Economic Policy Institute

1660 L STREET, NW • SUITE 1200 • WASHINGTON, DC 20036 • 202/775-8810 • FAX 202/775-0819 • www.epinet.org

Wage Inequality in the 1990s: Measurement and Trends

By Lawrence Mishel, Jared Bernstein, and John Schmitt

Economic Policy Institute
Washington, DC
December 28, 1998

Prepared for the 1999 ASSA Meetings, New York City
Session: "Wage Trends in the 1990s"

Much of the recent discussion of wage inequality draws exclusively on research and trends related to the 1980s. This is partly due to the lag between conducting research and actual publication. However, the continued focus on the 1980s is frequently based on the unexamined assumption that the pattern of wage growth prevailing in the 1980s continues in the 1990s. Another possible reason for relying on 1980's data is that various changes in the Current Population Survey, according to some analysts, have made it difficult to identify wage trends in the 1990s (Lerman 1997; Katz and Autor 1998).

This paper examines wage inequality trends in the 1990s using the CPS and assesses the sensitivity of measured trends to the choice of methods to deal with changes in the CPS, particularly changes in education coding and top-coding. We then contrast the trends of the 1990s and the 1980s and draw some implications for our understanding of wage inequality.

We find that wage trends in the 1990s differ from those of the 1980s in several key features and that measurement problems due to changes in the CPS are quite manageable (especially relative to problems with data for the 1970s or the 1950s and 1960s). In particular:

- Using various top-code adjustments, both the March and ORG data reveal that hourly wage inequality, as measured by the Gini coefficient, continued to grow in the 1990s, though at a slower rate than in the 1980s.
- Measured trends in the "college-high school" wage differential are more sensitive to endpoints and the inclusion of the "more than college" group than to changes in education coding.
- Trends in absolute and relative employment by education level, however, do show a marked discontinuity because of the 1992 coding change. Ignoring the discontinuities can generate misleading results.
- Changes in relative wages and employment by education suggest a significant slow-down in the growth of the relative demand for education among men in the 1990s (relative to the 1980s), and the slowest rate since the 1940s. This pattern generally does not accord with the story of an information-technology-driven new economy emerging in the 1990s, especially in the recovery.

- Wage inequality continued to widen in the top half of the distribution in the 1990s much as it did in the 1980s. However, the 50/10 differential has been flat or falling for at least ten years.
- Widening education differentials have played a smaller role in generating wage inequality in the 1990s than in the 1980s, while wage dispersion within education (and experience) groups has become more important.

The first part of the paper examines the 1990s trends in overall wage inequality, education differentials, education quantities, and associated measurement issues. The second part examines shifts in the character of wage inequality in the 1990s and examines the role of within group, or "residual" wage inequality. We summarize the implications of the findings in a conclusion.

I. Measurement Issues and Trends in Overall Wage Inequality and Education Differentials

A. Changes in the CPS

This paper employs the two CPS wages series-the Annual Demographic files (the March data) and the Outgoing Rotation Group (ORG) files-to analyze wage trends in the 1990s. We, therefore, have to adopt strategies to deal with two discontinuities in the data created by changes to the CPS. The first was a 1992 change in the education question and coding from "years of schooling" to "highest degree attained." The second was a difference in the way annual earnings from longest job (the key earnings variable from the March CPS) was recorded, beginning in the 1996 survey. Starting in that year, the top-code was lifted from $99,999 to $150,000, and the value in the field for persons with reported earnings above this level was a cell-based average of the variable, taken over individuals with earnings above the top-code (with cells partitioned on the basis of sex, race, and part-time/full-time status).

The timing of these changes in each CPS series is presented in Table 1 by "data year" (the year described by the data). Note that changes in the CPS survey in any year affect the March data describing the prior year and the ORG data describing the same year, i.e., a change in 1992 affects the March 1992 files which have retrospective earnings data from 1991 and affects the ORG data in 1992. We exploit this difference in timing to bridge trends over the discontinuity in education coding. A third measurement question relates to any possible bias introduced by the CPS redesign in 1994, including the introduction of computer-assisted interviewing.'

Top-Coding: We examine the effect of the top-coding change by comparing trends using four different approaches to dealing with the top-coding problem. Note that top-coding remains a challenge whether or not the actual top-code is changed. If the top-code remains constant and no adjustment (or a constant adjustment, i.e., each top-coded earner is assigned a multiple of the topcode) is made then the assumption is that the highest wage earners enjoy no nominal wage growth (and have real wage reductions). Also, as the wage distribution shifts rightward over time, an increasing share of earners become "top-coded." A change in the top-code corrects this problem but because the changes are infrequent they potentially introduce a spike in the wage data for top earners.

Our first approach is to compute the data trends with top-coding as presented in the public use files not only because some analysts have ignored top-coding (see Lerman 1997) but to show the affect of changes in top-codes.* A second strategy is "trimming,". which means identifying the share of earners that are unaffected by top-coding over the complete time period of interest, in our case 98% (97% in the ORG data), and limiting the sample to this group, i.e., throw out the upper 2% of the sample. This is the strategy used, for instance, by Karoly and Burtless (1992).

The last two strategies, ones which we prefer to the above, involve assigning new values to observations that are top-coded. In each case, the data are made consistent over time by suppressing the change in the top-code on earnings from longest job starting in the March files in data year 1995. That is, the new information is ignored and the old top-code of $99,999 is imposed on the data. The first type of adjustment is to estimate the degree to which the average top-coded value understates actual earnings in a particular year (e.g., by one-third) and adjust each year's data accordingly (e.g., multiply the top-code by 1.5). We refer to this as a "constant adjustment" method (see Katz and Murphy 1992, and Katz and Autor 1998). The drawback of a constant adjustment is that the top-code remains constant (no nominal change) over time and thereby understates

[1] Ryscavage (1995) examines the impact of the redesign on 1992-93 trends in the March data. He judges the impact to be "inconclusive," be he notes that the "forces for greater income inequality may have been particularly strong between 1992 and 1993."

[2] Census Bureau computations, such as those published in the P-60 Series, ***use*** higher "internal" top-codes.

growing inequality. The last strategy is a "Pareto adjustment" which assumes that the portion of the wage distribution above the top-code follows a Pareto distribution, and uses the shape of the distribution below the cutoff to estimate the average value of the truncated tail. This average, which we compute separately for males and females, changes each year, thus allowing nominal and real earnings to grow at the top. This is the approach we have used in our earlier work (Mishel et al, 1998).[3] The drawback is that when the official top-code remains fixed for a number of years, the adjustment becomes less accurate as the share of cases top-coded grows.

We employ each of these methods as we analyze summary measures of wage inequality and estimated education differentials using the March CPS data.

Education Coding: The second discontinuity, involving education coding, affects both the March and the ORG CPS data. We estimate regression-adjusted education wage differentials in two ways: First, for the years where the new "highest degree attained" coding is employed we convert the data to a "years of schooling" basis using imputations derived from a February 1990 CPS supplement, which has both types of questions. This way we can present a time series based on completed years of schooling, including, after the coding change, imputed years of schooling. Our second approach for dealing with the coding change is to use the new coding in the March data from 1991 forward, and from the 1992 on using the ORG data. We compare the trends in the actual/imputed years of schooling series with some "spliced" series to provide a range of estimates for the changes over the 1989-97 period.

We also present the employment shares by education level using the CPS data and show the importance of the discontinuity. We lack faith in the capacity of our imputation procedures to develop a reasonably consistent series on education quantities. We draw on independently collected data from higher education institutions to help discern trends in the supply of college graduates.

Last, we examine the information available on the possible effect of the CPS redesign on measured wage trends. Most importantly, we compare newly available trends in Social Security annual wage data (unaffected by any redesign) to those in the March

[3] Polivka (undated), using an uncensored data set, shows that this approach is quite effective, i.e., generates estimates close to the actual mean of the tail, for weekly earnings from the CPS ORG.

annual data. We find that wage inequality in the Social Security data actually rose faster than in the March data.

The sample for our analysis consists of wage and salary workers, age 18-64.[4] We excluded cases whose hourly wages were less than $0.50 or greater than $100.00 in 1989 dollars, considering them to be outliers. Most of our analysis focuses on hourly wages. For the March data, hourly wages are computed by dividing annual earnings by the product of weeks worked and usual hours worked per week. For the ORG data, we use the reported hourly wage from those who identify themselves as hourly workers; for others, we divide usual weekly earnings by usual weekly hours. See Webster (1998) for a detailed description of our construction of hourly wages from the ORG.

Table 1
Changes in the CPS in the 1990s

Data Year	**March CPS**	**ORG CPS**
1989		
1990		
1991	New Education Coding	
1992		New Education Coding
1993	New ***Survey*** Design	
1994		New Survey Design
1995	Raised Top-Code	
1996		
1997		

B. Overall Wage Inequality

This section reviews CPS trends in overall wage inequality by gender over the 1989-97 period. The measurement issues involve the treatment of top-coding, particularly the 1995 change in the March data, and any bias due to the CPS redesign.

Table 2 and **Figure 1** present Gini coefficients computed with the March CPS employing each of the four previously described strategies for dealing with top-codes. The one method we know is inappropriate is to make no top-code adjustments, a method which locates nearly the entire increase in male wage inequality in the 1994-95 period as the new, higher top-code was introduced. Other methods show a lesser increase in male

[4] The March data include the incorporated self-employed; they account for 3% (weighted) of the 1997 sample.

wage inequality, with both the Pareto and constant adjustment procedures yielding similar results. The “trimmed” series shows the least growth in wage inequality, establishing the importance of the trends among the upper two percent in generating the growth in wage inequality.

Trends in women’s wage inequality are less sensitive to the treatment of top-codes, for the simple reason that fewer women than men are very high earners. For instance, the series which makes no top-code adjustment shows a much smaller jump in 1995 for women than for men. Again, the two adjusted series show similar trends, with “trimmed” data showing the least growth.

The March data show a growth of wage inequality in the early portion of the recovery, from 1990 to 1994 among men and from 1992-94 among women. However, wage inequality, at least in a summary measure such as the Gini, was stable in the downturn and in the latter part of the recovery.

Similar trends are present in the CPS ORG series presented in **Table 3** and **Figure** 2. Because no top-code change was introduced in the ORG data in this time period (though a change was made in 1998) we omit the “no adjustment” strategy. We also rely on the Pareto adjustment and not the constant adjustment procedure because they show similar trends, at least over a period of a few years. Focusing on the Pareto-adjusted series, one again sees stable wage inequality in the downturn, growing wage inequality in the 1991-94 period and stability thereafter. There is a notable jump in the series, 1993-94, coinciding with the introduction of the CPS redesign, which we discuss below. The “trimmed” series reveals a similar trend but at lower levels.

A summary of these trends along with a comparison to the trends of the 1980s is presented in **Table 4,** where changes in Gini coefficients are expressed in percentage terms to put each series on a common scale. All of the series employ the Pareto adjustment of top-codes. Although it is clear that wage inequality has grown in the 1990s, it is equally clear that the growth has been less than that of the 1980s. If one adjusts the 1989-97 changes to ten-year rates, it appears that wage inequality has grown about half as fast in the 1990s as in the 1980s. In a later section, we show that this results from some equalizing trends in the bottom half of the wage structure accompanying widening wage dispersion in the upper half.

One other challenge is to assess whether, or by how much, these measured trends in wage inequality are an artifact of the CPS redesign in 1994, affecting the March data for 1993 and the ORG data for 1994. Competing views on this question were recently presented in the *Monthly Labor Review* (Bernstein and Mishel 1997; Lerman 1997). Lerman raises a number of issues with the March CPS, including the rising proportion of observations with imputed earnings and some slight revisions to the earnings questions. Lerman also contends that the **new** computer-assisted survey interviewing system, introduced with the 1994 redesign, may have stimulated more responses from high earners and biased wage inequality upwards. These factors lead Lerman to characterize March CPS trends as a "source and method" that is "subject to bias and error"(p. 22).

Bernstein and Mishel argue that Lerman has not offered evidence of any change in the CPS that leads to a particular bias and point out that the change in demographic weights in 1994 is a real phenomena (though not actually occurring in one year). Moreover, "trimmed" data, as shown above, still show a sizeable increase in inequality form 1992 to 1993, suggesting that an entry of very high earners into the CPS is not responsible for the growth of inequality in this period. Plus, the ORG data, unaffected by redesign until 1994, show a growth of wage inequality (though smaller than that of the March) over 1992-93.

A newly available wage series based on Social Security annual earnings data provides a reliable yardstick against which to measure March CPS trends (Utendorf 1998).[5] The Social Security data underwent no redesign over the analysis period, have no top-codes, have a much larger sample, and use no imputations. **Table 5** shows the Gini for annual earnings inequality based on Social Security data for the 1989-93 period (later years are not yet available) and the ratio of Ginis measured with the March CPS data to those measured with Social Security data.

Two results stand out: One is that wage inequality as measured by Social Security earnings data is 25% greater than March CPS measures. Second, the Social Security data show a marked jump in wage inequality in 1993, following several years of stable inequality. In fact, the Social Security data show a much larger jump in inequality in 1993 than the March CPS data, as seen by the decline in the ratios from 1992 to 1993 in

[5] These data include total annual earnings from a 1% sample of Social Security numbers for which wage and salary data were reported; they include total earnings, not simply Social Security taxable wages.

every series. This is very convincing evidence from an independent source that a real, not just a measured, increase in inequality occurred from 1992 to 1993.

C. Education Wage Differentials

An important dimension of the labor market, closely related to the growth in wage inequality, is the wage differentials between workers with varying degrees of education, usually illustrated by the level and change in the college-high school wage differential. The ability to track trends in education differentials in the 1990s is made more difficult by a change in the education questions asked in the CPS in 1992, from identifying "years of schooling completed" to "highest degree attained." The treatment of top-codes matters for measuring education differentials because those with the most education are more likely to have top-coded earnings. As we show below, however, measured trends in education differentials are more sensitive to the endpoints chosen and to whether one combines four-year college graduates with those with graduate or professional degrees (referred to as "more than college" below).

Some analysts have simply ignored the changes in the education coding, but this has potential pitfalls, particularly when measuring employment by education level. The 1990 February CPS, which has both coding formats, reveals that 15% of those who report having attained a bachelors degree under the new coding (highest degree attained) report completing year 17 or 18 under the old coding (years of school completed). While 18% of the population (15 years and above) report completed some college under the old coding, the analogous group under the new codes comprises 22% of the population. As we discuss below, ignoring the impact of these coding changes can lead to mistaken inferences about the demand for workers by education level.

The education differentials are derived from log wage regressions based on the same sample of workers as the previous section. Besides education dummies, the regressions include controls for marital status, region, race/ethnicity and experience or age, entered as a quartic. The coding change also presents a challenge to the conventional approach to measuring potential experience. Since the "Mincer" experience measure (age-education-6), cannot be reliably derived from the attainment variable, we use imputed years of education to calculate experience (meaning experience is measured as age-imputed education-6) for relevant years. When we shift to the new coding, we replace this variable with age. We discuss the impact of these choices below.

Table 6 presents estimated education differentials relative to a high school graduate using imputed education levels by top-code strategy, distinguishing (four-year) "college only" from "more than college" and the combination of the two categories. **Table 7** presents comparable estimates for 1991 forward, using the new coding of "highest degree attained."

One important result from these tables is that, particularly among men, the growth in education differentials (relative to high school graduates) has been greater among those with degrees beyond college than among those with a four-year degree. This is because, as we pointed out elsewhere (Mishel, et al 1998), the real wages of male college graduates has been flat or falling since 1986 or 1987 (with the exception of the last year or two), while real wages fell more slowly in the 1990s (relative to the 1980s) among male high-school graduates. In contrast, the real wages of those with advanced degrees continued to grow. Estimates that combine these two groups bury these differences and have incorrectly been used to imply higher returns to college graduates when, in fact, the estimates reflect higher returns to graduate and professional training.

Comparing Tables 6 and 7 leads to the impression that our imputation procedure (imputing years completed from information on highest degree attained) leads to less growth in the education differentials than when the new codes are used. For example, Table 6 shows that the male, Pareto-adjusted differential grew 0.4 percentage points, 1992-97; the analogous value from Table 7 is 1.9 points. However, this difference is actually due to our use of potential experience in the regression in Table 6 and our use of age in those in Table 7. When we replace experience with age in the Table 6 regressions (still using imputed years of education), the increase from 1992-97 is almost identical to that from Table 7.[6] Nevertheless, as we stress below, both formats lead to the conclusion that the growth rate of education differentials decelerated considerably in the 1990s.

The influence of combining "more than college" with "college only" into a "college or more" category is illustrated in the data presented in **Table 8.** For the period from 1989 to 1997 (or 1990-97), the change in the estimated male high school wage

[6] This result stems from the fact that our education imputations are mostly assigning cases to their attainment categories. There are, however, some differences, particularly among those with some college.

differentials are from two to four percentage points more in the "college or more" differential. This is a large difference relative to the size of the actual change in the "college only" differential which ranges from a small drop to a 2.9 percentage point increase, depending on which top-code method is applied. These distinctions are less important among women.

The specific endpoint selected also matters a great deal, as male education differentials rose significantly over the 1989-90 period but grew more slowly thereafter among "college only" men. Estimates using the new coding, with data from 1991-1997, allow us to examine the growth of differentials with consistent coding over the recovery. As shown in the bottom line of each panel in Table 8, the Pareto-adjusted, college-only premium grew by 1.9 percentage points, 1992-97, for males, and slightly slower for females.

In contrast to the selection of time period or the inclusiveness of the college group, estimated trends are not as sensitive to the particular top-code strategy chosen (although estimates can vary up to one percentage point or so). In the March data, education differentials among women seem to grow only slightly, if at all, regardless of particular measurement choices.

Table 9 presents education differential trends based on an analysis of the CPS ORG data, using both an imputed years of schooling series and a series from 1992-on for the new coding. Similar to the March CPS data, the male "college only" differential grows modestly in the 1990s and especially in the recovery. The ORG data, however, show a more substantial increase in women's "college only" differentials than does the March series (Tables 6 and 7).

Last, **Table 10** compiles several ways of computing the change in the "college only"-high school wage differential in the 1989-97 period and compares them to the 1979-89 trend. The first method is to "skip" the years where a discontinuity exists and assume it has the same trends as the average of the other years in the period. For instance, the change over the 1989-97 period can be obtained from the ORG by computing the change from 1989 to 1991 using the old coding and the 1992-97 change using the new coding, "skipping" 1991-92. These seven years of change are converted to a ten-year rate of change, to be compared to 1979-89, by multiplying by 10/7.

The “splice” method relies on the fact that the education coding change occurs in different “data years” in the ORG (1992) and the March (199 1) data. Thus, when discontinuities occur in one series there is a continuous trend in the other series. For instance, rather than skip 1991-92 for an ORG series, one can substitute the 1991-92 change in the differentials estimated with the March data (which had new coding in each year). Our third method is the previously discussed imputed years of schooling series.

These series presented in Table 10 provide a fairly wide range of estimates of the change in the college-high school wage differential over the 1989-97 period. Though the different methods of handling the coding change, as well as the two different data sets, yield different results, they all show a significant slowing in the growth of education differentials in the 1990s relative to the 1980s. Among men, the range extends from a “ten year” change of from 2.3 to 8.9 percentage points. These estimates, however, show a growth roughly one-third to one-half that of the 1979-89 period. We also know that the slowdown is even greater in the recovery, 1992-97, or if we start in 1990. Among women, there has been little growth in education wage differentials if one relies on March CPS trends; however, the ORG data yield growth rates that are 30 to 40 percent the rate of the previous decade.

There is, of course, the possibility that this deceleration in the college premium in the 1990s is due to accelerated growth in the relative supply of college-educated workers. As the next section shows, the coding change makes a notable difference here; nevertheless, our examination of various sources and methods fails to reveal an acceleration of relative supply.

D. Employment by Education Level

In our view, the most difficult time series to make consistent in the 1990s is the share of employment at differing education levels. This section reviews the discontinuities in the CPS series and draws on independent data from the Department of Education on degrees awarded in order to discern trends.

The problem is illustrated in **Table 11,** which shows the employment distribution by education level in 1991 and 1992 from the ORG CPS, the years where the switch in coding takes place. If the data are not adjusted, there appears among men to be a roughly *4.5* percentage-point shift from the share of workers with a high school degree or less to those with “some college.” The shift among women is 5.6 percentage points. Such a large one-year change is obviously an artifact of the coding change.

There is also a distortion in the trends for the share of workers with a (four-year) college degree. If one assumes those with 17 completed years of school in 1991 had a bachelors’ degree (rows 1991a) then the college share of employment fell from 199 1 to 1992 (0.7% for men, 0.9% for women). [7] If the 17 years of schooling group are considered to have advanced degrees (rows 1991b) then the share of workers with advanced degrees falls from 1991 to 1992. Neither trend is quite believable. The coding change may be affecting mostly the composition of the group of workers with a “college degree or more,” since the total share (college plus advanced) remained relatively constant from 1991 to 1992.

Ignoring these changes in coding can lead to large errors in trends. For instance, Johnson (1997) computes changes in the relative demand for education in the 1940s, 1950s, 1960s, 1970s, 1980s, and 1990s. One of his conclusions was that the growth of relative demand for college graduates accelerated in the 1990s over the 1980s and was the fastest of any decade in the post-war period. This finding of a 1990s acceleration, however, was driven in large part by the coding change because Johnson’s measure of college graduates includes half the workforce with “some college.“* As shown above, the share of employment with “some college” jumps five percentage points because of the coding change.[9] Thus, when Johnson’s data are recomputed with a measure of relative demand for college versus non-college-educated workers (placing all of the “some college” workers in the less educated category) the data show a marked deceleration of relative demand for education-a near halving of the 1980s rate and significantly slower than that of the 1950s, 1960s, and 1970s (Mishel et al 1997).

Another problem area is in the use of the data that Census reports on the percent of high school graduates (aged 14-24) who have enrolled in or completed some college. The measured trend shows a jump from 60.7% in 1991 to 65.6% in 1992, a 4.9 percentage point change almost as large as the entire growth over the 1979-89 period (of

[7] The 1990 February CPS, which has both education codes, shows that 9% of those who report a bachelors as their highest degree attained, report completing 17 years of education under the old coding.
[8] Part of this bias is reduced by the fact that Johnson lowers the “some college”-equivalent share by the wage differential between those with some college and those with bachelors degrees.

6.3 percentage points). The Census Bureau warns in a footnote that “the change in the educational attainment question and the college completion categories” caused this trend to jump approximately 5 percentage points in 1992 (the entire increase).

We assess the 1989-97 growth in the college-educated share of employment (limited to the “college-only” group) in **Table 12,** using the “skip” and “splice” methods described above in our analysis of education wage differentials in the 1990s. We omit any estimation of changes in education employment shares using an imputation procedure because we judge our imputations to be unreliable for estimating education quantity trends.”

As Table 12 shows, the skip and spice methods yield similar estimates, perhaps not surprisingly since there is only a difference in one year of ten (as the changes are converted to ten-year rates of change for comparison to the 1980s). The relative employment of college-educated males in the 1990s has been either somewhat slower (the ORG trends) or comparable (the March trends) than relative employment growth in the 1979-89 period. In contrast, the relative employment of college-educated women has grown in the 1990s at nearly double the rate among men in the 1990s and significantly faster than among women in the 1980s. This suggests a stable or slowing supply of college-educated men but an accelerating supply of college-educated women.

The notion of a steady or decelerating relative supply of college-educated men might seem to contradict the well-established fact that college enrollment rates have been rising rapidly since the late 1970s. In fact, the enrollment rate has grown faster in the 1990s than the 1980s, but this fact by itself does not generate an across-the-board increase in relative supply of college graduates, for several reasons. First, the trends have differed by gender, with a much faster growth in enrollment rates among women than men.” Second, the relatively small and shrinking size of the college age population (it

[9] On a more substantive note, under the new coding structure the majority (87%) of workers with “some college” have no degree past high school; other research has shown the wage trends of this group look more like those of high-school workers than college-educated workers (Katz and Murphy, 1992).

[10] Our imputations produce a similar jump in the share of employment with “some college.” This raises the reasonable question as to why we trust our education imputations on the wage side but not on the quantity side. Unlike the employment quantities, our examination of average wage trends by imputed education level show smooth, believable changes over the coding change.

[11] Using the same enrollment data shown below, the enrollment rate (the share of the high-school educated population, age 14-24, to enroll in college) for males grew 4.3 percentage points, 1989-96, while that of females grew 6.4 points. Note that these enrollment rates are not affected by the problematic “some college” category.

was 16% smaller in 1995 than 1979, Mishel et al, 1998, Table 3.54) lessens the impact. Third, recent changes in enrollment rates have not yet affected supply. Fourth, relative supply shifts also depend on the education levels of those exiting the labor force. If, as we suspect, the rate of labor force exits by more highly educated workers accelerated in the 1990s over the 1980s, more so than the inflow of college-educated workers, this would lead to a lesser increase in the relative supply of skill in the 1990s relative to the 1980s.

Figure 3, plots the enrollment rate by gender (the share of high-school graduates, 14-24, enrolled in college) lagged four periods against the share of bachelors' degrees awarded (from Department of Education data compiled from colleges), also as a share of high school graduates, age 14-24. While the lagged enrollment rate has climbed fairly steadily, the college completion rate begins to flatten in the early 1990s for both genders. (see Gladieux and Swail 1998 on these points). This divergence implies a slowing of the relative supply of college graduates in the 1990s relative to the 1980s.

E. The Relative Demand for Education

Having identified the changes in the increase in the relative price and relative supply of college-graduates in the 1990s, it is possible to draw some inferences about how relative demand for education has grown in the 1990s relative to the 1980s (or an earlier period).

Johnson (1997), Autor, Katz, and Krueger (1997) and others have used changes in relative wages and quantities to compute indices of relative demand for education. All use a reduced form of a CES production function with two factors, e.g., college and high-school educated workers, wherein relative demand is equal to the product of an elasticity of substitution between the two types of labor and their relative wages, plus relative supplies (see Katz and Murphy, 1992). Our analysis of the 1990s above has indicated that, among men, there was a substantial slowdown in the growth of the college-high school wage premium and steady or decelerating growth in relative supply in the 1990s relative to the 1980s. It is an inescapable conclusion, therefore, that the growth in the relative demand for college-educated men has decelerated in the 1990s relative to the 1980s, especially in the recovery years where education wage differentials flattened. Among women, however, there was a substantial slowdown in the growth of the college wage premium but an accelerated growth in relative supply. Thus, one can not rule out that accelerated supply led to a slower relative wage growth with an uncertain implication for relative demand growth.

Using our estimates of relative supply and relative wage growth in the 1980s and 1990s, and applying the substitution elasticity of 1.4 used in the studies noted above, we find that demand for college-educated workers decelerated for both males and females. For males, relative demand decelerated quite sharply from the 1980s to the 1990s, by 38%; for women, the decline was smaller, 15%.[12] As shown in Autor and Katz (Table 14, 1998), the 1990s register the smallest increase in relative demand of any decade since the 1940s.

Using the same reduced form index described above, various analysts (e.g., Autor, Katz, and Krueger, 1997) have argued that the primary cause of the acceleration of this index in the 1980s over the 1970s is evidence of skill-biased technological change (SBTC). Does the above evidence, then, imply a deceleration in SBTC in the 1990s? If so, this would certainly challenge a notion of the 1990s as a decade where the ever-increasing use of micro-processor technology has led to acceleration of skill demands, inevitably, but unavoidably, accompanied by increasing inequality.

We have argued elsewhere that this reduced form demand index reflects more than technology-related shifts in the demand for one type of labor relative to another (see Mishel et al 1997, p. 13-16). For instance, besides technology, the index also reflects the impact of institutional shifts, such as changes in the minimum wage and union density, as well the impact of international trade on relative wages and employment. As a simple example of the impact of a labor market institution, note that a decline in the value of the real minimum wage, such as occurred over the 1980s, will typically have the effect of raising the relative college wage, which will show up as an increase in relative demand. As regards trade, it is generally agreed that at least some part-at least 10% to 20%—of the growth in the skill premium in the 1980s over the 1970s is attributable to the increase of trade imbalances (including immigration). In this sense, the index really measures the combined impact of "non-supply" factors, i.e., demand (trade and technology) and institutions.

[12] Despite differences in endpoints, sample, and data source, our results are similar to those of Katz and Autor (1998) using the March data. They report a 36% deceleration in relative demand for "college equivalents" in the 1990s over the 1980s.

It is beyond the scope of this paper to engage in a full accounting of the roles of these factors in the 1990s. We note, however, that trade imbalances grew in the 1990s relative to the 1980s, which should have increased the rate of growth of skill premiums. The minimum wage, on the other hand, was increased twice over the 1989-97 period, which, as shown in the next section, compressed the wage gap between middle and low-wage workers, and perhaps played some role in dampening the increase in the college premium. The rate of unionization continued its long-term decline in the 1990s, though somewhat slower than in the previous decade. This slowing may also have played a role in the deceleration of the college premium.

Nevertheless, we do not see a clear institutional explanation for the deceleration in the relative demand index. Since there is no evidence that the rate of diffusion of micro-processor technology declined in the 1990s (to the contrary, Mishel et al 1997, show an acceleration in the within-industry investment in computerization per worker in the 1990s over the 1980s), we are left with the suspicion that either technological change was slower in this period, or technology itself had a lesser skill bias.

II. Character of Wage Inequality in the 1990s

In this section, we turn to an examination of changes in relative wage deciles, focusing primarily on the ratios of the 90th percentile to the 50th, and the 50th to the 10" (this section uses ORG data exclusively; for a similar analysis using the March data, see Bernstein and Mishel (1997); the March data yield similar results to the ORG). This approach has two advantages over the Gini analysis in the previous section. First, it allows us to look a changes in the patterns of inequality growth at different points in the distribution, and second, since the 90" percentile wage is always below the top-code, it is insensitive to that measurement issue. Finally, like the Gini ratios, but unlike the education differentials, these data represent changes in ***overall*** wage inequality, combining the effects of growing between- and within-group inequalities.

The 90/50, 50/10, and 90/10 log hourly wage ratios by gender, 1979-97, are shown in **Figure 4**.[13] For males, the figure shows the 90/50 trending upward through most of the period, while the 50/10 begins to decline in 1993. For females, the 50/10 falls 1989-92 and flattens thereafter, while the 90/50 trends upward. The log annual change

[13] The sample for these data is the same as that in Section I. The decile cutoffs are smoothed, as described in Webster (1998).

(see Table 14, second row) in the 90/10, which is simply the log difference in the ratios divided by 8 years (times 100), was 0.7 points for males and 0.5 for females. Note, however, that this change was driven fully by the growth of the 90/50 differential in both cases; the 50/10 differential actually compressed slightly over the period. Thus, the character of wage inequality for both genders in the 1990s was that of the top pulling away from the middle and bottom.

Figure 5, which extends these series back to 1979, reveals that this is a shift from the earlier pattern of wage inequality in the 1980s. The top panel, for males, shows that both the 90/50 and the 50/10 increased at similar rates through most of the 1980s. Towards the end of the decade, the male 50/10 began to decline, driven by real declines in the male median (the real male 10" percentile wage grew slightly in the late 1980s), but the 90/10 continued to climb. A similar pattern for women is evident in the bottom panel of Figure 5, though the reversal of the 50/10 started later and was less steep than the male case. At any rate, the figures demonstrate that in the 1980s wage dispersion occurred between the top, middle, and bottom among both males and females. In the 1990s, the top continued to pull away from the middle, while the 50/10 either flattened (females) or fell (males). As Table 13 (first row) reveals, the annualized growth rate of the 90/50 was about the same in both periods for males, and only slightly slower for females.

Figure 6 presents a more comprehensive perspective on relative wage changes by showing the annualized percent changes in real wages (deflated by the CPI-U-Xl) at each decile, by gender, 1979-89 and 1989-97. These figures show that the deceleration in the growth of overall male wage inequality in the 1990s was primarily driven by a hollowing out of the middle, with much smaller wage declines at the bottom of the wage scale, and similar losses among the middle deciles. Real hourly wages fell for 80% of the male workforce in the 1980s, but from the second decile upward, the rate of growth was monotonically less negative as we move up the wage scale. In the 1990s, wage rates for males declined at a slower rate in the bottom 30% than for those in the middle of the wage scale. And, unlike the 1980s, male wage decline at the seventh and eighth deciles occurred at a similar rate to that at the low end of the wage scale.

Table 4 showed that, as measured by the Gini, female hourly wage inequality in the 1990s ORG data grew at about 1/3 the rate of the 1980s. The bottom panel of Figure

6 shows that this deceleration in female wage inequality was driven by wage growth in the 1990s that was of a comparable magnitude (and generally much slower than the earlier decade) at the second through the seventh decile. A notable exception to this pattern occurred at the 10^{th} percentile female hourly wage, which fell at about 2% annually over the 1980s, but grew slightly over the 1990s (partly due to the increases in the minimum wage over this period).

Unlike the 1980s, the pattern of wage inequality in the 1990s, as shown in Figure 6, does not conform neatly to skill-based explanations. Under this explanation, the relative wages of more skilled workers (i.e., those with higher hourly wages) are expected to increase throughout the wage distribution, as occurred in the 1980s. In the 1990s, however, we observe the wage of the highest paid workers "pulling away from the pack," while wage inequality between workers in the middle and the bottom of the wage scale were flat (females) or falling (males). Only a "highly targeted" version of the demand for skill explanation, where it is only the relative wages of the highest skilled workers that are bid up, will serve to explain this growth pattern of inequality.

The fact that the 90/50 differential continued to grow in the 1990s as it did in the 1980s, in tandem with the previous section's finding of slowing education differentials in the 1990s, suggests that within-group inequality was a major contributor to the growth in the wage gap between middle and high wage workers in the 1990s. That is, the continued growth of the 90/50 in the1990s does not appear to be a function of growing education differentials between college and high-school educated workers. The implication is that wages within these groups became more dispersed over this period.

In fact, the variance of the error term from the above male (female) regressions did grow slightly from 0.466 (0.447) in 1989 to 0.476 (0.462) in 1997, but this scalar reveals nothing about the changes *within* the residual distribution. We are interested in changes in the residuals at the same quantiles analyzed above, i.e. the 90^{th}, 50^{th}, and 10^{th} percentiles.

To do so, we attached the difference between the actual and predicted wage from the above human capital regressions to each observation in the ORG (using imputed years of education, 1992 forward). We then ordered the observations by actual wage level and calculated the residual wage for the relevant quantiles in each **year.**[14] We then computed the percentage-point change in the 90-10, 90-50, and 50-10 differentials over time. This procedure forces the change in within-group inequality to reflect the experience of high, middle, and low-wage workers. It differs from the more common approach to residual analysis (as in Juhn, Murphy, Pierce 1993), which examines the residuals "disembodied" from their origin in the actual wage distribution (a case with a high residual wage could be a person with low actual earnings). A disadvantage with this approach is that it does not conform to an additive decomposition.

The prior table revealed continued growth in the overall 90/50 differential for both males and females in the 1990s, at rates similar to that of the 1980s (slightly slower for females). Earlier tables (e.g., Table 10) revealed the 1990s slowdown in the growth of education differentials, or between-group inequality. **Table 14,** using CPS ORG data, reveals that there was a clear acceleration in the 1990s in residual inequality, driven by the growth of the dispersion of residual wages between high (90" percentile) and middle wage (50^{th} percentile) workers. For males, this measure grew 0.4 percentage points per year in the 1990s; twice the rate of the 1980s, while the male residual differential at the 50/10 grew more compressed in the 1990s. For females, within group inequality between the 90/50 grew 0.2 points per year in the 1990s after remaining unchanged in the 1980s; as with males, after growing quickly in the 1980s, the within-group 50/10 declined in the 1990s.

III. Conclusion

It is often assumed that either the trends in wage inequality in the 1990s have been simply an extension of those in the 1980s, or that the measurement problems posed by changes in the CPS data render a reasonable comparison between the 1990s and earlier periods impossible. We argue that neither of these claims are correct.

Regarding measurement, the two most important changes are the lifting of the top-codes in the March 1995 data, and the 1992 (survey year) change in the education coding. We try various adjustment methods to deal with the top-code change, and argue that the Pareto adjustment offers a reasonable approximation of the mean earnings of that

[14] Actually, we approximated the relevant quantile by taking the averages over the following wage ranges:

segment of the distribution above the top-code. The change in the education coding, from years completed to highest degree attained, is, in our view, a more difficult problem. We attempt to adjust for the change by using the a regression-based imputation procedure that exploits the 1990 February CPS which has both education questions.

Our judgment is that this procedure leads to a consistent time series when estimating education-based wage-differentials. However, we also offer two other approaches to bridging the education coding change, and they yield quite different results. For example, the growth in the regression- adjusted, male college/high-school premium, 1989-97, was 2.5 times larger using the "skip" method (i.e., skipping the year of the coding change) than using the imputed years measure. Clearly, there is room for more research here.

Nevertheless, we do not feel that these changes introduce insurmountable inconsistencies into the CPS. After all, other inconsistencies from early CPSs (and decennial Censuses) are very commonly dealt with in the literature. A good example is the calculation of weekly or hourly earnings from the March (and decennial) earnings data prior to 1975, wherein hours and weeks worked data are in bracketed (as opposed to continuous) formats. Analysts such as Katz and Murphy (1992), have used similar techniques to ours to create believable time-series from these data sets.

Having made these adjustments to the March and ORG data, we find some interesting and potentially important differences in the trend in inequality between the 1980s and 1990s. First, while wage inequality continued to grow in the 1990s, it did so at about half the rate of the 1980s. Looking more closely at change in the hourly wage distribution, we find that for both genders, the 90/50 differential continued to expand in the 1990s at a similar rate as the 1980s (slightly slower for females), while the 50/10 actually compressed. Thus, while the 1980s was a period of widening dispersion throughout the distribution, the 1990s has been characterized by the top pulling away from both the middle and bottom of the wage scale.

Another important difference is the trend in education differentials. We stress that it is important here to separate out workers with a terminal four-year college degree from those with advanced degrees, particularly since the wage trends of these two groups have diverged for the past 10 years or so. Our focus of the college/high-school premium

---

85-94 for the 90^{th} percentile, 45-54 for the 50^{th}, etc. See Mishel, et al, 1998, Table 3.17.

reveals a clear slowdown in the college/high-school premium, which, for both men and women, grew ¼ to ½ as fast in the 1990s as in the 1980s, depending on which method we use to adjust for the education coding change.

In the 1980s, the college/high-school differential grew quickly relative to earlier periods, and this was widely interpreted as evidence of an acceleration in employers' skill demands, often linked to the advance of micro-processor technology. Do our findings for the 1990s suggest then, that these skill demands have decelerated? We examine supply trends to see if perhaps increased relative supplies can be blamed for the slowdown, and find no such evidence.[15] While there is some reason to think that changes in labor market institutions may have played a role in slowing the growth of differentials (specifically, the increase in the minimum wage and the deceleration of the decline in union density), large increases in trade imbalances over the period should have had the opposite effect. We conclude that the evidence suggests either a slowing in the rate of skill-biased technological change or that such change was slower in the 1990s than the 1980s.

There is clearly room for more research into these questions. Like many labor analysts who have examined these trends, our approach has been to dis-aggregate by gender, implicitly assuming different demand, supply, and institutional (including discrimination) structures facing each gender. We make a similar assumption regarding those with four-year college degrees and those with advanced degrees. Some researchers may not be so quick to accept these assumptions; Let-man (1997), for example, argues for examining wage inequality by combining males and females.

Also, while we judge our solutions to the coding challenges posed by changes in the CPS to yield a consistent time series, future work may discover better adjustment methods. We are confident, however, that the 1990s trends identified above are real and not artifacts of the coding changes.

[15] Our examination of these relative supply trends reveals the potential bias introduced by ignoring the change in education codes in the CPS.

Bibliography

Autor, David, Lawrence Katz, and Alan B. Krueger. 1997. Revised from July 1996. "Computing Inequality: Have Computers Changed the Labor Market?" Princeton University Working Paper No. 377. Princeton, NJ: Princeton University Press.

Bernstein, Jared and Lawrence Mishel. 1997. Has wage inequality stopped growing? *Monthly Labor Review.* Vol. 120, No. 12, pp. 3-16.

Gladieux, Lawrence E. and Watson Scott Swail. 1998. "Postsecondary Education: Student Success Not Just Access." In Samuel Halperin, Ed., *The Forgotten Half Revisited.* Washington, D.C.: American Youth Policy Forum.

Johnson, George E. 1997. Changes in earnings inequality: the role of demand shifts. *Journal of Economic Perspectives.* Vol. 11, No. 2, Spring, pp. 41-54.

Juhn, Chinhui, Kevin M. Murphy, and Brooks Pierce. 1993. Wage inequality and the rise in returns to skill. *Journal of Political Economy.* Vol. 101, June, pp. 410-442.

Karoly, Lynn A. and Bar-y Burtless. 1995. Demographic change, rising earnings inequality, and the distribution of personal well-being, 1959-1989. *Demography.* Vol. 32, no. 3, pp. 379-405.

Katz, Lawrence F. and David H. Autor. 1998. "Changes in the Wage Structure and Earnings Inequality." In Orley Ashenfelter and David Card, eds., *Handbook of Labor Economics.* May.

Lerman, Robert I. 1997a. "Is Earnings Inequality Really Increasing?" The Urban Institute. No. 1 in Series, "Economic Restructuring and the Job Market." Washington, D.C.: Urban Institute Press.

Lerman, Robert I. 1997b. Reassessing trends in U.S. earnings inequality. *Monthly Labor Review.* Vol. 120, No. 12, pp. 17-25.

Mishel, Lawrence and Jared Bernstein. 1992. "Declining Wages for High School and College Graduates." Economic Policy Institute, Briefing Paper No. 3 1. Washington, D.C.: EPI.

Mishel, Lawrence, Jared Bernstein, and John Schmitt. 1998. *The State of Working America, 1998-1999.* Economic Policy Institute Series. Ithaca, NY: Cornell University Press.

Mishel, Lawrence, Jared Bernstein, and John Schmitt. 1997. "Did Technology Have Any Effect on the Growth of Wage Inequality in the 1980s and 1990s?" Economic Policy Institute Technical Paper #230.

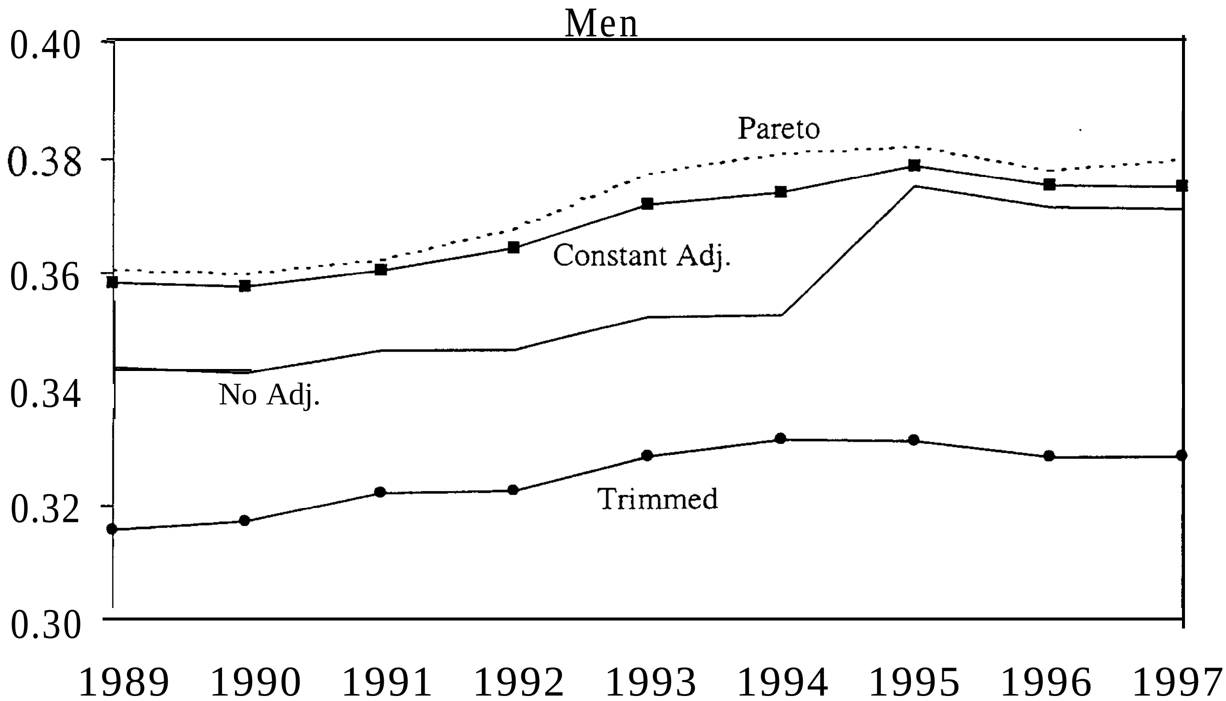
Polivka, Anne E. "Using earnings data from the Current Population Survey after the redesign." Bureau of Labor Statistics, Working Paper No. 306. Washington, D.C.: BLS.

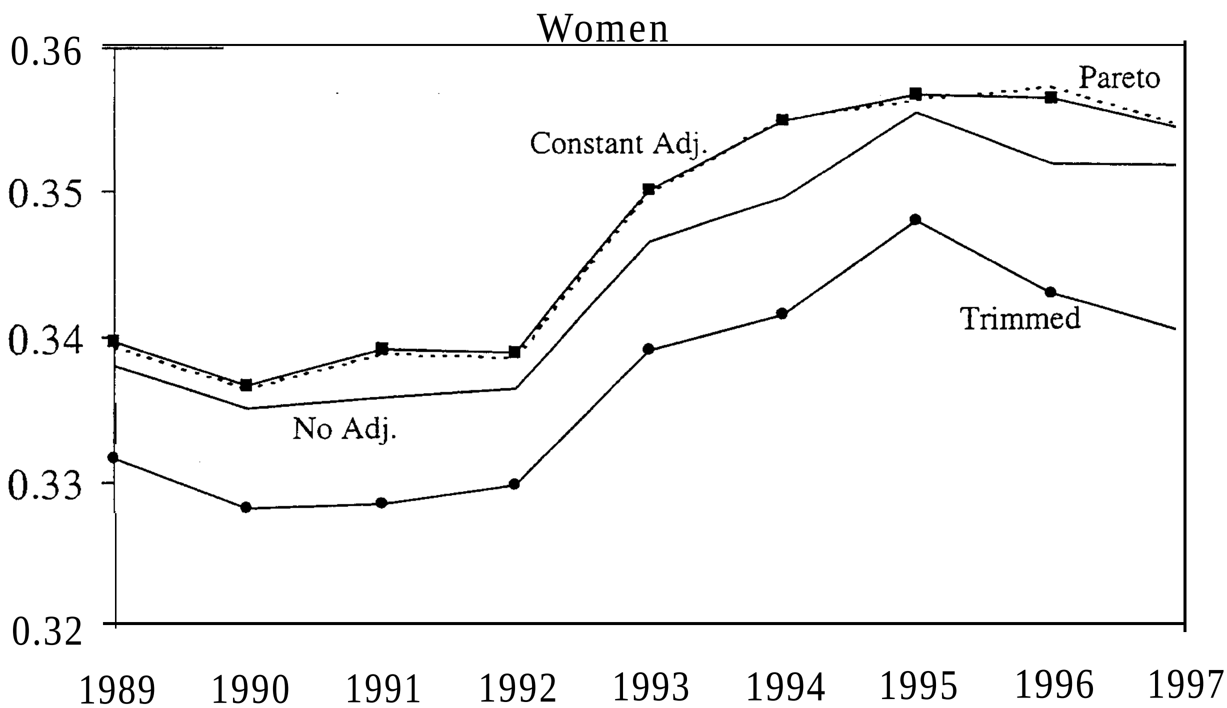
Ryscavage, Paul. 1995. A surge in growing income inequality? *Monthly Labor Review.* August, pp. 5 l-6 1.

Utendorf, Kelvin R. 1998. Recent changes in earnings distributions in the United States. *Social Security Bulletin.* Vol. 61, No. 2, p. 12-29. Washington, D.C.: Social Security Administration.

Webster, David E. 1998. "Wage Analysis Computations." In Lawrence Mishel, Jared Bernstein, and John Schmitt, *The State* **of** *Working America, 1998-1999.* Appendix B. Economic Policy Institute Series. Ithaca, NY: Cornell University Press.

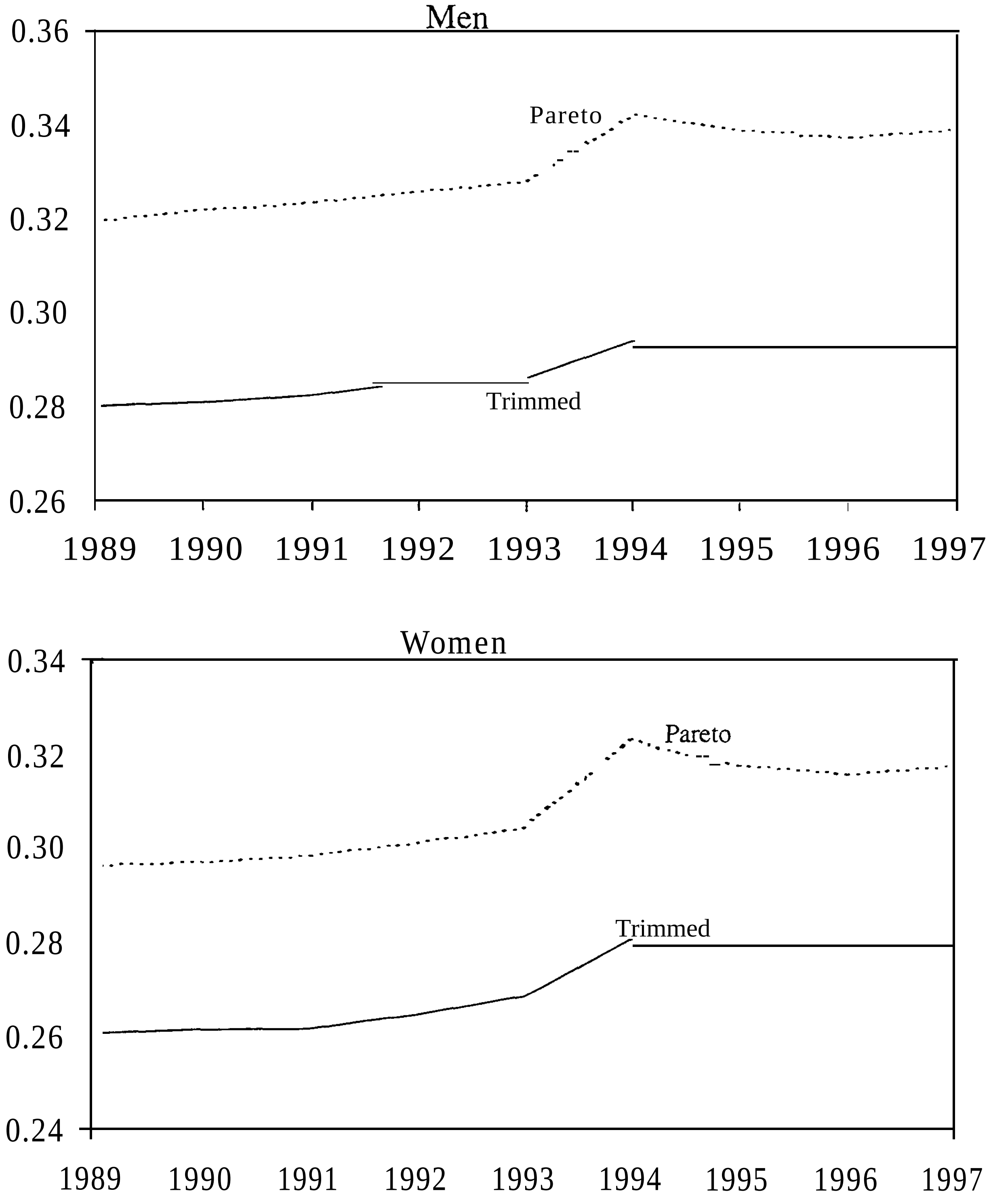
Figure 1
Gini Coefficients, Hourly Wages, March CPS, 1989-97





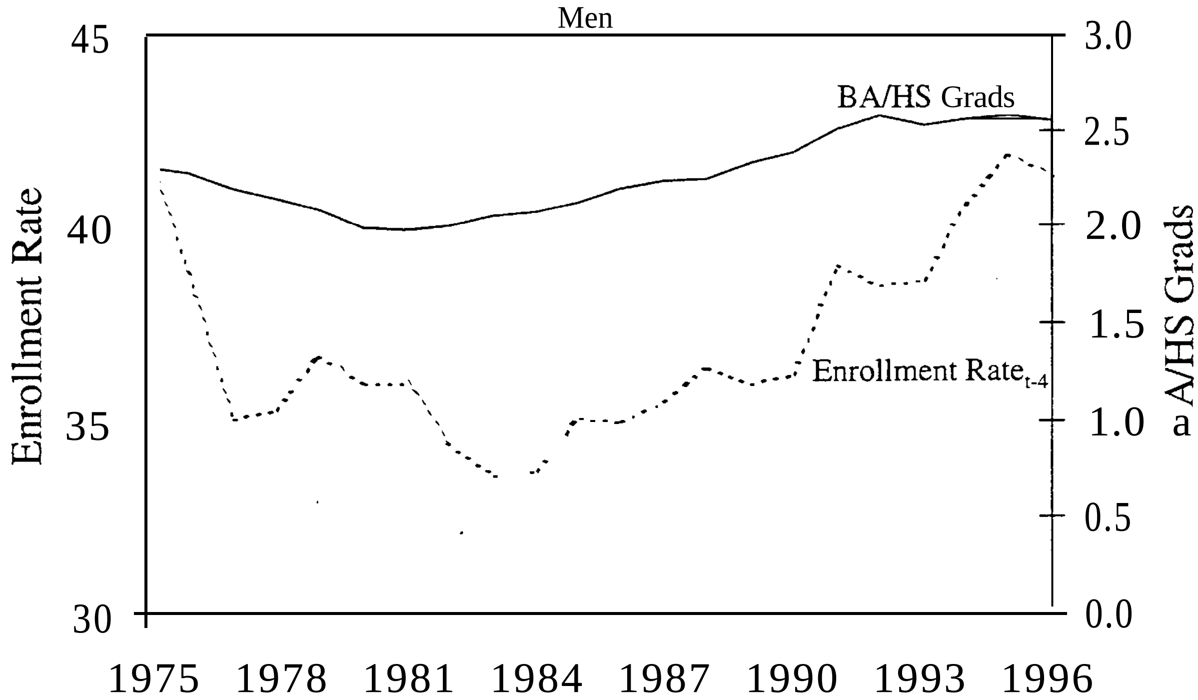
Source: Table 2

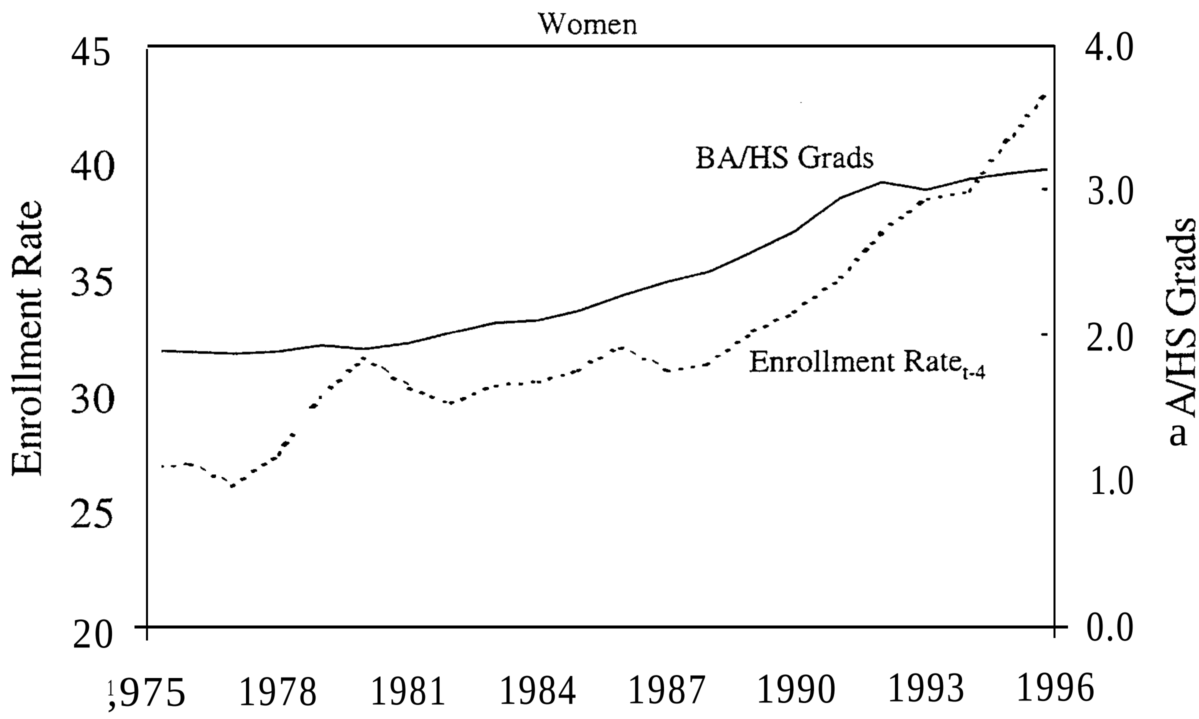
Figure 2
Gini Coefficients, Hourly Wages, ORG CPS, 1989-97



Source: Table 3

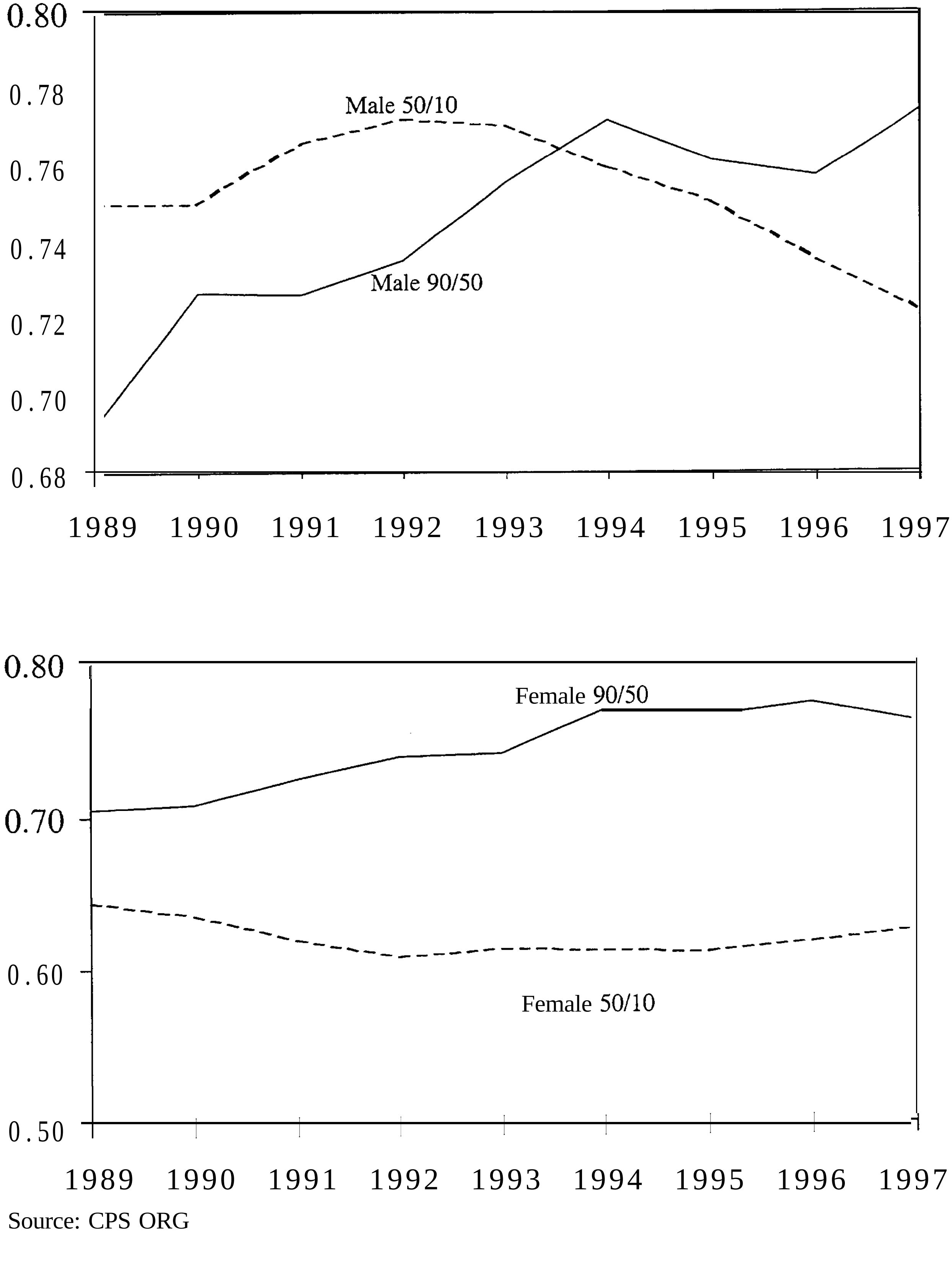
Figure 3
Enrollment, 4 & College Completion Rates, 1975-97





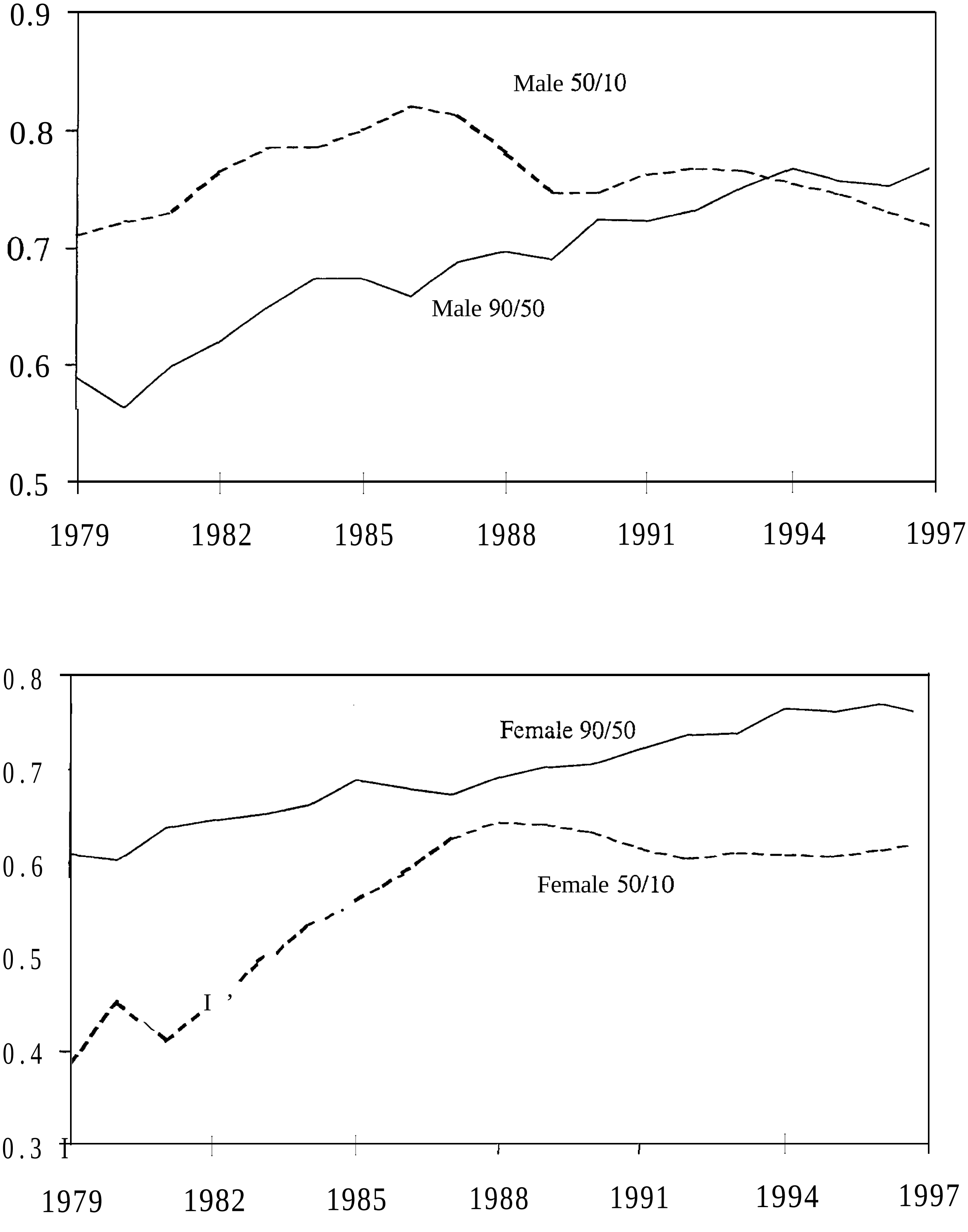
Source: Department of Education & CPS

Figure 4
Log Hourly Wage Differentials, 90/50, 50/10, 1989-97



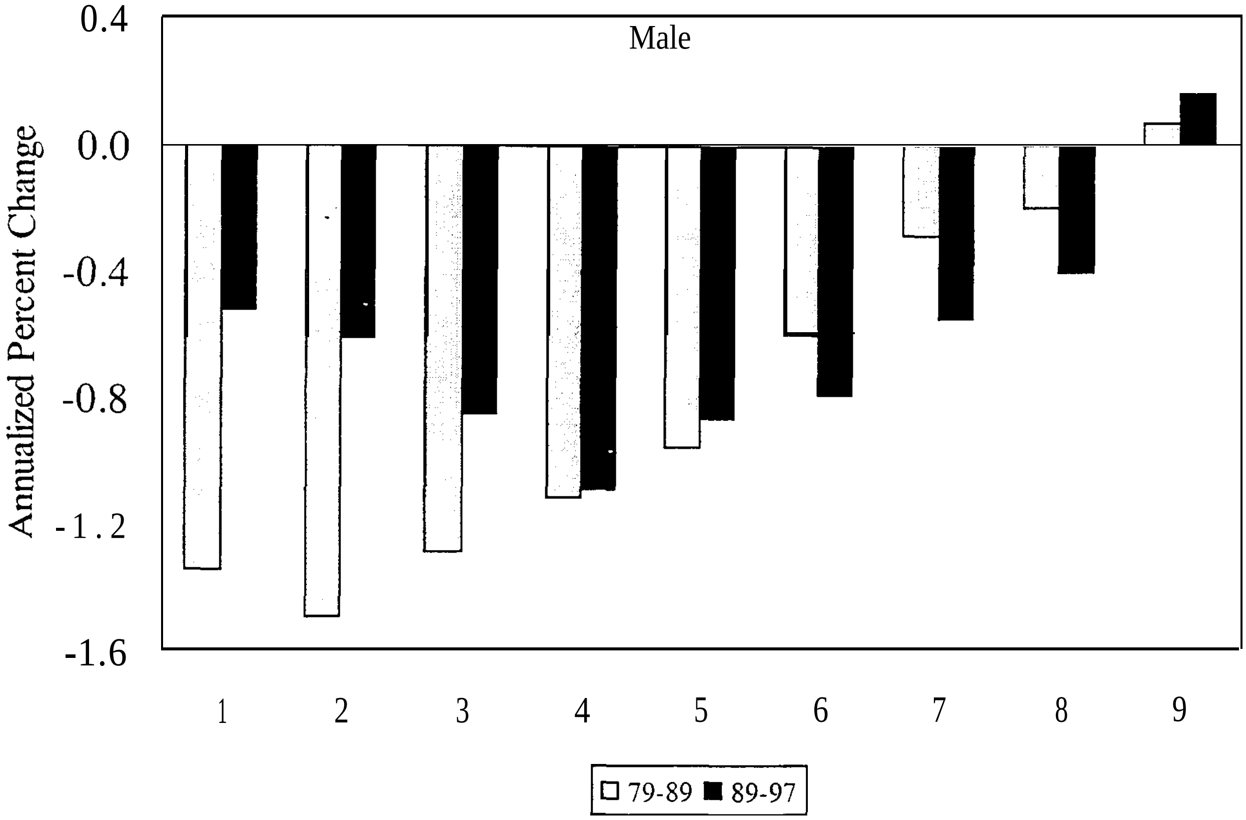
Source: CPS ORG

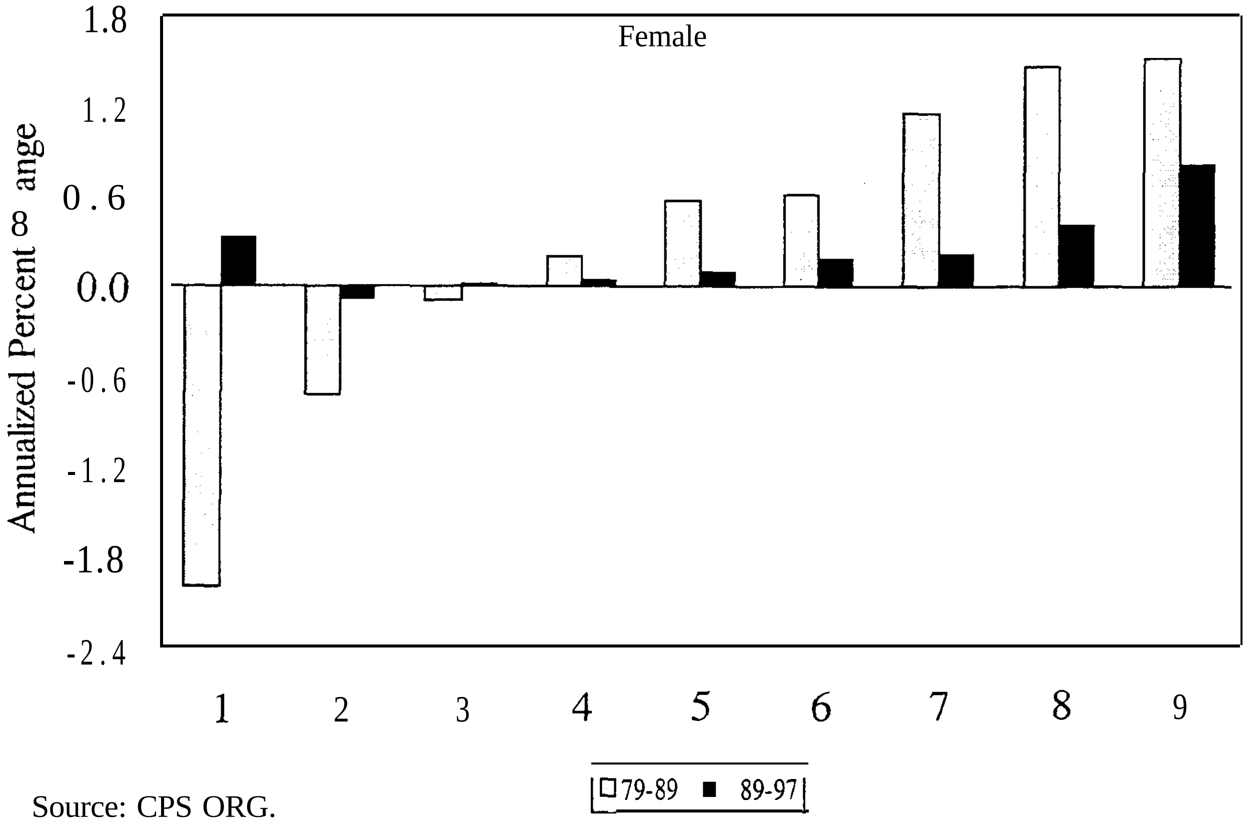
Figure 5
Log Hourly Wage Differentials, 90/50, 50/10, 1979-97



Source: CPS ORG

Figure 6
Annualized Percent Changes, Real Wage Deciles, 1979-97, 1989-97,





Source: CPS ORG.

Table 2
Gini Coefficients for Hourly Wage Inequality in March CPS, 1989-97

Gender/ Year	Gini for Hourly Earnings: Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Men				
1989	0.361	0.344	0.316	0.358
1990	0.360	0.343	0.317	0.358
1991	0.363	0.346	0.322	0.361
1992	0.368	0.346	0.323	0.364
1993	0.378	0.352	0.329	0.372
1994	0.381	0.353	0.331	0.374
1995	0.382	0.375	0.331	0.378
1996	0.378	0.371	0.328	0.375
1997	0.380	0.371	0.328	0.375
Change				
1992-97	0.012	0.024	0.006	0.011
1989-97	0.019	0.027	0.012	0.016

	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Women				
1989	0.339	0.338	0.332	0.340
1990	0.337	0.335	0.328	0.337
1991	0.339	0.336	0.328	0.339
1992	0.339	0.336	0.330	0.339
1993	0.350	0.346	0.339	0.350
1994	0.355	0.349	0.341	0.355
1995	0.356	0.355	0.348	0.357
1996	0.357	0.352	0.343	0.356
1997	0.354	0.352	0.340	0.354
Change				
1992-97	0.016	0.015	0.010	0.015
1989-97	0.015	0.014	0.008	0.014

Source: March CPS

Table 3
Gini Coefficients for Hourly Wage Inequality in ORG CPS, 1989-97

	Gini for Hourly Earnings			
	Men		Women	
Year	Pareto Adjustment	Trimmed*	Pareto Adjustment	Trimmed*
1989	0.319	0.280	0.296	0.261
1990	0.322	0.281	0.297	0.261
1991	0.324	0.282	0.298	0.261
1992	0.326	0.285	0.301	0.264
1993	0.329	0.286	0.304	0.268
1994	0.342	0.294	0.323	0.280
1995	0.339	0.293	0.318	0.280
1996	0.338	0.293	0.316	0.279
1997	0.339	0.291	0.317	0.278
Change				
1992-97	*0.013*	*0.007*	*0.016*	*0.014*
1989-97	*0.020*	*0.012*	*0.021*	*0.017*

• Bottom 97% of sample
Source: Outgoing Rotation Group Current Population Survey

Table 4
Hourly Wage Inequality

CPS Source	1979-89 Men	1979-89 Women	*1989-97* Men	*1989-97* Women	*1992-97* Men	*1992-97* Women
Percent Change *in Gini**						
ORG	15.4%	20.5%	6.2%	7.2%	4.1%	5.5%
March	12.3	13.0	5.3	4.4	3.3	4.4

* Using pareto adjustment to topcodes
Source: March and Outgoing Rotation Group CPS.

Table 5
Comparison of Wage Inequality Trends in Social Security and March CPS Data, 1989-93

Gender/ Year	Gini, March CPS Annual Earnings: Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Gini, Social Security Annual Earnings*	Ratio of March CPS to Social Security, Gini: Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Men									
1989	0.420	0.400	0.368	0.418	0.511	82.3%	78.3%	72.0%	81.7%
1990	0.420	0.399	0.368	0.417	0.511	82.1%	78.1%	72.0%	81.6%
1991	0.425	0.405	0.376	0.423	0.514	82.7%	78.8%	73.1%	82.2%
1992	0.435	0.411	0.382	0.431	0.513	84.8%	80.0%	74.4%	84.0%
1993	0.443	0.413	0.385	0.436	0.528	83.9%	78.1%	72.9%	82.6%
1994	0.444	0.411	0.385	0.436					
1995	0.439	0.450	0.380	0.439					
1996	0.437	0.451	0.378	0.434					
1997	0.437	0.449	0.378	0.433					

Gender/ Year	Gini, March CPS Annual Earnings: Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Gini, Social Security Annual Earnings*	Ratio of March CPS to Social Security, Gini: Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Women									
1989	0.431	0.428	0.421	0.431	0.477	90.3%	89.8%	88.2%	90.4%
1990	0.428	0.425	0.417	0.428	0.477	89.6%	89.1%	87.3%	89.7%
1991	0.430	0.427	0.417	0.431	0.477	90.2%	89.5%	87.5%	90.3%
1992	0.429	0.425	0.418	0.429	0.471	91.0%	90.3%	88.7%	91.1%
1993	0.436	0.432	0.423	0.437	0.488	89.4%	88.5%	86.6%	89.5%
1994	0.443	0.436	0.425	0.443					
1995	0.435	0.440	0.423	0.437					
1996	0.439	0.443	0.420	0.439					
1997	0.437	0.443	0.418	0.438					

Sources: March CPS and *Utendorf (1998).

Table 6
College-High School Wage Differentials in March CPS, 1989-97

March Imputed

	College Only				More Than College				College or More			
	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Men												
1989	40.9%	46.5%	34.2%	40.7%	56.6%	64.5%	42.3%	56.2%	46.2%	44.3%	36.7%	46.0%
1990	43.5	46.7	35.7	43.3	57.0	64.5	43.0	56.7	47.7	45.6	37.7	47.4
1991	41.8	46.0	35.3	41.6	60.5	67.3	45.4	60.2	48.2	46.1	38.5	48.0
1992	43.3	46.9	36.2	43.1	63.9	68.7	48.3	63.3	50.3	47.6	40.1	50.0
1993	44.7	48.4	37.7	44.3	68.0	69.9	49.3	64.7	51.7	48.3	41.2	51.0
1994	44.3	47.2	36.8	43.8	63.4	67.8	46.3	62.0	50.9	47.0	39.9	50.1
1995	43.3	47.2	35.8	42.4	67.9	71.8	51.6	66.7	51.7	50.1	40.9	50.6
1996	42.3	45.6	34.7	41.4	65.9	71.1	49.6	64.4	50.1	47.9	39.3	49.0
1997	43.8	46.0	37.0	42.4	68.5	71.7	52.4	66.9	51.8	49.5	41.5	50.3
1989-97	2.9	-0.5	2.8	1.6	11.9	7.2	10.1	10.7	5.5	5.2	4.8	4.3
1990-97	0.3	-0.7	1.3	-0.9	11.4	7.1	9.4	10.2	4.1	3.9	3.8	2.9
1992-97	0.4	-0.9	0.7	-0.7	4.6	3.0	4.1	3.6	1.4	1.9	1.4	0.3

	College Only				More Than College				College or More			
	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment	Pareto Adjustment	No Adjustment	Trimmed	Constant Adjustment
Women												
1989	49.3%	39.8%	48.6%	49.3%	68.4%	53.1%	66.3%	68.4%	55.2%	55.1%	54.1%	55.2%
1990	48.5	42.0	47.6	48.6	68.8	53.8	65.7	68.9	54.7	54.6	53.2	54.8
1991	48.5	40.6	47.9	48.5	72.6	56.6	69.5	72.6	55.6	55.4	54.0	55.6
1992	50.1	41.5	49.4	50.1	73.2	59.2	70.5	73.3	57.0	56.8	55.6	57.0
1993	52.2	42.6	51.4	52.2	76.0	59.9	72.3	76.1	59.1	58.7	57.1	59.1
1994	50.7	41.7	50.1	50.7	75.6	56.9	71.8	75.5	58.1	57.7	56.3	58.1
1995	51.3	42.0	50.8	51.5	73.7	65.8	70.6	73.5	58.1	58.1	56.8	58.2
1996	50.5	40.3	49.9	50.6	75.9	63.6	71.1	75.3	57.7	57.1	55.7	57.6
1997	50.0	41.8	48.7	49.9	74.0	65.8	69.3	73.7	57.0	56.7	54.6	56.8
1989-97	0.7	2.0	0.0	0.6	5.6	12.7	3.0	5.3	1.8	1.7	0.5	1.6
1990-97	1.5	-0.1	1.1	1.3	5.1	12.0	3.6	4.8	2.2	2.1	1.3	2.0
1992-97	0.0	0.3	-0.7	-0.2	0.7	6.6	-1.2	0.4	0.0	0.0	-1.1	-0.2

Source: Authors' analysis of March CPS. Education coding for 1991-97 is based on imputed years of schooling completed, with 16 years being "College Only" and 17 years or more being "More Than College." Estimates are from a simple human capital model wilh controls for experience as a quartic, marital status, region, and race/ethnicity and are presented relative to high school graduates.

Table 7
College-High School Wage Differentials in March CPS Using Degree-Attained Coding, 1991-97

Men	College Only: Pareto Adjustment	College Only: No Adjustment	College Only: Trimmed	College Only: Constant Adjustment
1991	33.6%	32.5%	27.1%	33.5%
1992	35.3	33.5	28.1	35.0
1993	36.6	34.5	29.6	36.2
1994	36.8	34.2	29.3	36.2
1995	36.5	35.3	29.1	35.6
1996	35.7	33.8	28.2	34.8
1997	37.2	35.3	30.5	35.6
1991-97	3.6	2.6	3.4	2.3
1992-97	1.9	1.8	2.4	0.8

Men	More Than College: Pareto Adjustment	More Than College: No Adjustment	More Than College: Trimmed	More Than College: Constant Adjustment
1991	51.2%	47.4%	36.3%	50.9%
1992	54.7	50.2	39.4	54.1
1993	56.8	51.1	40.6	55.7
1994	55.4	49.1	38.6	54.0
1995	60.5	58.6	44.6	59.3
1996	58.6	56.3	42.6	57.1
1997	61.2	58.5	45.1	59.6
1991-97	5.4	5.2	6.8	8.7
1992-97	6.5	8.3	5.7	5.5

Men	College or More: Pareto Adjustment	College or More: No Adjustment	College or More: Trimmed	College or More: Constant Adjustment
1991	39.6%	37.6%	30.1%	39.4%
1992	42.0	39.3	31.8	41.6
1993	43.3	40.0	33.0	42.7
1994	43.1	39.3	32.2	42.3
1995	44.5	43.0	34.0	43.5
1996	43.3	41.2	32.7	42.2
1997	45.0	42.8	34.9	43.6
1991-97	5.4	5.2	4.9	4.1
1992-97	3.1	3.6	3.2	2.0

Women	College Only: Pareto Adjustment	College Only: No Adjustment	College Only: Trimmed	College Only: Constant Adjustment
1991	41.7%	41.6%	41.1%	41.7%
1992	43.2	43.1	42.5	43.2
1993	45.5	45.4	44.7	45.5
1994	44.6	44.5	44.0	44.6
1995	45.1	45.2	44.6	45.3
1996	44.9	44.6	44.2	44.9
1997	44.6	44.5	43.3	44.5
1991-97	2.9	2.9	2.2	2.8
1992-97	1.5	1.4	0.8	1.3

Women	More Than College: Pareto Adjustment	More Than College: No Adjustment	More Than College: Trimmed	More Than College: Constant Adjustment
1991	66.6%	65.9%	63.2%	66.6%
1992	66.7	66.0	63.9	66.8
1993	69.0	68.1	64.9	69.0
1994	69.3	66.3	65.5	69.3
1995	68.3	68.3	65.5	68.2
1996	70.2	68.9	65.4	69.7
1997	68.4	68.0	63.8	68.2
1991-97	1.8	2.1	0.6	1.5
1992-97	1.7	2.0	-0.1	1.4

Women	College or More: Pareto Adjustment	College or More: No Adjustment	College or More: Trimmed	College or More: Constant Adjustment
1991	40.9%	40.7%	47.4%	48.9%
1992	50.1	49.9	48.8	50.2
1993	52.3	52.0	50.5	52.4
1994	52.0	51.6	50.3	51.9
1995	51.9	51.9	50.6	51.9
1996	52.1	51.5	50.0	51.9
1997	51.4	51.2	49.1	51.3
1991-97	2.5	2.6	1.7	2.4
1992-97	1.3	1.3	0.3	1.1

Source: Authors' analysis of March CPS data using "Highest Degree Attained" coding of education.
Estimates are from a simple human capital model with controls for age as a quartic, marital status, region, and race/ethnicity and are presented relative to high school graduates.

Table 8
March CPS College Education Differentials by Method, Period, and Category, 1989-97

	Pareto Adjustment		Trimmed		Constant Adjustment	
Period	College Only	College Or More	College Only	College Or More	College Only	College Or More
P e r c e n t a g e						
Men						
1989-97*	2.9%	5.5%	2.8%	**4.8%**	1.6%	4.3%
1990-97*	0.3	4.1	1.3	**3.8**	-0.9	2.9
1992-97**	1.9	3.1	2.4	**3.2**	0.8	2.0
Women						
1989-97'	**0.7%**	**1.8%**	**0.0%**	**0.5%**	**0.6%**	1.6%
1990-97*	**1.5**	**2.2**	1.1	**1.3**	**1.3**	**2.0**
1992-97"	1.5	**1.3**	**0.8**	**0.3**	**1.3**	**1.1**

***** Based on imputed years of schooling completed.

. * Based on new coding.

Source: Tables 6 and 7.

Table 9
College-High School Wage Differentials in ORG CPS, 1989-97

Year	Men: College Only	Men: College or More	Women: College Only	Women: College or More
Wage Differential Relative to High School:				
Imputed Years				
1989	41.8%	46.2%	46.0%	53.1%
1990	42.4	48.0	47.7	54.2
1991	42.3	48.3	46.5	53.4
1992	44.4	49.6	49.4	55.8
1993	45.1	50.4	49.9	56.7
1994	44.8	50.8	51.9	59.7
1995	44.1	50.3	52.2	59.2
1996	43.3	49.5	50.7	58.1
1997	44.1	50.1	51.1	58.0
Change				
1992-97	-0.3	0.5	1.7	2.1
1989-97	43.7	4.0	50.6	4.9
Highest Degree Attained*				
1992	36.9%	41.9%	43.7%	49.7%
1993	37.6	42.2	44.3	50.6
1994	37.5	42.8	46.3	53.6
1995	37.3	42.5	46.7	53.5
1996	36.8	42.2	45.4	52.7
1997	38.1	43.4	46.2	52.9
Change				
1992-97	37.7	43.0	45.7	52.4

Source: Authors' analysis of CPS ORG data. Pareto adjusted top-codes are used. Estimates are from a simple human capital model with controls for experience (or age*) as a quartic, marital status, region, and race/ethnicity and are presented relative to high school graduates.

Table 10
Estimates of Change in College-High School Wage Differential, 1979-97

Period	Men March	Men ORG	Women March	Women ORG
1979-89	14.3	14.5	16.7	15.2
1989-97 (Ten-Year Rate of Change)				
Skip*	8.9	3.0	3.1	4.9
Splice'*	7.6	4.6	1.3	6.1
Imputed***	3.6	2.9	0.9	6.4
1992-97	3.8	2.2	3.0	4.8

- Skip year where education coding changed, e.g.,1991-92 for ORG.

** Substitute trend in other CPS series to bridge trend.

*** Use imputed years of schooling rather than degree atttained.

Table 11
Change in Shares of Employment by Education in ORG CPS, 1991-92

Men	Less Than High School	High School	Some College	College	Advanced Degree	College & Advanced Degree	Total
1991[a]	14.7%	38.5%	21 . a%	17.2%	7.8%	25.0%	100.0%
1991[b]	14.7	38.5	21.8	14.7	10.3	25.0	100.0
1992	13.8	34.9	26.5	16.5	8.4	24.9	100.1
1992 vs. 1991[a]	-0.9	-3.6	4.7	-0.7	0.6	-0.1	0.1
1992 vs. 1991[b]	-0.9	-3.6	4.7	1.8	-1.9	-0.1	0.1

Women	Less Than High School	High School	Some College	College	Advanced Degree	College & Advanced Degree	Total
1991[a]	10.2%	41 .a%	24.5%	17.4%	6.1%	23.5%	100.0%
1991[b]	10.2	41 .a	24.5	14.8	a.7	23.5	100.0
1992	9.6	36.7	30.1	16.5	7.1	23.6	100.0
1992 vs. 1991[a]	-0.6	-5.1	5.6	-0.9	1.1	0.1	0.0
1992 vs. 1991[b]	-0.6	-5.1	5.6	1.7	-1.6	0.1	0.0

1991[a] has 17 years of schooling combined with "College."
1991 [b] has 17 years of schooling combined with "Advanced Degree."
In 1992, "Some College" includes those with "some college" and those with associates' degrees.
Source: Authors' analysis of CPS ORG data.

Table 12
Change in College-Educated Shares of Employment, 1979-97

	Percentage Point Change in Employment Share of College Graduates*			
	ORG		March	
	Men	Women	Men	Women
1979-89	2.7	3.4	–	--
1989-97	(Ten-Year Rates of Change)			
Skip	2.3	4.4	2.6	4.7
Splice	2.3	4.8	2.5	4.8

* "College Only" share of employment.

The "Skip" method omits the change between the years where an education coding discontinuity exists and scales the seven years of change to ten years.

The "Splice" method substitutes the alternative CPS change in share for the change in the period where the discontinuity exists and scales to ten years.

Table 13

Annualized Growth in Log Wage Differentials, 1979-97*

Males	90/50	50/10	90/10
1979-89	1.0	**0.4**	**1.4**
1989-97	1.0	**-0.4**	**0.7**
Females			
1979-89	1.0	**2.6**	**3.5**
1989-97	**0.7**	**-0.2**	**0.5**

Source: CPS ORG

* Difference in log wage differential, divided by number of years (*100).

Table 14

Annualized Changes in Within-Group Inequality, ORG CPS, 1979-97

Males	90/50	50/10	90/10
1979-89	0.2	0.5	0.6
1989-97	0.4	-0.6	-0.3
Females			
1979-89	0.0	1.5	1.5
1989-97	0.2	-0.7	-0.5

Source: Mishel et al, 1999, Table 3.17.