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Do Returns to Schooling Differ by Race and Ethnicity?

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Abstract

Using data from the U.S. Decennial Census and the *National Longitudinal Surveys*, we find little evidence of differences in the return to schooling across racial and ethnic groups, even with attempts to control for ability and measurement error biases. While our point estimates are relatively similar across racial and ethnic groups, our conclusion is driven in part by relatively large standard errors. That said, we find no evidence that returns to schooling are lower for African Americans or Hispanics than for non-minorities. As a result, policies that increase education among the low-skilled have a good possibility of increasing economic well-being and reducing inequality. More generally, our analysis suggests further research is needed to better understand the nature of measurement error and ability bias across subgroups in order to fully understand potential heterogeneity in the return to schooling across the population.

I. Introduction

Inequality in the U.S. has been increasing over the past 25 years. In 1979 workers in the bottom 10th percentile of the wage distribution earned $6.88 per hour (in 2003 dollars) while those at the 50th and 90th percentiles earned $14.58 and $30.19 per hour, respectively. By 2000 real wages at the 10th percentile had grown by roughly 5.4 percent to $7.25. At the same time real wages at the median and 90th percentile grew to $15.71 and $35.65 per hour, a growth of 7.8 percent at the median and 18.1 percent at the 90th percentile. Because differences in human capital (education and experience) account for approximately one-third of the variation in wages and because the mean economic return to schooling is estimated at 10 percent, many researchers and policy makers have appealed to education and training policies to bolster the wages of the lowest skilled workers (Heckman and Carneiro 2004, Krueger 2004).

Much less is known about how the estimated returns to schooling vary across the population. For example, Carneiro, Heckman and Vytlacil (2003) and Taber (2001) find that the return to education is higher for more able individuals. In contrast, neither Altonji and Dunn (1996) nor Ashenfelter and Rouse (2000) find consistent evidence that the returns to schooling are higher for individuals that come from more advantaged families. Further, Ashenfelter and Rouse (1998) provide some evidence that the return may be *higher* for more disadvantaged individuals, and this pattern would explain some of the empirical estimates of the economic benefit of schooling using instrumental variables techniques (Card 2001). And yet, much social policy hinges on what we believe to be the value of education for individuals. For example, policies aimed at improving the incomes of the lowest skilled members of society will not either improve their economic well-being or decrease inequality if their returns to schooling are low.

In this paper we provide further evidence on the variation in returns to schooling by examining whether the benefits vary by race and ethnicity of the individual. We do so by estimating the return to schooling using the U.S. Decennial Census as well as the *National Longitudinal Surveys of Young Men and Young Women*, and the *National Longitudinal Survey of Youth, 1979*. We find that the return to schooling is relatively constant across racial and ethnic groups, even controlling for ability and measurement error biases. The rest of the paper is organized as follows. In the next section we lay out the empirical framework, in Section III we describe the data, in Section IV we present the results, and we conclude in Section V.

II. Empirical Framework

A. Basic Specification

Following Mincer (1974) we estimate the relationship between schooling and income by regressing the (natural) logarithm of the hourly wage ($\ln w_{ij}$) of individual *i* of race or ethnicity *j* on years of completed schooling (S_{ij}) controlling for explanatory variables such as potential experience or age[1], sex, race, and geographic region of the country (X_{ij}). As such we estimate:

$$\ln w_{ij} = \alpha_j + \beta_j S_{ij} + X_{ij}\gamma_j + \varepsilon_{ij} \quad \textbf{(1)}$$

(where ε_{ij} is an error term). The coefficient on the schooling variable (β_j) is interpreted as the percentage increase in the hourly wage associated with one additional year of schooling and is referred to as the "return to schooling." Note that while we will refer to it as the "return to schooling," it is actually just the average percentage difference in mean earnings for each additional

[1] In the analyses using the *National Longitudinal Surveys* we control for age rather than potential work experience (age - education - 6) because of possible measurement error in education.

year of schooling. As Mincer (1974) shows, if foregone earnings are the only cost of school attendance this is the private rate of return to the investment in a year of schooling. A more detailed calculation of the "return" would incorporate the other costs of schooling, including tuition, as well.

Related to many of the econometric issues raised below is the question of why we may or may not expect to find differences in the estimated return to schooling by race or ethnicity; that is, why β_j may vary by *j*. If we begin by assuming that equation (1) represents the true relationship between wages and schooling and that differences in educational attainment occur randomly, then a constant β_j ($\beta_j = \beta$) implies that we should estimate the same return to schooling for any subgroup of the population. However, even if β_j is constant we may observe different estimates of the return to schooling for different subgroups if years of schooling is a poor proxy for human capital due to differences in school quality and if average school quality varies systematically by race.

Alternatively, β_j may not be constant. For example, the return to schooling may depend on the level of education. In this case, estimating returns to schooling for subgroups of the population with different levels of education (on average) will generate different estimates of the return to schooling. Further, it is important to keep in mind that differences in educational attainment do not occur randomly in the population but instead arise from individuals' decisions. A simple model of optimal schooling investment as in Becker (1967) and Card (2001) predicts that differences in optimal schooling choice arise from differences in the benefits and/or costs of obtaining additional schooling. As a result, individual differences in costs or expected benefits that vary systematically by subgroup may generate differing returns-to-schooling estimates for different subgroups. For example, assuming that education does not affect mortality rates, differences in mortality rates by race mean that for any given level of education, African Americans have fewer expected years than whites over which to receive the benefits of an additional year of schooling. If the costs of an additional year of education are the same for all individuals, then an African American who decides to invest in an additional year of education must expect to receive a larger increase in annual income than an otherwise similar white individual making the same decision. For these reasons, it is an empirical question whether the return to schooling is constant across the population.

B. Econometric Issues

More generally, much of the literature estimating returns to schooling is concerned with whether the ordinary least squares (OLS) estimates of β_j in equation (1) reflect the causal effect of education on wages. Specifically, does more schooling cause higher earnings or are more able people more likely both to get more schooling and to earn higher wages, even in the absence of additional schooling? In the latter case, the OLS estimates of β_j will likely be upward biased due to selection on ability.

Further, this ability bias may not be constant across the population. For example, if African Americans and Hispanics tend to attend schools of poorer quality, then those students who manage to get more schooling (particularly, perhaps, by going to college) may have unusually high ability. This would generate greater selection bias among African Americans and Hispanics suggesting that the cross-sectional estimate of the return to schooling is more severely upward biased for these populations.

In this paper we address the potential for selection bias by including controls for ability directly using test scores and using family relationships by studying siblings.[2] When considering family relationships, we characterize the wage equation as:

$$\ln w_{ijk} = \alpha'_j + \beta'_j S_{ijk} + X_{ijk}\gamma'_j + \mu_{jk} + \varepsilon'_{ijk} \qquad \textbf{(2)}$$

where ln w_{ijk} represents the log wage of individual *i* from racial/ethnic group *j* and family *k*, μ_{jk} represents the "family" ability, and the other variables are defined as before. Family ability may represent a number of things such as genetic endowment with respect to earnings capability or access to resources that facilitate both educational attainment and labor market earnings. If log wages are linear and separable in this ability, then one can address selection bias by relating the difference in wages between family members (or siblings) to the difference in their education. If there are no further unobserved differences between family members that are correlated with both schooling differences and differences in earnings, then this "within-sibling" estimator will provide an unbiased estimate of the return to schooling. In our analysis we can also directly control for an observable measure of "ability" by also including individual test scores in equation (2).

There have been many previous estimates of the return to schooling using this estimator to study the mean return to schooling (see, e.g., Gorseline (1932), Chamberlain and Griliches (1975, 1977)). More recently, Altonji and Dunn (1996) and Ashenfelter and Rouse (1998) use this estimator to study how the returns to schooling differ by family background. We are unaware of previous applications of this estimator to studying the return to schooling by race and ethnicity.

[2] Other researchers, such as Angrist and Krueger (1991), Kane and Rouse (1993), and Card (1995), address selection bias using instrumental variables estimators. In this strategy one must identify an "exogenous" event (the instrumental variable) that affects an individual's years of completed schooling but that is uncorrelated with the error term in the wage equation. Unfortunately, it is very difficult to identify valid instrumental variables and one usually requires large samples in order to get precise estimates.

Measurement error in reported schooling poses another econometric challenge. The attenuation caused by (classical) measurement error is exacerbated in within-sibling estimators, as identified by Griliches (1977) because sibling education levels are so highly correlated. As such the within-sibling estimator will generate a downward biased estimate of the return to schooling. If the measurement error is classical in nature (i.e., uncorrelated with the error term in the wage equation and with the true level of schooling), then an instrumental variables (IV) estimator using an independent report of the respondent's schooling as the instrumental variable will generate consistent estimates of β_j.[3]

Further, it is not clear why the measurement error need be constant across the population. The reliability (or "signal-to-noise") ratio is the proportion of the observed variance in schooling due to the variance in "true" schooling. If one has two independent estimates of an individual's schooling level, the correlation between the two measures provides an estimate of the reliability ratio.[4] Using NLSY79 data we estimate reliability ratios for self-reported schooling, both in levels (i.e., for each individual) and for the deviation from sibling-means. These estimates are reported in Table 1. Overall we estimate that 11 percent of the observed variance in schooling levels is due to measurement error. In addition, there is some variation by race/ethnicity. Nearly 20 percent of the observed variance in schooling for African Americans is due to error compared to 14 percent for Hispanics and 8 percent for whites. In contrast, there is not a lot of difference in the estimated

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[3] In this paper we assume classical measurement error in schooling. Kane, Rouse, and Staiger (1999) provide evidence that measurement error in schooling may not be classical. Unfortunately the sample sizes provided in our data are too small to implement their suggested estimator by race/ethnicity.

[4] See Ashenfelter and Krueger (1994) for an excellent discussion of measurement error models.

reliability ratios by sex. The results in column 2 of Table 1 indicate that sibling differences in educational attainment include more "noise" than individual measures of educational attainment. Overall, 26 percent of the variance in sibling-differences in education is due to measurement error, although the proportion due to error is one-third for African Americans. Based on these estimates, we expect the estimated returns to schooling for African Americans and to some extent Hispanics to be more downward biased than that for whites (non-African Americans/non-Hispanics).

III. Data

A. U.S. Decennial Census

We begin by using data from the 5 percent samples of the 1980, 1990, and 2000 Decennial Censuses. The samples included individuals aged 25-65 who were U.S. citizens and born in the U.S., who worked at least 1 week in the previous year, and who earned at least one-half of the minimum wage.[5] All wages and income are adjusted to 2003 dollars using the Personal Consumption Expenditures chain-type price index from the Bureau of Economic Analysis. The regression analysis is based on annual earnings.[6]

Because the schooling variable changed in 1990, we calculate the number of completed years

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[5] We constructed an hourly wage rate by adjusting annual wage and salary income by the number of weeks worked in the previous year and the usual number of hours worked each week. We used the minimum wage in effect in the year before the Census in question because the Census income and wages refer to the previous year.

[6] We get greater variation when we estimate the returns to schooling using hourly wages rather than annual earnings. This is because the relationship between greater schooling and more stable jobs is stronger for African Americans and Hispanics than for non-minorities. Whether this correlation is explained by access to more stable jobs or changes in labor supply decisions is an empirical question (Ashenfelter and Ham, 1979).

of schooling for 1990 and 2000 according to the recoding suggested by Park (1994). In addition, in 1980 and 1990 we identify 5 racial groups–White, Black, Native American, Asian, and Other–as well as people who identified themselves as Hispanic, regardless of their race. (Thus, the 6 racial and ethnic groups are *not* mutually exclusive.) While in the 1980 and 1990 Censuses individuals had to choose one race category, in the 2000 Census, individuals could choose multiple races. To make the 2000 data as consistent as possible with the previous data, we grouped those who self-identified as belonging to multiple racial groups into the "other" category.[7] Finally, all estimates using the U.S. Census are weighted by the individual weight assigned by the Census.

B. National Longitudinal Surveys: Young Men and Young Women Cohorts

Young Men and Young Women are two of the original cohorts of the *National Longitudinal Surveys* (NLS). Each cohort was chosen to be representative of Americans aged 14 to 24 in the initial survey year, 1966 for Young Men and 1968 for Young Women. Both include an over-sample of African Americans. We combine the Young Men and Young Women cohorts from the 1978 surveys to create a single data set of 7440 individuals. We restrict our estimation sample to those with hourly pay greater than one-half of the minimum wage in 2003 dollars and those who are not self-employed, not enrolled in school, and not in the military. Once we exclude those in 1978 with no hourly pay data and those with no information on highest grade completed, we are left with an estimation sample of 4802. The weighted means and standard deviations of this full sample are

[7] To judge the sensitivity of our results to how we categorized the 1.37 percent of individuals who selected multiple races, we tried alternative codings. Specifically, we tried running our regressions for whites counting anyone who selected white only and any other combination including white as "white" (and similarly for blacks and Asians). These alternative categorizations did not substantively change our results.

provided in Appendix Table 1a.

The NLS also provides information identifying respondent siblings who are also respondents in the Young Men or Young Women cohorts. Based on this information, we were able to identify 567 families with multiple respondents (1263 respondents) in our estimation sample. If we further restrict the sibling sample to have IQ scores we are left with 298 families (642 respondents). As shown in Appendix Table 1b, on average the sibling sample is somewhat younger than the full estimation sample. Otherwise, the mean characteristics are quite similar.

The last three columns in Appendix Tables 1a and 1b list descriptive statistics for the full and sibling samples conditional on the respondents having a non-missing IQ score. The underlying test scores used to construct the IQ score were only collected for individuals who had completed nine years of schooling as of the initial survey year resulting in a non-random sample of respondents for whom we have an IQ score. In particular, respondents with non-missing IQ scores are less likely to be African American or live in the South and have higher average wages and more years of schooling. Within the sibling sample, respondents with IQ scores are also more likely to be male.

C. National Longitudinal Survey of Youth 1979 (NLSY79)

The *National Longitudinal Survey of Youth 1979* (NLSY79) is a survey of youth aged 14 to 21 as of December 31, 1978 including a nationally representative sample of the civilian noninstitutionalized youths, an over-sample of civilian Hispanic, black, and economically disadvantaged non-black/non-Hispanic youth, and a small military sample of youths aged 17 to 21 years.[8] We use the 1993 survey of the NLSY79 in the analysis below and limit our sample to those with hourly pay greater than one-half of the minimum wage in 1993 and less than $300 per hour, as well as those who are not self-employed, enrolled in school or currently in the military.

An advantage of the NLSY79 is that in 1980 most survey participants were administered the ASVAB (Armed Services Vocational Aptitude Battery), a basic skills test, from which it is possible to construct an Armed Forces Qualification Test (AFQT) score. While researchers disagree about whether AFQT scores mostly reflect "innate intelligence" or also reflect skills acquired in school, most would agree that they reflect *some* information about the skills that individuals possess at the time of the test.[9] We use the AFQT as a measure of "observed" ability.[10]

As in the case of the NLS Young Men and Young Women surveys, many of the NLSY79 respondents have siblings who are also included in the survey. In 1979, 5914 of the civilian respondents lived in a household with at least one other sibling (*NLS Handbook,* Table 3.2, p. 35). And, in 1993 respondents were asked about their educational attainment as well as the educational attainment of up to 13 of their siblings – whether or not these other siblings were respondents in the

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[8] Much of the military sample is dropped after 1984, and the supplemental sample of economically disadvantaged youths is dropped after 1990.

[9] For example, Herrnstein and Murray (1994) argue that the AFQT has many of the properties of an IQ test – the scores do not just reflect specific knowledge that has been learned in school, rather they reflect more general factors of "intelligence." In contrast, Neal and Johnson (1996) argue that AFQT scores increase with years of schooling and therefore are not a good measure of IQ. Others, such as Rodgers and Spriggs (1996), argue that the AFQT is a racially biased test.

[10] In the estimates presented here, we simply control for AFQT and do not address the fact that individuals took the AFQT at different ages and had therefore completed differing years of schooling. We have also estimated our models controlling for the individual's education as of 1979 with qualitatively similar results.

original NLSY79.[11] As a result, we can obtain own-reported and sibling-reported measures of a respondent's education level for those with siblings in the original NLSY79 sample who also participated in the 1993 wave.

Once we additionally exclude those with no wage information and no information on education, our "full sample" includes 6119 men and women between the ages of 28 and 36. Our sibling sample contains information on 2419 individuals from 1062 households (for an average of 2.3 observations per household). Means and standard deviations for these samples are provided in Appendix Table 2.

IV. Results

A. Results Using the Decennial Census

Using data from the Decennial Censuses allows us to get very precise estimates of the relationship between education and wages by race and ethnicity. In addition, we look at how the relationships have changed between 1979 and 1999 when there have been large increases in both inequality and the returns to schooling. The primary drawback with the Census data is that we cannot examine the potential for ability bias or measurement error problems.

We present estimates of the returns to schooling for men and women by six race/ethnicity categories in Figure 1a.[12] In 1979, an additional year is associated with a 7.3 percent increase in

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[11] Respondents were also asked a few other questions about their siblings (e.g., age, sex).

[12] These returns to schooling were estimated from OLS regressions of the logarithm of annual earnings on years of schooling, indicators for 9 regions, a quadratic in potential experience, and in Figure 1a, an indicator for whether the individual was female. The regressions were weighted by the Census weight.

annual income for African Americans and a 8.5 percent increase in annual income for whites. Between 1979 and 1989 the estimated return to schooling increased dramatically for all races – especially African Americans – but remained in a relatively tight range from 10.7 percent for the "other" category to 12.3 percent for Asians and Pacific Islanders. Between 1989 and 1999 there was a much smaller increase in the estimated returns to education on average, but an increase in the range of estimates (10.2 percent for "other to 13.6 percent for Asians and Pacific Islanders). As shown in the contrast between Figures 1b (for women) and 1c (for men), this increase in the variation in the returns to schooling by race/ethnicity is particularly true for men.[13]

Based on estimates of the returns to schooling using Census data, we would conclude that the estimated return to schooling for African Americans is roughly the same as that for whites and that the return for Hispanics is somewhat lower. While previous selection- and measurement error-corrected estimates suggest OLS generates an estimate of the return to schooling that is roughly "right" overall (because the selection and measurement error biases balance one another) (e.g., Ashenfelter and Rouse (1998)), we do not know if this "rule of thumb" holds by race and ethnicity.

B. Results Using the NLS Young Men and Young Women

Using the NLS Young Men and Young Women cohorts we can similarly estimate returns to education overall and separately for whites and African Americans. While we cannot generate estimates for any other racial or ethnic group, we can use the NLS measure of ability and

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[13] For an analysis of the change in returns to schooling by race and ethnicity between 1979 and 2000 using the *Current Population Survey* see Bradbury (2002).

information on siblings to get some idea about the role of ability bias.[14] Later, when we turn to the NLSY79 cohort we will additionally be able to address measurement error issues.

Table 2a provides various estimates of the return to schooling (× 100) overall and separately for African Americans and whites using the NLS Young Men and Young Women cohorts. Each cell represents estimates from a separate regression and each column represents estimates from a different specification. All estimates are based on a regression of the natural logarithm of hourly pay on years of completed education, a third-order polynomial in age, an indicator for whether the individual is female, an indicator for whether the individual lives in the South, and a constant. We weight observations using the 1978 sampling weights. Estimates for the overall sample (shown in row 1) include indicators for whether the individual's race is African American or other.[15]

Using the entire sample we estimate that an extra year of education increases hourly pay by almost 6 percent. The separate estimates by race are quite similar. Although the estimate for African Americans is somewhat higher than for whites, the difference is not statistically significant. The estimates shown in column (1) of Table 2a do not control for the potential selection on ability problem discussed above. Because IQ scores are missing for a nonrandom subset of the sample, the estimates in column (2) are based on the sample of individuals who have an NLS measure of IQ but do not include IQ score in the regression. In column (3) we control for ability by including the IQ score in the regression. The column (2) estimates of the returns to education are somewhat smaller

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[14] In addition, because the sample sizes are so small we cannot estimate the returns to schooling for men and women separately.

[15] These specifications and those using the NLSY79 do not account for possible correlations across individuals within the same household. As a result the standard errors are conservative, although in the cross-sectional specifications allowing for such intra-household correlations makes little difference.

at 5 percent, but once again we find no strong evidence that the returns to education differ between African Americans and whites. Controlling for IQ score in column (3), we see some evidence that indeed those who get more education are more able as the estimated returns to education decline by almost one percentage point relative to the column (2) estimates.[16] Note, however, there is a slightly larger decline in the estimates between columns (2) and (3) for whites. This larger decrease may be indicative of more selection on ability for whites or the IQ score may be a noisier measure of ability for African Americans such that the column (3) estimates do not fully account for ability bias for this subgroup.

In Table 2b we turn to the NLS Young Men and Young Women sibling sample in order to allow for sibling fixed effects as well. The results in column (1) are based on the sample of siblings and are quite similar to the estimates based on the entire sample. The returns to education estimated from the sibling sample is about 5.5 percent overall; the estimate for African Americans is higher at roughly 7 percent although once again the difference is not statistically significant. When we allow for a sibling fixed effect in the column (2) estimates, the estimates decline by 10 percent for African Americans and 25 percent for whites. In column (3) we further restrict the sibling sample to those with non-missing IQ scores but do not directly control for IQ score in the regression. Again we see that the subsample with non-missing IQ scores is not a random subset of our estimation sample. Returns to schooling estimates for African Americans and whites are again slightly lower than for the full sibling sample. Finally, in column (4) we re-estimate returns to schooling while controlling for ability with IQ score and allowing for a sibling fixed effect. The estimated return to

[16] Estimates that include IQ score but include an indicator for IQ score is missing are quite similar to the estimates shown in column (2). The coefficient (standard error) estimates are 5.06 (0.25) overall, 5.93 (0.44) for African Americans, and 4.86 (0.30) for whites.

schooling for African Americans is little changed by including IQ score once we have already allowed for a sibling fixed effect. The estimate of the return to schooling for whites declines by 16 percent when controlling for a direct measure of ability.

In general we conclude from these surveys that the returns to schooling for African Americans and whites are roughly equal, even after controlling for ability bias. However, using these surveys we cannot correct for classical measurement error bias.

C. Results Using the NLSY79

Estimates of the return to schooling using the NLSY79 are presented in Table 3 overall and in Tables 4a and 4b for women and men separately. Each table has the following layout. Each cell represents the return to schooling (× 100) from a separate regression. The basic specification is an OLS regression of the natural logarithm of hourly pay on years of completed education, a third-order polynomial in age, an indicator for whether the individual is female, indicators for four geographic regions, and a constant. These regressions are unweighted, although results are similar if we weight. Further, as with the analysis using the older NLS surveys, we do not cluster the standard errors on the household such that the standard errors, especially those within sibling, are understated.[17]

Each column represents a different specification. The specifications in columns (1) and (2) use the full sample; those in columns (3) - (8) are restricted to the sibling sample. The estimates in the odd-numbered columns do not include the AFQT score while those in the even-numbered columns do. The estimates in columns (5) - (8) control for a sibling fixed effect, and those in

[17] In the cross-sectional specifications, the standard errors using "cluster" are very similar to those that do not use allow for such intra-household correlations; the standard errors for the within-sibling specifications presented in columns (5) and (6) in Table 3 are understated by about 50%.

columns (7) and (8) use the average of the sibling-reports of the respondent's education as an instrumental variable in an IV analysis. Thus, the estimates in columns (1) and (3) represent the cross-sectional estimates and those in columns (2) and (4) address for selection bias by controlling for the AFQT score. The estimates in column (5) control for selection by including a sibling fixed effect; those in column (6) control for both a sibling fixed effect *and* the AFQT score; and the estimates in columns (7) and (8) are similar to those in columns (5) and (6) but also correct for measurement error.

In Table 3 we estimate an overall cross-sectional return to schooling of about 9 percent. The estimate is highest among African Americans (10.4 percent) and lowest among Hispanics (7.6 percent). Only the estimated return for Hispanics is statistically different from that for the other two groups. In general, controlling for selection by including an AFQT score decreases the cross-sectional estimate of the return to schooling by about 3 percentage points (i.e., by comparing columns (2) and (1) or columns (4) and (3)). However, while controlling for the AFQT makes the biggest difference for the estimated returns to schooling (i.e., decreases the coefficient by the most) for African Americans and Hispanics in the full sample (columns (1) and (2)), it makes the biggest difference for non-minorities in the sibling sample (columns (3) and (4)). Overall, based on the cross-sectional estimates, we find little difference in the return to schooling by race/ethnicity.

Estimates that account for a sibling fixed effect are presented in columns (5) - (8). As also found with the older NLS surveys, a comparison of the estimates in columns (3) and (5) (or columns (4) and (6)) suggest that controlling for a sibling fixed effect makes a bigger difference for non-minorities than for minorities, especially African Americans. In fact, the within-sibling estimate of the return to schooling for African Americans is at most 1percentage point lower than the cross-sectional estimate. The fact that controlling for siblings makes a smaller difference for minorities than non-minorities may reflect less selection bias in the cross-sectional returns to schooling. Or, it may suggest that controlling for a sibling fixed effect is less effective for some populations than others. Why might this occur? While we limit our sample to "siblings" (excluding other household relationships, such as spouses, parents, foster siblings, step-siblings, and adopted siblings), we cannot distinguish between "full siblings" and "half-siblings." If African Americans and Hispanics are more likely to live with half-siblings than are non-minorities, then the family fixed effect may not be a good proxy for unobserved family "ability."[18]

Finally, the overall measurement error corrected (IV) estimate of the within-sibling return to schooling increases to 9.16 percent in column (7) relative to an estimate of 7.6 percent in column (5) suggesting an attenuation bias of about 20 percent. Further, we find that correcting for measurement error has the greatest effect on the estimated returns to schooling for African Americans, as expected based on the reliability ratios in Table 1. Generally, while we continue to estimate a larger point estimate for African Americans than for Hispanics and non-minorities, the differences across race/ethnicity are not statistically significant.[19]

While we find that returns to schooling, overall, do not appear to vary much by race or ethnicity, in Tables 4a and 4b we examine whether this pattern also holds separately for men and

[18] As evidence consistent with this explanation, the within-sibling coefficients in column (6) (that control for AFQT) decrease the most relative to column (5) (that do not control for AFQT) for African Americans. This pattern of results would be expected if the AFQT controls for other aspects of "ability" that are not captured by the sibling fixed effect.

[19] In results not presented here, we have also estimated these specifications using annual earnings rather than hourly wages. While the estimated returns to schooling are alittle higher, there are no differences by race or ethnicity.

women. The results for women are reported in Table 4a and those for men in Table 4b. These tables have a similar structure to Table 3. We continue to estimate no significant differences in the returns to schooling across African Americans, Hispanics, and non-minorities for both men and women. While this result may partially obtain because of smaller sample sizes (resulting in less precise estimates), the point estimates are also quite similar.

V. Conclusion

Alarmed by the increasing wage and income inequality in the U.S., many researchers and policymakers who are concerned that low-income individuals are losing ground have turned to policies aimed at increasing educational attainment. And because African Americans and Hispanics are disproportionately among the low-income, they are also disproportionately the focus of such policies. Yet, we know little about the magnitude of the economic benefit from the increased education for these subgroups of the population. Using data from the Census and the *National Longitudinal Surveys*, we find little evidence of differences in the return to schooling across racial and ethnic groups, even with attempts to control for ability and measurement error biases. While we find point estimates that are relatively similar across racial and ethnic groups, we also partly conclude this because of relatively large standard errors in some specifications due to small sample sizes. That said, we find no evidence that returns to schooling are lower for African Americans or Hispanics than for non-minorities. As a result, policies that increase education among the low-skilled have a good possibility of increasing economic well-being and reducing inequality.

We find some evidence that measurement error and selection bias may differ by race and ethnicity. For example, self-reported levels of schooling are “noisier” for African Americans than for other groups. And, we find less evidence of ability bias among African Americans and Hispanics than among non-minorities. The finding of less ability bias among minorities may arise because there is indeed less selection among these groups. Or, estimators that attempt to address self-selection may be less effective for some subgroups. More generally, our analysis suggests further research is needed to better understand the nature of measurement error and ability bias across subgroups in order to fully understand potential heterogeneity in the return to schooling across the population.

References

Altonji, Joseph and Thomas Dunn. "The Effects of Family Characteristics on the Return to Education." *Review of Economics and Statistics*, 78 (1996), 692-704.

Angrist, Joshua D. and Alan B. Krueger. "Does Compulsory Schooling Affect Schooling and Earnings?" *Quarterly Journal of Economics*, vol. 106 (November 1991): 979-1014.

Ashenfelter, Orley and John Ham. "Education, Unemployment, and Earnings." *Journal of Political Economy* 87 (October 1979), no. 5 part 2: S99-S116.

Ashenfelter, Orley and Alan Krueger. "Estimating the Returns to Schooling Using a New Sample of Twins." American Economic Review 84 (December 1994): 1157-1173.

Ashenfelter, Orley and Cecilia Rouse. "Schooling, Intelligence, and Income in America" in Kenneth Arrow, Samuel Bowles, and Steven Durlauf, eds. *Meritocracy and Economic Inequality*, Princeton, NJ: Princeton University Press, 2000, 89-117.

Ashenfelter, Orley and Cecilia Rouse. "Income, Schooling, and Ability: Evidence from a New Sample of Twins." *Quarterly Journal of Economics* 113 no. 1 (February 1998), pp. 253-284.

Becker, Gary S. *Human Capital: A Theoretical and Empirical Analysis, with Special Refernce to Education.* New York: Columbia University Press, 1967.

Bradbury, Katharine L. "Education and Wages in the 1980s and 1990s: Are All Groups Moving Up Together?" *New England Economic Review*, First quarter 2002, pp. 19-46.

Bureau of Labor Statistics, U.S. Department of Labor. *NLS Handbook, 2003*. Retrieved January 13, 2005, from http://www.bls.gov/nls/handbook/nlshpdf.htm.

Card, David. "Earnings, Schooling, and Ability Revisited" in Solomon Polachek, ed., *Research in Labor Economics*, Greenwich, CT: JAI Press (1995), 23-48.

Card, David, "Estimating the Return to Schooling: Progress on Some Persistent Econometric Problems," *Econometrica* 69, no. 5 (September 2001): 1127-1160.

Carneiro, Pedro and James J. Heckman. "Human Capital Policy" in *Inequality in America: What Role for Human Capital Policies?* Benjamin M. Friedman (ed.) (Cambridge, MA: The MIT Press, 2003), pp. 77-239.

Carneiro, Pedro, James J. Heckman, and Edward Vytlacil. "Understanding What Instrumental Variables Estimate: Estimating Marginal and Average Returns to Education." University of Chicago working paper, 2003.

Chamberlain, Gary and Zvi Griliches. "Unobservables with a Variance-Components Structure: Ability, Schooling and the Economic Success of Brothers." *International Economic Review* 16 (June 1975): 422-449.

Chamberlain, Gary and Zvi Griliches. "More on Brothers." In *Kinometrics: The Determinants of Socio-economic Success within and between Families*, edited by Paul Taubman. (Amsterdam: North-Holland), 1977.

Gorseline, D.W. *The Effect of Schooling Upon Income*. (Bloomington: Indiana University Press), 1932.

Griliches, Zvi. "Estimating the Returns to Schooling: Some Econometric Problems." *Econometrica* 45, no. 1 (January 1977): 1-22.

Herrnstein, Richard J. and Charles Murray. *The Bell Curve: Intelligence and Class Structure in American Life* (New York: The Free Press), 1994.

Kane, Thomas, Cecilia Rouse, and Douglas Staiger. "Estimating the Returns to Schooling When Schooling is Misreported" National Bureau of Economic Research Working Paper No. 7235, July 1999.

Krueger, Alan. "Inequality, Too Much of a Good Thing" in *Inequality in America: What Role for Human Capital Policies?* Benjamin M. Friedman (ed.) (Cambridge, MA: The MIT Press, 2003), pp. 1-75.

Mincer, Jacob. *Schooling, Experience, and Earnings* (New York: Columbia University Press), 1974.

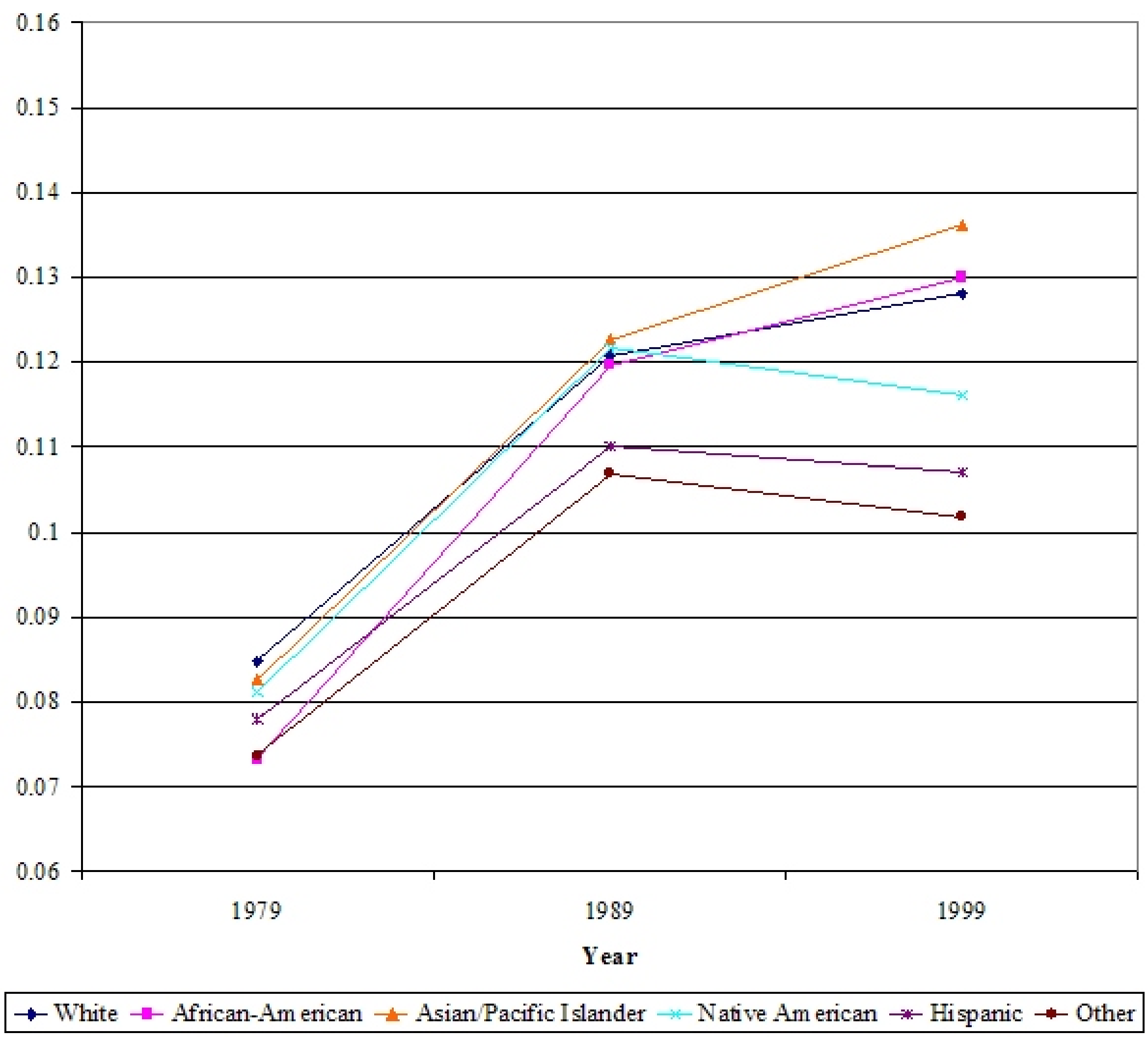
Neal, Derek A., and William A. Johnson. "The Role of Pre-Market Factors in Black-White Wage Differences." *Journal of Political Economy* 104 (October 1996), pp. 869-895.

Park, Jin Huem. "Estimation of Sheepskin Effects and Returns to Schooling Using the Old and the New CPS Measures of Educational Attainment." Industrial Relations Section Working Paper #338 *(*August 1994).

Rodgers, William and William Spriggs. "What Does AFQT Really Measure: Race, Wages, Schooling and the AFQT Score." *The Review of Black Political Economy*, 24 (1996), pp. 13-46.

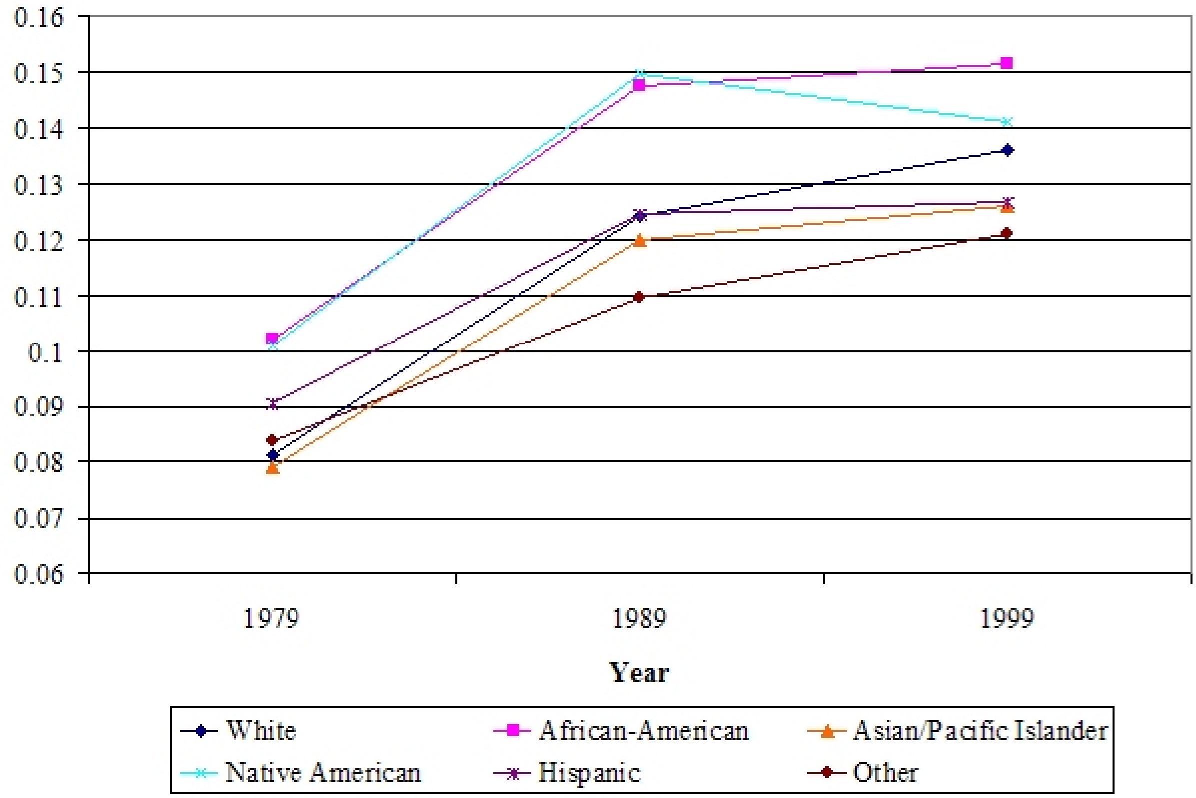
Taber, Christopher. "The Rising College Premium in the Eighties: Return to College or Return to Unobserved Ability?" *Review of Economic Studies* 68, no. 3 (July 2001), pp. 665-691.

Figure 1a: Returns to Education for Women and Men



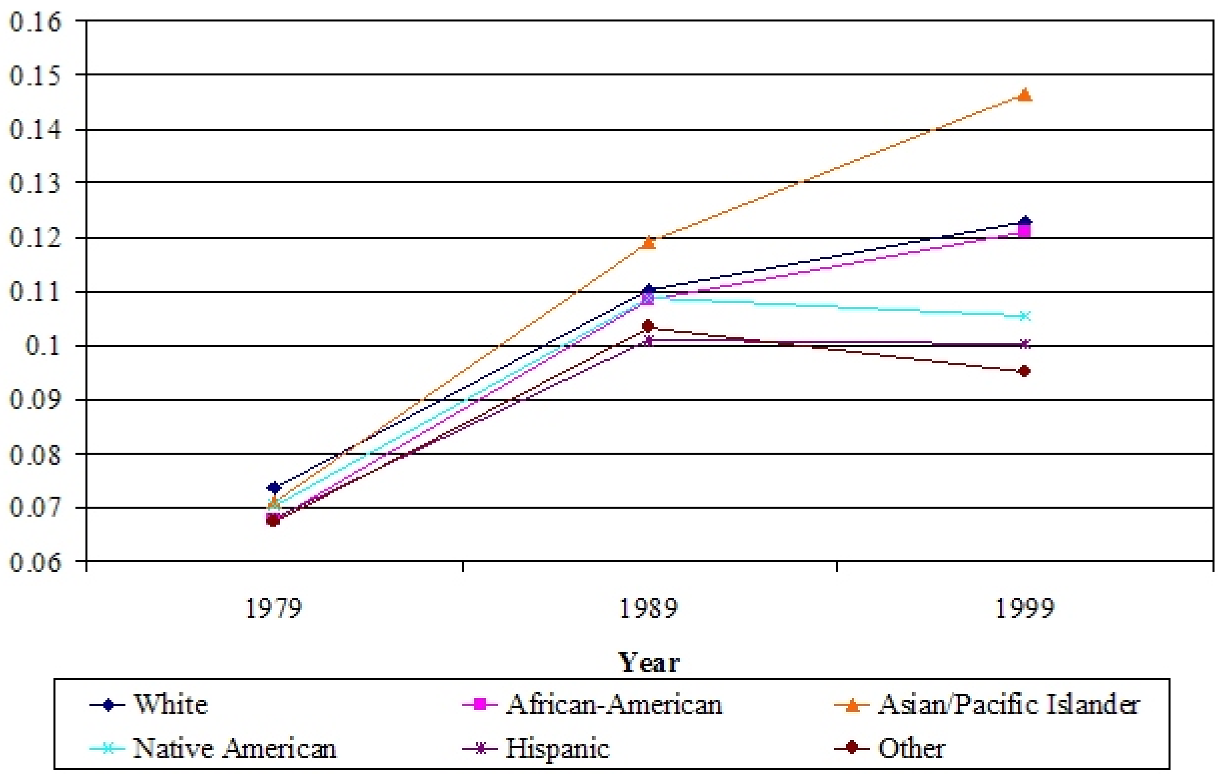
Notes: Estimates use U.S. Decennial Census Data from 1980, 1990, and 2000 and are based on a regression of the logarithm of annual earnings on years of schooling, indicators for 9 regions, a quadratic in potential experience, and an indicator for whether the individual was female. Observations are weighted using Census individual weights.

Figure 1b: Returns to Education Over Time for Women



Notes: Estimates use U.S. Decennial Census Data from 1980, 1990, and 2000 and are based on a regression of the logarithm of annual earnings on years of schooling, indicators for 9 regions, and a quadratic in potential experience. Observations are weighted using Census individual weights.

Figure 1c: Returns to Education Over Time for Men



Notes: See notes for Figure 1b.

Table 1
Estimated Reliability Ratios for Schooling Levels and Within-Siblings,
by Race, Ethnicity, and Sex of Individual

	Levels	Within-sibling
	All	
All	0.89	0.74
African American	0.81	0.67
Hispanic	0.86	0.77
Not African American/Not Hispanic	0.92	0.78
	Women	
All	0.88	0.76
African American	0.83	0.68
Hispanic	0.82	0.76
Not African American/Not Hispanic	0.91	0.81
	Men	
All	0.89	0.73
African American	0.79	0.65
Hispanic	0.88	0.77
Not African American/Not Hispanic	0.92	0.76

Note: Estimates are based on sibling data from the NLSY79. The within-sibling education is the deviation of the individual's schooling level from the mean education of his or her siblings.

Table 2a
Estimates of the Returns to Schooling Using the NLS Young Men and Young Women Cohorts

	(1)	(2)	(3)
Overall	5.69 (0.23)	5.11 (0.30)	4.18 (0.35)
African American	6.24 (0.41)	6.03 (0.78)	5.32 (0.85)
White	5.51 (0.27)	5.02 (0.33)	4.11 (0.38)
Drop observations missing IQ score	N	Y	Y
Include IQ score	N	N	Y

Notes: Estimates of the return to schooling (× 100) based on regressions of the natural logarithm of hourly pay on years of completed schooling, an indicator for sex equals female, a third-order polynomial in age, an indicator for living in the South, an indicator for region is missing, and a constant. Observations are weighted using the NLS provided sampling weights for 1978. The "overall" estimates reported in row 1 also include indicators for race is African American and race is other. Estimates in column (3) include the NLS measure of IQ.

Table 2b
Estimates of the Returns to Schooling Using Siblings from the NLS Young Men and Young Women Cohorts

	(1)	(2)	(3)	(4)
Overall	5.49 (0.47)	4.63 (0.85)	3.65 (1.32)	3.14 (1.38)
African American	7.05 (0.73)	6.28 (1.23)	5.73 (3.34)	5.54 (3.43)
White	5.09 (0.59)	3.80 (1.09)	3.27 (1.46)	2.76 (1.54)
Drop observations missing IQ score	N	N	Y	Y
Include IQ score	N	N	N	Y
Sibling fixed effect	N	Y	Y	Y

Notes: Estimates of the return to schooling (× 100) based on regressions of the natural logarithm of hourly pay on years of completed schooling, an indicator for sex equals female, a third-order polynomial in age, an indicator for living in the South, an indicator for region is missing, and a constant. Observations are weighted using the NLS provided sampling weights for 1978. The "overall" estimates reported in row 1 also include indicators for race is African American and race is other. Estimates in columns (2), (3), and (4) include a family fixed effect. Estimates in column (4) include the NLS measure of IQ.

Table 3
OLS and IV Estimates of the Returns to Schooling Using the NLSY 1979 Cohort

	Full Sample		Sibling Sample					
	Cross-section				Within Siblings			
	OLS	OLS	OLS	OLS	OLS	OLS	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Overall	9.29 (0.24)	6.29 (0.30)	9.36 (0.40)	6.15 (0.48)	7.60 (0.69)	5.21 (0.76)	9.16 (0.88)	6.74 (1.02)
African American	10.43 (0.49)	7.15 (0.58)	9.97 (0.73)	7.14 (0.89)	9.84 (1.23)	6.60 (1.38)	12.66 (1.71)	9.87 (2.12)
Hispanic	7.63 (0.55)	4.22 (0.67)	9.18 (1.05)	5.92 (1.21)	7.21 (1.60)	5.32 (1.76)	8.52 (1.91)	6.45 (2.20)
Not African American/Not Hispanic	9.43 (0.33)	6.69 (0.40)	9.17 (0.54)	5.64 (0.66)	6.30 (1.00)	4.44 (1.07)	7.54 (1.24)	5.65 (1.38)
Include AFQT score	N	Y	N	Y	N	Y	N	Y
Sibling Fixed Effect	N	N	N	N	Y	Y	Y	Y

Notes: Estimates of the return to schooling (× 100) based on regressions of the natural logarithm of hourly pay in 1993 on years of completed schooling, an indicator for sex equals female, a third-order polynomial in age, indicator for regions, an indicator for region is missing, and a constant. The "overall" estimates reported in row 1 also include indicators for race is African American and ethnicity is Hispanic. Estimates in columns (2), (4), (6), and (8) include the AFQT score. The average of the sibling-reports of the respondent's education is used as the instrumental variable in columns (7) and (8). All estimates are unweighted.

Table 4a
OLS and IV Estimates of the Returns to Schooling Using the NLSY 1979 Cohort, Women

	Full Sample		Sibling Sample					
	Cross-section				Within Siblings			
	OLS	OLS	OLS	OLS	OLS	OLS	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Overall	10.45	7.37	9.79	6.01	9.87	6.62	10.83	7.37
	(0.35)	(0.42)	(0.80)	(0.92)	(1.31)	(1.44)	(1.59)	(1.83)
African American	11.45	7.54	12.08	7.37	10.63	5.39	13.93	8.26
	(0.67)	(0.78)	(1.35)	(1.56)	(2.42)	(2.59)	(3.38)	(4.06)
Hispanic	9.48	5.52	11.88	7.17	9.04	7.36	9.71	8.27
	(0.74)	(0.91)	(2.04)	(2.38)	(6.04)	(3.65)	(3.54)	(4.51)
Not African American/Not	10.37	7.90	8.10	5.35	9.13	6.61	9.54	6.75
Hispanic	(0.50)	(0.59)	(1.14)	(1.31)	(1.95)	(2.13)	(2.23)	(2.52)
Include AFQT score	N	Y	N	Y	N	Y	N	Y
Sibling Fixed Effect	N	N	N	N	Y	Y	Y	Y

Notes: Estimates of the return to schooling (× 100) based on regressions of the natural logarithm of hourly pay in 1993 on years of completed schooling, a third-order polynomial in age, indicator for regions, an indicator for region is missing, and a constant. The "overall" estimates reported in row 1 also include indicators for race is African American and ethnicity is Hispanic. Estimates in columns (2), (4), (6), and (8) include the AFQT score. The average of the sibling-reports of the respondent's education are used as the instrumental variable in columns (7) and (8). All estimates are unweighted.

Table 4b
OLS and IV Estimates of the Returns to Schooling Using the NLSY 1979 Cohort, Men

	Full Sample		Sibling Sample					
	Cross-section				Within Siblings			
	OLS	OLS	OLS	OLS	OLS	OLS	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Overall	8.23 (0.34)	5.28 (0.42)	8.01 (0.70)	5.15 (0.85)	5.69 (1.26)	3.32 (1.39)	8.16 (1.66)	5.91 (1.98)
African American	9.43 (0.72)	6.66 (0.87)	8.66 (1.31)	7.60 (1.61)	8.05 (2.10)	4.88 (2.47)	8.20 (2.99)	4.89 (3.92)
Hispanic	6.05 (0.79)	2.90 (0.96)	6.31 (1.94)	3.73 (2.20)	2.92 (3.55)	0.52 (3.96)	8.61 (4.27)	6.97 (5.04)
Not African American/Not Hispanic	8.66 (0.44)	5.67 (0.55)	8.22 (0.93)	4.35 (1.13)	5.06 (1.71)	3.30 (1.85)	7.79 (2.24)	6.23 (2.56)
Include AFQT score	N	Y	N	Y	N	Y	N	Y
Sibling Fixed Effect	N	N	N	N	Y	Y	Y	Y

Notes: Estimates of the return to schooling (× 100) based on regressions of the natural logarithm of hourly pay in 1993 on years of completed schooling, a third-order polynomial in age, indicator for regions, an indicator for region is missing, and a constant. The "overall" estimates reported in row 1 also include indicators for race is African American and ethnicity is Hispanic. Estimates in columns (2), (4), (6), and (8) include the AFQT score. The average of the sibling-reports of the respondent's education are used as the instrumental variable in columns (7) and (8). All estimates are unweighted.

Appendix Table 1a
Mean Characteristics for the 1978 National Longitudinal Surveys
Young Men and Young Women Cohorts

	Full Sample			Sub-sample with IQ Scores		
	Overall	African American	White	Overall	African American	White
Hourly pay	15.39	11.84	15.88	16.17	12.79	16.42
	[10.13]	[5.82]	[10.50]	[10.12]	[5.78]	[10.33]
Years of completed schooling	13.30	12.20	13.44	13.73	13.00	13.78
	[2.53]	[2.50]	[2.48]	[2.29]	[2.00]	[2.30]
Female	0.45	0.49	0.44	0.43	0.51	0.43
	[0.50]	[0.50]	[0.50]	[0.50]	[0.50]	[0.49]
Age	29.66	29.42	29.70	30.01	29.66	30.04
	[3.30]	[3.28]	[3.30]	[3.06]	[2.94]	[3.07]
Southern region	0.34	0.65	0.30	0.30	0.56	0.28
	[0.48]	[0.48]	[0.46]	[0.46]	[0.50]	[0.45]
Number of observations	4802	1263	3486	3240	558	2646

Notes: Standard deviations are in brackets. All means and standard deviations are calculated using the 1978 survey weights. Hourly pay is in real 2003 dollars. Men's age equals age in the initial survey year plus 12; women's age equals age in the initial survey year plus 10. The overall sample is 12 percent African American, 0.8 percent other, and 87 percent white. Conditional on having an IQ score, the overall sample is 7.4 percent African American, 0.7 percent other, and 92 percent white.

Appendix Table 1b
Mean Characteristics for Siblings from the 1978 National Longitudinal Surveys
Young Men and Young Women Cohorts

	Full Sample			Sub-sample with IQ scores		
	Overall	African American	White	Overall	African American	White
Hourly pay	15.11	11.59	15.71	16.37	13.98	16.46
	[12.60]	[5.43]	[13.42]	[10.66]	[6.51]	[10.82]
Highest grade completed	13.48	12.16	13.72	14.09	13.02	14.16
	[2.48]	[2.67]	[2.43]	[2.36]	[1.90]	[2.36]
Female	0.44	0.49	0.43	0.39	0.43	0.39
	[0.50]	[0.50]	[0.50]	[0.49]	[0.50]	[0.49]
Age	28.41	28.39	28.39	28.96	28.91	28.91
	[2.73]	[2.85]	[2.68]	[2.56]	[2.49]	[2.50]
Southern region	0.35	0.70	0.29	0.30	0.62	0.27
	[0.48]	[0.46]	[0.45]	[0.46]	[0.49]	[0.44]
Number of observations	1263	434	819	642	124	512

Notes: See notes for Appendix Table 1a. The overall sample is 16.3 percent African American, 0.6 percent other, and 83 percent white. Conditional on having an IQ score, the overall sample is 8.4 percent African American, 0.9 percent other, and 91 percent white.

Appendix Table 2
Mean Characteristics for the 1993 National Longitudinal Survey of Youth, 1979

	Full Sample				Sibling Sample			
	Overall	African American	Hispanic	White	Overall	African American	Hispanic	White
Hourly pay	13.67	11.52	13.22	15.06	13.94	11.41	13.79	15.67
	[9.97]	[7.66]	[10.30]	[10.76]	[11.43]	[8.37]	[14.08]	[11.76]
Years of completed schooling	13.11	12.89	12.43	13.48	13.21	12.87	12.57	13.68
	[2.35]	[2.01]	[2.48]	[2.42]	[2.30]	[2.00]	[2.11]	[2.46]
Female	0.48	0.48	0.47	0.47	0.46	0.45	0.43	0.47
	[0.50]	[0.50]	[0.50]	[0.50]	[0.50]	[0.50]	[0.49]	[0.50]
Age	31.51	31.55	31.39	31.54	31.17	31.23	31.00	31.19
	[2.25]	[2.23]	[2.25]	[2.26]	[2.05]	[2.04]	[2.04]	[2.07]
Northeast region	0.17	0.14	0.14	0.19	0.17	0.12	0.13	0.21
	[0.37]	[0.35]	[0.35]	[0.39]	[0.37]	[0.32]	[0.33]	[0.41]
North Central region	0.23	0.17	0.07	0.33	0.25	0.14	0.10	0.38
	[0.42]	[0.37]	[0.26]	[0.47]	[0.43]	[0.34]	[0.29]	[0.49]
Southern region	0.41	0.62	0.31	0.32	0.41	0.71	0.31	0.25
	[0.49]	[0.48]	[0.46]	[0.46]	[0.49]	[0.45]	[0.46]	[0.44]
AFQT score	64.97	51.62	57.83	75.25	63.97	50.03	58.42	75.29
	[21.71]	[18.73]	[19.82]	[18.52]	[21.63]	[17.96]	[18.56]	[18.51]
Number of observations	6119	1819	1142	3158	2419	786	446	1187

Notes: Standard deviations are in brackets. All means and standard deviations are calculated using the 1993 survey weights. Hourly pay is in real 2003 dollars. Age equals age in the initial survey year plus 14.

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