

Education and Labor Market Discrimination

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Using a model of statistical discrimination and educational sorting, we explain why blacks get more education than whites of similar cognitive ability, and we explore how the Armed Forces Qualification Test (AFQT), wages, and education are related. The model suggests that one should control for both AFQT and education when comparing the earnings of blacks and whites, in which case a substantial black-white wage differential emerges. We reject the hypothesis that differences in school quality between blacks and whites explain the wage and education differentials. Our findings support the view that some of the black-white wage differential reflects the operation of the labor market. (JEL I21, J15, J24, J31, J71)

Models of statistical discrimination often imply that black Americans—or others suffering from discrimination—invest less in themselves than otherwise comparable whites do, because blacks receive a lower return from investment in human capital than whites. The critical assumption underlying these models is that the market receives imperfect information about productivity and no direct information about investments in human capital (see, for example, Shelly J. Lundberg and Richard Startz 1983; Stephen Coate and Glenn C. Loury 1993). Although many human-capital investments are undoubtedly unobservable, others, such as educational attainment, are readily observed. Observable investments can signal productivity as in Michael A. Spence (1973), and the value of the signal will be greater, the less reliable is the direct observation of productivity by employers. Moreover, evidence suggests that employers find it more difficult to evaluate black job candidates than to evaluate white candidates. Therefore, when observable investments in human capital are available, it is plausible that statistical discrimination will induce blacks to invest in themselves more than whites do, not less.

To formalize this intuition, we construct a model of statistical discrimination in which a worker's race and educational attainment are observable by prospective employers, but a worker's ability is not observable except by the worker himself. We analyze the model from a theoretical perspective, and then we apply it to the data.

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In our empirical analysis, we proxy ability by performance on the Armed Forces Qualification Test (AFQT). In many cases, we limit the sample to those too young to have completed school as of the time of testing. We find that educational attainment, conditional on AFQT, is higher among blacks than it is among whites, a result that is consistent with our signaling hypothesis. We also predict that whites and blacks will have similar earnings at very low and very high levels of education (not controlling for ability) but that blacks will have lower earnings at intermediate levels of education. We do not test this for women because of selection issues (Derek Neal 2004), but we do confirm the hypothesis for men. Our particular formalization implies that blacks would have levels of education similar to those of whites at low and high levels of ability and levels exceeding those of whites at intermediate levels of ability. This is confirmed only for men (black women have more education than whites at all ability levels but the lowest).

Of course, there are other plausible explanations for these findings. In particular, if blacks attend lower-quality schools than whites do, and if AFQT performance is mostly determined by school inputs, then the AFQTs of blacks would benefit less from schooling than those of whites. For a given amount of schooling, blacks will have lower AFQTs, and, conversely, for a given AFQT score, blacks will have more education than whites. We test this hypothesis directly by controlling for measurable differences in school quality and find that school quality cannot explain the education differential. We also show that our results are robust to controlling for a number of other factors that influence educational attainment.

We consider AFQT performance to be a measure that reflects both innate ability and ability acquired up to the time of testing, though our approach requires that AFQT performance be exogenous to later educational attainment. We do not attempt to explain the behavior of children and adolescents prior to taking the AFQT, nor do we assume such behavior takes account of the value of investment in human capital. But we do assume that after the administration of the AFQT, students act as rational agents.

Our model and findings have important implications for the debate over the role of markets and premarket factors in explaining black-white wage differentials. In a highly influential article, Neal and William Johnson (1998; see also June O'Neill 1990) show that the black-white wage differential is dramatically reduced, and in some cases eliminated, by controlling for AFQT. Moreover, Neal and Johnson (hereafter, NJ) show that the effect of AFQT on the earnings of blacks is at least as large as on the earnings of whites,¹ so that labor-market discrimination ought not reduce black investment in the cognitive skills that the AFQT measures. Thus, they conclude that “the black-white wage gap primarily reflects a skill gap” generated by premarket factors, not by anticipated labor-market discrimination.

In what follows, we reopen this question. We argue that when examining black-white wage differentials, it is inappropriate to control only for AFQT performance.

¹Note, however, that if AFQT is influenced by investments in human capital, then we should expect the observed returns to AFQT to be equated to the agent's discount rate in equilibrium. If black and white workers have similar discount rates, they would also have similar equilibrium rates of return to AFQT. But it is perfectly possible that at a given AFQT score, the returns would differ between blacks and whites.

Blacks get more education than whites with the same AFQT do. So absent discrimination, we would expect blacks to be rewarded for their greater education. The similar earnings of blacks and whites when we control only for AFQT suggest that blacks are not so rewarded. We show that when we control for both education and AFQT, wage differentials between blacks and whites are substantially larger than when we control for AFQT alone.

Our model suggests that if we control for AFQT but not education, blacks will earn more than whites do at all but the highest and lowest levels of ability.² We reject this hypothesis for men. The failure of this hypothesis could reflect either missing variables or labor-market discrimination. This is an old debate and not one we pretend to resolve. We explore whether the wage differential could be explained by differences in the quality of schools attended by blacks and by whites, but we find no evidence to support that possibility.

I. Education and Race: A Signaling Model

In this section, we argue that statistical discrimination creates incentives for blacks to signal ability through education. Ethnographic evidence shows that blacks see education as a means of getting ahead. Blacks in low-skill jobs in Harlem view education as crucial to getting a good job, and those with low levels of education have difficulty obtaining even jobs that we would not normally think of as requiring a high school diploma (Katherine S. Newman 1999). Employers are particularly circumspect in their assessment of low-skill blacks (Joleen Kirschenman and Kathryn M. Neckerman 1991), a finding consistent with our approach.

Our theoretical model merges a standard statistical-discrimination model (e.g., Dennis J. Aigner and Glen G. Cain 1977) with a conventional sorting model. In a sense, it stands Lundberg and Startz on its head, by dealing with observable rather than unobservable investments. As is standard in the statistical discrimination literature, we assume that the productivity of blacks is less easily observed than the productivity of whites. We provide a detailed justification of this supposition after we present the basic model.

Furthermore, in accordance with our reading of the ethnographic literature, we add a nonstandard feature to our model of observability: as education levels increase, the ability of firms to assess the productivity of both black and white workers improves, until, for sufficiently high levels of education, firms assess productivity for the two races equally well. This feature is not required for our principal result, that blacks generally get more education than equivalent whites do, but it is needed for some of the other predictions. In addition, we assume that firms observe the productivity of the highly educated perfectly, although our results require only no asymmetry of information between highly educated workers and firms. This is consistent with Peter Arcidiacono, Patrick Bayer, and Aurel Hizmo (2008) who find that when workers first enter the labor market, college graduates but not high school graduates are rewarded for their ability. The market either observes the ability of a college graduate directly or infers it from grades, college

²The only exception occurs when education at the margin functions as a pure signal with no effect on productivity, a model variant described below.

attended, informal networks, etc., whereas the market learns this information about high school graduates only gradually.

We first develop a game-theoretic signaling model of educational attainment and apply it separately to blacks and whites, setting parameters for each group consistent with the characteristics described above. Then we compare the equilibrium outcomes of the two groups. The principal result is that, because employers have greater difficulty directly observing the productivity of blacks than of whites, they put more weight on education (an observable trait related to productivity) when making offers to blacks than they do when making offers to whites. In response, blacks get more education than do whites of similar ability. However, because observability changes, this effect disappears at high levels of education.

A. The Signaling Game among Racially Homogeneous Workers

Employers observe race and are free to separate workers by race when evaluating them. Workers of a given race differ in ability and educational attainment, both of which increase productivity. Only the worker observes his own ability, but potential employers observe the worker's educational attainment precisely and productivity directly but with error. The worker then has an incentive to use education as a signal of productivity. The greater the error in the employer's direct observation, the more weight he places on the education signal.

In the standard job-market signaling model, signaling is costly, and the cost falls as ability, and productivity, increase. If there is a separating equilibrium in which high-ability agents buy more of the signal than low-ability agents do. In our model, however, the opportunity cost of additional education is the same for all agents earning the same wage. Therefore, if equilibrium education fully revealed productivity and determined the wage, low-ability agents would replicate the signal of high-ability agents, and the equilibrium would unravel. To surmount this problem, we make the very realistic assumption that the worker cannot deduce his productivity from knowledge of his ability and education alone. Instead, productivity depends not only on characteristics known to the worker, but on a random element that the worker does not observe. Consequently, the employer always places a positive weight on his direct observation of the worker's productivity, which has a higher expected value for a higher-ability agent. A low-ability person who chose more schooling in order to pool with a high-ability person would therefore have a lower expected return than his high-ability counterpart. This allows a separating equilibrium to exist.

Consider, now, a game between a continuum of workers of different ability levels a , where a is continuously distributed over some fixed interval. Each worker must choose a level of education s . Because we assume that education and ability are complementary inputs in the creation of productivity (in a sense defined below), we search for a separating equilibrium in which the workers' strategy profile is described by a continuous and differentiable function $S(a)$, strictly increasing in a , where $s = S(a)$ is the education obtained by a worker of ability a . Firms in our model simply follow the rules of a competitive labor market—they play no strategic role. (But we require their beliefs about $S(a)$ in equilibrium to be correct.)

Suppose that a worker's productivity p^* , conditional on his education level s and ability a , has the log-normal distribution given by

$$(1) \quad p^* = Q(s, a)\hat{\varepsilon},$$

where $Q(s, a)$ is a deterministic function of education and ability and where $\varepsilon \equiv \ln \hat{\varepsilon}$ is a normal random variable with mean 0 and variance σ_ε^2 . Letting $q(s, a) \equiv \ln Q(s, a)$ denote the mean of $\ln p^*$, we can write log productivity as

$$(2) \quad \ln p^* = q(s, a) + \varepsilon.$$

We assume that the effect of additional education on log productivity is diminishing ($q_{ss} < 0$) but that ability complements the productivity-increasing effects of education ($q_{sa} > 0$).

A potential employer can observe a worker's education level s but not his true productivity p^* . However, the employer obtains a direct productivity signal p given by

$$(3) \quad \ln p = \ln p^* + u,$$

where u is a random error. The error term u has variance $\sigma_u^2(s)$, which is common to all firms, continuous, and decreasing in s . We assume that ε and u are independently distributed.

Let $\lambda(s) \in [0, 1]$ be given by

$$\lambda(s) \equiv \frac{\sigma_\varepsilon^2}{\sigma_\varepsilon^2 + \sigma_u^2(s)}.$$

For given $\sigma_\varepsilon^2 > 0$, if $\lambda(s)$ is near 0, then $\sigma_u^2(s)$ must be large, in which case the employer's ability to observe worker productivity directly is poor. Conversely, if $\lambda(s) = 1$, then $\sigma_u^2(s) = 0$, and the employer observes worker productivity perfectly. In the latter case, workers have no incentive to signal their productivity to employers and obtain the efficient level of education.

The Equilibrium Competitive Wage.—In the candidate-separating equilibrium described by the workers' strategy profile S , an employer can infer a worker's ability a from his knowledge of the worker's education s . If $\hat{q}(s)$ denotes the employer's equilibrium inference about the value of $q(s, a)$ conditional on s , it follows that $\hat{q}(s) \equiv q(s, A(s))$, where $A \equiv S^{-1}$.

PROPOSITION 1: *From the point of view of an employer who has observed a worker's productivity signal p and education level s , the conditional mean and variance of the unobservable random element ε is given by*

$$(4) \quad E[\varepsilon | p, s] = \lambda(s)(\ln p - \hat{q}(s))$$

and

$$(5) \quad \sigma^2[\varepsilon | p, s] = (1 - \lambda(s))\sigma_\varepsilon^2.$$

PROOF:

Because the values of $\ln p - \hat{q}(s)$ and s uniquely determine p , we know that any expectation conditioned on p and s will remain unchanged if conditioned on $\ln p - \hat{q}(s)$ and s instead. Therefore, we can write

$$(6) \quad E[\varepsilon | p, s] \equiv E[\varepsilon | \ln p - \hat{q}(s), s].$$

Moreover, (2) and (3) imply that

$$(7) \quad \ln p - \hat{q}(s) = u + \varepsilon$$

in equilibrium. The proposition now follows from (6) and from standard results for the sum of independent normal random variables.

In a competitive labor market, an employer will offer the wage $\hat{w}(p, s) \equiv E[p^* | p, s]$ to a worker with observed characteristics p and s . We show:

PROPOSITION 2: *The log of the equilibrium competitive wage is given by*

$$(8) \quad \ln \hat{w}(p, s) = \lambda(s)\ln p + (1 - \lambda(s))(\hat{q}(s) + 0.5\sigma_\varepsilon^2).$$

PROOF:

We calculate the expected values of the terms of equation (2) conditional on the observed p and s . This yields

$$(9) \quad E[\ln p^* | p, s] = \hat{q}(s) + E[\varepsilon | p, s].$$

Applying Proposition 1 give us

$$E[\ln p^* | p, s] = \lambda(s)\ln p + (1 - \lambda(s))\hat{q}(s)$$

and

$$\sigma^2[\ln p^* | p, s] = (1 - \lambda(s))\sigma_\varepsilon^2.$$

A lognormally distributed random variable x satisfies

$$\ln E[x] = E[\ln x] + (1/2)\sigma^2[\ln x],$$

which, applied to $\hat{w}(p, s) \equiv E[p^* | p, s]$, yields the proposition.

Workers' Equilibrium Strategies.—Each worker knows his own ability a , but must choose education s before ε and u are realized. When other workers have the strategy

profile $S(a)$, a designated worker's expectation of his wage, conditional on his own s and a , is given by $E_{\varepsilon,u} [\hat{w}(p,s)]$, where $E_{\varepsilon,u}$ integrates over ε and u .

As a first step in deriving the best response to $S(a)$ of a worker with ability a , we compute the value of $\ln E_{\varepsilon,u} [\hat{w}(p,s)]$. From (8), (2), and (3), we see that

$$\ln \hat{w}(p,s) = \lambda(s)(q(s,a) + u + \varepsilon) + (1 - \lambda(s))(\hat{q}(s) + 0.5\sigma_\varepsilon^2),$$

which is a normally distributed random variable with mean

$$E_{\varepsilon,u} [\ln \hat{w}(p,s)] = \lambda(s)q(s,a) + (1 - \lambda(s))(\hat{q}(s) + 0.5\sigma_\varepsilon^2)$$

and variance

$$\sigma^2[\ln \hat{w}(p,s)] = \lambda(s)^2 (\sigma_\varepsilon^2 + \sigma_u^2(s)) = \lambda(s)\sigma_\varepsilon^2.$$

Again, from the standard properties of log-normal random variables, we have

$$\ln E_{\varepsilon,u} [\hat{w}(p,s)] = E_{\varepsilon,u} [\ln \hat{w}(p,s)] + \frac{1}{2}\sigma^2 [\ln \hat{w}(p,s)],$$

so that

$$(10) \quad \ln E_{\varepsilon,u} [\hat{w}(p,s)] = \lambda(s)q(s,a) + (1 - \lambda(s))\hat{q}(s) + 0.5\sigma_\varepsilon^2.$$

This confirms the intuition that a designated worker's expected wage depends both on his actual ability and the ability level inferred by the employer, which in turn depends on $S(a)$.

Workers maximize expected discounted net income. Assume that the only cost of education is lost income while in school. If r is the worker's discount rate, the expected present value at time $t = 0$ of the future income of a worker with characteristics (s,a) is given by

$$v(s,a) \equiv \int_s^\infty e^{-rt} E_{\varepsilon,u} [\hat{w}(p,s)] dt \equiv \frac{1}{r} e^{-rs} E_{\varepsilon,u} [\hat{w}(p,s)],$$

or

$$\ln v(s,a) \equiv -\ln r - rs + \ln E_{\varepsilon,u} [\hat{w}(p,s)].$$

The first-order condition for maximizing $v(s,a)$ with respect to s is

$$(11) \quad \frac{\partial}{\partial s} \ln E_{\varepsilon,u} [\hat{w}(p,s)] = r.$$

This restates the well-known proposition that when the only cost of schooling is the student's opportunity cost, the worker gets education until the return equals the discount rate r . We restrict the class of equilibria we consider to those for which (11) has a unique solution.

We can now describe a separating equilibrium of the wage/education game among the class of strategy profiles $S(a)$ that are “well behaved” (continuous, differentiable, strictly increasing, and specifying a unique best response for every worker type). Throughout we use the term “equilibrium” to refer to a perfect-Bayesian equilibrium of the signaling game.

PROPOSITION 3: *If the support of worker abilities is the interval $[a_0, a_1]$, then any well-behaved separating equilibrium S has the property that the education level $S(a_0)$ of the lowest-type worker must be efficient and not influenced by signaling.*

PROOF:

In an equilibrium with $S(a)$ strictly increasing in a , the employer would infer that a worker with education $S(a_0)$ has ability a_0 , the lowest level in the support. If $S(a_0)$ were inefficiently high, the worker of ability a_0 could safely deviate to the lower efficient level of education without lowering the employer’s inference of his ability, and so raise his payoff. If $S(a_0)$ were inefficiently low, the worker of ability a_0 would deviate to $s > S(a_0)$ even without consideration of the positive payoff from signaling.

We can now provide a complete description of any well-behaved equilibrium.

PROPOSITION 4: *Suppose $[a_0, a_1]$ is the support of worker abilities. If a worker’s equilibrium strategy profile $S(a)$ is well behaved, then its inverse $A(s)$ must satisfy the differential equation*

$$(12) \quad q_s + (1 - \lambda)q_a A' = r.$$

For $0 \leq \lambda < 1$, this equation is equivalent to

$$(13) \quad S' = \frac{(1 - \lambda)q_a}{r - q_s}.$$

For $\lambda = 1$, the equilibrium condition is given by the solution for s of the equation

$$(14) \quad q_s(s, a) = r.$$

The solution of (14) defines the efficient level of education, which we denote by $S^*(a)$. Furthermore, we have $S(a_0) = S^*(a_0)$ for any function $\lambda(s)$. Therefore each $\lambda(s)$ corresponds to exactly one well-behaved equilibrium.

PROOF:

Substituting (10) into (11) yields the differential equation

$$\frac{\partial}{\partial s} (\lambda(s)q(s, a) + (1 - \lambda(s))\hat{q}(s) + 0.5\sigma_e^2) = r,$$

or

$$(15) \quad \lambda'(s)q(s, a) + \lambda(s)q_s(s, a) - \lambda'(s)\hat{q}(s) + (1 - \lambda(s))(q_s(s, A(s)) + q_a(s, A(s))A'(s)) = r.$$

This equation implicitly defines the best response s of a worker with ability a to the strategy profile S . Consequently, $a = A(s)$ and $q(s, a) = \hat{q}(s)$ in equilibrium, and (15) reduces to (12). Equation (13) follows from the fact that the derivative of S is the reciprocal of the derivative of A . Proposition 3 implies that $S(a_0) = S^*(a_0)$.

The left-hand side of equation (12) represents the worker's rate of return to a marginal unit of education, which, given appropriate concavity conditions and the strategy-profile-inverse A , must equal r . This rate of return reflects the direct and indirect effects of education.

First, consider the direct effect of additional education on the employer's inference of productivity when inferred ability is held constant. The direct effect works through two channels. For a given productivity signal p , additional education leads the employer to infer higher productivity by $(1 - \lambda)q_s$. But additional education also increases the expected value of the p signal, and the increase in p raises expected productivity by λq_s . These effects sum to q_s , the first term of (12).

Second, in equilibrium, increasing education causes the employer to increase inferred ability. The rate of increase of inferred ability with respect to education is A' , the effect of increased ability on expected log productivity is q_a , and the weight the employer puts on this inference (as opposed to his signal) is $1 - \lambda$. The second term, $(1 - \lambda)q_a A'$, is the product of these effects.

In (12), the term $(1 - \lambda)q_a A'$ is always nonnegative so that for any equilibrium $S(a)$, we have $q_s(S(a), a) \leq r$. Because we have assumed that q_{ss} is negative, and because the efficient level of education $S^*(a)$ is defined by $q_s(S^*(a), a) = r$, the following proposition holds:

PROPOSITION 5: *Let $S(a)$ describe any separating equilibrium of the workers' signaling game. Then for all $a \in [a_0, a_1]$, $S(a) \geq S^*(a)$.*

Because an equilibrium strategy profile satisfies $q_s(s, a) = r$ whenever $\lambda(s) = 1$ (see Proposition 4), and that equation characterizes a full-information level of education, we have:

PROPOSITION 6: *Let s^* be the lowest value of s such that $\lambda(s) = 1$ for all $s \geq s^*$, and let $a^* = A(s^*)$. Then, for $a \geq a^*$, $S(a)$ is the same as in the case where information about productivity is perfect at all levels of education.*

B. Ability as the Capacity to Be Educated: An Illustrative Example

If ability is defined simply as the capacity to be educated, we obtain an easily analyzed variant of the one-race signaling model. Let productivity p^* be given by

$$p^* = \min \{s, a\} \hat{\varepsilon},$$

where $\hat{\varepsilon} = \exp(\varepsilon)$ is a lognormal random variable. Then, the mean of $\ln p^*$ is given by

$$(16) \quad q(s, a) = \min \{\ln s, \ln a\}.$$

In this example, additional education is productive only when $s < a$. But additional ability is not productive in this region, so that the worker has no incentive to use additional education to signal ability. When $s > a$, additional ability is productive but additional education is not, so if the worker obtains additional education, signaling can be his only purpose. Therefore, the example decouples the productivity and signaling effects of education.

Because q is not differentiable at $s = a$ and $q_{ss} = 0$ for $s > a$, the propositions of the previous section do not necessarily apply to this example. However, because of its simplicity, we can derive unique closed-form equilibria for some specifications of $\lambda(s)$.

We begin by finding the efficient level of education. From (16), we have

$$q_s(s, a) = \begin{cases} 1/s & \text{for } s < a \\ 0 & \text{for } s > a \end{cases},$$

which defines the social rate of return to education. Additional education is efficient so long as $q_s(s, a) > r$. This means that the efficient level of education is given by

$$(17) \quad S^*(a) = \min \left\{ \frac{1}{r}, a \right\}.$$

This will be the equilibrium when information is perfect ($\lambda = 1$).

Let $S(a)$ be the equilibrium level of education associated with ability a . Consider, first, an interval of a on which $S(a) > a$. Because education is not productive in that region, we must have equilibrium signaling, which makes sense only if $\lambda(s) < 1$. Suppose the equilibrium is separating, so that S is strictly increasing there. Let A denote the inverse of S . The direct effect of more education on productivity is zero, and the rate of return of additional education mediated by the change in perceived ability is $(1 - \lambda) q_a A'$, as in equation (13). This yields the following necessary condition for equilibrium on the $S(a) > a$ interval:

$$(18) \quad S'(a) = \frac{1 - \lambda(s)}{r} \frac{1}{a}.$$

Suppose, now, that on a different interval of a we have $S(a) \leq a$. In that region, expected log productivity is determined by s alone, which the employer observes perfectly. Therefore, there is no incentive to signal and $S(a) = S^*(a)$. However, S^* cannot be an equilibrium in an interval of the form $[a_0, \hat{a}]$, because a worker with ability slightly greater than a_0 would always choose to signal (his rate of return to added education via inferred ability would exceed r). Therefore, $S(a) = \tilde{S}(a)$ near a_0 , where $\tilde{S}(a)$ is the solution of (18) with $\tilde{S}(a_0) = S^*(a_0)$ (Proposition 3 applies to equilibria that are separating in a neighborhood of a_0). Because $S(a) = S^*(a)$ whenever $S(a) \leq a$, it follows that the equilibrium strategy profile will have the form $S(a) = \max \{S^*(a), \tilde{S}(a)\}$.

For Figure 1, we specify $r = 0.0625$ ($1/r = 16$) and set $a_0 = 1$. From (17) we see that the efficient level of education $S^*(a)$ increases along the 45-degree line until

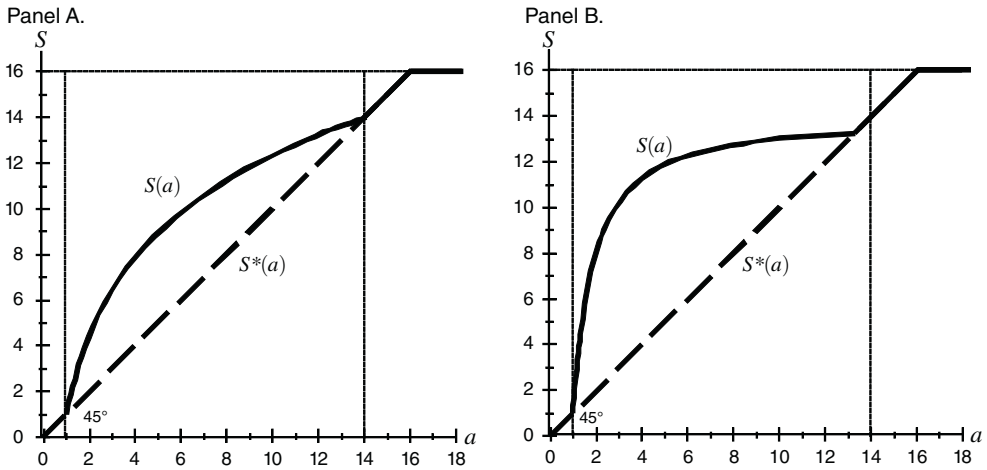


FIGURE 1. EDUCATION AS A FUNCTION OF ABILITY WITH AND WITHOUT SIGNALING

$s = 16$ and is constant at 16 thereafter. For panel A, we hold $\lambda(s)$ constant at λ_0 , in which case the appropriate solution of (18) is

$$(19) \quad \tilde{S}(a) = \frac{1 - \lambda_0}{r} \ln a + 1.$$

This describes the equilibrium in the region $s > a$ (above the 45-degree line). The function $S(a)$ is graphed with $\lambda_0 = 0.692$, a value calibrated to cross the diagonal at $s = 14$.

In panel B, we illustrate the situation in which $\lambda(s) = s/b$ (λ increases linearly in s and reaches 1 at $s = b$). In that case, the differential equation for an equilibrium in the region $s > a$ becomes

$$\tilde{S}'(a) = \frac{b - s}{br} \frac{1}{a},$$

and with $\tilde{S}(a_0) = S^*(a_0)$, its unique solution is

$$\tilde{S}(a) = b + (1 - b)a^{-\frac{1}{br}}.$$

The equilibrium profile $S(a)$ is graphed in panel B for $b = 14$. Note that $S(a)$ first equals $S^*(a)$ before $a = 14$, the point at which information is perfect. This is because the potential wage increase from additional education, here an entirely non-productive signal, would not justify its cost.

C. Race-Based Statistical Discrimination

In order to analyze race-based statistical discrimination, we apply our signaling model to white and black workers, with parameters that differ between the two groups. The literature on statistical discrimination suggests that firms observe the productivity of low-education blacks less accurately than that of low-education whites. This generalization can be justified on three grounds. First, networks are

extremely important in the market for low-skill jobs (Judith K. Hellerstein, Melissa McInerney, and David Neumark 2008); and because race and neighborhood affect networks, blacks have poorer networks than do otherwise comparable whites (Michael A. Stoll, Stephen Raphael, and Harry J. Holzer 2004; Bayer, Stephen Ross, and Giorgio Topa 2005).

Second, considerable research shows that blacks and whites use different nonverbal listening and speaking cues, which can lead to miscommunication (Lang 1986). While nonverbal code-switching is common among educated blacks who are “bilingual,” low-skill blacks are more likely to be confined to listening and speaking cues that are dominant in the black speech community. Moreover, blacks who use African American Vernacular English (AAVE) in survey interviews rather than Standard American English (SAE) receive lower wages than other apparently similar blacks (Jeffrey Grogger 2008). The use of AAVE may interfere with the communication of information about productivity (the mechanism on which we focus) or may be perceived as a signal of low productivity (a factor not included in our model).

Third, in the literature on learning in the labor market (Henry S. Farber and Robert Gibbons 1996; Joseph G. Altonji and Charles R. Pierret 2001), authors assume that the market cannot observe AFQT directly but that AFQT is related to productivity. The coefficient on AFQT in a wage equation is therefore an indicator of how well the market knows productivity. Joshua C. Pinkston (2006) finds that, unlike whites, blacks entering the labor market are not rewarded for their AFQT, but that the coefficient on AFQT rises more rapidly with experience for blacks than for whites. Moreover, we show below that AFQT does not affect the earnings of black school dropouts even after considerable experience in the labor force, though it does affect the earnings of analogous whites.³

Since firms, in our model, observe an applicant’s race, the differences in the accuracy of productivity observations induce firms to put more weight on education and less on observed productivity for black workers than for white workers. Therefore, education is a more valuable signal of ability for blacks than it is for whites, which leads blacks to obtain more education than whites of equal ability. It follows that at any level of education, blacks will be of lower ability and have lower wages. However, at any level of ability, since blacks get more education, they should have higher wages if we do not hold education constant. We derive these results formally below.

Let the subscripts b and w denote black and white workers. As before, let (s_0, s^*) represent the interval of educational attainment over which employers are uncertain about productivity (i.e., for $s_0 < s < s^*$, $\lambda(s) < 1$), and let (a_0, a^*) be the corresponding interval of ability in equilibrium. If black productivity is observed less accurately than white productivity, then $\lambda_b(s) < \lambda_w(s)$ there. The following proposition shows that blacks will get more education than whites of equal ability for all such intermediate ability levels.

PROPOSITION 7: *Given $\lambda_b(s) < \lambda_w(s)$ for all $s < s^*$, we have $S_b(a) > S_w(a)$ for all $a \in (a_0, a^*)$ in equilibrium.*

³ Arcidiacono, Bayer, and Hizmo (2008) conclude that the market learns the productivity of black and white high school graduates at similar rates. Above, we assume that λ reaches one for blacks and whites at the same level of educational attainment. Our argument would be similar if the λ for blacks and whites converged at a value less than one.

PROOF:

From (13) we know that for $\lambda_i(s) < 1$, the equilibrium S_b and S_w are characterized by

$$(20) \quad S'_i(a) = \frac{(1 - \lambda_i(s)) q_a(s, a)}{r - q_s(s, a)},$$

where i is either b or w . If for $s < s^*$ blacks and whites have the same values of a and s , then from $\lambda_b(s) < \lambda_w(s)$ we know that $S'_b(a) > S'_w(a)$. By the continuity of S_b and S_w and the fact that $S_b(a_0) = S_w(a_0)$, we can infer that $S_b(a) > S_w(a)$ in a neighborhood of a_0 . If \hat{a} is the smallest value of a greater a_0 at which $S_b(a) = S_w(a)$, it must be true that $S'_b(\hat{a}) \leq S'_w(\hat{a})$, because $S_b(a)$ is converging to $S_w(a)$ from above. But by (20), this is possible only if $\lambda_b(s) = \lambda_w(s)$, which implies that $\hat{a} = a^*$. The proposition follows.

We can now show that at any education level (except the lowest) at which black productivity is observed less accurately than white productivity, the expected equilibrium earnings of blacks are less than those of whites with the same level of education.

PROPOSITION 8: *In equilibrium, for $s \in (s_0, s^*)$, $E_{\varepsilon,u}[\hat{w}_b(p, s)] < E_{\varepsilon,u}[\hat{w}_w(p, s)]$.*

PROOF:

Equation (10) implies that in equilibrium we have

$$(21) \quad \ln E_{\varepsilon,u}[\hat{w}_i(p, s)] = \lambda_i(s) q(s, A_i(s)) + (1 - \lambda_i(s)) \hat{q}_i(s) + 0.5\sigma_\varepsilon^2,$$

which reduces to

$$(22) \quad \ln E_{\varepsilon,u}[\hat{w}_i(p, s)] = \hat{q}_i(s) + 0.5\sigma_\varepsilon^2,$$

because $\hat{q}_i(s) \equiv q(s, A_i(s))$. From the previous proposition, we know that $\hat{q}_b(s) < \hat{q}_w(s)$ for $s \in (s_0, s^*)$ and the proposition follows.

D. Empirical Implications of the Model

Our model generates the following predictions:

- (i) Blacks with $a_0 < a < a^*$, that is, those with low (except for the very lowest) or intermediate levels of ability obtain more education than their white counterparts. Blacks with $a \geq a^*$ (high ability) obtain the same levels of education as comparable whites. Thus, overall blacks get more education than do whites of similar ability.
- (ii) Let $E_b(w(s))$ and $E_w(w(s))$ be the mean wage of black and white workers with education s . Then,

$$(23) \quad \frac{E_b(w(s)) - E_b(w(s_0))}{s - s_0} < \frac{E_w(w(s)) - E_w(w(s_0))}{s - s_0}$$

and

$$(24) \quad \frac{E_b(w(s^*)) - E_b(w(s))}{s^* - s} > \frac{E_w(w(s^*)) - E_w(w(s))}{s^* - s}$$

for all $s_0 < s < s^*$. This means that the measured return to education, not controlling for ability, should be lower for blacks than for whites at low levels of education and higher at higher levels. Of course, this conclusion refers to the measured return. Any worker's expected marginal private return to education is the common interest rate r .

- (iii) Since, relative to whites with the same ability, blacks with intermediate levels of ability get more education, we have $E_b(w(a)) > E_w(w(a))$ for all $a_0 < a < a^*$ and $E_b(w(a)) = E_w(w(a))$ for $a = a_0$ or $a \geq a^*$. In other words, at intermediate ability levels, blacks earn more than whites do, although if the market is otherwise perfect, they are worse off because they overinvest in education to a greater extent than do whites. At low and high ability levels, blacks and whites have similar earnings. Thus, the return to ability (not controlling for education) should be higher for blacks than for whites at low levels of ability and lower for blacks than for whites at higher ability levels.

II. Data

Although our initial focus is on differences in educational attainment, not wages, later in the paper we compare our results with those of Neal and Johnson. Therefore, to a large extent, we mimic their procedures. Following NJ, we rely on data from the National Longitudinal Survey of Youth (NLSY79). Since 1979 the NLSY has followed individuals born between 1957 and 1964, annually at first, and every other year more recently. In the period we use, the NLSY oversamples blacks and Hispanics, but oversamples of people from poor families and the military had been discontinued. We use sampling weights to generate representative results.

Education is measured by the highest grade completed as of 2000. For those missing the 2000 variable, we used highest grade completed as of 1998, and for those missing 1998 as well, we used the 1996 figure. Where available, we used the 2000 weight. For observations missing the 2000 weight, we imputed the weight from the 1996 and 1998 weights using the predicted value from regressions of the 2000 weights on the 1996 and/or 1998 weights.

We determined race and sex from the subsample to which the individual belongs. Thus, all members of the male Hispanic cross-section sample were deemed to be male and Hispanic regardless of how they were coded by the interviewer.

In 1980, the NLSY administered the Armed Services Vocational Aptitude Battery (ASVAB) to members of the sample. A subset of the ASVAB is used to generate the AFQT score. The AFQT is commonly viewed as an aptitude test comparable to other measures of general aptitude or ability. Like other such measures, it reflects both environmental and hereditary factors. The AFQT was recalibrated in 1989. The NLSY data provide the 1989 AFQT measure. Following NJ, we regressed the AFQT score on age (using the 1981 weights) and adjusted the AFQT score by subtracting

age times the coefficient on age. We then renormed the adjusted AFQT to have mean zero and variance one.

In the later part of the paper, we also examine wages. Because addressing differential labor force participation of black and white women is difficult (Neal 2004), we limit our estimates to men when we examine wages. To minimize missing data, we used hourly earnings from the 1996, 1998, and 2000 waves of the survey. Next we took all observations with hourly wages between \$1 and \$100 in all three years and calculated (unweighted) mean hourly earnings for this balanced panel. We used the average changes in hourly wages to adjust 1996 and 2000 wages to 1998 wages. Note that this adjustment includes both an economy-wide nominal wage growth factor and an effect of increased experience. We then used the adjusted 1996, 1998, and 2000 wages for the entire sample to calculate mean adjusted wages for all respondents.

We limited ourselves to observation/years in which the wage was between \$1 and \$100. If the respondent had three valid wage observations, we used the mean of those three. If he had two observations, we used the average of those two. For those with only one observation, the wage measure corresponds to that adjusted wage. There were 237 men who were interviewed in at least one of the three years but did not have a valid wage in any of those years. In the quantile regressions, we impute low wages for these individuals except for five cases coded as missing, for which the reported wage in at least one of the years exceeded \$100 per hour and for which there was no year with a valid reported wage.

III. Differences in Educational Attainment

Most labor economists know that average education is lower among blacks than among whites. In our sample blacks get about three-quarters of a year less education than do whites. It is less well known that conditional on AFQT, blacks get more education than whites do. As shown in the first row of Table 1, black men get about 1.2 years more education than do white men with the same AFQT.⁴ Among women, the difference is about 1.3 years. There are also smaller but statistically significant differences between Hispanics and non-Hispanic whites (not shown).

A. School Quality

The difference between blacks and whites in educational attainment conditional on AFQT is predicted by our theoretical model. But, as always, there are many other potential explanations for this finding. One is that AFQT is largely determined by schooling, and that since blacks attend lower-quality schools, they gain fewer cognitive skills on average from a given level of education. Under this view, blacks have more schooling given their AFQT because they require more schooling to reach a

⁴Steven G. Rivkin (1995) finds that conditional on math and reading scores, blacks are more likely to remain in high school and begin college. Stephen V. Cameron and James J. Heckman (2001) also use the NLSY and find that blacks get more education than whites conditional on measures of family background, and note that AFQT has a particularly strong effect on reversing the education differential.

TABLE 1—EDUCATIONAL ATTAINMENT OF BLACKS RELATIVE TO NON-HISPANIC WHITES

Birth years	Other controls	Men		Women	
		Black-white gap	N	Black-white gap	N
All		1.17 (0.10)	4,060	1.30 (0.09)	4,337
All	School inputs	1.16 (0.13)	2,302	1.25 (0.14)	2,326
All	School composition	1.11 (0.16)	2,336	1.29 (0.16)	2,385
All	Family background	1.20 (0.11)	3,323	1.40 (0.10)	3,558
All	School inputs, school composition, family background	1.16 (0.20)	1,603 (0.20)	1.32	1,618
1962+	Grade in 1980	0.92 (0.14)	1,719	1.21 (0.15)	1,665
1963+	Grade in 1980	0.94 (0.18)	1,106	1.30 (0.19)	1,054
1964+	Grade in 1980	0.72 (0.26)	508	1.43 (0.26)	474
1962+		0.98 (0.14)	1,737	1.29 (0.15)	1,683
1963+		0.99 (0.18)	1,116	1.35 (0.19)	1,062
1964+		0.78 (0.26)	514	1.40 (0.26)	478
1962+	Grade in 1980, school inputs	0.87 (0.21)	913	1.26 (0.23)	862
1962+	Grade in 1980, school composition	1.04 (0.25)	914	1.26 (0.26)	889
1962+	Grade in 1980, family background	1.01 (0.17)	1,385	1.36 (0.17)	1,355
1962+	Grade in 1980, school inputs, school composition, family background	1.05 (0.31)	630	1.30 (0.33)	592

Notes: Standard errors in parentheses. All estimates control for age. School inputs: In enrollment, In no. of teachers, In no. of guidance counselors, In library books, percent teachers with MA/PhD, percent teachers left during the year, average teacher salary. School composition: percent disadvantaged, daily attendance rate, dropout rate, percent students Asian, percent students black, percent students Hispanic. Family background: mother's education, father's education, no. of siblings, born in US, lived in US at age 14, lived in urban area at age 14, mother born in US, father born in US.

given level of cognitive skills. If so, when we regress education on AFQT, we are, in effect, estimating a reverse regression.

To take a simple example, suppose that AFQT is proportional to quality of education multiplied by years of education (ability does not matter). Then, if school quality is lower for blacks than for whites, a black and a white could have the same AFQT only if the black had more education years. Therefore, when regressing education on AFQT, it might appear, on the one hand, that failing to control for school quality biases the coefficient on *black* upward. On the other hand, failing to include school quality should be to generate a downward bias on *black*. In standard theoretical models the sign of the effect of school quality on years of schooling is ambiguous. The data, however, suggest a positive correlation between school quality and years of schooling (e.g., David Card and Alan B. Krueger 1992a, b; Eric A. Hanushek, Victor Lavy, and Kohtaro Hitomi 2006; and the discussion in Hanushek and Ludger

Woessmann 2008). In this case, the lower school quality faced by blacks should lead them to get fewer, not more, years of schooling than whites with the same ability.

To summarize, if AFQT is heavily influenced by education and if most sample members had completed their education at the time that they took the AFQT, then school quality differences would provide a plausible explanation for the higher education among blacks given their AFQT. If school quality has little effect on AFQT or if most sample members had not completed schooling, then we would expect blacks to get less education given their AFQT (or given their AFQT and educational attainment at the time they took the test). Our view is that the AFQT measures skills that are more heavily affected by preadolescent and early adolescent education so that the endogeneity of AFQT to ultimate educational attainment and the consequent link to school quality is not likely to be a major issue. However, others certainly disagree. Therefore we address the question empirically.

Our first approach is to measure the education differential conditional on secondary school inputs (log enrollment, log number of teachers, log number of guidance counselors, log library books, proportion of teachers with an MA or PhD, proportion of teachers who left the school during the year, average teacher salary). Almost none of the individual coefficients is statistically significant (see the online Appendix). Among men, attending a school with more highly educated teachers is associated with greater educational attainment. Among women this variable and attending a school with more library books is associated with greater educational attainment. In part, the paucity of individually significant factors reflects multicollinearity among the measured inputs. For both men and women, the coefficients on the school inputs are jointly significant. More important, as shown in the second row of Table 1, controlling for these factors has almost no effect on the estimated education gaps, and some of the change is due to a change in the sample rather than additional controls.

Because inputs may be a very poor proxy for school quality, we control for school composition and student behavior (percent disadvantaged students, dropout rate, attendance rate, racial and ethnic composition of the school) in the third row of Table 1. These variables are designed to capture some of the elements that people think about when they think about struggling schools. The results are very similar to those in the first two rows. Moreover, consistent with Cameron and Heckman (2001), as shown in the fourth row, the findings are robust to including measures of family background (mother's and father's education, number of siblings, whether each parent was born in the United States, respondent born in the United States, lived in an urban area at age 14). Finally, the fifth row provides our most complete set of controls using the full sample. Controlling simultaneously for school inputs, school composition and family background costs us about 60 percent of the sample but has a negligible effect on the black/white education differential.

B. *Younger Cohorts*

The remaining rows of Table 1 restrict the sample to younger cohorts. In this way, we control for the endogeneity of AFQT by limiting the sample to individuals who would not have completed their education at the time the NLSY administered the AFQT. Regardless of whether AFQT is endogenous to education, for these cohorts we have a measure of ability prior to their completion of schooling. We experiment

with different age cutoffs and with using only AFQT and both AFQT and educational attainment at the time of testing as “ability” controls.

Only about 5 percent of those born after 1961 had left school when they took the AFQT. While their education up to this point may have influenced their AFQT, skills acquired up to this point should affect future education, and not vice versa. The estimated differential is somewhat smaller using the younger cohorts, particularly for men, but remains substantial and is not greatly affected by whether we control for grade completed at the time of testing, measures of school quality, or family background. The last specification is the most restrictive we estimate. It limits the sample to those born in 1962 or later and controls for school inputs, school composition, and family background. The estimated differentials are very close to those reported for the full sample with no controls.

In short, there is a robust result that among men of equal ability as measured by the AFQT, blacks get a year or more education than do non-Hispanic whites. Among women this differential is about 1.25 years.

Before we move on, it is important to make it clear what we are *not* claiming. As stated in the introduction, we are also not claiming that AFQT is innate or even unaffected by education and school quality. We are also not claiming that school quality is unrelated to educational attainment or that school quality does not differ between blacks and whites. To the contrary, we believe that average school quality is lower for blacks and that individuals who attend higher quality schools both have higher AFQTs and get more education. It is beyond the scope of the paper to address whether these last two relations are causal. From our perspective, however, the simplest and most probable explanation for our results is that the effect of school quality on AFQT and the effect of school quality on educational attainment roughly cancel so that, given educational attainment, AFQT is roughly independent of school quality.

C. Affirmative Action in Education

Finally, we address an additional explanation frequently raised in seminars: affirmative action in college admissions. Since affirmative action is practiced at only a small number of elite colleges (Thomas J. Kane 1998), it cannot account for the large difference in educational attainment we observe. Moreover, we show in the next section that the education difference is found even at lower AFQT scores unlikely to be associated with students attending such elite colleges. In fact, the education gap disappears at AFQT scores above roughly 1.5 standard deviations above the mean, the level at which affirmative action might play a role.

IV. Further Evidence

While our model predicts that, conditional on ability, blacks will on average get more education than whites, also it makes stronger predictions, that we examine in this section.

A. Education and Ability

Our model suggests that blacks should get more education than whites at intermediate levels of ability but not at very low or very high levels of ability. (Note the

TABLE 2—REGRESSIONS OF EDUCATION ON AFQT WITH RACE/ETHNICITY INTERACTIONS, BY SEX

	Men			Women		
	All	Young cohort		All	Young cohort	
Constant	12.15 (0.23)	13.82 (0.77)	14.47 (0.77)	12.10 (0.23)	12.62 (0.80)	13.41 (0.80)
AFQT	1.64 (0.04)	1.62 (0.06)	1.44 (0.07)	1.67 (0.04)	1.70 (0.08)	1.43 (0.09)
AFQT ²	0.56 (0.04)	0.42 (0.06)	0.47 (0.06)	0.31 (0.04)	0.23 (0.07)	0.36 (0.07)
Black interactions						
Constant	1.29 (0.13)	1.02 (0.19)	0.96 (0.19)	1.40 (0.12)	1.32 (0.19)	1.19 (0.19)
AFQT	-0.13 (0.12)	-0.18 (0.19)	-0.16 (0.19)	-0.01 (0.13)	0.03 (0.21)	0.13 (0.21)
AFQT ²	-0.41 (0.10)	-0.26 (0.16)	-0.29 (0.16)	-0.28 (0.11)	-0.16 (0.20)	-0.14 (0.20)
Interaction equals 0	-1.94, 1.63	-2.34, 1.64	-2.13, 1.57	-2.24, 2.22	-2.73, 2.93	-2.51, 3.45
Hispanic interactions						
Constant	0.68 (0.17)	0.29 (0.26)	0.30 (0.25)	1.00 (0.15)	0.72 (0.26)	0.77 (0.26)
AFQT	0.09 (0.13)	0.02 (0.21)	0.04 (0.21)	0.02 (0.15)	-0.06 (0.25)	-0.05 (0.25)
AFQT ²	-0.48 (0.12)	-0.06 (0.21)	-0.08 (0.21)	-0.66 (0.12)	-0.38 (0.22)	-0.45 (0.23)
Interaction equals 0	-1.10, 1.29	-2.07, 2.41	-1.66, 2.11	-1.21, 1.24	-1.46, 1.30	-1.36, 1.25
N	4,060	1,733	1,715	4,337	1,687	1,669
Other controls	Age	Age	Age, education in 1980	Age	Age	Age, education in 1980

Notes: Standard errors in parentheses. Interaction equals zero solves the quadratic $a + bAFQT + cAFQT^2 = 0$, where a is the race/ethnicity constant term (e.g., 1.29), b is the coefficient on the race/ethnicity interaction with AFQT (e.g., -0.13), and c is the race/ethnicity interaction with $AFQT^2$ (e.g., -0.41). Weights for education results are described in the text.

contrast between this prediction and an explanation based on affirmative action: the latter would suggest that the education differential would be found only at relatively high levels of ability.) Table 2 shows the relation between education and AFQT with interaction terms between race/ethnicity and a quadratic in AFQT. Results are shown separately for men (the left panels) and women (the right panels). Within each sex, the younger cohorts, who had not completed school at the time that they took the AFQT, are shown separately in the second and third columns of each panel, both without (the second column) and with (the third column) a control for educational attainment at the time they took the test. In each column, therefore, the specification takes the form

$$\begin{aligned}
 s_i = & a_0 + a_1AFQT_i + a_2AFQT_i^2 + b_0black_i + b_1black_i \times AFQT_i \\
 & + b_2black_i \times AFQT_i^2 + c_0Hispanic_i + c_1Hispanic_i \times AFQT_i \\
 & + c_2Hispanic_i \times AFQT_i^2 + \mathbf{Z}_i\Gamma + \varepsilon_i,
 \end{aligned}$$

where s is educational attainment measured as highest grade completed in years and \mathbf{Z} is a vector including a small number of additional regressors.

In every specification, the interaction of race and the AFQT-squared term has its predicted negative sign. This is true for Hispanics as well as for blacks. Although the individual interaction terms tend to be statistically insignificant when we limit the sample to the younger cohorts, the differences between the younger and older cohorts are also insignificant at the 0.05 level. This is the case whether we test each coefficient individually, the three black interaction terms jointly, the three Hispanic interaction terms jointly, or the six interaction terms jointly. Thus the results are not driven by the causal impact of education on AFQT.

For all men, the black-white education differential is maximized at approximately 1.3 years at an AFQT about one-sixth ($-0.13/(2 \times 0.41)$) of a standard deviation below the mean. Educational attainment is equal (shown in the table as “interaction equals 0”) for blacks and whites at almost two standard deviations below the mean and at one-and-two-thirds standard deviations above the mean. For women, the black-white education differential is maximized at a value of approximately 1.4 years at very near the mean AFQT. The education levels of black and white women are estimated to be equalized near the extremes of the AFQT distribution. We note that the affirmative-action hypothesis would imply that the differences in education would be largest at the high levels of AFQT associated with application to selective colleges. Our results are not consistent with that explanation.

Figure 2 shows the smoothed relation between education and AFQT for men. The nonparametric approach confirms the parametric approach.⁵ Education levels for blacks and whites converge around a standardized AFQT of -2 and a little above 1.5. Figure 3, analogous to Figure 2 but for women, is less consistent with the parametric estimates. It shows that education levels converge at a standardized AFQT between -2 and -2.5 . However, education levels for black women remain higher than for white women even at very high AFQT levels. One potential explanation for this difference is the very high rate of labor-force participation of high-skill black women relative to white women, discussed in Neal (2004).

Figures 4 and 5 show the distribution of education for selected AFQT deciles (all deciles are described in the online Appendix). Among those in the lowest decile, blacks are more likely to graduate high school and less likely to drop out. This difference comes from those in the fifth to tenth percentiles. Black and white men below the fifth percentile of the AFQT distribution have similar education distributions. Black women in the lowest decile are more likely than their white counterparts to attend college and less likely to drop out. By the fourth and seventh deciles, both black men and black women are noticeably more likely than are whites to attend college. In the highest decile, there is no observable difference between the education distributions of black and white men. Black women appear more likely to get postgraduate education, although neither the difference in mean education nor the distribution (as tested using Goodman and Kruskal’s gamma) is significant at conventional levels.

⁵The figures use unweighted data and do not control for age. We could have controlled for age by plotting the residuals of regressions of education on age and those of AFQT on age against each other. However, except for a change in the scale of education, the plots would be visually indistinguishable from those we show. Our plots present local weighted sum of squares estimates (the Stata “lowess” command) and a bandwidth of 30 percent of the observations.

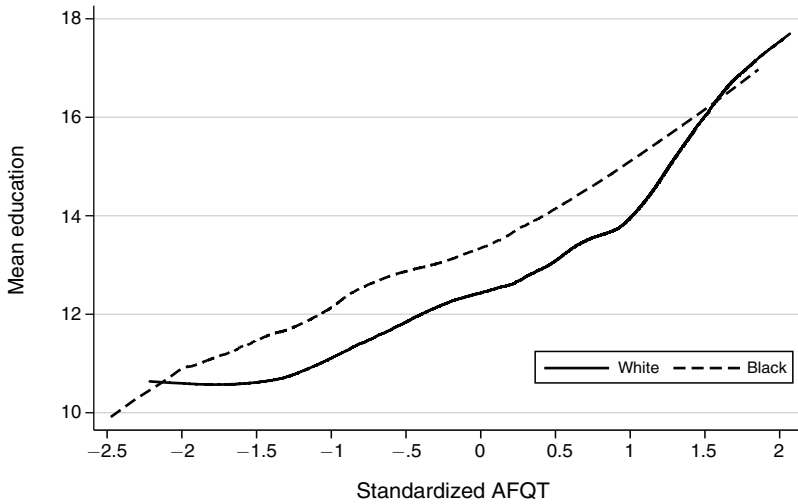


FIGURE 2. EDUCATION AND AFQT BY RACE: MEN

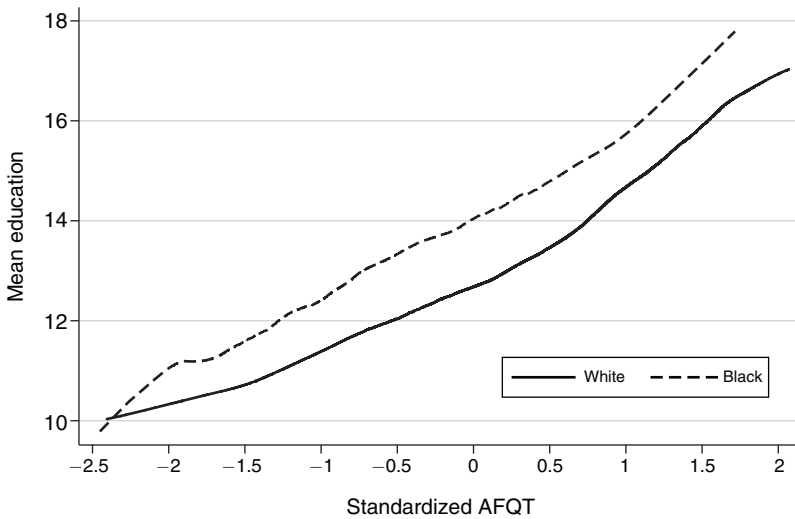


FIGURE 3. EDUCATION AND AFQT BY RACE: WOMEN

B. Education and Earnings

Because of the differences between the labor-force participation of black and white women, we restrict our discussion of wage predictions to men. Our model implies that the wages of blacks and whites are similar at low and high levels of education, but blacks will receive lower wages at intermediate levels. To test this prediction, we regress the log wage on education and its square and interactions with race and ethnicity, as well as direct effects of age, race, and ethnicity. Table 3 shows the results.

For all four specifications (all/young cohorts, with/without controls), as predicted, the return to education is initially lower for blacks than for whites and then

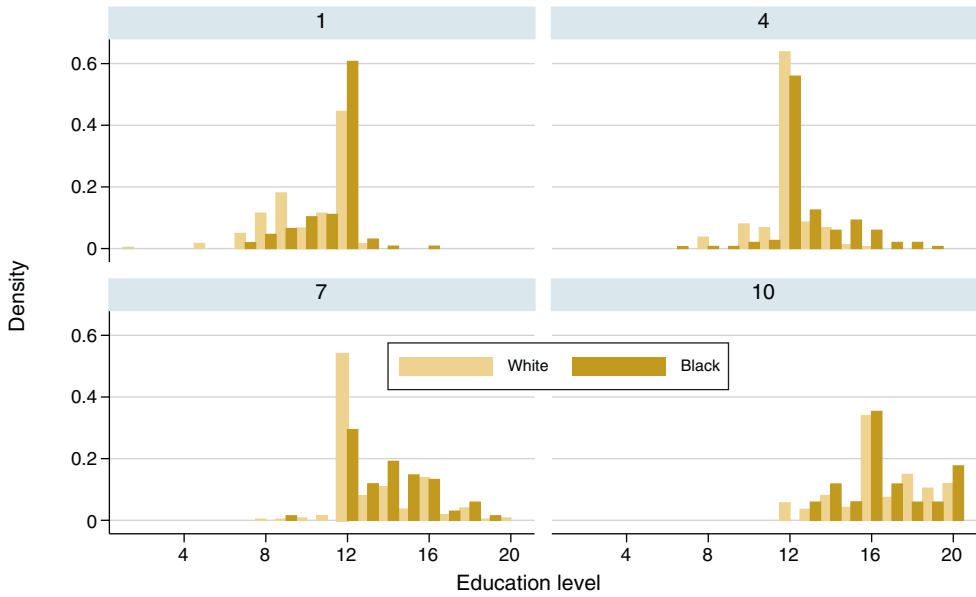


FIGURE 4. EDUCATION DISTRIBUTION BY AFQT DECILE: MEN

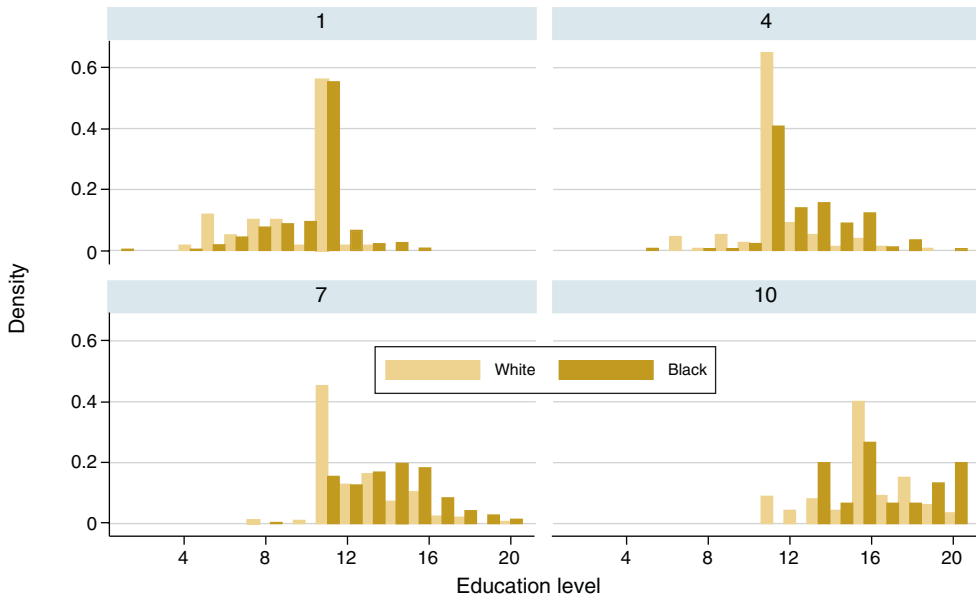


FIGURE 5. EDUCATION DISTRIBUTION BY AFQT DECILE: WOMEN

turns more positive. With the full sample and no controls, wages for blacks and whites are estimated to be equal for those with a fifth-grade education and for those with 19 years of completed education, although these points of equality are estimated very imprecisely. In the most restrictive specification (only the young cohorts and a complete set of controls), we estimate that wages are equal for blacks and

TABLE 3—BLACK-WHITE EARNINGS DIFFERENTIALS BY EDUCATIONAL ATTAINMENT FOR MEN

	All		Young cohort	
	No controls	Controls	No controls	Controls
Black	0.51 (0.48)	1.20 (1.26)	1.83 (0.91)	3.27 (2.14)
Black × education	-0.14 (0.07)	-0.24 (0.18)	-0.34 (0.14)	-0.54 (0.31)
Black × education ² /100	0.59 (0.26)	0.98 (0.66)	1.30 (0.49)	2.12 (1.14)
Grade at which wages are equal	4, 19	7, 17	8, 18	10, 16
N	4,041	1,601	1,661	607

Notes: Standard errors in parentheses. Also controls for age, school inputs, school composition, and family background (see Table 1 for details) and, for young cohorts only, grade in 1980.

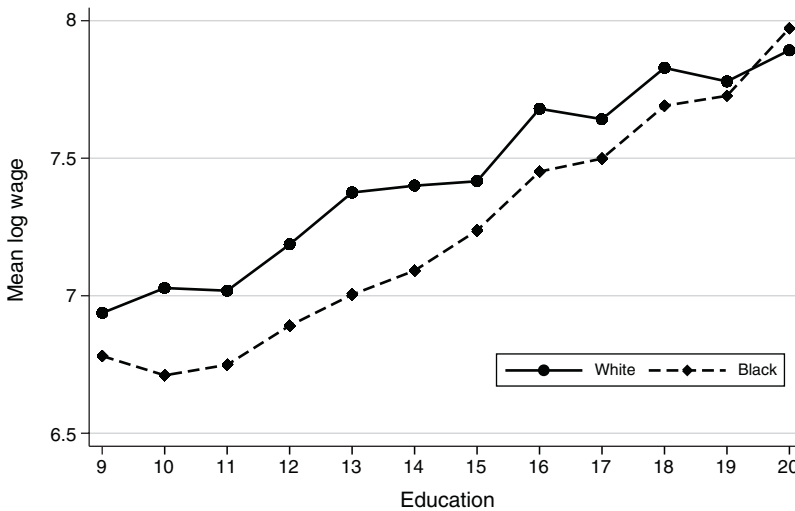


FIGURE 6. EDUCATION AND WAGES BY RACE: MEN

whites who have completed college or who have completed only grade 10. However, these estimates are even more imprecise than those using the whole sample and no controls.

Figure 6 shows the relation between education and earnings nonparametrically for the full sample. It plots average log wages for men by education and race. Few individuals have no high school education and few blacks have more than 18 years of education. As predicted by the model, wages are similar for blacks and whites at low and high levels of education.

C. Wages and AFQT

Our model also predicts that blacks should earn more than equally able whites, except at very high and very low levels of ability. The test of this prediction is shown in Table 4. The table follows the same format as Table 3. With no controls, this prediction is soundly rejected both for all cohorts and for the young cohorts. In

TABLE 4—BLACK-WHITE EARNINGS DIFFERENTIALS BY AFQT SCORE

	All		Young cohort	
	No controls	Controls	No controls	Controls
Black	-0.08 (0.03)	-0.05 (0.07)	-0.14 (0.05)	-0.05 (0.10)
Black \times AFQT	0.04 (0.03)	0.03 (0.05)	0.05 (0.05)	-0.02 (0.08)
Black \times AFQT ² /100	0.60 (2.60)	0.02 (4.66)	2.65 (4.04)	0.87 (8.03)
F-test on all 3 coefficients	4.83	0.47	5.18	0.11
F-test on interaction terms	1.12	0.22	0.58	0.04
<i>N</i>	3,841	1,534	1,634	600

Notes: Standard errors in parentheses. Also controls for age, school inputs, school composition, and family background (see Table 1 for details) and, for young cohorts only, grade in 1980.

both cases, there is no evidence of an interaction between race and AFQT, although the point estimates suggest that black and white wages are equalized at an AFQT about 1.4 standard deviations above the mean. Below that level, blacks earn less, not more, than whites with the same AFQT. When we add controls for family background, all of the coefficients on black and its interaction terms become individually and jointly insignificant. Nevertheless, the estimates using all cohorts imply that the black-white wage differential declines from about 10 percent at an AFQT two standard deviations below the mean to zero at 1.4 standard deviations above the mean. The estimates using only the young cohorts and controls are too imprecise to be meaningful. If one accepts the results with controls, then they are consistent with the model when education is a pure signal at the margin. If, instead, we rely on the results without controls, then the model must be supplemented with some other explanation for wage differentials.

Labor economists generally agree that education is rewarded in the labor market. This implies that in the absence of labor-market discrimination, blacks would earn more than whites with the same AFQT. Given the education differential, the absence of a wage differential favoring blacks when we control only for AFQT suggests that blacks are not rewarded fully for their skills, a point to which we now turn.

V. Premarket versus Market Discrimination

There is a heated debate among labor economists about the extent to which black-white wage differentials can be ascribed to premarket factors (including discrimination outside the labor market) rather than to labor market discrimination.⁶ One of the critical issues in this debate is for which factors we should control.

Table 5 presents OLS estimates of wage differentials in the tradition of this research. In the first column, we use only those cohorts born in 1962 or later; in the second column, we use all cohorts. In no case are the differences between estimates

⁶In addition to the papers by O'Neill and Neal and Johnson discussed in the introduction, see Johnson and Neal (1998) and the critiques in William M. Rodgers III and William E. Spriggs (1996, 2002) and William A. Darity Jr. and Patrick L. Mason (1998), and the reply by Heckman (1998).

TABLE 5—BLACK-WHITE WAGE DIFFERENTIALS

Young cohort	All	Median regression	AFQT	Other controls	
				Education	Family background, school input
−0.37 (0.04) 1,634	−0.36 (0.03) 3,841	−0.42 (0.03) 4,055			
−0.13 (0.03) 1,634	−0.09 (0.02) 3,841	−0.10 (0.03) 4,055	✓		
−0.18 (0.03) 1,634	−0.15 (0.02) 3,841	−0.18 (0.03) 4,055	✓	✓	
−0.07 (0.06) 731	−0.06 (0.04) 1,876	−0.05 (0.03) 1,955	✓		✓
−0.11 (0.06) 731	−0.11 (0.04) 1,876	−0.11 (0.04) 1,955	✓	✓	✓

Notes: Standard errors in parentheses. All estimates control for age. Estimates for young cohorts control for education completed in 1980.

for the restricted and full samples statistically significant at even the 0.1 level, and the substantive interpretations of the results are similar. Therefore, we concentrate on the more precisely estimated results in the second column and remind readers concerned by the potential endogeneity of AFQT to employment or schooling that the results for the younger cohorts are similar. In the third column, we present the results of median regressions using the full sample. This addresses selection issues, since black men are noticeably more likely than are white men not to have a wage.

The first row shows the very large differential that exists when we control only for age. However, we focus on the wage differential after we control for ability in the form of AFQT and its square. This is shown in the second row. Consistent with our findings in Table 4 that the race-AFQT interaction terms are statistically insignificant, we drop these terms in Table 5, which simplifies interpretation. The second row suggests much more modest wage differentials, although they are not trivial and are somewhat higher than in Neal and Johnson's study of the younger cohorts in the early 1990s.⁷

In the fourth row, we also control for family background and school inputs.⁸ The estimated wage differential becomes noticeably smaller and statistically insignificant in all three specifications. Somewhat surprisingly, the median differential is slightly smaller than that obtained using a standard regression.

⁷Derek Neal was very helpful, supplying us with the code to replicate his and William Johnson's results. The modest difference in our results derives from a number of differences including NJ's use of the "class of worker" variable and our use of a later time period. Pedro Carneiro, Heckman, and Dimitriy V. Mastrov (2005) explore the issue of time variation in the black-white wage differential using various specifications, including those used by NJ. See also the discussion of this issue in Steven Haider and Gary Solon (2006).

⁸We dropped school composition after pretesting showed that these variables were highly insignificant but that their inclusion reduced the sample sufficiently to noticeably increase the imprecision of the estimate of the black-white wage differential. The decision to include or exclude school composition has no effect on the interpretation of Table 4.

The important point in Table 5, however, is the comparison of the second and third rows and of the fourth and fifth rows. In each case the latter specification is identical to the former except that it also controls for educational attainment. In each case the estimated differential increases substantially when we include education in the equation.⁹

Although Neal and Johnson explore the effect of controlling for education,¹⁰ they explicitly reject including education in their main estimating equation. They provide two arguments for their position. First, they maintain that we should examine black-white wage differentials without conditioning on education, because education is endogenous. Their argument would be much more compelling if blacks obtained less education than equivalent whites. In that case, we might argue that blacks get less education because they expect to face discrimination in the labor market, so that controlling for education understates the importance of discrimination.

If blacks obtain *more* education, however, because they anticipate labor market discrimination, *failing to control* for education understates the impact of discrimination. Consider the following example: the market discriminates against blacks by paying them exactly what it would pay otherwise equivalent whites with exactly one less year of education. Then, to a first approximation,¹¹ all blacks will get one year more education than otherwise equivalent whites. Controlling only for ability, blacks and whites will have the same earnings, but controlling for education as well as ability, blacks will earn less than whites by an amount equal to the return to one year of education. Even if the higher educational attainment among blacks reflects premarket factors, it may still be appropriate to control for education when measuring discrimination in the labor market. After all, we would still expect the labor market to compensate blacks for their additional education.

The second argument that NJ make is that education is a poor proxy for skills. In particular, on average, blacks attend lower-quality schools than do whites. Whites will have more effective education than do blacks with the same nominal years of completed education. We have already noted that this argument is incomplete. Holding ability constant, students who attend lower-quality schools tend to get less education. Therefore, if blacks attend lower-quality schools, they will have higher ability for any given level of education. If in the regression of wages on education and AFQT, AFQT only partially controls for ability, blacks will tend to have higher ability than do whites with the same education and AFQT. If the selection effect outweighs the direct effect of higher quality education on earnings, the coefficient on *black* would be spuriously positive; otherwise, it would be spuriously negative.

The last row in Table 5 shows that controlling for both family background and school inputs somewhat reduces the estimated differential relative to the estimate in the third row (which controls for education but not for these additional variables). Although this seems to suggest that it is important to control for school quality,

⁹Carneiro, Heckman, and Masterov (2004) use a specification similar to that in row (4), but adjust AFQT for schooling completed at the time the respondent took the AFQT. They find similar results.

¹⁰In part because the return to education was lower in the period they studied, the results were similar to those based on their principal specification.

¹¹This statement is precise if all workers maximize the present discounted value of lifetime earnings, lifetimes are infinite, there are no direct costs of education, and the return to experience is zero.

TABLE 6—DETERMINANTS OF LOG WAGES USING CONTROLS FOR SCHOOL QUALITY: N&J WAGES

Inputs		Student composition/behavior	
Black	-0.14 (-0.03)	-0.14 (-0.04)	Black
Hispanic	-0.02 (-0.05)	-0.01 (-0.06)	Hispanic
Age/10	0.13 (0.04)	0.14 (0.04)	Age/10
Education	0.06 (0.01)	0.06 (0.01)	Education
AFQT	0.14 (0.01)	0.15 (0.01)	AFQT
log of enrollment	-0.08 (0.04)	-0.08 (0.06)	Proportion disadvantaged
log number of teachers	0.02 (0.05)	-0.05 (0.07)	Proportion daily attendance
log number of counselors	0.10 (0.04)	-0.10 (0.05)	Proportion dropout
log number of library books	0.01 (0.01)	0.08 (0.06)	Proportion students black
Proportion of teachers MA/PhD	0.17 (0.05)	-0.07 (0.10)	Proportion students Hispanic
Teacher salary \$0,000s	0.02 (0.01)	0.82 (0.43)	Proportion students Asian
Teacher who left/100	-0.30 (0.13)		
<i>N</i>	2,194	2,223	<i>N</i>

Note: Standard errors in parentheses.

the reduction in the black-white wage differential actually results from the family-background controls.

Table 6 shows the results when we control only for school quality as measured by either inputs or composition. The results are similar when we control for both simultaneously. Most of the coefficients have the anticipated sign. Holding other resources constant, larger schools are associated with lower wages. Holding enrollment constant, having more guidance counselors, more teachers, and more library books is associated with higher wages. Having more educated teachers and more highly paid teachers is associated with higher student earnings, while teacher turnover has a negative effect.

Yet, controlling for inputs has almost no effect on the measured black-white wage differential. The difference between the coefficients with and without school quality controls reflects differences in the sample rather than the effect of adding the controls. Using the observations for which we have school input measures, we obtain a coefficient on *black* of -0.14. At least as measured by inputs, differences in school quality do not account for the black-white wage differential.

The right side of Table 6 controls for measures of student composition and behavior. Perhaps surprisingly, this effort is in some ways less successful than the estimation using school inputs. While higher fractions of disadvantaged students and dropouts are associated with lower wages, average absenteeism and the fraction of students who are black are not. The results are again quite similar to those obtained without controls for school quality.

Thus, we find no evidence that school quality accounts for the wage and education differentials. Note that the absence of evidence for the role of these premarket factors does not require a causal interpretation of the relation between education quality and outcomes. Even if the school dropout rate predicted individual dropout only because it captures unmeasured characteristics of the individual, we would expect controlling for the dropout rate to lower the black-white education differential. The fact that it does not supports the view that such premarket differences do not explain the wage and education differentials.

It is also important to be clear about what we are *not* saying. Our results do not mean that school quality is unimportant or that it is uncorrelated with race. It is entirely possible (we think likely) that school quality is important and correlated with race but that its direct effect is offset by the negative correlation between unmeasured student quality and measured school quality among individuals with the same level of completed education.

Neither Table 5 nor Table 6 captures the full complexity of the relation between wages, on the one hand, and race, AFQT, and education, on the other. In the Appendix to their paper, Neal and Johnson estimate separate wage equations for blacks and whites. None of the coefficients differs significantly between blacks and whites, but the point estimates suggest that the wage gap declines with education and disappears for those with two or three years of college.

In specifications presented in the online Appendix, we explore the relation among wages, race, education, and AFQT more fully. We estimate the equivalent of the second specification in Table 5 by education category (dropout, high school graduate, some college, college graduate, more than college) and allow the effect of AFQT to differ by race, although we do not follow Neal and Johnson in allowing the other coefficients to differ by race.

Many of our standard errors are large, so point estimates of zero should be treated with caution. Nevertheless, our point estimates suggest a complex relation among the variables that is consistent with our theoretical model. Our point estimates suggest no wage differential between white and black college graduates regardless of AFQT. There is a notable wage differential between blacks and whites who are either high school graduates or have some college, and this differential is not strongly related to AFQT either. The point estimates also suggest a nontrivial wage advantage for blacks with postcollege education, but the estimates are highly imprecise and statistically insignificant.

Perhaps most strikingly, while the relation between AFQT and earnings is similar for blacks and whites at all other levels of education, increases in AFQT are associated with higher wages for white dropouts but not for black dropouts. This is consistent with our assumption that the market has more difficulty assessing the productivity of low-skill blacks than of low-skill whites, but we are surprised that this difference remains so long after the workers enter the labor market. Our point estimates for the entire sample imply that black and white high-school dropouts with a normalized AFQT of about -2.3 have similar wages, consistent with our prediction about the wages of the least skilled black and white workers.

VI. Discussion and Conclusion

While some of the principal predictions of our theory are consistent with the data, the combination of statistical discrimination and educational sorting that we discuss

cannot fully explain the data. Our model implies that, conditional on ability, relative to whites, blacks get more education than whites do, so that conditional on AFQT, blacks ought to earn more than whites. Neither our results nor any that we are aware of support that conclusion for men.

One explanation is that education is a pure signal at the margin, as in the model variant in which ability represents the capacity to be educated. In that variant, in order to signal their ability, most workers invest in education beyond the point at which it increases their productivity. However, we view this model as extreme.

Our model and the supporting evidence identifies statistical discrimination as one source of differences in outcomes for blacks and whites.¹² We have focused on only one effect: increased investment in the observed signal. Blacks may also invest less in unobservable skills as in Lundberg and Startz, which could lead to them have lower wages even conditional on AFQT. In addition, Marianne Bertrand and Sendhil Mullainathan (2004) find that applicants with African American names are less likely to receive calls for interviews than are similar applicants with names common among whites. If evaluating workers is costly, statistical discrimination may prevent large numbers of African American workers from consideration for many jobs. This suggests to us that statistical discrimination is particularly important in the presence of search frictions. We expect that in this setting our principal results would hold: blacks would have greater incentives to signal their productivity and would earn less conditional on their education. However, they might also earn less conditional on their ability.

In our view, the results in this paper cast doubt on an emerging consensus that the origins of the black-white wage differential lie in premarket rather than labor-market factors. Blacks earn noticeably less than whites with the same education and cognitive score. Controlling for measures of school quality (both inputs and composition) leaves this result unaltered. Although it is still possible that the earnings gap reflects differences between blacks and whites that are not captured by our controls, our results should lead us to question whether the gap is solely due to premarket factors and ask whether at least some of the black-white wage differential reflects differential treatment in the labor market.

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¹² Our results stand somewhat in contrast to Altonji and Pierret (2001), who find that the black-white differential does not decline with experience when one controls for hard-to-observe measures of productivity. However, our model implies that the estimated return to education changes differentially with experience for blacks and whites. This is permitted in Pinkston (2006), whose results are consistent with our prediction.

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