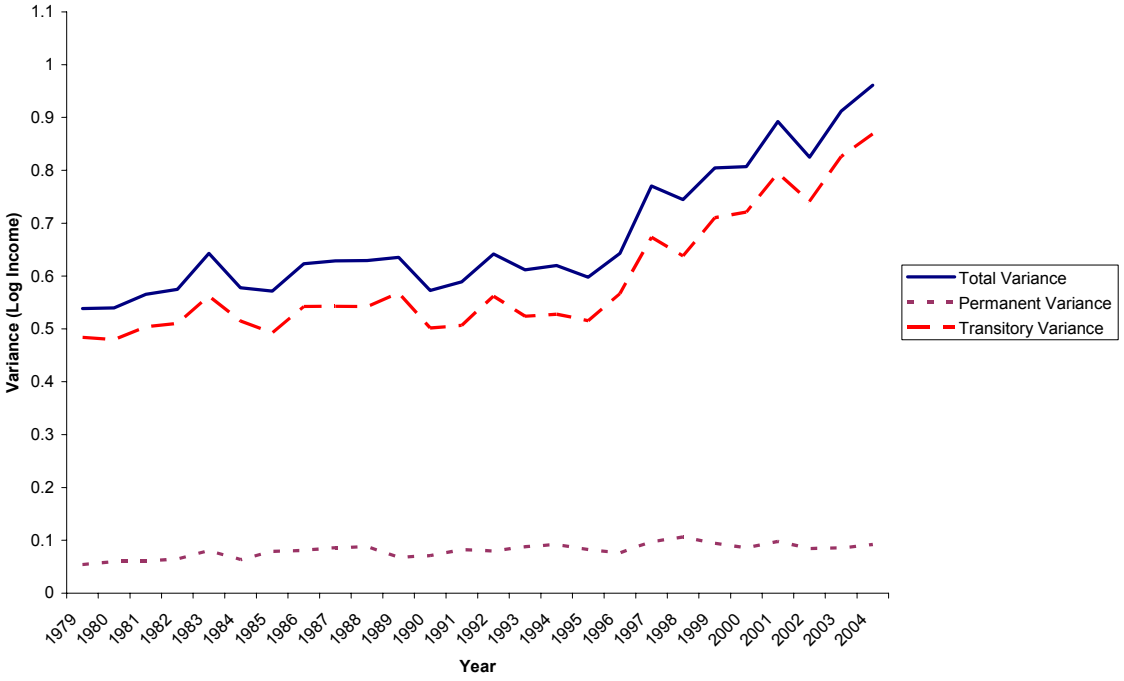

Figure 17: Total Income Volatility of Single Mothers under Age 35

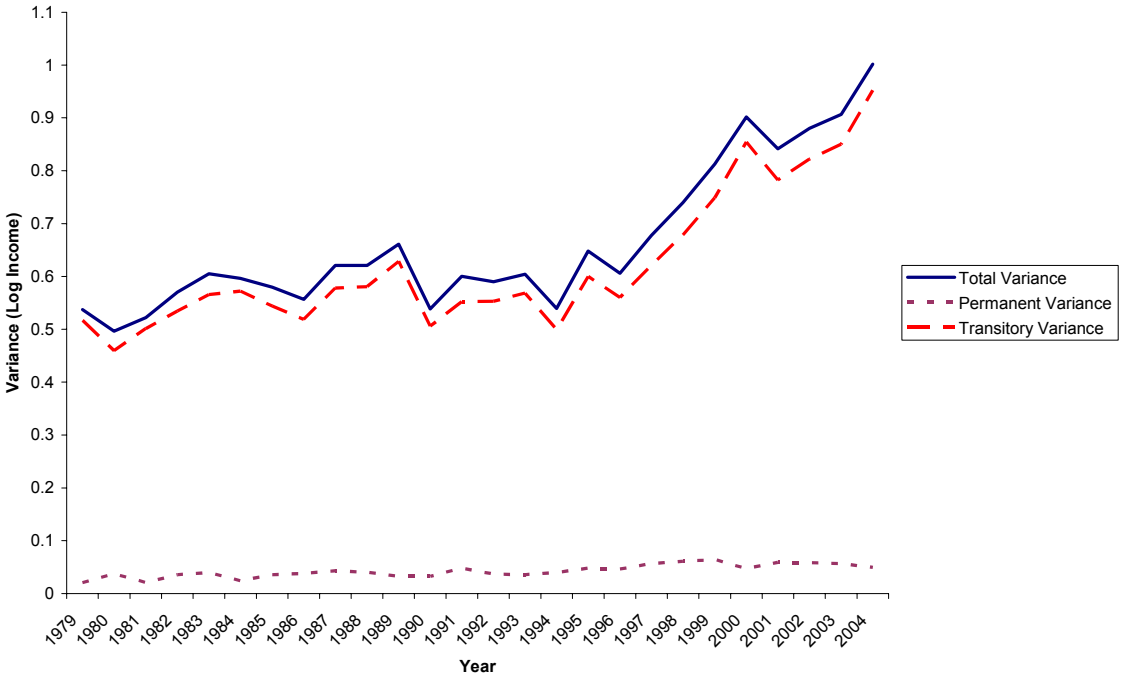

Figure 18: Total Income Volatility of Single Mothers, Ages 35 to 45

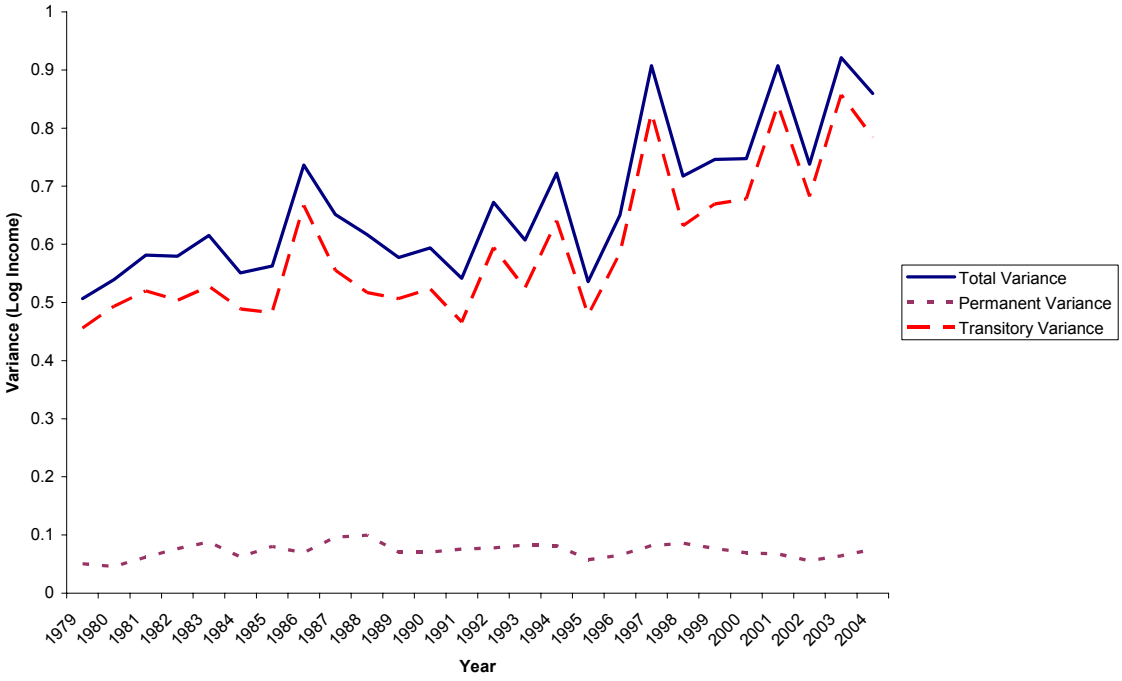

Figure 19: Total Income Volatility of Single Mothers with Less Than High School Education

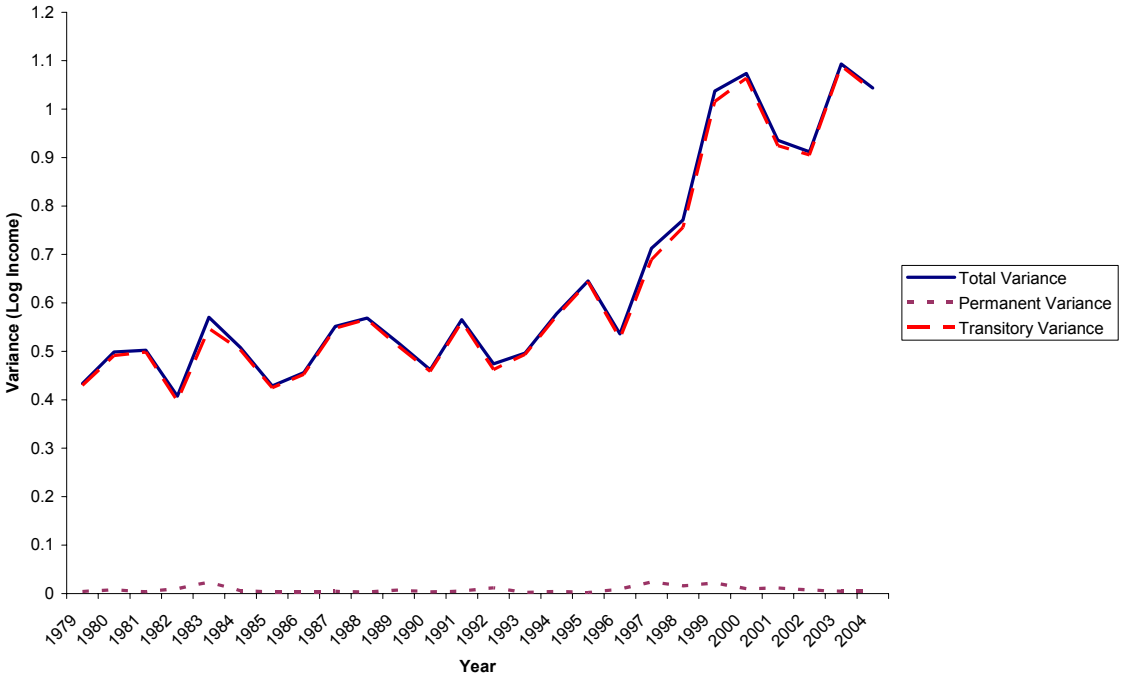

Figure 20: Total Income Volatility of Single Mothers with More than High School Education

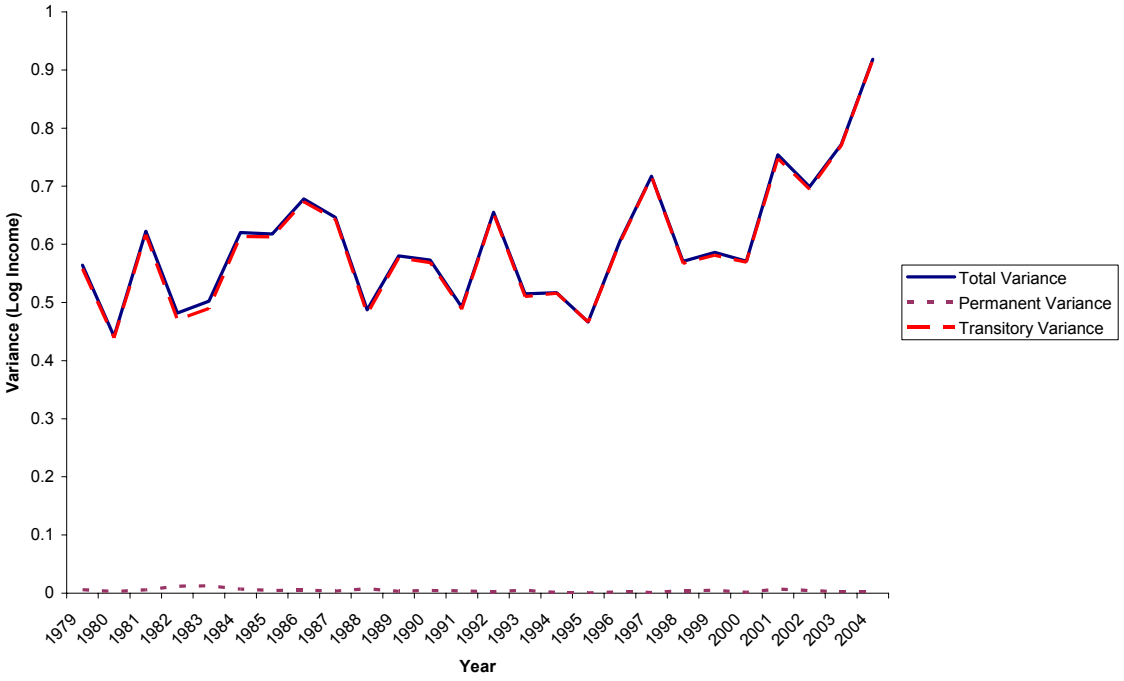

Figure 21: Total Income Volatility of Never-Married Single Mothers

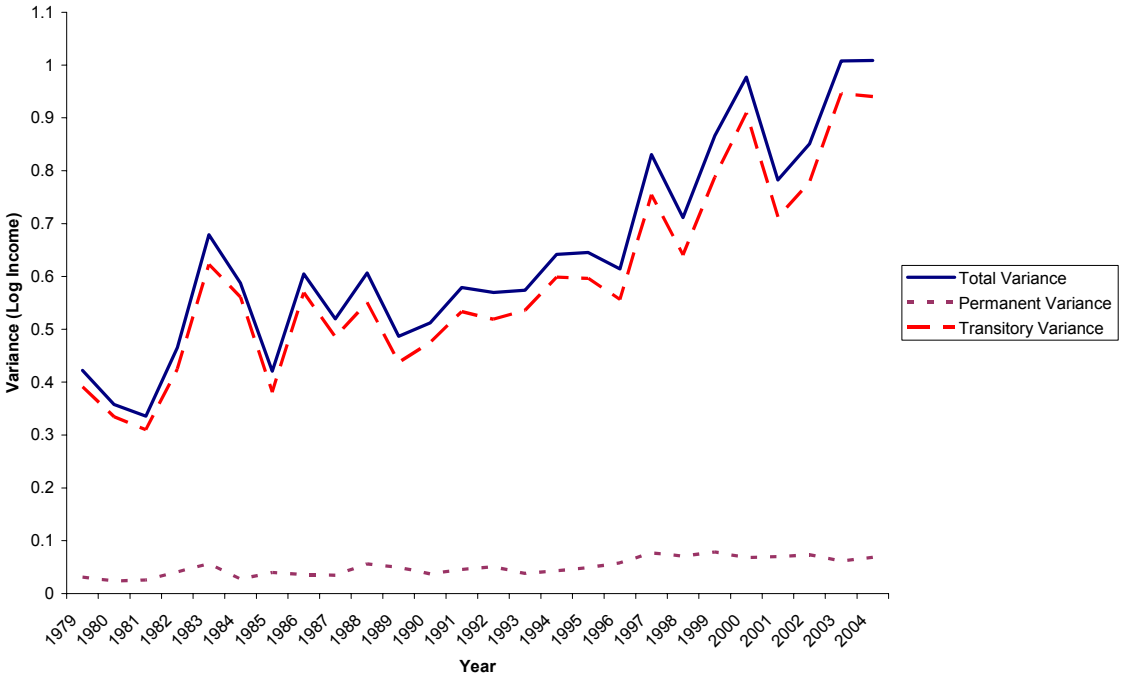

Figure 22: Total Income Volatility of Previously Married Single Mothers

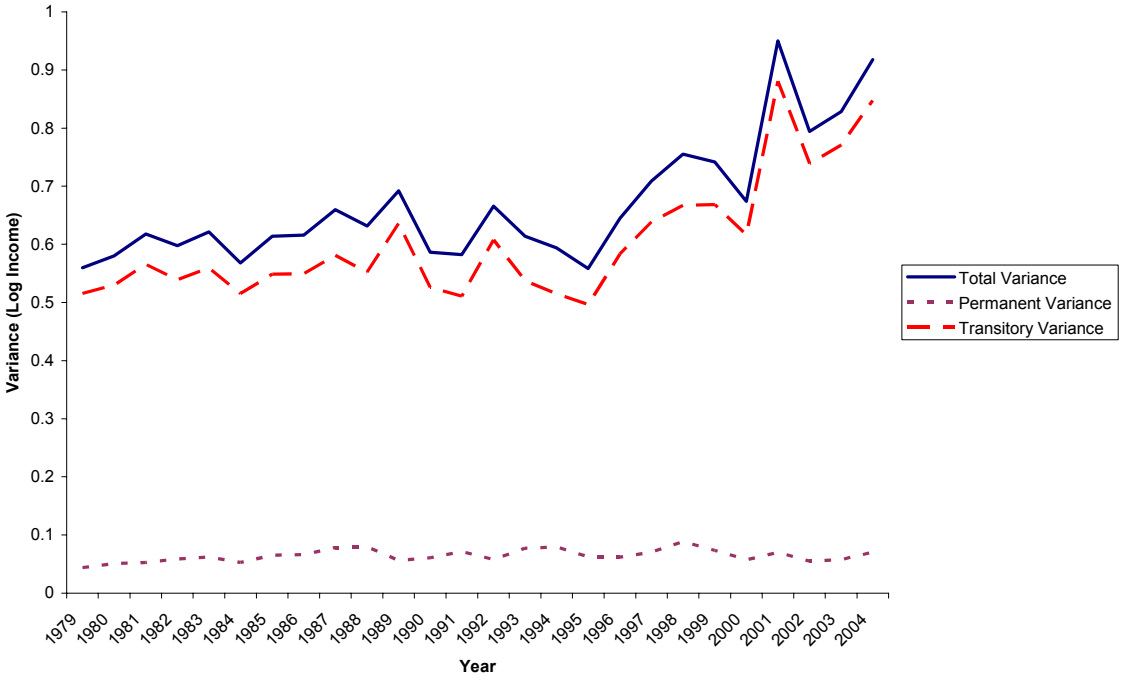

Figure 23: Earnings Volatility of Single Mothers, Participants and Nonparticipants

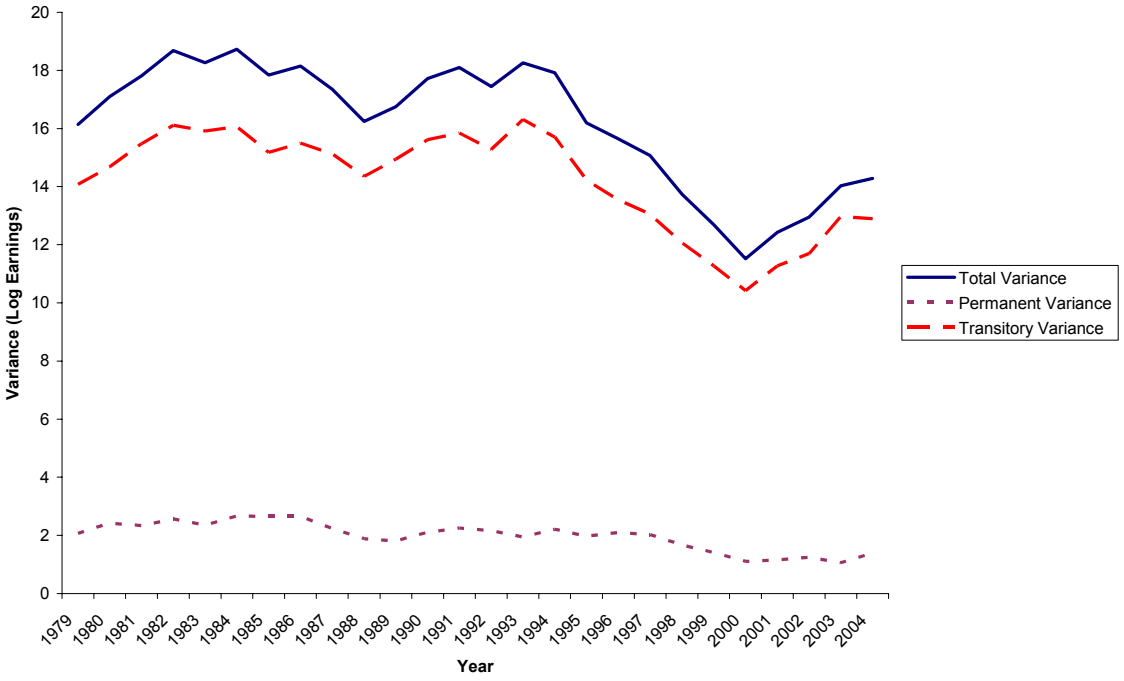

Figure 24: Earnings Volatility of Single Mothers, Participants Only

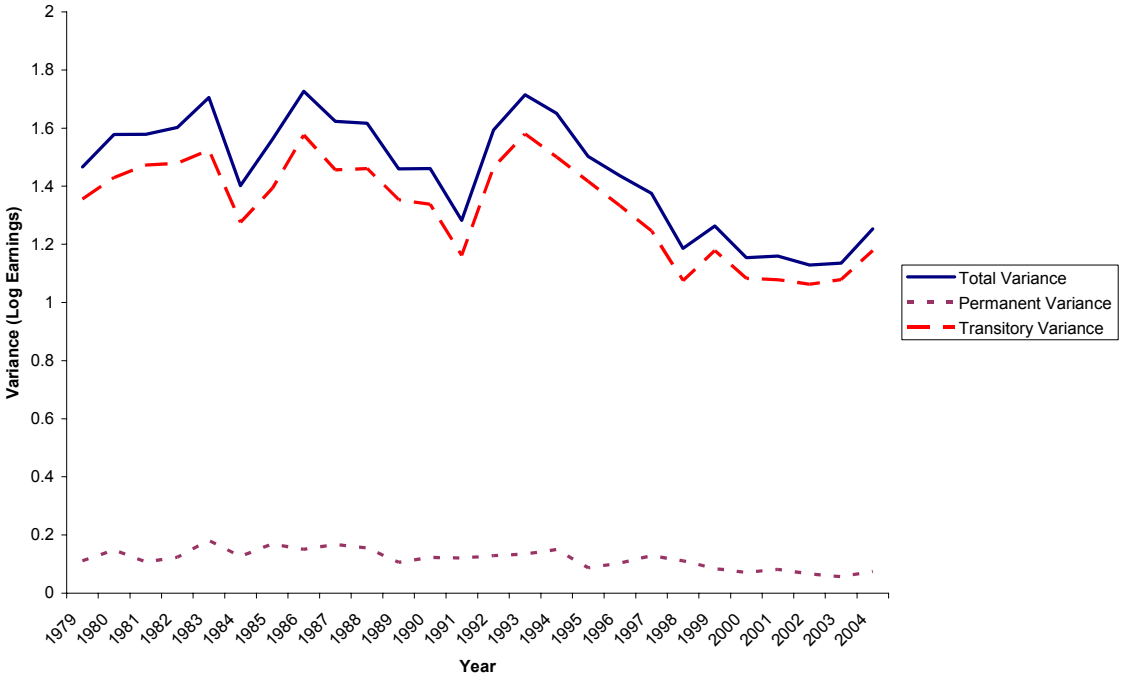

Figure 25: EITC Income Volatility of Single Mothers, Participants and Nonparticipants

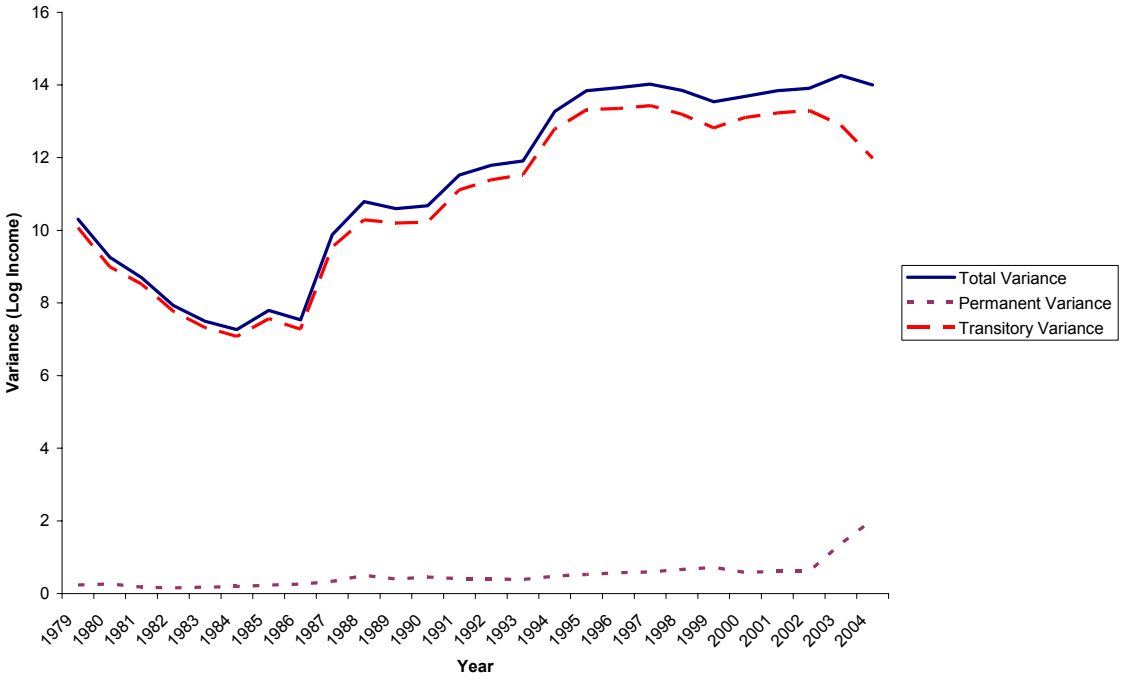

Figure 26: EITC Income Volatility of Single Mothers, Participants Only

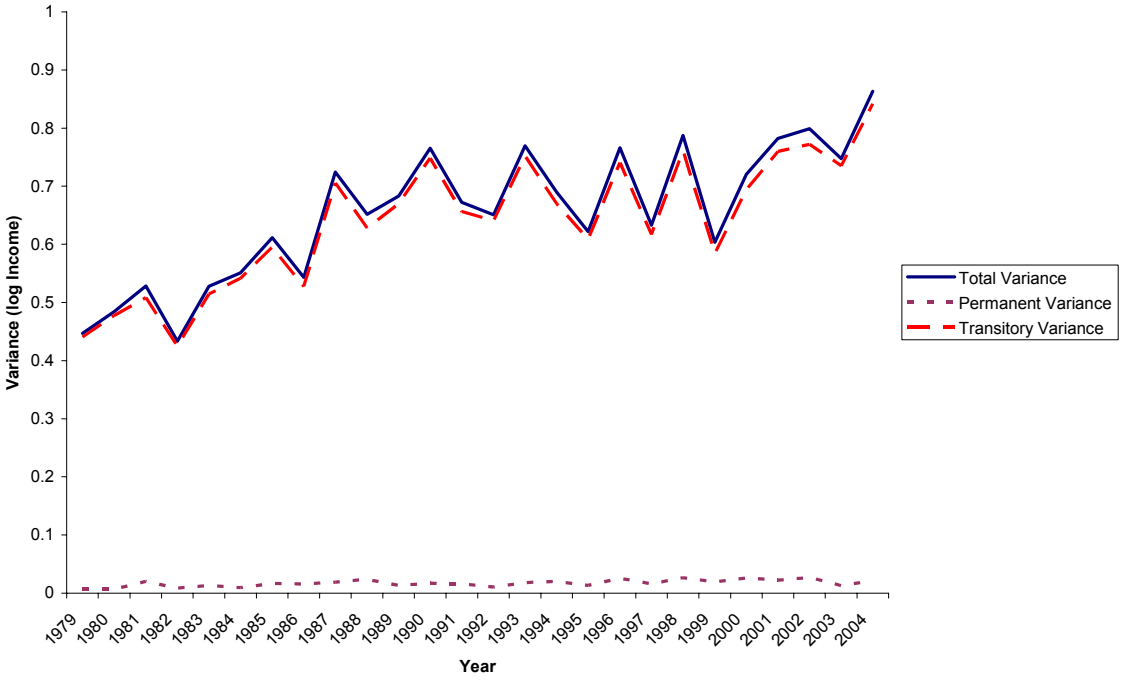

Figure 27: TANF Income Volatility of Single Mothers, Participants and Nonparticipants

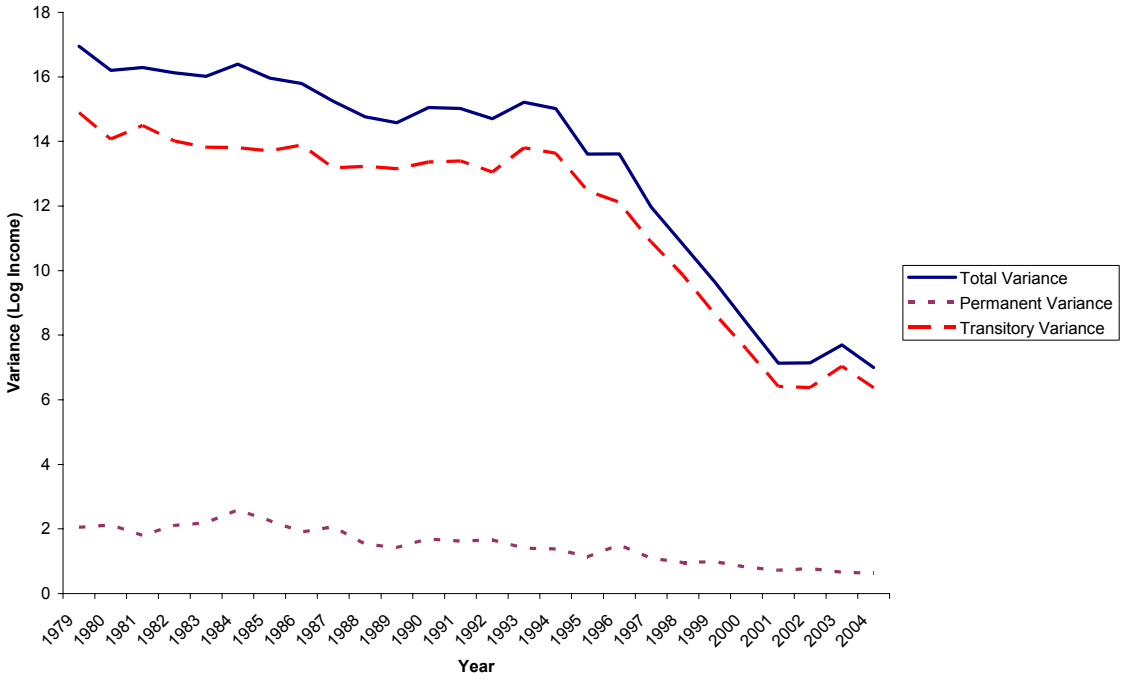

Figure 28: TANF Income Volatility of Single Mothers, Participants Only

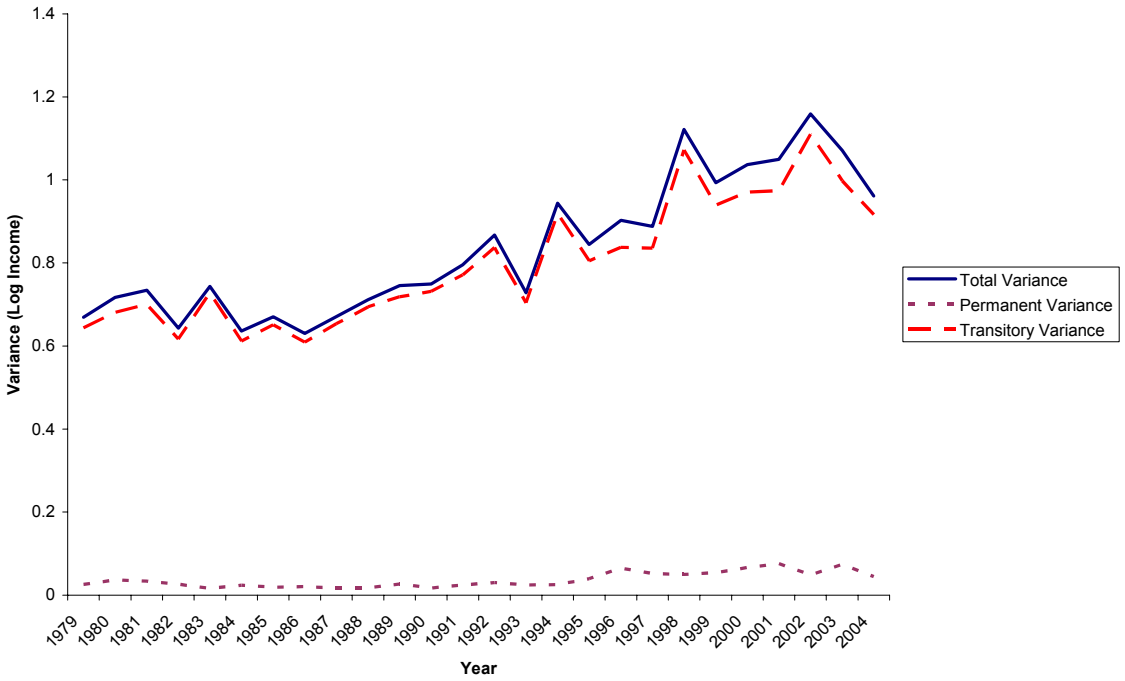

Figure 29: SSI Income Volatility of Single Mothers, Participants and Nonparticipants

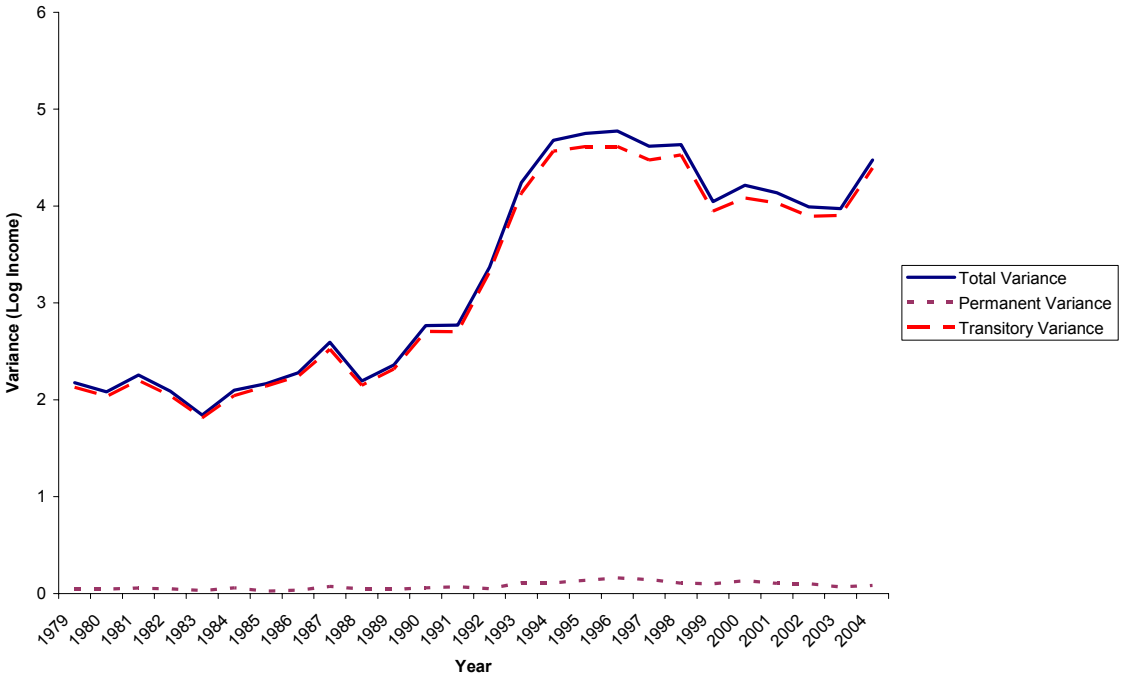

Figure 30: SSI Income Volatility of Single Mothers, Participants Only

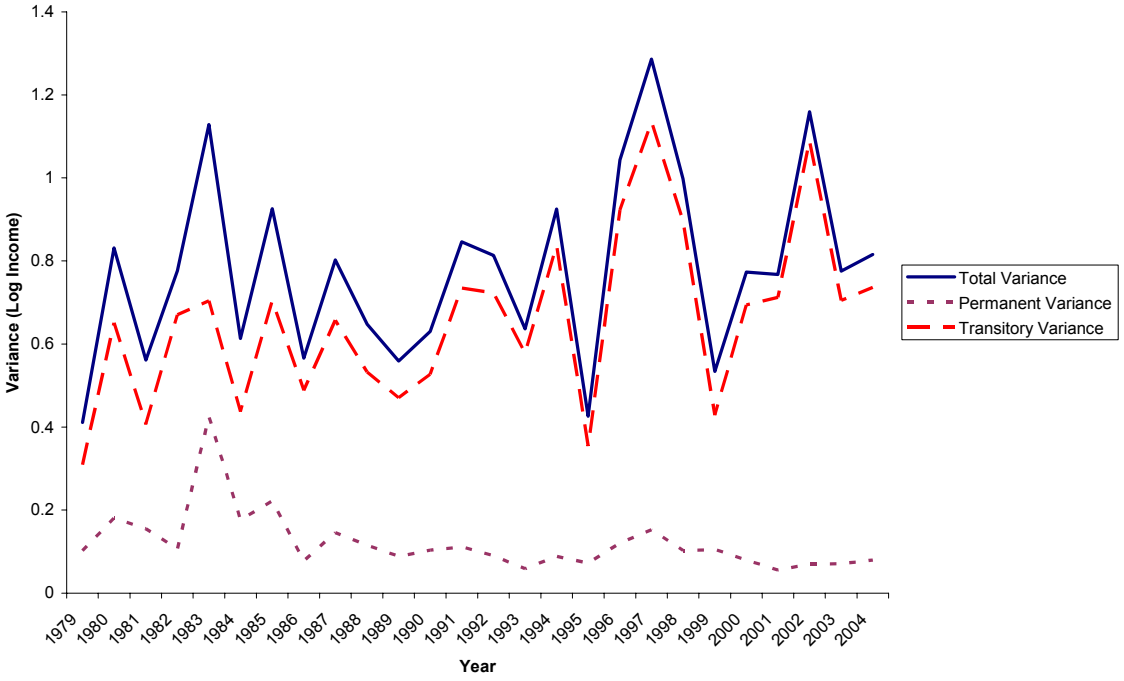

Figure 31: Social Security and Disability Income Volatility of Single Mothers, Participants and Nonparticipants

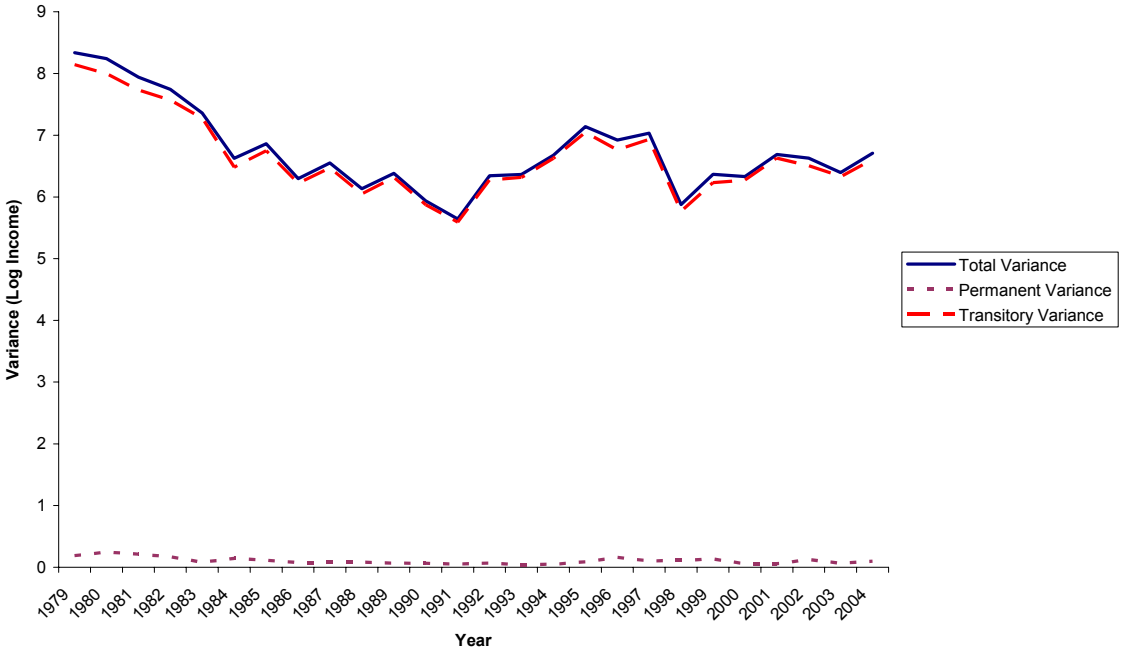

Figure 32: Social Security and Disability Income Volatility of Single Mothers, Participants Only

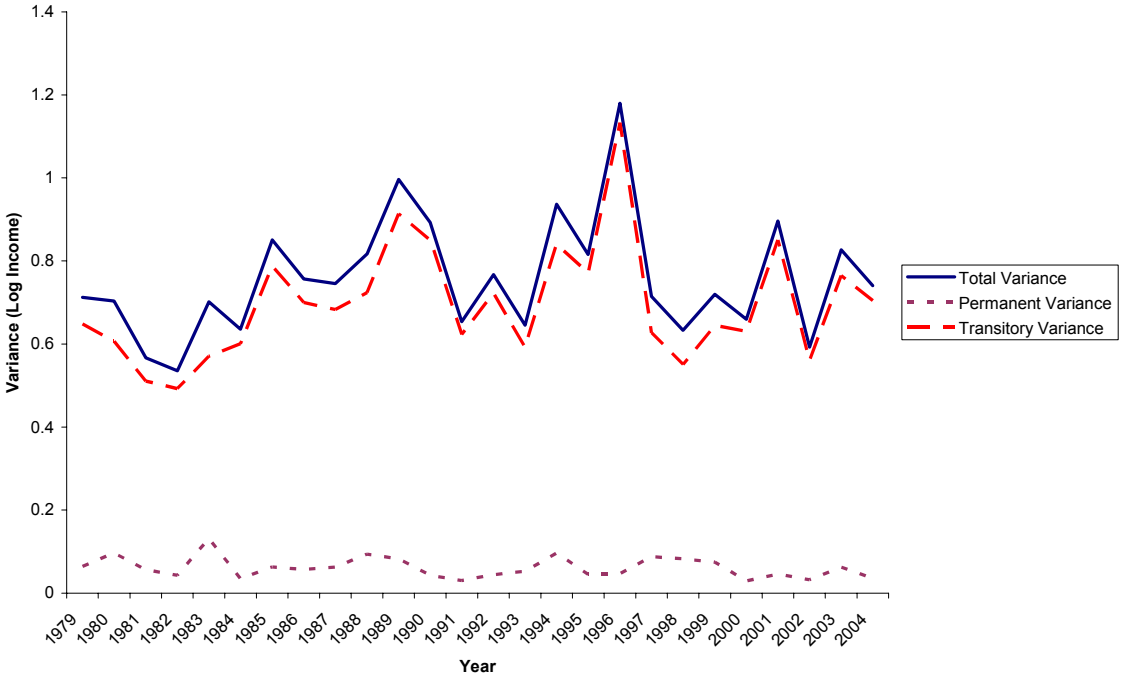

Figure 33: Food Stamp Income Volatility of Single Mothers, Participants and Nonparticipants

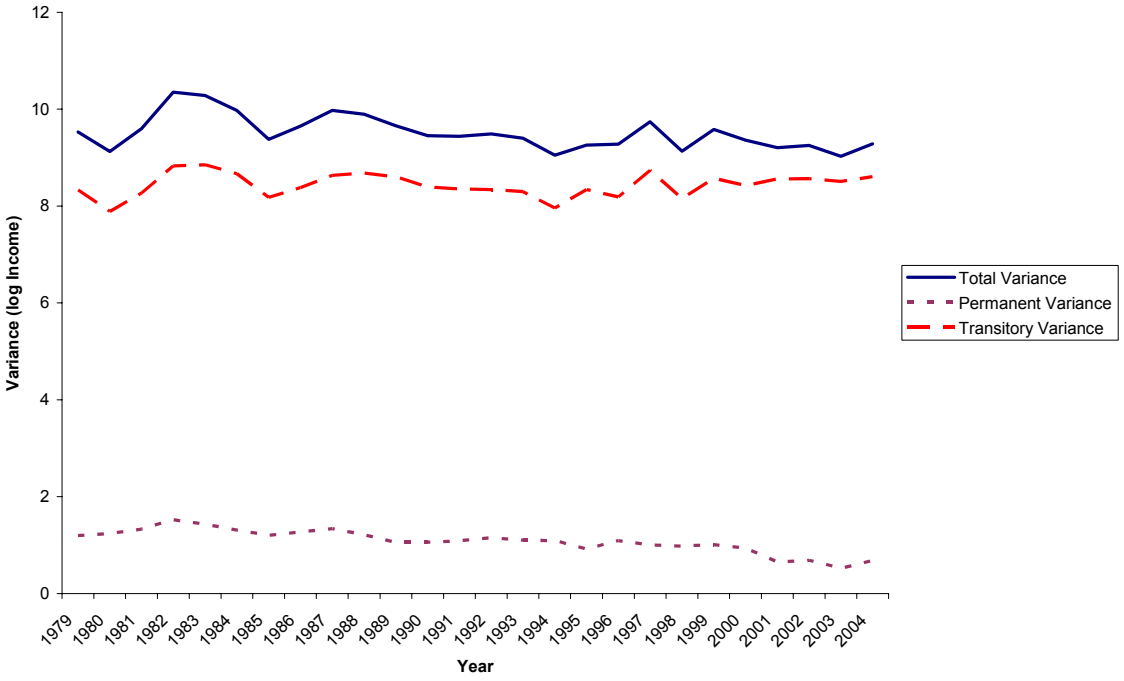

Figure 34: Food Stamp Income Volatility of Single Mothers, Particpants Only

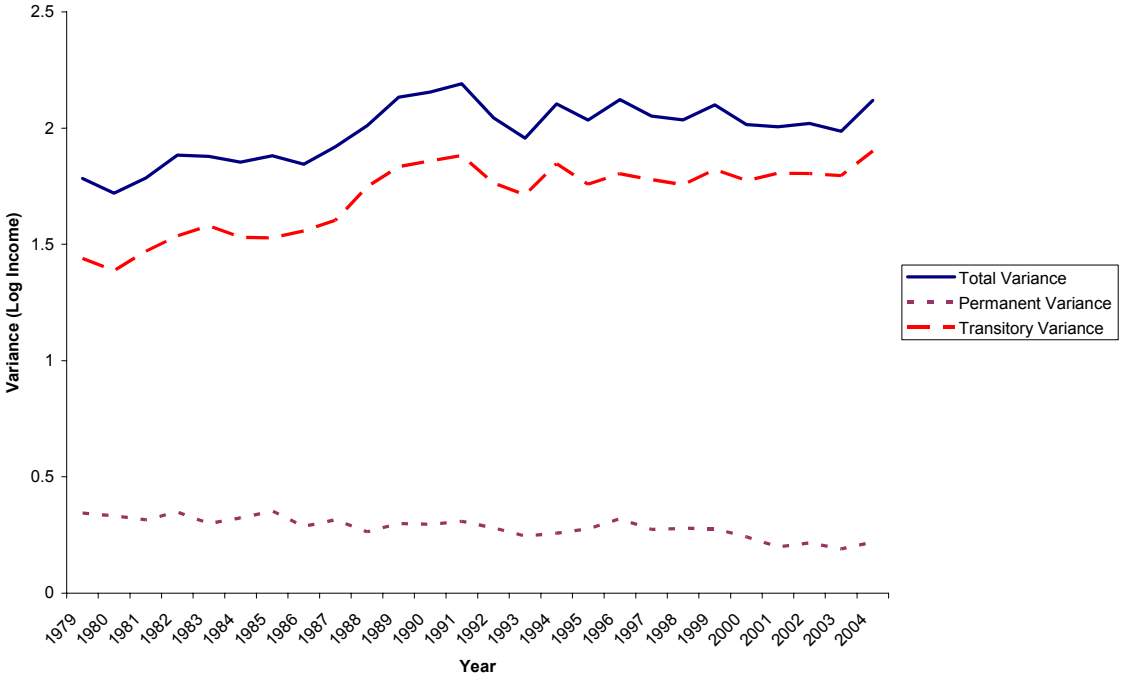

Figure 35: Other Income Volatility of Single Mothers, Participants and Nonparticpants

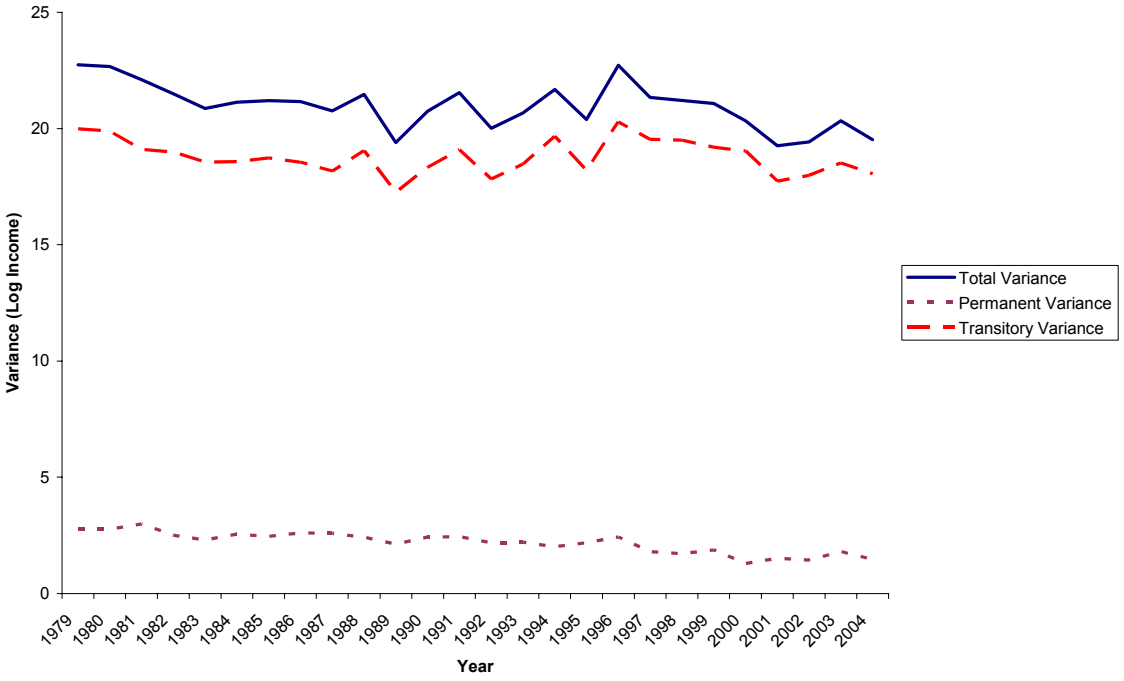

Figure 36: Other Income Volatility of Single Mothers, Participants Only

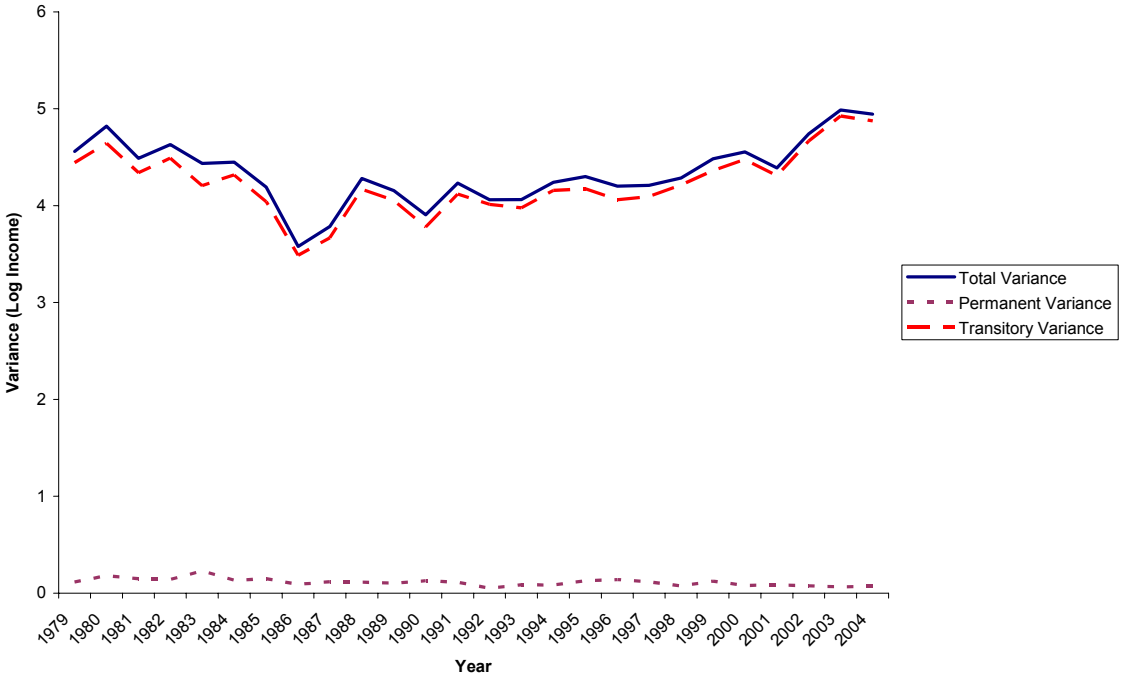

Table 1: Quandt-Andrews Tests of Structural Change of Unknown Breakdate in Income Volatility of Single Mothers, 1979–2004

Series	supWald Statistic	Break Year	CI-width (Years)	Sign of Trend Coefficient (pre-break, post-break)
Total Income	80.94***	1996	+/- 2	+,+
Earnings (all)	27.64***	1997	+/-1	–,0
Earnings (participants)	14.79**	1991	+/- 4	0,–
EITC (all)	53.56***	1986	+/-1	–,+
EITC (participants)	9.35	1990	+/-6	+,+
TANF (all)	143.26***	1996	+/-1	–,–
TANF (participants)	12.95**	1997	+/-2	+,0
SSI (all)	86.51***	1992	+/-1	+,–
SSI (participants)	4.47	1983	+/- 4	+,0
SSDI (all)	67.33***	1991	+/-1	–,0
SSDI (participants)	10.02	1996	+/-3	+,0
Food Stamps (all)	16.04***	1983	+/- 3	+,–
Food Stamps (participants)	44.90***	1988	+/-2	+,–
Other Income (all)	13.00**	1995	+/-2	–,–
Other Income (participants)	71.59***	1985	+/-1	–,+

Note: The 0 in the last column indicates that the coefficient on the trend variable is statistically zero at least at the 10% level. For the Wald statistics *** = reject at the 1% level; ** = reject at the 5% level; *=reject at the 10% level.