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ABSTRACT

Large increases in educational attainment have resulted in dramatic shifts in the composition of educational groups. Utilizing 1960-90 Decennial Census and other data, this paper uses educational ranks (cohort-specific relative rankings in educational attainment) as a control for changes in the composition of educational groups. This approach assumes that people in different cohorts with the same educational rank have about the same level of ability. The paper also examines a second approach to controlling for changes in the composition of educational groups, within cohort comparisons. For native white males between 1969-89, accounting for changes in the composition of educational groups: (1) explains about half of the increase in the college-high school weekly earnings differential; (2) results in increases in weekly earnings for the less educated; and (3) doubles the increases in experience differentials for high school graduates who are less educated. The paper questions the common research strategy of using educational groups as a proxy for skill groups over long time periods, noting that estimates of the returns to skill using education differentials are likely to present a misleading portrait of the labor market and arguing that this misleading portrait has been significant over time. (Contains 30 references and 9 tables.) (SM)

Ability, Educational Ranks, and Labor Market Trends:
The Effects of Shifts in the Skill Composition of Educational Groups

JCPR Working Paper 146

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ABSTRACT
Large increases in educational attainment have resulted in dramatic shifts in the composition of educational groups. Utilizing the 1960-1990 Decennial Census and other data sources, I account for these changes in composition using educational ranks – cohort-specific relative rankings in educational attainment. For native white males between 1969 and 1989 accounting for changes in the composition of educational groups (1) explains about half of the increase in the college/high school weekly earnings differential, (2) results in increases in weekly earnings for the less educated, and (3) doubles the increases in experience differentials for high school less educated. These findings raise questions about the common research strategy of using educational groups as a proxy for skill groups over long periods of time.

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I. Introduction

"Perhaps the most important change in the labor market over the past 25 years has been the increase in the demand for more educated workers" (*Economic Report of the President*, 1998, p. 149). Findings of widening differentials in earnings by education along with "wages for less educated workers [that] deteriorated . . . in the mid-1970s and 1980s" (p. 148) color the way in which both policymakers and researchers think about the world. For that reason it is important that this picture of the U.S. labor market be as accurate as possible.

This paper argues that increases in education differentials have been overstated and that the wages of less educated workers have not "deteriorated" if we take into account one simple fact. Over the past few decades, educational attainment has increased and with it, the composition of educational groups has changed, implying that using educational groups is a particularly bad way to examine how both inequality and the returns to skill have changed over time. Accounting for these changes in the composition of educational groups results in a very different picture of the U.S. labor market, including (1) increases between 1969 and 1989 in the college/high school weekly earnings differential that are about half as large, (2) weekly earnings for high school graduates and dropouts that increase between 1969 and 1989 rather than decrease, and (3) a doubling of the 1969-1989 increases in the differences in weekly earnings between older and younger high school graduates and dropouts.[1]

This new picture of the U.S. labor market is important for both policymakers and the vast literature in economics in search of explanations for the rising U.S. wage inequality over the past few decades.[2] In fact, these findings challenge one of the bodies of evidence for increases in

[1] These results are for native white males with 1-30 years of experience.
[2] See Levy and Murnane (1992) and Katz and Autor (1998) for a review of the wage inequality literature. See Gottschalk and Smeeding (1997) for a discussion of international wage inequality. Explanations for

wage inequality – increases in education differentials. Other recent research raises concerns about the other body of evidence for wage inequality – increases in within group inequality.[3]

This literature has tended to ignore changes in the composition of educational groups resulting from the dramatic increases in educational attainment over the past few decades. One particularly extreme example comes from Autor, Katz, and Krueger (1998) which compare the wages of college and high school graduates between 1940 and 1996. In 1940, college graduates made up the upper 6.4 percentiles of the educational attainment distribution, while high school graduates were between the 68th and 87th percentiles (p.1174). By 1996, college graduates made up the upper 28 percentiles, while high school graduates were between the 9th and 43rd percentiles (p. 1174). If educational attainment and ability (or unobserved skill) are positively correlated, then comparisons such as these require heroic assumptions in order to be interpretable as changes in only the return to skill.[4] Given that the decline in educational rankings for high school graduates has fallen much faster for high school graduates than it has for college graduates (due to the expansion of the some college group), one might expect that changes in the composition of educational groups would have resulted in a large increase in the college/high school wage differential between 1940 and 1996, even in the absence of any change in the returns

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the increases in U.S. wage inequality explore most of the institutions shaping the U.S. labor market, including skill-biased technological change (Bound and Johnson, 1992; Autor, Katz, and Krueger, 1998), increased globalization (Wood, 1994), changes in the relative supplies of skilled labor (Katz and Murphy, 1992; Card and Lemiex, 1999), and a deterioration of the labor market institutions that benefit less skilled workers (DiNardo, Fortin, and Lemieux, 1996).

[3] Hoxby and Terry (1999) suggest that almost half of the increase in within group inequality among college graduates may be due to changes in the demographics of college graduates and the sorting among colleges. Haider (1998) argues that increases in the instability of earnings explains much of the increase in wage inequality in the 1970s.

[4] Ability refers to attributes or skills which are not acquired through education, but are correlated with both educational and labor market outcomes. Consequently, ability incorporates not only traditional ability attributes, such as those measured by IQ tests, but also attributes like family background, physical

to ability and schooling. In this light, the six percentage point increase in the college/high school wage differential between 1940 and 1996 (p. 1174) likely indicates that wage inequality *decreased* rather than *increased* between 1940 and 1996.

This paper uses educational ranks – cohort-specific relative rankings in educational attainment – as a control for changes in the composition of educational groups.[5] This approach assumes that individuals in different cohorts with the same educational rank, i.e. those in the same place in their cohort's educational attainment distribution, have about the same level of ability. Thus, the ability (or unobserved skill) of high school graduates in 1940 was more like that of college graduates in 1996 than their counterpart high school graduates.

As mentioned above, accounting for these changes in the composition of educational groups profoundly changes our picture of the U.S. labor market. Increases in the college/high school differential are considerably smaller but more concentrated among younger workers, and it is more difficult to argue that the earnings of less educated groups have deteriorated.

One of primary advantages of using educational ranks as a control for changes in the composition of educational groups is that it applicable to the large, nationally representative data sets that span long time periods and multiple cohorts, such as the Decennial Census (Census) which is used in this paper. This applicability is in contrast to another approach, using test scores

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attractiveness, determination, perseverance, and congeniality that affect both educational and labor market outcomes.

[5] Note that when computing educational ranks, all persons are placed into four mutually exclusive educational groups: high school dropouts (persons without a high school diploma or GED), high school graduates (persons with a high school diploma or a GED), some college (persons who attended college but did not receive a college degree), and college graduates (persons with a college diploma, including those with advanced college degrees). Conditional on being in a given educational group, educational ranks differ by cohort and year. See Section 3 for further details on educational categories and the procedure for constructing educational ranks.

as a control for ability, which typically are available in surveys that span only a short time period or only a few cohorts.[6] A second advantage of this approach stems from educational ranks being a function of educational attainment. Hence, educational ranks are an ideal control for the variation in ability (or unobserved skill) that is correlated with educational attainment, precisely the variation which biases estimates of changes over time in the returns to skill.

This paper also examines a second approach of controlling for changes in the composition of educational groups – within cohort comparisons.[7] One would expect that compositional shifts would be negligible within cohort, yet the evidence presented below finds substantial increases in educational attainment even as cohorts enter their 40s and 50s, thereby raising questions about the validity of this approach.

There is more at stake in this paper than just estimates of earnings patterns for different educational groups. In essence, any research that uses educational groups to delineate different skill groups or conditions on education over long periods of time runs the risk of confounding changes in composition with changes in the parameter of interest. For example, estimates of changes over time in race differentials often compare blacks and whites with a given educational attainment. Consequently, estimates of changes over time in race differentials may simply reflect racial differences the patterns of educational attainment.

2. Educational Ranks and Compositional Shifts

[6] One notable exception is Murnane, Willet, and Levy (1995), which compares samples in two different surveys from different time periods. However, differences in the achievement tests and the surveys make these comparisons somewhat difficult to interpret.
[7] See Murphy and Welch (1993); Juhn, Murphy, and Pierce (1993); MaCurdy and Mroz (1995); Duncan, Boisjoly, and Smeeding (1996), and Juhn, Kim, and Vella (1998) for other papers applying within cohort comparisons.

This section presents a reduced form model of wage determination to illustrate how educational ranks can be used to control for changes over time in the composition of educational groups. Given that the purpose of this paper is to investigate the ramifications of assumptions made in previous work using similar reduced form models, this approach seems appropriate.

Equation (2.1) presents a simple econometric model where wages are determined by both observed and unobserved skills:

(2.1) $w_{it} = \beta_t s_{it} + \phi_t a_i + \varepsilon_{it}$,

where w_{it} is the natural log of weekly wages for individual i at time t, s_{it} is the observed skills (i.e. educational attainment) of individual i at time t, a_i is the unobserved skills (i.e. ability) of individual i, and ε_{it} is a combination of measurement error and other random determinants of wages.[8] The model allows both the returns to education (β_t) and ability (ϕ_t) to vary over time. In order to capture changes in the composition of educational groups, the expected value of ability (i.e. unobserved skills) conditional on education is allowed to vary across cohorts and time.[9]

Equation (2.2) gives the change over time in the difference in expected wages between college graduates *col* and high school graduates *hs*, conditional on age (*t-c*):

(2.2) $$\begin{aligned}&[E(w_{it}|J=col,C=c,T=t) - E(w_{it}|J=hs,C=c,T=t)] \\ &- [E(w_{it}|J=col,C=c-1,T=t-1) - E(w_{it}|J=hs,C=c-1,T=t-1)] \\ &= (\beta_t - \beta_{t-1}) + (\phi_t - \phi_{t-1})(\tilde{A}_{col} - \tilde{A}_{hs}) + \phi[(A_{col,c,t} - A_{hs,c,t}) - (A_{col,c-1,t-1} - A_{hs,c-1,t-1})],\end{aligned}$$

where $A_{j,c,t}$ gives the expected value of ability conditional on being in educational group j and cohort c at time t, $\tilde{A}_{col} = (A_{col,c,t} + A_{col,c-1,t-1})/2$, $\phi = (\phi_t + \phi_{t-1})/2$, and the difference in schooling between college and high school graduates has been normalized to be one. Equation (2.2)

[8] This model is easily generalized to include returns for other observed skills or to include returns to education that differ by race, gender, or experience.

[9] For simplicity, the expected value of observed skills conditional on schooling is not allowed to vary over time or by cohort, i.e. schooling quality is constant over time and across cohorts.

indicates that conventional estimates of changes over time in education differentials incorporate changes over time in both the returns to education $(\beta_t - \beta_{t-1})$ and ability $(\phi_t - \phi_{t-1})(\tilde{A}_{col} - \tilde{A}_{hs})$.

Unlike changes in the first two terms which indicate changes over time in the price for skill, changes in education differentials due to shifts in the last term, the ability composition term $\phi[(A_{col,c,t} - A_{hs,c,t}) - (A_{col,c-1,t-1} - A_{hs,c-1,t-1})]$, provide absolutely no information about how prices may have changed. It simply reflects changes over time in the quantity or quality of unobserved skill for particular educational groups. Hence, it is important to identify assumptions under which this term can be ignored or to use methods which account for it. Two assumptions which restrict this term to be zero are: (1) the return to ability equals zero and (2) the difference in ability between educational groups (e.g. college and high school graduates) is the same for all cohorts and time periods. The first assumption is difficult to reconcile with the findings of large increases in within educational group inequality, which suggest that the return to unobserved skills has increased over time.[10] The dramatic increases in educational attainment over the past few decades along with a positive correlation between education and ability make the second assumption equally difficult to justify.

In order to relax these assumptions, I turn to using educational ranks as a proxy for the ability composition term.[11] The condition under which educational ranks can be used to control for shifts in the composition of educational groups is given by:

A1. $E(a_i|J=j,C=c,T=t) = g(R_{jct})$,

where R_{jct} is the educational rank for group *j* in cohort *c* at time *t*. In other words, conditional on educational group, cohort, and time, the expected value of ability is a function of the educational

[10] See Taber (1996) and Chay and Lee (1997).

rank. Note that g(·) does not vary over time or by cohort, i.e. it is not $g_{c,t}(\cdot)$. Thus, the same function can be used to parameterize ability for all cohorts and time periods.

Using educational ranks as a control for ability relies upon assumptions about how the relationship between educational ranks and ability (or unobserved skills) varies over time and across cohorts. First, educational ranks are a proxy for unobserved skill that does not vary across cohorts or over time, implying that the *unconditional* distribution of unobserved skills is the same for all cohorts and time periods.[12] Second, using educational ranks implies that the relationship between ability (or unobserved skills) and relative rankings in educational attainment, i.e. "cognitive sorting," does not vary much across cohorts or over time.[13] However, differences across cohorts in family income, parents' education, the prevalence of two-parent families, college tuition costs, etc. could affect both the unconditional distribution of unobserved skills across cohorts and the amount of "cognitive sorting."

To examine these issues, I use the General Social Survey (GSS), which since 1974 has regularly administered a 10-item verbal ability test (WORDSUM) to a cross-section sample of adults. I create samples for the two cohorts of native white males who would have had 6-25 years of potential experience in 1969 and 1989.[14] This procedure results in 1,357 observations

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[11] Section 5 uses a second method to control for compositional shifts – within cohort analyses of changes over time in education differentials.

[12] Note that a shift in the average level of skill that affects all educational groups similarly does not affect comparisons between educational groups.

[13] Hernstein and Murray (1994) argue that "at every level of cognitive ability, the links between IQ and the probability of going to college became tighter and more regular" (p. 34). On the other hand, Hauser and Huang (1997) find no support for increased cognitive sorting after examining the evidence used in Hernstein and Murray (1994) along with more convincing evidence from the General Social Survey verbal ability test.

[14] First, the sample is limited to native white males aged 24-59. Second, those with missing values in the race, age, immigration, education, or sampling weight fields are deleted. This gives a sample size of about 400 in each of the 13 survey years between 1974 and 1996. Third, a WORDSUM score is created by adding the scores (correct answer = 1, wrong answer = 0, no answer = 0.2) from each of the ten items

for the 1969 cohort and 2,436 observations for the 1989 cohort. The WORDSUM scores suggest that the average level of ability (or unobserved skill) is about the same for these two cohorts. Conditional on educational group and educational rank, the WORDSUM score is only 1/45th of a standard deviation smaller for the 1989 cohort. The evidence also suggests that "cognitive sorting" has decreased slightly over time. Conditional on educational group and cohort, a one standard deviation increase in the educational rank increases WORDSUM scores by 0.30 standard deviations for the 1969 cohort and 0.22 standard deviations for 1989 cohort.[15] Note that these coefficients indicate that even after conditioning on education, there is a strong positive relationship between educational ranks and this admittedly crude measure of ability, the WORDSUM score. In fact, educational ranks explain about half of the difference in WORDSUM scores between high school graduates and dropouts. Thus, these findings provide suggestive evidence that educational ranks are reasonable proxies for ability (or unobserved skill) and that the relationship between educational ranks and ability (or unobserved skill) has been fairly stable over the past few decades.

3. Data and Empirical Methodology

The data used in this paper come primarily from the 1% samples of the 1960-1990 Decennial Census.[16] First, the sample is limited to *native* white males aged 18-59.[17] Immigrants are

on the WORDSUM test. (Those who fail to answer any of the ten test items are deleted.) Fourth, this WORDSUM score is adjusted for the age at which the test was taken by subtracting the mean value of WORDSUM conditional being in one of the following age groups: 24-29, 30-39, 40-49, or 50-59). Fifth, weights are used so that the experience distribution is the same for each cohort.

[15] These two coefficients as well as their difference are statistically significant at the 5% level.

[16] More specifically, the Census data are from the 1% Integrated Public Use Microdata Series (IPUMS) samples obtained from the University of Minnesota. See Ruggles and Sobek (1997) for details about these data. Since the earnings questions ask about earnings in the prior year, samples are referred to using the preceding year. For example, the data from 1960 Census is referred to as the 1959 sample.

excluded, because educational opportunities differ for immigrants and natives, making it unlikely that the mapping from ability to educational ranks is the same for immigrants and natives.

Educational ranks, conditional on education, vary by cohort and year, so the sample of native white males is divided into cells defined by year and cohort (single birth years).[18] Within each cell, the weighted fractions are calculated for each of the four educational groups – high school dropouts, high school graduates, some college, and college graduates.[19] These four educational groups are used for two reasons. First, these four educational groups are the only ones which can be consistently identified in all of the sample years (except for some high school dropout sub-groups). Second, all of the later econometric specifications use indicator variables to distinguish these four educational groups. By restricting educational ranks to be functions of only these four educational groups, the identification of educational ranks does not come through variation in educational attainment not included in the indicator variables.

Suppose the fractions in each of these educational groups for a given cell are 0.20, 0.40,

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[17] In the Census, the place of birth variable is used to identify immigrants, since no other variable is available in all five sample years.

[18] Cohorts from different sample years are not pooled, because people may acquire more education as they age (see Section 5.1) Also, note that the sample includes both workers and non-workers when educational ranks are computed. Thus, educational ranks are not affected by labor force participation.

[19] Between the 1980 and 1990 Census years and the 1991 and 1992 March CPS years, the education question changed from asking about the number of years of school attended and completed to asking about educational attainment. Educational groups are defined using the taxonomy suggested by Jaeger (1997). Before the change, educational groups are as follows: high school dropouts (completed less than 12 years); high school graduates (completed exactly 12 years), some college (attended at least 13 years but completed less than 16 years), and college graduate (completed 16 or more years). After the change, educational groups are as follows: high school dropouts (less than first grade through 12[th] grade no diploma); high school graduates (high school graduate, diploma or equivalent (GED)), some college (some college or associate degree), and college graduate (bachelor's, master's, professional, or doctorate degree). Note that both the Census and March CPS data do not distinguish those with high school diplomas from those with GEDs, even though the earnings patterns of those with GEDs typically are found to be more like high school dropouts than like high school graduates (Cameron and Heckman, 1993). Using data from the March CPS, Rosenbaum (1998) shows that the change in the education question did not have much of an effect on these four educational groups.

0.30, and 0.10, respectively. Ordering by educational attainment and normalizing the ranks to be between 0 and 1, these fractions result in educational rank ranges from 0-0.20 for high school dropouts, 0.20-0.60 for high school graduates, 0.60-0.90 for some college, and 0.90-1 for college graduates. Specific educational rank values are calculated under assumptions on the distribution of ability and their corresponding functions of $g(\cdot)$ (see assumption A1 in section 2). For the uniform (0,1) distribution, $g(R_{jct}) = R_{jct}$. For the standard normal distribution, $g(R_{jct}) = \Phi^{-1}(R_{jct})$. Once the distribution function and its associated $g(\cdot)$ have been chosen, educational ranks are calculated by integrating $g(\cdot)$ over the range of educational ranks using the probability density function. Note that these average values of $g(\cdot)$, i.e. educational ranks, are identical for those in a given cell with a particular educational attainment.

The average values of $g(\cdot)$ under the uniform (0,1) distribution are referred to as the *uniform rank*, which end up being the midpoint of the educational rank range. The average values for the standard normal distribution are referred to as the *normal rank*. The normal rank is expressed in units with a mean of zero and a standard deviation of one. In the regressions, I parameterize $g(\cdot)$ in a more flexible manner using a five segment continuous spline (one for each of the quintiles of the educational rank distribution). This spline allows for a separate linear relationship between the uniform rank and weekly earnings in each of the five quintiles of the educational rank distribution.

Once the educational ranks are computed, the sample is restricted to persons not attending school and with 1 to 30 years of potential experience.[20] Finally, note that the wage measures used

[20] Potential experience is calculated as age - 17 for high school dropouts, age - 18 for high school graduates, age - 20 for those with some college, and age - 23 for college graduates, where age equals age at the time of the interview minus one. Throughout this paper, potential experience and experience are used interchangeably.

in this paper include both wage and salary plus farm and self-employment income.[21]

4. Main Results

Section 4 presents the main results in this paper. It describes the dramatic shifts in educational attainment during the past few decades, analyzes how compositional shifts among educational groups have affected the wage patterns of different educational groups, and examines how education differentials change once controls for educational ranks are included.

4.1 Shifts in Educational Attainment

Table 1 illustrates the dramatic increases in educational attainment over the past few decades for native white males with 1-30 years of potential experience.[22] The two most eye-catching results in this table are (1) the decline in the fraction of high school dropouts from 44 percent in 1959 to just 14 percent in 1989 and (2) the increase in the fraction of those with some college but no bachelors degree from 15 percent in 1969 to almost 28 percent in 1989.

These increases in educational attainment affect the mapping from a given educational attainment to one's relative ranking in the educational attainment distribution. For example, the uniform rank metric reveals that between 1959 and 1989 the typical high school graduate fell from the 57th to the 29th percentile of his cohort's educational attainment distribution. Using the normal rank metric, this shift in the composition of high school graduates can be interpreted as being nearly eight tenths of a standard deviation from 0.2 standard deviations above the mean in 1959 to 0.6 standard deviations below the mean in 1989. Under the assumption that educational

[21] Earnings fields are topcoded at different amounts in different survey years. Topcoded values are multiplied by 1.50. All earnings values are expressed in 1996 PCED dollars.

[22] For all of the following tables and regressions, weights have been used so that the experience distribution is the same in each year. Thus, comparisons across time are not sensitive to fluctuations in

ranks proxy for ability (or unobserved skill), Table 1 implies that high school graduates in 1989 had considerably less ability than their counterparts in 1959.

Table 1 clearly shows that educational ranks have fallen over time for all educational groups.[23] The effect on education differentials, i.e. differences between educational groups, depends on which educational groups have experienced the most severe declines. One might expect that the large increases in the some college group might have resulted in an increase in the difference in educational ranks between college and high school graduates. This is precisely what Table 1 shows. Between 1969 and 1989, the difference in educational ranks between college and high school graduates increased by ten percentiles (using the uniform rank as a metric) or by over two tenths of a standard deviation (using the normal rank as a metric).[24]

4.2 Educational Groups vs. Educational Quintiles

Table 2 shows how average weekly wages have changed over time both for educational groups and for quintiles in the distribution of educational attainment.[25] The sample includes natives white males with 1-30 years of experience and weekly earnings greater than $100.[26] The patterns by educational group reproduce the widely documented patterns in the education

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the sizes of particular cohorts. Appendix Table 1 presents movements in educational ranks disaggregated by experience.

[23] Data from the General Social Survey exhibit a similar pattern. Comparing cohorts in 1969 and 1989, WORDSUM verbal ability scores declined for all educational groups with the largest declines for high school dropouts.

[24] Note that among younger cohorts the fraction with bachelors degrees actually decreased between 1979 and 1989 while the fraction with some college but no bachelors degree increased. This resulted in an increase in college/high school difference in educational ranks that was larger for younger cohorts than for older cohorts.

[25] Appendix Table 2 presents movements in weekly earnings disaggregated by experience.

[26] Note that the $100 weekly earnings restriction is not innocuous. Between 1969 and 1989 the percentage not meeting this criterion rises from 10.4% to 20.5% for high school dropouts, 4.8% to 7.5% for high school graduates, and 2.1% to 2.9% for college graduates.

differentials literature. Dramatic growth in weekly earnings for all educational groups in the 1960s was followed by declines for most groups in the 1970s, which led to further declines for all groups except college graduates in the 1980s. Consequently, the high school graduate/high school dropout differential rose between 1969 and 1989, while the college/high school differential dropped in the 1970s and then rose sharply in the 1980s.

Interestingly, when the data are sorted by educational quintile, a remarkably different pattern emerges.[27] For example, between 1969 and 1989 high school graduates and dropouts experienced *declines* in their weekly earnings of 2.4 and 8.7 percent, respectively, while the weekly earnings of those in the lowest three quintiles *rose* by an average of 8.5 percent. These findings suggest that changes in the composition of the less educated may explain their widely documented declines in earnings. Changes in composition also appear to affect differences between educational groups. For example, the log difference between college and high school graduates *increased* by 8.2 percentage points between 1969 and 1989, while the difference between the highest and middle quintiles *decreased* by 5.4 percentage points. Consistent with the finding in Table 1 of increasing differences in ability (or unobserved skill) between college and high school graduates, accounting for changes in composition appear to explain much of the recent increase in earnings differentials between college and high school graduates.

4.3 Controlling for Educational Ranks

These educational group/educational quintile comparisons are suggestive but may be misleading for a couple of reasons. First, the difference in years of schooling between educational quintiles may change over time. For example, the difference in years of schooling

[27] Means for educational quintiles are calculated by using weights for the probability than a given individual is in a particular quintile of educational attainment distribution.

between the highest and middle quintiles decreased by a little over a year during this time period. Second, educational attainment is measured using only four educational groups, implying that analyses of educational quintiles may be sensitive to small changes in the distribution of educational attainment.

Consequently, this section uses educational ranks to control for changes in the composition of educational groups. Table 3 estimates the following log earnings equation:

(4.1) $W = S\beta + A\phi + X\gamma + \varepsilon$,

where W is a vector of log weekly earnings, S is a matrix of educational groups interacted with year groups or individual years, A is a matrix of educational rank controls, X is a matrix of single year experience levels and ten-year experience groups interacted with educational groups, and ε is a normally distributed error term. High school graduates are the excluded group, so the coefficients for interactions between college graduate and year give estimates for college/high school differentials and (after multiplying by -1) the coefficients for interactions between high school dropout and year give high school graduate/dropout differentials.

Specifications (1) and (2) employ a simple OLS framework, where weekly earnings under \$100 are set to \$100. These specifications assume that workers with earnings less than \$100 would have earned \$100 per week had they chosen to work.[28] Specification (1) presents estimates without educational rank controls and exhibits a pattern similar to that in the first panel of Table 2. The college/high school differentials increased from 56 percentage points in 1969 to 66 percentage points in 1989, while the high school graduate/dropout differential increased from 25 to 37 percentage points.

[28] Tobit models with a \$100 truncation point give very similar results to specifications (1) and (2).

Specification (2) is identical to (1), except that a five segment continuous spline (one for each of the quintiles of the educational rank distribution) is included. This spline allows for a separate linear relationship between the uniform rank (see Section 3) and weekly earnings in each of the five quintiles of the educational rank distribution.[29] The relationship between the uniform rank and weekly earnings appears to be roughly quadratic with a very large positive coefficient in the lowest quintile, small positive coefficients in the middle three quintiles, and a negative coefficient in the highest quintile. The coefficients imply that moving from the 10^{th} percentile to the median increases weekly earnings by 23 percentage points, while moving from the median to the 90^{th} percentile increases weekly earnings by only 1.5 percentage points.

The education differential estimates change once educational ranks are included with the college/high school differential typically being about one sixth smaller and the high school graduate/dropout being about two thirds smaller. These estimates can be interpreted as the education differential estimate less any ability composition effects and less the average effect of ability over the entire time period.[30] Note, however, that changes over time in education differentials still reflect changes in the returns to both education and ability.

Once educational ranks are included, the increase between 1969 and 1989 in the college/high school differential shrinks from 10.2 to 4.4 percentage points. Between 1959 and 1989, the change is even more dramatic as the increase shrinks from 16.9 to 6.4 percentage points. For the high school graduate/dropout differential, the increase between 1969 and 1989

[29] I also have tried many other specifications of $g(\cdot)$ including a quadratic function of the uniform rank, a five quintile step function, a ten decile step function, and a ten decile spline. These specifications tend to give similar or stronger results than those reported in this paper.

[30] These results suggest a signaling value of education. See Riley (1979) and Kroch and Sjoblom (1994) for theoretical and empirical papers, respectively, on the signaling value of education.

drops from 12.2 to 6 percentage points after controlling for shifts in composition of educational groups. Finally, the difference in log weekly earnings between college graduates and high school dropouts jumps from 73 percentage points in 1959 to 103 percentage points in 1989, an increase of almost 30 percentage points. However, once an adjustment for the shifts in the composition of educational groups has been made, the increase shrinks to a mere 11 percentage points. Changes in the composition of educational groups matter.

Specifications (3) and (4) exclude non-workers and those with earnings less than $100, making these specifications like many in the literature on education differentials. These specifications produce unbiased estimates of education differentials under the assumption that the unobservables for non-workers and low earners do not differ from the unobservables for workers with similar observable characteristics and weekly earnings over $100. After excluding non-workers and a similar group of low earners, Buchinsky (1994) warns that "one should be cautious, however, in interpreting the results because of possible selection bias introduced by this procedure" (407). Excluding non-workers and low earners does change the results somewhat.[31] The increase between 1969 and 1989 in the college/high school differential now shrinks from 7.6 to 6.9 percentage points once educational ranks are included. On the other hand, excluding non-workers and workers with low earnings reduces the change over time in the high school graduate/dropout differential, even without including educational ranks.

Excluding non-workers (and low earners) results in samples with varying degrees of selection over time. In 1969 education differentials estimated using only those earning more than $100 per week include 97.9 percent of college graduates, 95.2 percent of high school graduates,

[31] Chandra (2000) presents a very persuasive case that ignoring non-workers has biased estimates of changes over time in the racial wage gap.

and 89.6 percent of high school dropouts. In 1989 these respective percentages change to 97.1, 92.5, and 79.5 percent. If these non-workers would have earned less than workers with the same observable characteristics, then the rising fraction of excluded non-workers from less educated groups likely has resulted in a downward bias for estimates of changes over time in education differentials. On the other hand, given the log scale's magnification of small changes near zero, specifications (1) and (2) place a lot of weight on non-workers. Also, assuming that non-workers would have earned $100 is a strong and relatively arbitrary assumption. Thus, in specifications (5) and (6), a Heckman 2-step procedure is used where the first stage dependent variable is whether a person earned more than $100 per week. (Those without earnings are coded as zero.) The state unemployment rate is used as an exclusion restriction – a variable that affects employment but does not affect weekly earnings. The Mills ratio is positive and highly significant in both specifications, indicating the non-workers likely would have earned less than workers with similar observable characteristics.

As expected, controlling for selection results in considerably larger education differential estimates in specification (5) versus those in either specifications (1) or (3). The changes over time in education differentials also are larger. In specification (5), the college/high school differential increases by 12.6 percentage points between 1969 and 1989, while the high school graduate/dropout differential increases by 19.2 percentage points. Once educational ranks are included (in both the first and second stages), these increases in the college/high school and high school graduate/dropout differentials shrink to 4.1 and 16.9 percentage points, respectively. So once again, controlling for changes in composition among educational groups appears to be quite important.[32]

Overall, Table 5 suggests that accounting for changes in the composition of educational groups affects the time pattern of the college/high school differential, but has no consistent effect on the high school graduate/dropout differential. This is not surprising. Table 1 shows that while educational ranks have fallen over time for all educational groups, the decreases were much slower for college graduates (mostly due to the large increases in the some college group). The mushrooming of the some college group increases the differences in educational ranks between college and high school graduates. In Table 5, these increasing educational rank differences translate into smaller increases in the college/high school differential once educational ranks are included. Also, the decreases in educational ranks were about the same for high school graduates and dropouts, suggesting that accounting for educational ranks should not have any consistent effect on the difference in earnings between high school graduates and dropouts, which is exactly what Table 3 shows.

5. Additional Evidence

5.1 Within Cohort Comparisons

Another method of controlling for changes in the composition of educational groups is through the use of within cohort comparisons. Within cohorts the composition of educational groups should not change much over time, except when large fractions of the cohort are still attending school. The major disadvantage of within cohort comparisons is that changes in

[32] I have also tried median regression specifications (using one quarter of the sample), implying that the identifying assumption is that those without earnings would have earned less than the median. The estimates suggest a smaller role for changes in the composition of educational groups, though the larger standard errors make the results difficult to interpret.

education differentials over time are difficult to distinguish from differences between educational groups in experience differentials. For example, if the returns to experience are higher for college graduates than for high school graduates, then the college/high school differential, as measured by a within cohort analysis, may increase even if the returns to education and ability do not change. However, comparisons of multiple cohorts do allow one to control for differences by education in experience differentials.

Table 4 presents education differentials by year and cohort along with the normal rank differences between the educational groups. Note that the normal rank differences indicate that even within cohorts, compositional shifts are quite large, even for older cohorts. For example, for the cohort born between 1935 and 1944, the college/high school normal rank difference increases by about one ninth of a standard deviation between ages 25-34 and 45-54 with most of the increase occurring between ages 35 and 54. This increase is over half the size of the increase for the experience-constant samples in Table 1.

These within cohort shifts in the composition of educational groups raise questions about the validity of this approach in controlling for changes in composition. In particular, the evidence of shifts in composition for older cohorts is particularly curious, given that older cohorts are unlikely to acquiring additional schooling in large enough numbers in order to make these shifts possible. Higher mortality rates for the less educated could be part of the explanation, but another potential cause is simple misreporting of true educational attainment. As educational attainment for the entire population increases, the stigma of having a low education likely increases. It is quite possible that people respond to this stigma by reporting educational attainments that they never truly acquired. In fact, this stigma may be the strongest for workers with higher than average earnings for their given education, which would bias within cohort education differential estimates towards the experience-constant without educational rank estimates.

The education differential estimates in Table 4 come from an OLS regression of log weekly earnings onto full interactions of educational group, cohort, and year. The cohort-year-specific education differentials are then adjusted so that college/high school and high school graduate/dropout differentials do not vary with age. Consequently, if educational group-specific experience differentials do not vary over time (and the composition among educational groups does not change over time), then comparisons within a cohort should reflect time effects.

Comparisons of education differential estimates from the same cohort in different years suggest that the college/high school differential increased almost 16 percentage points between 1959 and 1989, while the high school graduate/dropout differential increased by 10.4 percentage points. These increases are a little smaller than those in specification (1) of Table 3, indicating some role for changes in the composition of educational groups. However, the within cohort shifts in the composition of education groups make these results very difficult to interpret.

5.2 Comparing the Census and March CPS

All of the previous results come from the Decennial Census, yet the March Current Population Survey (March CPS) is the most pervasively used source to examine earnings patterns. One may wonder whether these strong effects of the educational ranks from the Census hold in the March CPS. The short answer is that the effects of educational ranks tend to be larger in the March CPS for the high school graduate/dropout differential and smaller for the college/high school differential.

The inability to identify immigrants in all years and the exclusion of those in institutions makes the March CPS less suited to using educational ranks as a control for changes in the composition of educational groups. The mapping from ability (or unobserved skill) to educational ranks is likely to be very different for natives and immigrants, resulting in measurement error in the March CPS due to the necessary inclusion of immigrants in most years. Including immigrants in the Census explains about half of the difference between Census and March CPS estimates for college/high school differentials, while the difference for high school graduate/dropout differentials widen.[33] In addition, individuals in institutions are unlikely to be randomly distributed across educational groups, which biases the estimates of educational ranks in the March CPS. Excluding those in institutions from the Census explains the other half of the difference between the samples in the college/high school differential, while not having much of an effect on the high school graduate/dropout differential.

5.3 Employment and Annual Earnings Education differentials

One might wonder whether the strong results for the effects of changes in the composition of educational groups on weekly earnings education differentials hold for other outcomes, such as employment and annual earnings. In fact, the effects of including educational ranks are considerably stronger for these outcomes. For annual employment, education differentials are slightly *negative* once educational ranks are included suggesting that employment differentials are more closely related to a person's educational rank than their educational attainment.[34] Moreover, accounting for changes in composition of educational groups results in change in the 1969-1989 college/high school annual employment differential switching from a 1.3 percentage point *increase* to a 0.9 percentage point *decrease*.

[33] Appendix Table 3 presents regression results comparing the Census and March CPS.

[34] Appendix Table 4 presents probits for annual employment and OLS regressions for annual earnings.

Similarly, the results for annual earnings are quite strong. For example, the difference in annual earnings between college graduates and high school dropouts increases by 32.6 percentage points between 1959 and 1989. After accounting for changes in the composition of educational groups, this increase falls to a dramatically smaller 11.5 percentage points. Without question, accounting for changes in the composition of educational groups generates a very different picture of the U.S. labor market.

6. The Counterfactual: No Changes in Educational Attainment

One way to evaluate the overall importance of changes in the composition of educational groups is to compare the actual movements in earnings patterns to those in a counterfactual world where the educational attainment distribution did not change over time.[35] Figure 1 allows such comparisons by presenting weekly earnings patterns where the educational attainment distribution is the same in all years and for all cohorts (the distribution is normalized to be the overall 1989 distribution). In other words, all college graduates in all cohorts and years have the same educational rank, but this educational rank is different than the educational rank for high school graduates or dropouts. Note that this procedure eliminates educational rank differences both over time and between experience groups.

Figure 1 illustrates the dramatic differences in weekly earnings patterns and education differentials once changes in the composition of educational groups are accounted for. Between 1969 and 1989 the weekly earnings of high school graduates and dropouts *declined* by 2.4 and 8.7 percent, respectively (also see Table 2). Accounting for changes in the composition of

[35] This analysis ignores general equilibrium effects.

educational groups suggests that between 1969 and 1989 the weekly earnings of high school graduates and dropouts *increased* by 7.4 and 5.6 percent, respectively.

These striking differences in weekly earnings patterns also hold for education differentials. Note that for the college/high school differential in Figure 1, 1989 is no longer an outlier once educational ranks are included, since both 1959 and 1969 are within a few percentage points. Conversely, when changes in the composition of educational groups are ignored, the college/high school differential for 1989 is more than 12 percentage points higher than any other year. Consequently, these pictures suggest that the decline of the college/high school differential in the 1970s is the real mystery; the 1980s may simply indicate a return to the labor market of the 1960s.

Accounting for compositional shifts is important in another dimension as well. Experience differentials typically compare younger and older workers with the same educational attainment. Changing educational attainment patterns suggest that the average level of ability (or unobserved skills) is lower for younger workers, implying that experience differentials are biased downwards. The amount of bias could vary over time. Appendix Table 1 shows that educational rank differences have fallen over time between older and younger cohorts with less education. For example, between 1969 and 1989 the difference in educational ranks between workers with 21-30 years of experience and 1-10 years of experience fell about half a standard deviation for both high school graduates and dropouts. Thus, we would expect that once controls for educational ranks are included, the increases in experience differentials should be larger for these groups. In Appendix Table 5, the 1969-1989 increases in experience differentials double for both high school graduates and dropouts once controls for educational ranks are included. Accounting for changes in the composition of educational groups results in a new picture of the U.S. labor market where increases in education differentials tend to be much less dramatic but are much more heavily concentrated among younger workers.

7. Conclusion

This paper makes two important arguments. First, over long periods of time educational groups are bad proxies for skill groups. The composition of educational groups has changed quite dramatically over time, implying that the distribution of skill for educational groups has also changed. Hence, changes over time in education differentials confound changes in the returns to skill with changes in the composition of the educational groups used to measure these returns to skill. In short, estimates of the returns to skill using education differentials are likely to present a very misleading portrait of the labor market.

The second argument is that this misleading portrait of the labor market has been significant. Accounting for changes in the composition of educational groups results in (1) increases between 1969 and 1989 in the college/high school weekly earnings differential that are about half as large, (2) weekly earnings for high school graduates and dropouts that increase between 1969 and 1989 rather than decrease, and (3) a doubling of the 1969-1989 increases in the differences in weekly earnings between older and younger high school graduates and dropouts. These "stylized facts" about the U.S. labor market play an important role in both policy and research, making the issue of compositional changes that much more important.

This paper provides a wide range of evidence that ignoring changes in the composition of educational groups is problematic, but the evidence on the precise magnitude of the problem is less clear. Adjusting for changes in the composition of educational groups is messy, implying that measuring changes in the returns to skill using educational groups likely is less reliable than other approaches, such as analyses of changes over time in within skill group inequality. Note, however, that changes in the composition of educational groups, i.e. skill groups, also has implications for within skill group inequality. Moreover, recent research by Hoxby and Terry (1999) and Haider (1998) raises questions about the common interpretations of this body of evidence.

Estimates of changes over time in education differentials shape the way in which both policymakers and economists think about the U.S. labor market. For policymakers these "stylized facts" generate support for particular policies. For economists they often serve as the barometers that test the validity of one theory versus another. For these reasons it is crucial that we distinguish changes in the returns to skill, such as the returns to education and ability, from changes in the composition of the educational groups used to measure these returns to skill.

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Table 1
Average Educational Ranks by Education
For Native White Males with 1-30 Years of Experience

Educational Attainment	1959	1969	1979	1989
	Fraction in Educational Group			
High school dropout	0.439	0.314	0.191	0.143
High school graduate	0.312	0.369	0.377	0.321
Some college	0.121	0.149	0.202	0.276
College graduate or more	0.128	0.168	0.231	0.260
	Average Uniform Ranks			
High school dropout	0.222	0.156	0.093	0.069
High school graduate	0.574	0.456	0.349	0.285
Some college	0.801	0.729	0.629	0.578
College graduate or more	0.933	0.910	0.872	0.860
	Uniform Rank Differences			
HS graduate - HS dropout	0.352	0.300	0.256	0.216
College graduate - HS graduate	0.359	0.454	0.523	0.575
	Average Normal Ranks			
High school dropout	-0.896	-1.149	-1.452	-1.602
High school graduate	0.194	-0.117	-0.409	-0.592
Some college	0.861	0.622	0.339	0.204
College graduate or more	1.622	1.464	1.264	1.206
	Normal Rank Differences			
HS graduate - HS dropout	1.090	1.032	1.043	1.010
College graduate - HS graduate	1.428	1.581	1.673	1.798

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school.

Notes: The standard deviation for uniform ranks is approximately 0.3. Normal ranks are distributed standard normal. Weights are used so that the experience distribution is the same in each year. See text for details.

Table 2
Weekly Earnings by Educational Group and Quintile For Native White Males with 1-30 Years of Experience

Educational Groups	**1959**	**1969**	**1979**	**1989**
High school dropout	467	597	635	545
High school graduate	530	668	724	652
Some college	651	797	825	804
College graduate or more	902	1,206	1,186	1,278
Educational Group Log Differences				
HS graduate - HS dropout	0.127	0.112	0.130	0.178
College graduate - HS graduate	0.532	0.591	0.494	0.673
Educational Quintiles				
Lowest quintile (0-20%)	456	577	638	601
Mid-low quintile (20-40%)	462	611	719	665
Middle quintile (40-60%)	507	682	759	764
Mid-high quintile (60-80%)	579	789	920	995
Highest quintile (80-100%)	832	1,178	1,200	1,249
Educational Quintile Log Differences				
Middle - Lowest	0.106	0.167	0.175	0.240
Highest - Middle	0.495	0.546	0.458	0.492

Source: The data are from the 1960-1990 1% samples of the Decennial Census (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in schooll. It excludes workers without earnings or weekly earnings under $100.

Notes: *Educational quintiles* are calculated by ranking all persons in a given cohort and year by their educational attainment. Earnings are expressed in 1996 PCED dollars. Weights are used so that the experience distribution is the same in each year. See text for details.

Table 3
Weekly Earnings Education Differentials
For Native White Males with 1-30 Years of Experience
Coefficient (Standard Error)

	Dependent Variable = Log Weekly Earnings					
	OLS				Heckman 2-Step	
Explanatory Variable	(1)	(2)	(3)	(4)	(5)	(6)
College/HS*1959	0.4915	0.4733	0.4731	0.5786	0.5541	0.4354
	(0.0047)	(0.0132)	(0.0038)	(0.0108)	(0.0071)	(0.0198)
College/HS*1969	0.5582	0.4937	0.5142	0.6030	0.6277	0.4403
	(0.0043)	(0.0126)	(0.0034)	(0.0103)	(0.0070)	(0.0190)
College/HS*1979	0.4736	0.3702	0.4210	0.4992	0.5401	0.2965
	(0.0037)	(0.0121)	(0.0030)	(0.0099)	(0.0062)	(0.0186)
College/HS*1989	0.6604	0.5372	0.5906	0.6725	0.7538	0.4812
	(0.0036)	(0.0123)	(0.0029)	(0.0101)	(0.0071)	(0.0190)
HS/Dropout*1959	0.2429	0.0830	0.1617	0.1815	0.3754	0.1392
	(0.0037)	(0.0084)	(0.0031)	(0.0072)	(0.0087)	(0.0130)
HS/Dropout*1969	0.2484	0.0681	0.1586	0.1703	0.3620	0.1137
	(0.0038)	(0.0095)	(0.0032)	(0.0081)	(0.0084)	(0.0147)
HS/Dropout*1979	0.3263	0.1016	0.1820	0.1691	0.4654	0.1913
	(0.0039)	(0.0117)	(0.0033)	(0.0102)	(0.0105)	(0.0184)
HS/Dropout*1989	0.3705	0.1280	0.1946	0.1628	0.5543	0.2831
	(0.0039)	(0.0128)	(0.0034)	(0.0112)	(0.0128)	(0.0208)
Mills Ratio	.	.	.	.	1.5524	1.7383
	.	.	.	.	(0.0496)	(0.0551)
Exclude Earnings < $100	No	No	Yes	Yes	Yes	Yes
Educational Rank Controls	No	Yes	No	Yes	No	Yes
Number of Observations	1,290,763		1,201,367		1,201,367	

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school. Specifications (3)-(6) exclude workers with zero earnings or weekly earnings less than $100.

Controls: Each specification includes the following controls: indicators for year, year interacted with some college, individual years of experience, and ten-year experience groups interacted with education groups. Specifications (2), (4), and (6) include a five-segment spline (one for each quintile of the uniform rank distribution). The state unemployment rate is used in the first stage of specifications (5) and (6).

Notes: In specifications (1) and (2) weekly earnings under $100 are set to $100. In specifications (5) and (6) the first stage dependent variable is an indicator for whether a person has weekly earnings greater than $100. Earnings are expressed in 1996 PCED dollars. Weights are used so that the experience distribution is the same in each year. See text for details.

Table 4
Weekly Earnings Schooling Differentials
Cohort Comparisons for Native White Males Aged 25-54
Coefficient (Standard Error) [Normal Rank Difference]

	1959	1969	1979	1989
	College/High School			
Birth Year, 1915-24	*0.4295*	0.5805	.	.
	(0.0078)	(0.0078)	.	.
	[1.356]	[1.380]	.	.
Birth Year, 1925-34	0.4348	*0.4969*	0.3851	.
	(0.0074)	*(0.0069)*	(0.0067)	.
	[1.385]	*[1.378]*	[1.437]	.
Birth Year, 1935-44	.	0.4799	*0.4545*	0.5230
	.	(0.0066)	*(0.0060)*	(0.0062)
	.	[1.556]	*[1.585]*	[1.663]
Birth Year, 1945-54	.	.	0.4132	*0.5876*
	.	.	(0.0050)	*(0.0051)*
	.	.	[1.767]	*[1.838]*
	High School/Dropout			
Birth Year, 1915-24	*0.2964*	0.3983	.	.
	(0.0053)	(0.0055)	.	.
	[1.155]	[1.142]	.	.
Birth Year, 1925-34	0.2674	*0.2759*	0.2701	.
	(0.0056)	*(0.0056)*	(0.0058)	.
	[1.091]	*[1.087]*	[1.071]	.
Birth Year, 1935-44	.	0.2976	*0.3530*	0.3455
	.	(0.0058)	*(0.0062)*	(0.0068)
	.	[1.087]	*[1.069]*	[1.005]
Birth Year, 1945-54	.	.	0.3849	*0.4408*
	.	.	(0.0062)	*(0.0068)*
	.	.	[1.037]	*[0.974]*

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males aged 25-54 who are not in school. It includes those with no weekly earnings.

Controls: Controls for all interactions of year, birth year, and educational group are included.

Notes: The *normal rank difference* [in brackets] gives the difference in normal rranks between the two educational groups. Coefficients are adjusted so that the average differential does not vary by age. Coefficients in italics are for 35-44 year-olds. Coefficients above or to the right of the italicized coefficients are for 45-54 year-olds, and those below or to the left are for 25-34 year-olds. Weekly earnings are expressed in 1996 PCED dollars, and those under $100 are set to $100. Weights are used so that the experience distribution is the same in each year. See text for details.

Appendix Table 1
Average Normal Ranks by Education and Experience
For Native White Males with 1-30 Years of Experience

Educational Attainment	1959	1969	1979	1989
	1-10 Years of Experience			
High school dropout	-1.048	-1.370	-1.524	-1.534
High school graduate	0.042	-0.317	-0.468	-0.516
Some college	0.754	0.510	0.296	0.325
College graduate or more	1.530	1.404	1.186	1.286
HS graduate - HS dropout	1.090	1.053	1.056	1.018
College graduate - HS graduate	1.488	1.721	1.654	1.802
	11-20 Years of Experience			
High school dropout	-0.914	-1.175	-1.566	-1.663
High school graduate	0.205	-0.067	-0.511	-0.646
Some college	0.840	0.633	0.270	0.123
College graduate or more	1.643	1.437	1.258	1.123
HS graduate - HS dropout	1.119	1.108	1.055	1.017
College graduate - HS graduate	1.438	1.504	1.769	1.769
	21-30 Years of Experience			
High school dropout	-0.757	-0.966	-1.301	-1.626
High school graduate	0.436	0.145	-0.210	-0.633
Some college	1.072	0.789	0.517	0.164
College graduate or more	1.804	1.604	1.382	1.218
HS graduate - HS dropout	1.193	1.111	1.091	0.993
College graduate - HS graduate	1.368	1.459	1.592	1.851

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school.

Notes: Normal ranks are distributed standard normal. Weights are used so that the experience distribution is the same in each year. See text for details.

Appendix Table 2
Weekly Earnings by Education and Experience
For Native White Males with 1-30 Years of Experience

Educational Attainment	1959	1969	1979	1989
	1-10 Years of Experience			
High school dropout	372	480	473	409
High school graduate	424	545	566	487
Some college	507	631	651	597
College graduate or more	710	920	862	935
Log(HS graduate/HS dropout)	0.133	0.127	0.180	0.174
Log(College graduate/HS graduate)	0.514	0.523	0.421	0.653
	11-20 Years of Experience			
High school dropout	490	618	660	579
High school graduate	580	718	777	698
Some college	711	867	895	828
College graduate or more	1,015	1,313	1,325	1,354
Log(HS graduate/HS dropout)	0.168	0.149	0.164	0.187
Log(College graduate/HS graduate)	0.561	0.604	0.533	0.663
	21-30 Years of Experience			
High school dropout	515	658	740	680
High school graduate	635	795	862	826
Some college	797	974	1,006	1,017
College graduate or more	1,159	1,477	1,470	1,613
Log(HS graduate/HS dropout)	0.210	0.189	0.153	0.195
Log(College graduate/HS graduate)	0.601	0.620	0.533	0.669

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school. It excludes workers without earnings or weekly earnings under $100.

Notes: Earnings are expressed in 1996 PCED dollars. Weights are used so that the experience distribution is the same in each year. See text for details.

Appendix Table 3

OLS Estimates of Weekly Earnings Schooling Differentials, Census and March CPS Comparisons

Coefficient (Standard Error)

	Dependent Variable = Natural Log of Weekly Earnings									
	Decennial Census								March CPS	
Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
College/HS*1969	0.5547	0.4478	0.4579	0.5502	0.4752	0.5288	0.5040	0.5026	0.4494	0.6063
	(0.0045)	(0.0157)	(0.0156)	(0.0044)	(0.0153)	(0.0044)	(0.0152)	(0.0146)	(0.0083)	(0.0284)
College/HS*1979	0.4706	0.3370	0.3497	0.4706	0.3713	0.4569	0.4174	0.4106	0.4364	0.5650
	(0.0039)	(0.0140)	(0.0139)	(0.0038)	(0.0137)	(0.0038)	(0.0136)	(0.0129)	(0.0077)	(0.0258)
College/HS*1989	0.6569	0.5086	0.5264	0.6566	0.5489	0.6392	0.5969	0.5893	0.6049	0.7303
	(0.0038)	(0.0140)	(0.0138)	(0.0037)	(0.0136)	(0.0037)	(0.0136)	(0.0128)	(0.0072)	(0.0255)
HS/Dropout*1969	0.2448	0.0861	0.0951	0.2429	0.1106	0.2342	0.1604	0.1635	0.2742	0.1963
	(0.0041)	(0.0100)	(0.0101)	(0.0040)	(0.0096)	(0.0041)	(0.0096)	(0.0094)	(0.0092)	(0.0196)
HS/Dropout*1979	0.3230	0.1285	0.1440	0.3260	0.1651	0.3117	0.2125	0.2254	0.3417	0.2108
	(0.0042)	(0.0124)	(0.0124)	(0.0041)	(0.0117)	(0.0041)	(0.0117)	(0.0109)	(0.0085)	(0.0223)
HS/Dropout*1989	0.3672	0.1551	0.1775	0.3715	0.1993	0.3620	0.2501	0.2666	0.4182	0.2629
	(0.0043)	(0.0137)	(0.0132)	(0.0039)	(0.0123)	(0.0040)	(0.0124)	(0.0116)	(0.0081)	(0.0242)
Percentage Explained	0.1022	40.5%	33.0%	0.1064	30.7%	0.1103	15.7%	21.4%	0.1555	20.2%
	0.1223	43.5%	32.6%	0.1287	31.0%	0.1277	29.8%	19.3%	0.1439	53.7%
Include Immigrants	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Include Group Quarters	Yes	Yes	Yes	Yes	Yes	No	No	No	No	No
Ranks with Immigrants	N/A	No	Yes	N/A	Yes	N/A	Yes	Yes	N/A	Yes
Ranks using March CPS	N/A	No	No	N/A	No	N/A	No	Yes	N/A	Yes
Number of Observations	1,019,666			1,088,055		1,063,563			230,727	

Sources: The data are from the 1970-1990 1% samples of the Decennial Census (*IPUMS Version 2.0*) and the 1969-71, 1979-81, and 1989-91 survey years of the March Current Population Survey.

Sample Restrictions: All samples include white males with 1-30 years of experience who are not in school. Specifications (1)-(3) exclude immigrants. Specifications (6)-(10) exclude those living in group quarters.

Controls: Each specification includes the following controls: indicators for year, year interacted with some college, individual years of experience, and ten-year experience groups interacted with education groups.

Notes: Specification (2) use educational ranks computed with samples that exclude immigrants, whereas the educational ranks in (3), (5), (7), (8), (10) are computed with samples that include immigrants. Specifications (2), (3), (5), and (6) use data from the Census to compute educational ranks, while specifications (8) and (10) use the March CPS. Weekly earnings are expressed in 1996 PCED dollars, and those under $100 (including zeros) are set to $100. Weights are used so that the experience distribution is the same in each year. See text for details.

Appendix Table 4
Employment and Annual Earnings Schooling Differentials For Native White Males with 1-30 Years of Experience
Average Derivative (Standard Error)

	Probit		OLS	
	Annual Employment		Log Annual Earnings	
Explanatory Variable	**(1)**	**(2)**	**(3)**	**(4)**
College/HS*1959	0.0196	-0.0161	0.5230	0.4645
	(0.0019)	(0.0050)	(0.0048)	(0.0137)
College/HS*1969	0.0427	-0.0076	0.5826	0.4721
	(0.0019)	(0.0046)	(0.0044)	(0.0130)
College/HS*1979	0.0397	-0.0229	0.5198	0.3665
	(0.0015)	(0.0042)	(0.0038)	(0.0125)
College/HS*1989	0.0558	-0.0166	0.7170	0.5402
	(0.0014)	(0.0042)	(0.0037)	(0.0127)
HS/Dropout*1959	0.0447	-0.0185	0.3128	0.1125
	(0.0012)	(0.0025)	(0.0038)	(0.0087)
HS/Dropout*1969	0.0471	-0.0181	0.3079	0.0859
	(0.0011)	(0.0026)	(0.0039)	(0.0098)
HS/Dropout*1979	0.0581	-0.0112	0.4043	0.1316
	(0.0011)	(0.0030)	(0.0040)	(0.0121)
HS/Dropout*1989	0.0634	-0.0025	0.4447	0.1523
	(0.0010)	(0.0031)	(0.0041)	(0.0132)
Schooling Rank Controls	No	Yes	No	Yes
Number of Observations	1,290,763		1,290,763	

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school. It includes those with no annual earnings.

Controls: Each specification includes the following controls: indicators for year, year interacted with some college, individual years of experience, and ten-year experience groups interacted with education groups. Specifications (1) and (2) include the state unemployment rate. Specifications (2) and (4) include a five-segment spline (one for each quintile of the uniform rank distribution).

Notes: In specifications (1) and (2) the average derivative gives the average change in the probability of working when the explanatory variable increases by one. In specifications (3) and (4) the *average derivative* is simply the OLS coefficient estimate. In specifications (1) and (2) weekly earnings under $100 are set to $100. Earnings are expressed in 1996 PCED dollars. Weights are used so that the experience distribution is the same in each year. See text for details.

Appendix Table 5
Weekly Earnings Experience Differentials
For Native White Males
21-30 Years of Experience Minus 1-10 Years of Experience

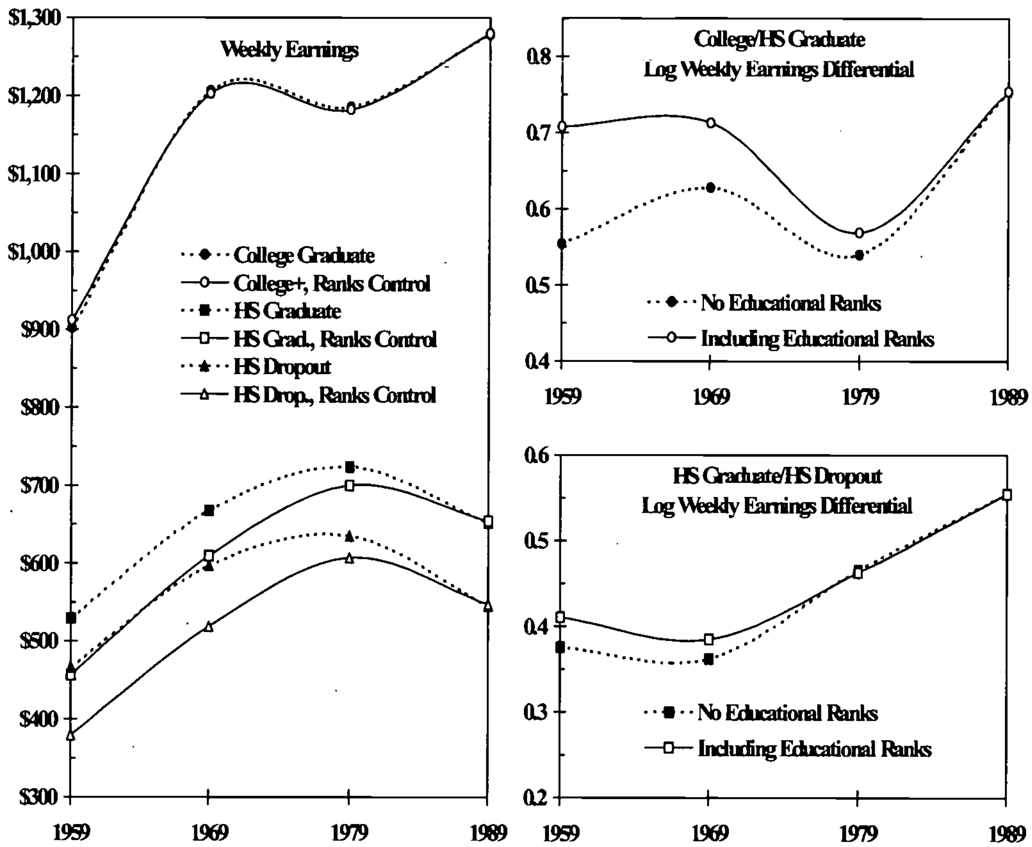
Educational Attainment	1959	1969	1979	1989
	No Educational Ranks			
High school dropout	0.334	0.336	0.440	0.484
High school graduate	0.400	0.392	0.437	0.512
Some college	0.385	0.408	0.417	0.481
College graduate or more	0.386	0.398	0.488	0.475
	With Educational Ranks			
High school dropout	0.266	0.218	0.364	0.510
High school graduate	0.335	0.305	0.387	0.537
Some college	0.358	0.345	0.379	0.505
College graduate or more	0.409	0.416	0.477	0.480

Source: The data are from the 1960-1990 Decennial Census 1% samples (*IPUMS Version 2.0*).

Sample Restrictions: The sample includes native white males with 1-30 years of experience who are not in school.

Notes: Experience differentials are derived from parameters estimated in specifications (5) and (6) of Table 3. Earnings are expressed in 1996 PCED dollars. Weights are used so that the experience distribution is the same in each year. See text for details.

Figure 1
Log Weekly Earnings and Log Weekly Earnings Differentials by Educational Group
With and Without Schooling Rank Controls



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