Eliminating Race Differences in School Attainment and Labor Market Success

Michael P. Keane, *New York University*

Kenneth I. Wolpin, *University of Pennsylvania*

In this article, we provide quantitative evidence on the effects of monetary incentive schemes designed to reduce racial differences in school attainment and earnings. Our analysis is based on the structural estimation of a dynamic model of schooling, work, and occupational choice decisions over the life cycle. We consider two recent proposals that, although not specifically targeted to blacks, can be expected to have differential racial impacts. One proposal, suggested by Robert Reich, provides a high school graduation bonus to youths from lower-income families. The other, suggested by Edmund Phelps, provides wage subsidies to low-wage workers.

I. Introduction

Between 1940 and 1980, the average school attainment of black men increased by about 1.5 years every decade, about .5 years more than for white men.[1] Similarly, over the same period real income more than quadrupled for black men, while increasing two-and-one-half times for white men.[2] However, considerable differences remain. In 1993, the percentage of 25–29-year-old black males who were college graduates was

Support for this research from National Institutes of Health grant HD30156 and from the University of Minnesota Supercomputer Institute is gratefully acknowledged.

[1] See table 9 in Smith and Welch (1989).
[2] Smith and Welch (1989), table 1.

[*Journal of Labor Economics*, 2000, vol. 18, no. 4]

0734-306X/2000/1804-0003$02.50

only half that of white males of the same age (12.6% relative to 24.4%), and earnings among 25–34-year-old black males were only 83% of that of white males of the same age.[3] There is still much debate over the causes of the relative improvement in black status (see Smith and Welch 1989; Donahue and Heckman 1991) as well as considerable uncertainty as to the efficacy of interventions that would ameliorate the remaining differences.

In this article, we provide quantitative evidence on the causes of these racial differences and on the effects of monetary incentive schemes designed to reduce them. Our analysis is based on the structural estimation of a dynamic model of schooling, work, and occupational choice decisions over the life cycle. The structural approach is exploited in two ways: (1) it is used to distinguish among a number of possible explanations for the current racial disparities in schooling and earnings and (2) it is used to assess the impact of "new" policy interventions designed to increase the relative schooling of blacks and to increase the relative earnings of blacks. The model allows for two sources of racial differences in schooling and labor market outcomes: differences in the price of market skills, that is, wage discrimination, and differences in unobserved "endowments" of market, nonmarket, and schooling (learning) skills measured at age 16. These endowments capture both innate traits as well as parental and own investments in skills up to age 16 in a way that is indistinguishable in our model.

Our estimation uses a cohort of 13–16-year-old black and white males drawn from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience (NLSY), which we follow for 11 years. Modeling unobserved heterogeneity as a mixture of finite types, our estimation identifies four distinct types for each race in terms of their skill endowment mix. We find that a specification in which blacks and whites differ only in their occupation-specific wage constants and in their proportionate representations among the four types cannot be rejected by the data. Blacks not only have lower wage constants, but there are relatively fewer blacks than whites among the high-market skill types. Given that, equalizing type proportions is found to account for the greatest difference in school completion levels (60% of the difference), in earnings (60%), and in welfare (65%). On the other hand, wage discrimination can account for at most about 30% of the completed schooling differential, 40% of the lifetime earnings differential, and 25% of the lifetime utility differential.

The structural framework permits an evaluation of the potential success of alternative interventions to influence schooling and labor market outcomes that have been proposed but not previously implemented. In

[3] Kominski and Adams (1994), tables 1 and 9, March 1993.

particular, we consider two recent proposals that, although not specifically targeted to blacks, can be expected to have differential racial impacts. One proposal by Robert Reich (1998), directed toward school attainment, provides a high school graduation bonus to youths from lower-income families. The other by Edmund Phelps (1997), directed toward labor market outcomes, provides wage subsidies to low-wage workers. Note that although each is directed toward a single outcome, in a dynamic, forward-looking model in which schooling and work are choices, each of these interventions will affect both outcomes.

Our main finding with respect to the school bonus is that it can induce substantial responses in school completion levels. However, because such an incentive scheme does not reduce the large estimated age-16 "endowment" differences that account for much of the racial difference in school attainment and subsequent labor market success, bonus programs can do little to reduce racial differences in lifetime earnings or welfare even if they are successfully targeted to black youths. On the other hand, wage subsidies of the magnitude suggested by Phelps (1997) can significantly affect the earnings of low-wage workers. However, such schemes reduce the return to schooling, and thus completed schooling levels fall with their introduction. In the article, we provide quantitative estimates of the impact of both school-bonus and wage-subsidy schemes on schooling, earnings, and welfare.

The format of the article is as follows. In Section II, we present the model, first in a rudimentary specification to fix ideas and then in the augmented version we empirically implement. Given that the model is estimated structurally, we also integrate the discussion of estimation method. We then discuss the data and provide descriptive statistics. Section III presents the estimation results, including discussions of model fit, the effects of endowment heterogeneity and its correlation with family background and a measure of IQ, and the impact of school bonus and wage-subsidy schemes on schooling and labor market outcomes. Section IV presents a number of conclusions.

II. Model

A. A Basic Human Capital Formulation

We begin with a basic human capital formulation. Each individual has a decision horizon of 50 years, beginning at age 16. At each age, $a = 16, \ldots, 65$, an individual chooses among five mutually exclusive and exhaustive alternatives: work in either a blue- or white-collar occupation, work in the military, attending school, or engaging in home production.[4]

[4] Primarily for computational reasons, we do not allow for joint activities, e.g., going to school and working.

Let $d_m(a) = 1$ if alternative m is chosen ($m = 1, \ldots, 5$) at age a and equal zero otherwise. Per-period utility at any age a is given by

$$U(a) = \sum_{m=1}^{5} R_m(a)d_m(a), \tag{1}$$

where $R_m(a)$ is the per-period reward associated with the mth alternative. These rewards contain all of the benefits and costs associated with each alternative.

1. *Working alternatives [$d_m(a) = 1$; $m = 1, 2, 3$].*—The current period return to working in occupation $m(R_m(a))$ is the wage $w_m(a)$. An individual's wage in an occupation is the product of the occupation-specific market (equilibrium) rental price (r_m) times the number of occupation-specific skill units possessed by the individual, $e_m(a)$.[5] The latter will depend on the technology of skill production. In a standard human capital formulation, the level of skill accumulated up to any age in an occupation depends on the number of years of schooling (successfully) completed, $g(a)$, and on work experience in that occupation, $x_m(a)$. Letting $e_m(a)$ be the number of skill units possessed at age a, $e_m(16)$ the skill endowment at age 16, and $\varepsilon_m(a)$ a skill technology shock,

$$e_m(a) = \exp[e_m(16) + e_{m1}g(a) + e_{m2}x_m(a) - e_{m3}x_m^2(a) + \varepsilon_m(a)], \quad m = 1, 2, 3;\ a = 16, \ldots, 65. \tag{2}$$

This specification leads to a standard (ln) wage equation in which the constant term is $(\ln(r_m) + e_{mk}(16))$, the sum of the ln rental price and the age-16 skill endowment.

2. *Nonwork alternatives—attending school ($d_4(a) = 1$) or remaining at home ($d_5(a) = 1$).*—In a strict human capital interpretation, the current period reward to school attendance ($R_4(a)$) is the effort cost (measured in dollars) necessary to complete a grade level. The reward to remaining home ($R_5(a)$) is the value of home production (measured in dollars). Each of these rewards is determined by a "skill" endowment at age 16 and is subject to a technology shock that varies over time (age).

Thus, the structure of rewards is given by

[5] This formulation can be motivated by an aggregate technology in which within-occupation skill units are perfect substitutes. In that case the rental prices are equal to occupation-specific skill marginal products. See Roy (1951), Heckman and Sedlacek (1985), and Willis (1986) for further discussion.

$$R_m(a) = w_m(a)$$
$$= r_m \exp[e_m(16) + e_{m1}g(a) + e_{m2}x_m(a) + e_{m3}x_m(a)^2 + \varepsilon_m(a)],$$
$$m = 1, 2, 3,$$
$$R_4(a) = e_4(16) + \varepsilon_4(a), \tag{3}$$
$$R_5(a) = e_5(16) + \varepsilon_5(a).$$

To close the model, productivity shocks are assumed to be joint normal, $N(0, \Omega)$, and serially uncorrelated; initial conditions are the given level of schooling completed at age 16, $g(16)$; and the accumulated work experience at age 16 in each occupation, which is assumed to be zero ($x_m(16) = 0$). It is convenient to define the age-16 endowment vector $e(16) = \{e_1(16), e_2(16), e_3(16), e_4(16), e_5(16)\}$, the work experience vector $x(a) = \{x_1(a), x_2(a), x_3(a)\}$, and the technology shock vector $\varepsilon(a) = \{\varepsilon_1(a), \varepsilon_2(a), \varepsilon_3(a), \varepsilon_4(a), \varepsilon_5(a)\}$.

At any age the individual's objective is to maximize the expected present value of remaining lifetime rewards. Defining $V(S(a), a)$, the value function, to be the maximal expected present value of lifetime rewards at age a given the individual's state $S(a)$ and given discount factor δ (dropping the endowment-type subscript for convenience),

$$V[S(a), a] = \max_{d_m(a)} E[\sum_{\tau=a}^{65} \delta^{\tau-a} \sum_{m=1}^{5} R_m(a)d_m(a)|S(a)]. \tag{4}$$

The state space at age a, $S(a) = \{e(16), g(a), x(a), \varepsilon(a)\}$, consists of all factors, known to the individual at a, that affect current rewards or the probability distribution of future rewards. Thus, $S(a)$ contains the relevant history of choices that enter the current period rewards, the endowment vector, and the realizations of all shocks at a, $\varepsilon_m(a)$ for $m = 1, \ldots, 5$.[6] In addition, the individual knows all relevant prices and functions (occupation-specific rental prices, the reward functions, the skill technology functions, direct schooling costs, and the distributions of shocks and endowments). The maximization in equation (4) is achieved by choice of the optimal sequence of control variables $\{d_m(a): m = 1, \ldots, 5\}$ for $a = 16, \ldots, 65$.

The value function can be written as the maximum over alternative-specific value functions, each of which obeys the Bellman equation:

[6] Past realizations of the shocks are also known but do not enter the state space because of the assumption of serial independence.

$$V[S(a), a] = \max_{m \in M}\{V_m[S(a), a]\}, \tag{5}$$

where $V_m(S(a), a)$, the alternative-specific value functions, are given by

$$V_m[S(a), a] = R_m[S(a), a] + \delta E\{V[S(a + 1), a + 1]| \\ S(a), d_m(a) = 1\}, \quad a < 65, \tag{6}$$

$$V_m[S(65), 65] = R_m[S(65), 65].$$

The expectation in equation (6) is taken over the distribution of the random components of $S(a + 1)$ conditional on $S(a)$, that is, over the unconditional distribution of $\varepsilon(a + 1)$ given serial independence. The predetermined state variables such as schooling and occupation-specific work experience evolve in a Markovian manner that is (conditionally) independent of the shocks: $x_m(a + 1) = x_m(a) + d_m(a)$ $(m = 1, 2, 3)$ in the case of occupation-specific work experience and $g(a + 1) = g(a) + d_4(a)$ in the case of schooling.

The individual's decision process is as follows: beginning at age 16, given $e(16)$ and $g(16)$, the individual draws five random shocks from the joint $\varepsilon(16)$ distribution, uses them to calculate the realized current rewards and thus the (five) alternative-specific value functions, and chooses the alternative that yields the highest value. The state space is then updated according to the alternative chosen, and the process is repeated through age 65. The solution of the optimization problem at each age a can be represented by the set of regions in the five-dimensional $\varepsilon(a)$ space over which each of the alternatives would be optimal, that is, would have the highest alternative-specific value function. There is no closed-form representation of the solution. Numerical solution is carried out by backward recursion using the approximation method developed and analyzed in Keane and Wolpin (1994).

The solution of the optimization problem serves as the input into estimating the parameters of the model given data on choices and, possibly, some of the rewards. To see this, consider having data on a sample of individuals from the same birth cohort who are assumed to be solving the model described above and for whom choices are observed over at least a part of their lifetimes. In addition, assume, as is the case, that wages are observed only in the periods in which market work is chosen and only for the occupation that is chosen. Thus, for each individual, $n = 1, \ldots, N$, the data consist of the set of choices and rewards $\{d_{nm}(a), w_{nm}(a) d_{nm}(a) : m = 1, \ldots, 3\}$ and $\{d_{nm}(a) : m = 4, 5\}$ for all ages in the given range [16, 65]. Let $c(a)$ denote the choice-reward combination at age a and let $\bar{S}(a) = \{e(16), g(a), x(a)\}$ denote the predetermined components of the state space, that is, $S(a)$ net of the technology shocks. Serial independence of the shocks implies that the probability of any sequence of choices and rewards can be written as follows:

$$\Pr[c(16), \ldots, c(65) | g(16), e(16)] = \prod_{a=16}^{65} \Pr[c(a) | \bar{S}(a)]. \qquad (7)$$

The sample likelihood is the product of the probabilities in equation (7) over the N individuals. The solution to the individual's optimization problem provides the choice probabilities that appear on the right-hand side of equation (7). Estimation is an iterative process between solving numerically the dynamic programming problem and computing the likelihood function. Estimation is by simulated maximum likelihood as described in Keane and Wolpin (1994).

B. An Augmented Model

The parsimonious human capital model described above is able to capture important qualitative features of behavior. For example, as in the Ben Porath (1967) deterministic model, there is an incentive to specialize in learning (school) early in the life cycle. However, the basic model cannot fit quantitatively the fall in school attendance that occurs at the ages of 18 and 22 because of the drop-off after high school and college graduation. Moreover, in the stochastic setting with independently and identically distributed productivity shocks and a 50-year horizon, the model cannot concentrate school attendance in the short period that is observed (and be consistent with the observed wage variance). To fit these and other features of the data, it was necessary to augment the basic human capital model in plausible directions (see also Keane and Wolpin 1997). We briefly discuss these augmentations (the full analytical representation is provided in the appendix).[7]

1. *Work Alternatives*

Skill technology functions ($e_m(a)$).—With respect to the two civilian occupations, in addition to completed schooling and (the quadratic in)

[7] The final specification is the outcome of an iterative process that involves augmenting the model to correct for specific deficiencies in the fit of the model to a set of basic characteristics of the data, namely, to the age-choice distribution and the choice transitions. All of the parameters we add, in our view, have coherent interpretations. Although this method of iterating between model specification and data fit clearly contaminates statistical measures of model fit, it would seem that such a strategy is unavoidable given the number of behaviors we are trying to model simultaneously.

occupation-specific work experience, the level of skill accumulated up to any age depends (linearly) on the total number of periods worked in each of the other two occupations and on whether the individual worked in the same occupation in the previous period (skill depreciation effect) or ever before, that is, a first-year experience effect.[8] The military skill function includes neither cross-experience effects nor depreciation effects. We allow for a linear age effect and for an age-less-than-18 effect in all three occupations.

Mobility and job search costs.—There exists a cost of finding a job in an occupation in which one did not work in the previous period. There is an additional cost if the person has no prior work experience in that occupation.

Nonpecuniary rewards plus indirect compensation.—We also allow for nonwage aspects of employment. Specifically, the current period reward for each civilian occupation is augmented by the net (positive or negative) monetary-equivalent value of working conditions or the indirect compensation associated with that occupation. These parameters can be interpreted as including occupation-specific fixed costs of work. The nonpecuniary reward associated with military employment is treated differently than for civilian occupations. We assume that the demand for military labor (skill units) is perfectly inelastic. In this case, nonpecuniary aspects of military employment must be fully compensated in their rental price (wage), and, in equilibrium, their existence cannot affect an individual's choice.

2. *School Attendance and Remaining at Home*

The effort cost of schooling is more generally interpreted to include a consumption value of school attendance and is allowed to depend systematically on age. In addition, college and graduate school tuition costs, net of their consumption value, are separately estimated, as are additional costs of reentry into high school and into postsecondary school.[9]

3. *Remaining at Home*

The payoff to remaining home is allowed to depend systematically on age.

[8] Allowing for lagged dependence is not a substitute for serial correlation in wage draws, which would be considerably more computationally burdensome. We allow for serial correlation through permanent unobserved heterogeneity.

[9] One possible reason for a reentry cost is that because of knowledge depreciation, effort may have to increase if school attendance is not continuous. Alternatively, there may be a psychic cost of attending school with a younger school entry cohort.

4. *Common Returns*

All alternatives also contain a common set of rewards. There is a psychic value of having earned a high school diploma (β_1) and an additional psychic value of a college diploma (β_2). Also, there is a cost of leaving the military prematurely, that is, without having remained there for at least two years, which enters the rewards of all alternatives except the military (β_3). Thus, choosing any alternative other than the military given that the individual has exactly 1 year of military experience leads to a loss in the reward associated with the alternative.

5. *Endowment Heterogeneity*

We allow for the possibility that individuals do not have identical endowments. Specifically, define a type k individual, $k = 1, \ldots, K$, by their endowment vector $e_k(16) = (e_{mk}(16) : m = 1, 2, 4, 5)$. Thus, individuals may have comparative advantages in different alternatives, including schooling and home production. Types are common knowledge so that wage offers always reflect actual skill level. We assume that while endowment heterogeneity is unobserved by us, we know there to be K types.[10] Denote π_k as the proportion of the kth type in the population. In this case, the likelihood function is a mixture of the type-specific likelihoods, that is, $\Pi_{n=1}^{N} \{\Sigma_{k=1}^{K} \pi_k L_{nk}\}$, where L_n is the likelihood of person n's observed choice sequence and rewards if person n is of endowment type k, and where the parameter vector is augmented to include the endowment vectors for the K types and the type probabilities.

If initial schooling (at age 16) is not exogenous, conditioning the likelihood on it as if it were nonstochastic is problematic. One remedy would be to specify the optimization problem back to the age when initial schooling was zero for everyone (say, age 3) and to solve for the correct probability distribution of attained schooling at age 16 (conditional on age-3 endowments). However, such a model would have to focus on parental decision making with respect to investments in children (including fertility decisions) and would be very demanding in many dimensions (modeling, computation, data). The alternative we follow is to assume that initial schooling is exogenous conditional on the age-16 endowment vector.[11] The likelihood contribution for the nth individual is thus

[10] We obviously do not know K a priori. We arrive at K empirically in the same iterative process we use to augment the other aspects of the model, namely, we stop when we are satisfied with the fit of the model.

[11] Another alternative is to treat the unobserved heterogeneity as an incidental parameters problem (Heckman 1981). However, estimating individual-specific endowments is not computationally feasible.

$$\Pr[c_n(16), \ldots, c_n(65) | g_n(16)] = \sum_{k=1}^{K} \pi_{k|g_n(16)} \Pr[c_n(16), \ldots, c_n(65) | g_n(16), \text{type} = k]. \quad (8)$$

Note that the type proportions are now conditioned on initial schooling, $g(16)$.

III. Data

The data are from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience (NLSY). The NLSY consists of 12,686 individuals, approximately half of them men, who were 14–21 years old as of January 1, 1979. The sample consists of a core random sample and an oversample of blacks, Hispanics, poor whites, and the military. This analysis is based on the white males in the core random sample and the black males in the core and oversample who were age 16 or less as of October 1, 1977. Interviews were first conducted in 1979 and have been conducted annually to the present. We follow each individual in the subsample defined above from the first year they reach age 16 as of October 1 of that year through September 30, 1988.

The NLSY collects schooling and employment data as an event history retrospective to the preceding interview. Schooling data include the highest grade attended and completed at each interview date, monthly enrollment in each calendar month, school leaving dates, and the dates of diplomas and degrees. Employment data include the beginning and ending dates (to the calendar week) of all jobs (employers), all gaps in employment within the same job, usual hours worked on each job, the usual rate of pay on each job, and the three-digit occupation for each job. In the 1979 interview, employment data were collected back to January 1, 1978.

The behavioral model is implemented on annual aggregates of the discrete alternatives beginning as of the first time the individual was age 16 on October 1 of that year. The sample consists of 2,232 males: 1,373 whites and 857 blacks. Definitional arbitrariness in assigning observations to alternatives is unavoidable given that the assumed decision period is longer (annual) than the weekly observation period. The assignments were made as follows.

A. Enrolled in School

To simplify the determination of school enrollment, we looked at an individual's activity in the fortieth week of each year (October 1), the first week of each year (January 1), and the fourteenth week of each year (April 1), beginning with January 1, 1978. An individual is considered to be enrolled in school during the year if the individual was enrolled in any of the 3 weeks and reported completing one grade level by October 1 of the next year.[12]

B. Worked

The work assignment used data on work status in 9 weeks, again to simplify the classification, between October 1 and June 30.[13] An individual is considered to have worked during the year if the individual was not enrolled in school and was employed in at least two-thirds of the weeks for at least 20 hours per week on average.[14]

Occupation classification.—A working individual is assigned to one of three occupations: blue-collar (BC), white-collar (WC), and the military (ML). The occupation that is assigned is the one in which the individual worked the most weeks during the year (based on the same 9 weeks used to determine work status).[15] Aggregating occupations into just two categories implies that the disaggregated occupations within each category utilize the same type of skill units. Although a finer disaggregation would

[12] There are a considerable number of observations (as many as 20%) with longitudinally inconsistent enrollment and highest grade completed data. The records of all observations with inconsistent data were carefully scrutinized. In most cases we were able to reconstruct a reasonable school enrollment and grade completion history or at least a partial history based on the different pieces of information that are reported in the NLSY, i.e., the monthly enrollment calendar; survey date enrollment, highest grade attended, and highest grade completed; dates of school leaving; dates of diplomas; and the highest grade completed as of May 1 ("key" variable created by the Center for Human Resource Research). In determining highest grade completed, an individual who obtained a General Equivalency Diploma (GED) was not considered to have completed 12 years of schooling; instead highest grade completed is the number of years that were actually attended and successfully completed. This treatment is consistent with recent work by Cameron and Heckman (1993). The rule was a bit more complicated because of missing enrollment data. If 2 weeks or more of enrollment data was missing, then the only determinant of school enrollment was whether a grade was completed. If highest grade completed was missing as of October 1, in any 2 consecutive years, then the observation is truncated at that period.

[13] The 9 weeks were the first, the seventh, and the thirteenth of each of the 3 calendar quarters spanning the period. We ignored the summer quarter so as not to count summer jobs of those in school.

[14] If work status is missing for less than two-thirds of the weeks, then the work criterion is the same based on the nonmissing weeks. If work status is missing for more than two-thirds of the weeks, then the observation is dropped from then on.

[15] Occupational categories are based on one-digit codes. Blue-collar occupations are (i) craftsmen, foremen, and kindred; (ii) operatives and kindred; (iii) laborers, except ones on farms; (iv) farm laborers and foremen; and (v) service workers. White-collar occupations are (i) professional, technical, and kindred; (ii) managers, officials, and proprietors; (iii) sales workers; (iii) farmers and farm managers; and (iv) clerical and kindred.

probably be desirable, a nontrivial number of year-to-year transitions between finer occupational categories (even between one-digit codes) appears to be spurious.[16] Moreover, the computational burden would increase significantly with more occupations.

Real wages.—Real (occupation-specific) wages are obtained by multiplying the average real weekly wage for the weeks worked in the occupation (assigned as above) times 50 weeks. The wage is, therefore, a "full-time" equivalent.[17]

C. Home

An individual is classified as being at home during the year if the individual neither enrolled in school nor worked during the year, according to the above definitions. In actuality, some individuals would be classified as being at home if they were enrolled even for the full year but did not successfully complete a grade level, or if they worked during the year but did not satisfy the weeks and hours criterion.

Tables 1 and 2 show the choice distribution by age for whites and blacks. As already noted, there are 1,373 whites and 857 blacks in the sample at age 16; the number declines slightly over the first 8 years primarily because of sample attrition.[18] Over the last 3 years the sample size falls because part of the sample never reaches the older ages during the sample period. As the tables show, approximately 86% of whites and

[16] With one-digit occupation codes, the transitions between the calendar quarters surrounding the interview date are significantly higher than between any other quarters. Individuals, even those with the same employer, appear to report verbatim characterizations of their jobs that coders, who are trained to classify the verbatim responses into appropriate three-digit codes, interpret as occupation changes that are not real. This problem essentially disappears in the white- and blue-collar classification scheme.

[17] The wage is deflated by the gross national product deflator, with 1987 as the base year. Ideally, we would like measured annual earnings to reflect the annual wage offer that an individual is presumed to receive in our model. We normalized annual earnings for differences in weeks worked (recall that two-thirds of the weeks must be spent employed to be defined as working) on the grounds that most wage offers are expressed as full-year offers, i.e., part-year jobs are an exception. Jobs that last less than a full year reflect either job-changing behavior or unanticipated unemployment. We did not normalize for differences in weekly hours (recall that the work classification requires that weekly hours be 20 or greater) on the grounds that wages are likely permanently different for part-time and full-time jobs.

[18] Given the sample restrictions, namely, male respondents in the core component of the survey in a particular age group, there would have been at most 1,401 whites and 870 blacks observed at age 16 without any loss of observations because of missing data. Effective attrition is minimized in the NLSY by obtaining the retrospective employment and schooling information for respondents who return to the sample after an attrition spell.

Table 1
Choice Distribution—White Males, Ages 16–26

	Choice					
Age	School	Home	White-Collar	Blue-Collar	Military	Total
16	1,178	145	4	45	1	1,373
	(85.8)	(10.6)	(.3)	(3.3)	(.1)	(100.0)
17	1,014	197	15	113	20	1,359
	(74.6)	(14.5)	(1.1)	(8.3)	(1.5)	(100.0)
18	561	296	92	331	70	1,350
	(41.6)	(21.9)	(6.8)	(24.5)	(5.2)	(100.0)
19	420	293	115	406	107	1,341
	(31.3)	(21.9)	(8.6)	(30.3)	(8.0)	(100.0)
20	341	273	149	454	113	1,330
	(25.6)	(20.5)	(11.2)	(34.1)	(8.5)	(100.0)
21	275	257	170	498	106	1,306
	(21.1)	(19.7)	(13.0)	(38.1)	(8.1)	(100.0)
22	169	212	256	559	90	1,286
	(13.1)	(16.5)	(19.9)	(43.5)	(7.0)	(100.0)
23	105	185	336	546	68	1,240
	(8.5)	(14.9)	(27.1)	(44.0)	(5.5)	(100.0)
24	65	112	284	416	44	921
	(7.1)	(12.2)	(30.8)	(45.2)	(4.8)	(100.0)
25	24	61	215	267	24	591
	(4.1)	(10.3)	(36.4)	(45.2)	(4.1)	(100.0)
26	13	32	88	127	2	262
	(5.0)	(12.2)	(33.6)	(48.5)	(.81)	(100.0)
Total	4,165	2,063	1,724	3,762	645	12,359
	(33.7)	(16.7)	(14.0)	(30.4)	(5.2)	(100.0)

NOTE.—Number of observations (percentages in parentheses).

82% of blacks are in school at age 16. The largest decline in enrollment occurs at the natural high school graduation age, between 17 and 18. Fewer black males are enrolled at each age, and the difference grows with age. The propensity to work increases monotonically for both races; however, by age 25 over 85% of white males and less than 70% of black males are working. Further, among those working, almost two-thirds of blacks, but only about one-half of whites, are in blue-collar occupations. Perhaps more striking is that at age 25, almost three of every 10 black males in this sample are neither in school (i.e., attending and completing a grade level) nor working ("full-time"). The same is true for only one in every 10 white males.

Table 3 reports age-specific average real wages overall and by occupation. Real wages rise with age in all occupations for both black and white males. However, except for the military, black males earn less than white males essentially at every age. At age 23, the black-white wage ratio is .81 for those who are working in white-collar occupations and .76 for those in blue-collar occupations. Notice of course that the individuals in these categories have both chosen to work and to work in a particular occupation.

Table 2
Choice Distribution—Black Males, Ages 16–26

Age	School	Home	White-Collar	Blue-Collar	Military	Total
16	699	129	5	23	1	857
	(81.6)	(15.1)	(0.6)	(2.7)	(0.1)	(100.0)
17	553	232	3	49	15	852
	(64.9)	(27.2)	(0.4)	(5.8)	(1.8)	(100.0)
18	317	311	19	129	66	842
	(37.7)	(36.9)	(2.3)	(15.3)	(7.84)	(100.0)
19	176	338	41	186	94	835
	(21.1)	(40.5)	(4.9)	(22.3)	(11.3)	(100.0)
20	110	327	48	245	98	828
	(13.3)	(39.5)	(5.8)	(29.6)	(11.8)	(100.0)
21	68	295	67	295	90	815
	(8.3)	(36.2)	(8.2)	(36.2)	(11.0)	(100.0)
22	55	279	97	296	76	803
	(6.9)	(34.7)	(12.1)	(36.9)	(9.5)	(100.0)
23	39	217	116	329	66	767
	(5.1)	(28.3)	(15.2)	(42.9)	(8.6)	(100.0)
24	18	160	102	257	48	585
	(3.1)	(27.4)	(17.4)	(43.9)	(8.2)	(100.0)
25	9	112	72	168	25	386
	(2.3)	(29.0)	(18.7)	(43.5)	(6.5)	(100.0)
26	4	50	44	83	12	193
	(2.1)	(25.9)	(22.8)	(43.0)	(6.2)	(100.0)
Total	2,048	2,450	614	2,060	591	7,763
	(26.4)	(31.6)	(7.9)	(26.5)	(7.6)	(100.0)

NOTE.—Number of observations (percentages in parentheses).

Table 3
Average Real Wages by Occupation: White and Black Males

	Mean Wage							
	All Occupations		White-Collar		Blue-Collar		Military	
Age	White	Black	White	Black	White	Black	White	Black
17	11,036	9,413	*	*	11,572	9,597	*	*
18	12,060	10,043	11,775	*	12,603	10,480	10,171	9,230
19	12,246	11,424	12,376	15,226	12,949	11,400	9,714	9,971
20	13,635	11,982	13,824	11,631	14,363	12,547	10,852	10,823
21	14,977	12,696	15,578	14,309	15,313	12,524	12,619	12,033
22	17,561	13,880	20,236	15,577	16,947	13,322	13,771	13,920
23	18,719	14,390	20,745	16,850	17,884	13,536	14,868	14,453
24	20,942	15,750	24,066	17,686	19,245	15,019	15,910	15,536
25	22,754	16,665	24,899	18,371	21,473	16,050	*	*
26	25,390	17,044	32,756	19,952	20,738	15,489	*	*

* Fewer than 25 observations.

School attainment at age 16 (as of October 1 of the first calendar year in which the respondent is age 16) is concentrated at two values. Within the two groups, 87% of whites and 82% percent of blacks had attained either 9 or 10 years of schooling by that age. Specifically, the percentage of blacks who had completed 11, 10, 9, 8, and 7 years of schooling at age 16 were 8%, 52%, 30%, 8%, and 2%, respectively. For whites, the analogous percentages for these ages are 8%, 67%, 20%, 4%, and 1%, respectively. Initial schooling differences are amplified with age for both blacks and whites. White males with 9 years of schooling at age 16 had completed 11.5 years of school on average by age 23, while those with 10 years of schooling completed at age 16 had 13.3 years of schooling at age 23. Similarly, blacks who had completed 9 years of schooling at age 16 had 11.3 years of schooling at age 23, while those with 10 years of schooling at age 16 had completed 12.6 years at age 23.

IV. Estimation Results

A. Parameter Estimates

The model was first estimated on blacks and whites separately. Four types for each race group provided an adequate fit of the model to the data.[19] Comparing individual parameters for the separate races led us to perform a likelihood-ratio test for a restricted model in which all of the parameters except the constant terms in the occupation-specific wage functions and the type proportions were equal. We found that this restriction did not compromise fit according to the likelihood ratio criterion.[20] The full parameter estimates from this constrained specification (and standard errors) are reported in appendix tables A1–A3.

Table 4 presents estimates by race of the unconditional type proportions ($\Sigma_{g(16)}\ \pi_{k|g(16)} \cdot \pi_{g(16)}$, for $k = 1, \ldots, 4$) and of the type-specific endowment vectors. The latter include initial schooling, the occupation-specific (ln) wage constants ($\ln(r_m) + e_{mk}(16)$), the consumption (net of effort cost) values of schooling ($e_{4k}(16)$), and the home production skills

[19] Except insofar as we test the overall fit of the model (see below), we do not perform tests on the number of types. Such tests are problematic because the existence of more or fewer types involves tests of parameters on the boundaries of the parameter space. However, the estimation strategy we outlined above involved adding types only as we saw significant quantitative improvements in model fit. Actually, moving from type 2 to type 3 improved the fit a great deal, while moving from type 3 to type 4 improved the fit modestly.

[20] Twice the difference in the log-likelihood values was 51.8, which does not exceed the critical χ^2 value with 74 degrees of freedom, 95.1. The unrestricted log-likelihood values were −14,033 for whites and −8,628 for blacks, while the restricted value was −22,687. In the restricted specification, the separate log-L values were −14,045 for whites and −8,642 for blacks.

Table 4
Race Differences in "Endowments"

	Type 1	Type 2	Type 3	Type 4	Mean
Sample proportion:					
White	.135	.297	.365	.203	. . .
Black	.041	.133	.410	.416	. . .
School attainment at age 16, 10 years or more:					
White	.833 (1)	.564 (2)	.468 (3)	.337 (4)	.746
Black	.892 (1)	.551 (2)	.545 (3)	.409 (4)	.600
White-collar skill endowment + (ln) skill rental price:					
White	8.86 (1)	8.78 (3)	8.28 (4)	8.80 (2)	8.52
Black	8.78 (1)	8.70 (3)	8.20 (4)	8.72 (2)	8.38
Blue-collar skill endowment + (ln) skill rental price:					
White	8.80 (2)	9.12 (1)	8.61 (4)	8.72 (3)	8.81
Black	8.75 (2)	9.07 (1)	8.56 (4)	8.67 (3)	8.72
Consumption value of school attendance net of effort cost:					
White	17,876 (1)	10,956 (3)	5,694 (4)	11,220 (2)	10,023
Black	17,876 (1)	10,956 (3)	5,694 (4)	11,220 (2)	9,192
Value of home production:					
White	22,437 (1)	20,179 (2)	8,745 (4)	15,998 (3)	18,866
Black	22,437 (1)	20,179 (2)	8,745 (4)	15,998 (3)	13,844

NOTE.—Numbers in parentheses provide the rank order over the types.

($e_{5k}(16)$). Given that the constant terms in the (ln) wage functions represent the sum of the (ln) skill rental prices and skill endowments, it is not possible to distinguish whether race differences in within-type, occupation-specific wage offers are because of discrimination (blacks and whites of equal skill receive different wage offers) or of within-type skill endowment differences at age 16.[21] The last column in table 4 provides the implied mean endowments (by race) taken over types. The numbers in parentheses represent the within-race rank of the type with respect to the endowment.

Endowment rankings are independent of race, which, except for initial schooling, must be the case in the restricted pooled black and white specification. None of the four types has either an absolute advantage or

[21] This specification implies that black-white differences in skill rental prices, if they exist, are independent of type. Although it is possible that tastes for discrimination may in fact be type specific, that specification is implicitly rejected as part of the null that we tested. That restriction is critical to the identification of the upper-bound estimate of the effect of discrimination on the wage differential.

an absolute disadvantage in endowment levels. Type 1s have the highest endowment in every category except blue-collar skill, where they rank second. At the other extreme, type 3s have the lowest endowment in every category, except in initial schooling, where they rank third. The black type distribution is significantly more concentrated than is the white type distribution; 82% of blacks are either of type 3 or 4, and only 4% are of type 1, while comparable figures for whites are 56% and 14%.

With respect to race differences, as the last column in table 4 shows, an average (over types) 16-year-old white male from this cohort has a higher skill endowment plus skill rental price in both white- and blue-collar occupations relative to a similar average 16-year-old black male. Within each type, black males receive wage offers (for the same schooling and work experience) that are 8% lower in white-collar occupations and 5% lower in blue-collar occupations. However, given the differences in their type distributions, the average black male receives offers that are 14% lower in white-collar and 9% lower in blue-collar occupations. The level of schooling and home productivity endowments, as noted, were found not to differ by race. However, because there is a higher percentage of low endowment types among blacks relative to whites, their net (of effort cost) consumption value of schooling is about 10% lower and their home productivity is 36% lower. Finally, initial school attainment at age 16 varies by type similarly for both races. At the extremes, over 83% of whites and 89% of blacks who are type 1s have at least 10 years of schooling at age 16, but the same is true of only 34% of whites and 41% of blacks who are type 4s.

B. Model Fit

The model fits the choice data well. Figures 1–5 for whites and 6–10 for blacks show the actual and predicted choice frequencies for both the within-sample period up to age 26 and for the entire "lifetime" through age 65. In addition, the figures contrast the fit for a static model, that is, for a specification in which the discount factor is equal to zero instead of its estimated value of .93.[22] Both the dynamic programming model and the static models give similar within-sample predictions for whites and blacks. However, the static model gives quite unusual out-of-sample predictions, with the entire sample being employed as white-collar workers by age 50. In addition, the static model predicts white-collar wage offers to rise extremely rapidly (to about $250,000 by age 50), extrapo-

[22] The static model is fit separately on blacks and whites, which improves its within-sample fit relative to the restricted dynamic model. The log-L values for the static model, restricting the discount factor to be zero, were −14,068 for whites and −8,630 for blacks.

lating the trend from the actual data. On the other hand, because such extremely large future wage offers would adversely affect the fit of the dynamic model to observed choices, the dynamic programming model predicts a more credible out-of-sample pattern of white-collar wage growth (to about $35,000 at age 50).[23]

With respect to school completion levels specifically, table 5 shows that mean schooling at age 23 predicted by the model is 12.8 years for whites and 11.9 years for blacks, while actual completed schooling at that age is 12.7 and 11.9 years, respectively.[24] However, the model does not fit the overall schooling distribution as well as it fits the mean. It overpredicts the percentage of high school dropouts at age 23 (by somewhat less than 3 percentage points for each race group) and underpredicts the percentage of those with exactly 12 years of education (by about 6 percentage points for each race group). The model thus also overpredicts the percentage who attend or graduate from college, but the only substantial divergence is the overstatement of black college graduates (by 2.9 percentage points).

By age 30, the model predicts that 25.1% of white males and 12.6% of black males will be college graduates. Using NLSY data from the 1993 survey year, highest grade completed at age 30, based on the "key variable" created by the Center for Human Resource Research, has college completion rates for this sample of 26.1% for whites and 10.0% for blacks.[25] Current Population Survey estimates, as cited in the introduction, give comparable figures of 24.4% and 12.6%, respectively.[26]

C. Endowment Effects

Initial conditions matter for educational attainment, as tables 6 and 7 demonstrate. Table 6 shows completed schooling (predicted at age 30) by initial schooling (at age 16) and by type. Proportions of the sample in each joint initial schooling-type category are shown in parentheses. For both blacks and whites, the mean schooling for those having type 1 endowments is above 16 years regardless of initial schooling; there is about a one-half year difference between those having completed 9 or less years

[23] See Keane and Wolpin (1997) for further evidence on the out-of-sample fit of the dynamic model.

[24] All of the predictions and simulations in the tables that follow are based on simulating a random sample for each race of 5,000 individuals from age 16 through age 65 based on the estimated model parameters.

[25] It should be noted that this measure differs from the highest grade completed we have generated ourselves through the 1989 round of the NLSY both because it assigns 12 years of schooling to those with a GED and because the Center for Human Resource Research measure is not longitudinally consistent.

[26] Comparisons with respect to summary statistics based on grade levels of 12 or less are not possible without accounting for GEDs because the CPS does not distinguish GEDs from regular high school diplomas.

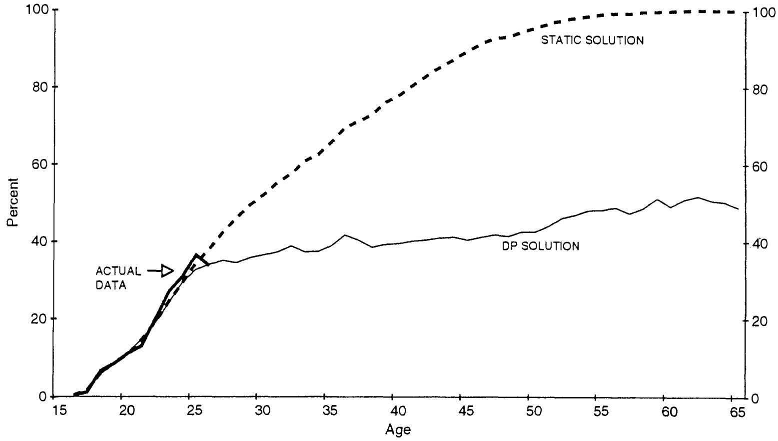

FIG. 1.—White-collar by age: whites (%)

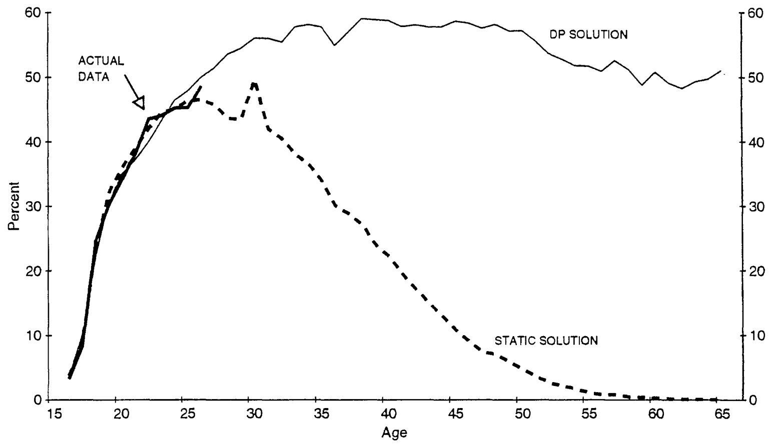

FIG. 2.—Blue-collar by age: whites (%)

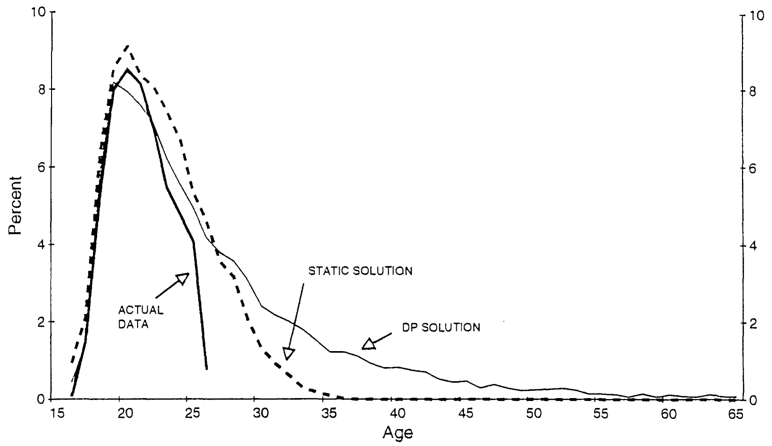

FIG. 3.—In military by age: whites (%)

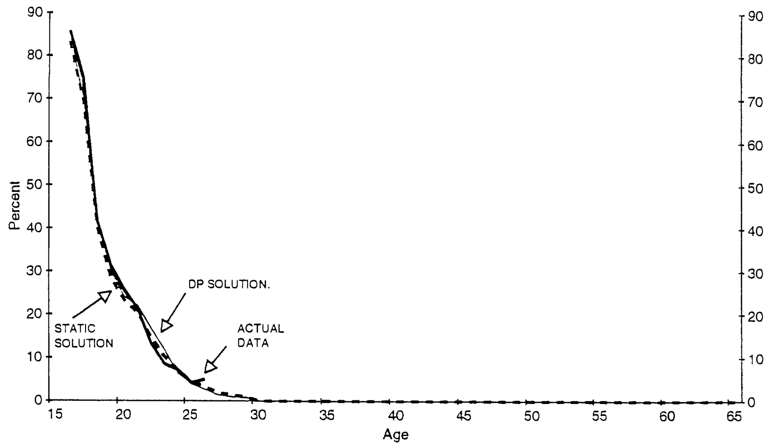

FIG. 4.—In school by age: whites (%)

of schooling by age 16 and those having completed 10 or more years. However, if one is a type 2, 3, or 4, initial schooling matters more for completed schooling than does type. Regardless of endowment type or race, being in the lower initial schooling group reduces completed schooling by between 1.5 and 2 years. Thus, for these three types, a random shock to school completion levels prior to age 16 (e.g., poor performance in one year due to a transitory health condition) has a permanent impact on future schooling that is quantitatively large.

Table 7 shows what would happen, according to the model's predic-

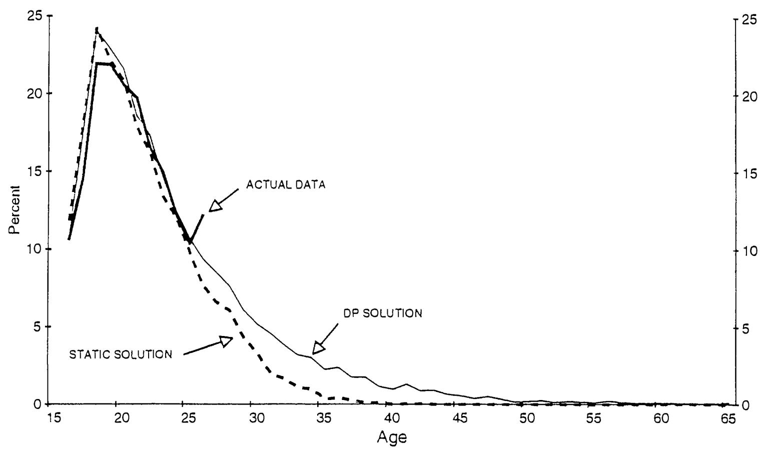

FIG. 5.—At home by age: whites (%)

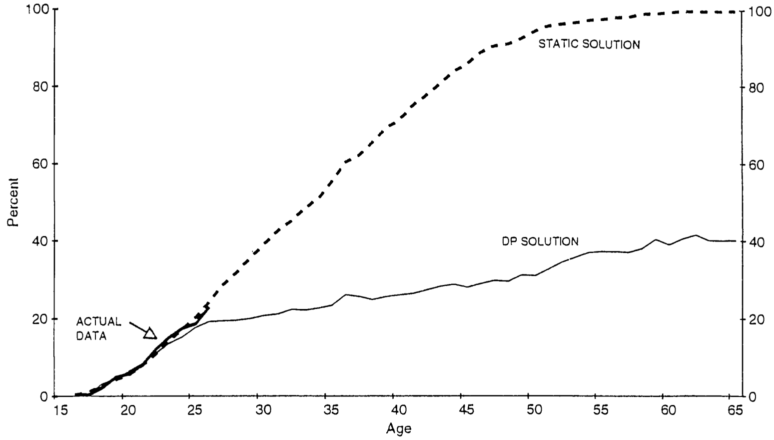

FIG. 6.—White-collar by age: blacks (%)

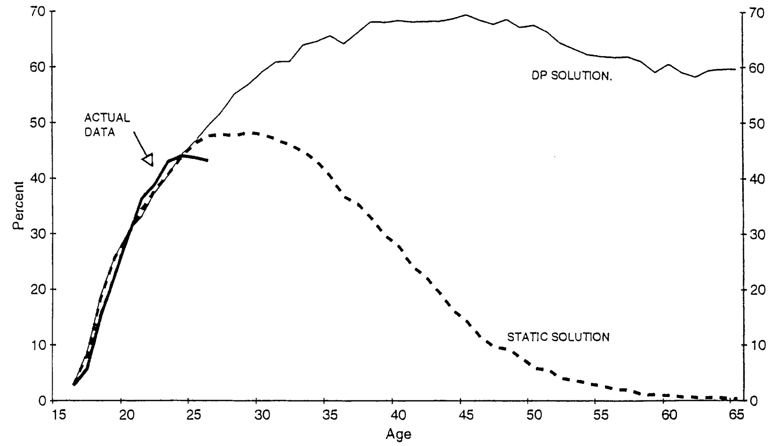

FIG. 7.—Blue-collar by age: blacks (%)

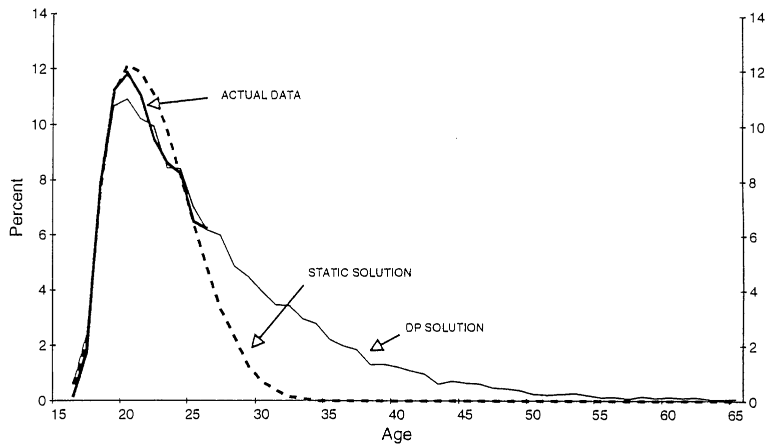

FIG. 8.—In military by age: blacks (%)

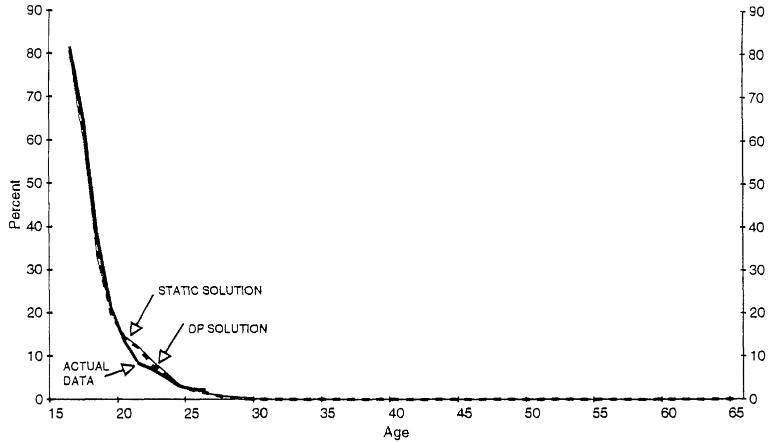

FIG. 9.—In school by age: blacks (%)

tions, to race differences in school attainment and later labor market success if initial schooling, type proportions, and (within-type) skill endowment levels plus skill rental prices were equalized. We consider these effects with respect to school attainment first. The first column presents the (simulated) baseline situation. It is apparent that the completed schooling distributions for blacks and whites differ mainly in the percentage who are high school dropouts (12% more blacks than whites) and the percentage who are college graduates (12% more whites than blacks). In addition, mean educational attainment is 1 year less for blacks.

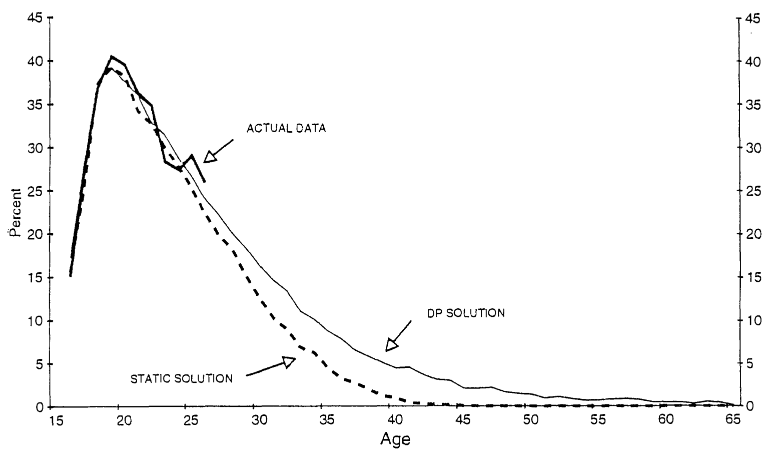

FIG. 10.—At home by age: blacks (%)

Table 5
Actual and Predicted School Attainment by Race

	White			Black		
	Actual at Age 23	Predicted at Age 23	Predicted at Age 30	Actual at Age 23	Predicted at Age 23	Predicted at Age 30
Highest grade completed < 12 (%)	24.0	26.8	26.3	35.9	38.6	38.4
Highest grade completed = 12 (%)	36.2	30.0	29.5	37.7	31.7	31.3
Highest grade completed ≥ 13 and ≤ 15 (%)	18.7	22.6	19.1	19.6	20.0	17.7
Highest grade completed ≥ 16 (%)	21.3	20.6	25.1	6.8	9.7	12.6
Mean highest grade completed (%)	12.7	12.8	13.0	11.9	11.9	12.0

The second column in table 7 reports an experiment in which blacks had the same initial schooling distribution as whites, but in which type proportions and wage offer constants were the same as the baseline. Our estimates imply that equalizing the initial schooling distribution would cause the mean schooling difference to fall by 30%. The next experiment (col. 3) assumes that blacks have the same type proportions as whites within initial schooling groups but that the initial schooling distribution is the same as in the baseline. In this experiment, the mean schooling differential is reduced by 60%. The third experiment combines the first

Table 6
Predicted Mean Completed Schooling by Initial Schooling and Endowment Type and by Race

	White	Black
Schooling at age 16 ≤ 9 years:		
Type 1	16.9 (.70)	16.6 (.26)
Type 2	10.9 (5.9)	10.8 (4.6)
Type 3	11.0 (10.3)	10.7 (14.3)
Type 4	11.1 (8.0)	10.6 (19.4)
Schooling at age 16 ≥ 10 years:		
Type 1	17.3 (12.7)	17.2 (3.9)
Type 2	13.0 (23.8)	12.6 (8.7)
Type 3	12.4 (26.2)	12.4 (26.7)
Type 4	13.0 (12.3)	12.7 (22.2)

NOTE.—% of sample in parentheses.

Table 7
Predicted Effects of "Endowment" Equalization and School Performance-Based Bonuses on Completed Schooling and Labor Market Success

	Baseline		Blacks with White Initial Schooling Distribution	Blacks with White Type Proportions	Blacks with White Initial Schooling and Type Proportions	Blacks with White Wage Offer Function	Bonus Scheme
	White	Black					
Mean highest grade completed	13.0	12.0	12.3	12.6	12.7	12.3	12.8
Highest grade completed < 12 (%)	26.2	37.9	32.3	34.1	30.8	32.7	26.0
Highest grade completed = 12 (%)	29.6	31.5	33.2	28.5	29.7	31.6	31.6
Highest grade completed ≥ 13 and ≤ 15 (%)	19.1	17.9	19.5	16.4	17.4	20.6	16.2
Highest grade completed ≥ 16 (%)	25.1	12.8	15.0	21.1	22.1	15.3	26.1
Earnings at age:							
30	22,795	15,818	16,418	19,987	20,188	18,327	16,526
40	32,635	24,626	25,262	29,559	29,877	27,375	25,414
50	44,014	34,120	34,880	39,776	40,108	37,563	35,143
Working in white-collar occupation at age (%):							
30	36.5	20.8	22.9	28.9	30.1	27.8	27.1
40	39.5	26.1	28.0	34.3	35.0	32.0	30.0
50	42.2	31.0	32.5	38.5	38.9	37.4	32.3
Working in blue-collar occupation at age (%):							
30	55.9	59.1	58.3	57.7	56.7	60.1	55.1
40	58.7	68.3	66.8	62.2	61.5	65.3	65.1
50	57.0	67.4	66.2	60.4	60.3	64.1	66.3
% home at age:							
20	21.6	37.6	35.3	28.8	28.0	29.7	33.1
30	5.2	16.2	14.6	10.0	9.7	9.3	13.9
40	1.0	4.4	3.9	2.5	2.4	1.9	3.6
50	.2	1.4	1.1	.8	.7	.4	1.1
Mean present discounted value of earnings	285,400	210,258	216,268	254,228	256,457	239,219	213,885
Utility	316,898	264,886	267,608	298,552	299,921	278,602	269,402

two, that is, blacks are assumed to have the same joint initial schooling-type distribution as whites. Given that our estimates imply significant nonindependence of initial schooling and type (the more highly endowed also have higher initial schooling), the reduction in the mean schooling differential is not additive. This experiment leads to a reduction in the mean schooling differential of 70%.[27]

In contrast to altering initial schooling and type distributions, the next column considers an experiment that equalizes the wage offer constants. As noted, one can think of these differences as arising from a combination of rental price differences because of discrimination against blacks and race differences in within-type, occupation-specific skill endowments. If black-offered wages were drawn from the same distribution as were white-offered wages, that is, if all within-type wage offer differences were because of discrimination and were remediated entirely through government policies, the mean schooling differential would fall by 30%, the same as equalizing initial schooling. Obviously, eliminating all differences, the joint initial schooling-type distribution and wage-offer functions, would completely erase race differences in educational attainment.[28]

Table 7 also reports on a number of measures of labor market success: mean annual earnings (zero for nonworkers) at age 30, 40, and 50; the percentage working in white-collar occupations at age 30, 40, and 50; the percentage working in blue-collar occupations at age 30, 40, and 50; the percentage neither working nor in school (i.e., at home) at age 20, 30, 40, and 50; and the present discounted value of earnings and of utility (from age 16 through age 65). According to the baseline estimates, the black-white earnings ratio increases with age but is still only .77 at age 50. This increase with age arises from the relatively larger increase in the proportion of white-collar to blue-collar employment for blacks as well as from the larger overall decline in the proportion at home (because nonworkers are assigned zero earnings).[29] Our estimates imply that over one-third of black males (only one-fifth of whites) from this cohort are neither working (full-year) nor in school at age 20; by age 50, the percentage at home is only 1.4 (.2).

In terms of lifetime welfare measures, the present discounted value of earnings (assuming zero earnings for those at home) for black males of this cohort is estimated to be only 73.7% that of whites. Based on utility

[27] The importance of initial conditions in explaining black-white schooling differentials is consistent with the findings in Cameron and Heckman (1998).

[28] Obviously, in this case, discounted lifetime wages and utilities would also be equalized.

[29] The black-white earnings ratio is approximately constant at .78 if based only on those with positive earnings.

(normalized to dollars), that figure increases to 83.6%. Considering these two summary measures, the experiments in table 7 indicate that equalizing type proportions, as in the case of school attainment, would have a larger impact on the black-white present discount value (PDV) of earnings or utility differential than would equalizing initial schooling or wage-offer constants. Such a change would increase the black-white percentage to 89.1% for earnings and to 94.2% for utility, closing 59% and 65% of the respective gaps. On the other hand, eliminating discrimination (equalizing wage offers) would close at most only 39% and 26% of the respective gaps.

Accomplishing these changes in practice would require policies that eliminate race differences in age-16 endowments or discriminatory behavior. Many government programs exist for just these purposes, from subsidizing prenatal and infant care Women, Infants, and Children to compensatory educational programs (Head Start, Title I) to antidiscrimination and affirmative action policies (Equal Employment Opportunity Commission, Office of Federal Contract Compliance). Yet, race differences in school attainment and later labor market success persist. We thus turn to two issues: (1) what are the correlates of endowment differences that could possibly be manipulated by policy to reduce racial schooling and earnings differentials and (2) what would be the impact on those differentials of direct government subsidies that alter the rewards to schooling or to work.

D. Correlates of Endowments

All permanent heterogeneity is assumed in the behavioral choice model to be captured by the existence of the four endowment types. As noted, blacks are underrepresented relative to whites among the two high skill groups (types 1 and 2), and this difference accounts for a substantial proportion of the black-white schooling and earnings gap. Given this specification of permanent heterogeneity (as of age 16), family background factors and innate talents are related to decisions only as they are related to an individual's type.[30] Within the random-effects framework, it is not possible to estimate each individual's endowment type. However, given the estimated parameters, it is possible to use Bayes's rule to estimate type probabilities for each individual in the sample, which can then be correlated with observable factors.

[30] An alternative to modeling all heterogeneity as unobserved would be to have allowed some parameters of the model to depend directly on measures of family background and ability found in the NLSY. The specification we adopt is more parsimonious in parameters and, subject to the model adequately fitting the data, still allows us to explore the relationship of family background and ability to endowment heterogeneity.

Table 8
The Relationship between Black-White Differences in Endowment Types and Family Background and AFQT

	Probability of Type 1	Probability of Type 2	Probability of Type 3	Probability of Type 4
Control variables:				
No controls	−.094	−.164	.045	.213
	(.010)	(.014)	(.017)	(.014)
Family background*	−.037	−.163	.015	.195
	(.014)	(.018)	(.022)	(.018)
AFQT	.014	−.156	.023	.119
	(.012)	(.016)	(.020)	(.016)
Family background + AFQT	.023	−.158	.011	.123
	(.014)	(.020)	(.024)	(.019)

NOTE.—AFQT = Armed Forces Qualifying Test. The table reports the coefficient on an indicator variable for black in regressions in which the dependent variables are type probabilities. Standard errors are in parentheses.
* Includes father's schooling, mother's schooling, residence in nuclear family at age 14, residence in the South at age of 14, and number of siblings.

Table 8 presents race differentials in type probabilities based on a sequence of ordinary least squares regressions in which there are first no other controls, then in which the only controls are a set of family background variables, then only the Armed Forces Qualifying Test (AFQT) score, and, finally, both family background variables and the AFQT score. These regressions show the extent to which black-white differences in type probabilities can be accounted for by available measures of family background and ability. As the table shows, family background characteristics (parents' schooling, residence in a nuclear family at age 14, residence in the South at age 14, and number of siblings) and the AFQT together account entirely for the race differential in the probability of being a type 1, the type with the largest white-collar skill endowment (also the largest home productivity and schooling endowment), and of being a type 3, the type with the poorest endowment vector (last row). However, together they account for almost none of the race differential in the probability of being a type 2, the type with the highest blue-collar skill endowment, and about one-half of the differential in the probability of being a type 4, a mixed type in terms of the white- and blue-collar endowment.

These results need to be interpreted cautiously. Endowments at age 16 reflect the prior accumulation of investments by parents and youths as well as innate abilities. Although family background characteristics such as parental schooling in most cases are determined prior to the birth of the NLSY youths, their relationship to child investments may reflect heterogeneity in preferences for human capital or in abilities, that is, parental decisions about how much to invest in themselves and in their children are determined, in part, by this heterogeneity. It could be quite wrong to conclude that policies that equalize family background characteristics will also alter the race differential in type probabilities as depicted in table 8.[31]

It has recently been argued that the AFQT score is a powerful measure of skills and that it can account for much of the race difference in earnings (Neal and Johnson 1996). Now, in our model, there is not a unidimensional market skill but, rather, separate skills associated with white- and blue-collar occupations (as well as home production and schooling). In that context, it would seem implausible that a single measure of skills, like the AFQT, would adequately represent a multidimensional skill vector. The results in table 8 indeed indicate that although the AFQT is a powerful measure of white-collar skills (accounting for the entire black-white differential in the type 1 probability), it is a rather poor measure of blue-collar skills (accounting for almost none of the differential in the type 2 probability).[32]

Neal and Johnson (1996) found that controlling for AFQT reduced the racial wage gap (measured at age 27) from about 24% to 7%; they interpreted the latter as the maximum potential effect of discrimination. The estimates presented in table 7 showed that equalizing wage offers would reduce the racial gap in the present value of lifetime earnings from 26% to 16%, implying (at most) a 10-percentage-point gap because of discrimination. That these estimates are close is reassuring because the Neal and Johnson (1996) methodology that includes only endowments as wage determinants is itself consistent with the underlying behavioral model we estimate in which endowments and productivity shocks determine wages at any age. However, both of these estimates would understate the extent of discrimination if parental investments in children, which determine the AFQT score in the Neal and Johnson case or the endowment type proportions in ours, are lower for blacks as a result of the discrimination-induced reduction in skill rental prices faced by blacks.

A potentially important problem with interpreting the AFQT as an (age-16) endowment measure is that for most of the sample the test was not administered prior to age 16.[33] The AFQT score, thus, would be

[31] As an example, according to the quality-quantity fertility model (Rosenzweig and Wolpin 1980), parents who have preferences for higher-quality children will also have fewer children. Thus, a finding that youths with fewer siblings are more likely to be type 1s does not imply that a policy that reduced fertility would necessarily increase the percentage of type 1s. Neal and Johnson (1996), discussed below, do not make this distinction in estimating the "determinants" of AFQT.

[32] This finding would seem to be consistent with the usual interpretation of AFQT as an IQ measure. It is also not inconsistent with the finding in Neal and Johnson (1996) that it accounts for a large part of the wage differential, given that white-collar jobs have higher wages.

[33] The Armed Services Aptitude Battery (ASVAB), from which the AFQT is

influenced by whatever human capital was accumulated, either in school or while at work, between age 16 and the age at which the test was taken. As Neal and Johnson (1996) argue, in order to assess the relative importance of discrimination in the determination of wage differentials, initial skills must be measured prior to schooling and work decisions, which led them to restrict their sample to those who had taken the test prior to age 19 (a restriction similar to ours). However, even for that age-restricted sample, there was considerable variation in school attainment at the time of the test administration; thus, AFQT scores reflect that differential attainment as well as initial skills. But, as they argue and as is clear from our model, school attainment is itself affected by the existence of discrimination.[34] Thus, Neal and Johnson (1996) understate the extent of discrimination.[35]

E. School Performance-Based Subsidies

Schooling is costly, both in direct tuition costs for postsecondary education and in opportunity costs associated with employment and home production. One possible mechanism for increasing black male school attainment is to reduce the cost of schooling through targeted subsidies. Recently, Robert Reich, former secretary of labor in the Clinton administration, proposed a $25,000 bonus on high school graduation targeted at youths from families with incomes less than 120% of the median.[36] Although not directly targeted to black youths, eligibility would be highly skewed toward them. In the NLSY sample used in this

formed, was administered to approximately 94% of the NLSY sample in the summer of 1980. Our sample respondents were then age 16–19.

[34] Indeed, they report that an additional year of schooling, estimated for their 18-and-under sample, improves AFQT scores by about one-quarter of a standard deviation.

[35] Our method of controlling for endowment heterogeneity at age 16 avoids that particular problem.

[36] This proposal is not new. Performance-based monetary incentives have already been implemented on a small scale. For example, a philanthropic foundation, the "I Have a Deam Foundation," provided a guarantee of free college tuition to a sixth-grade class in New York. Although that program was reported to be successful, to our knowledge there was no scientific evaluation. In addition, Fischer (1987) proposed that college aid be made "graduation contingent." See Manski (1989) for a theoretical discussion of the impact of that proposal on college completion. More recently, the University of Arizona offered free tuition to 101 third-grade students in a South Tucson elementary school, provided that they graduate from high school and meet the University of Arizona's admission requirements. According to the Arizona Daily Wildcat (February 9, 1998), "The UA hopes the free education project will encourage the children to stay in school."

Table 9
The Impact of a High School Graduation Bonus on School Attainment

	Baseline		$25,000 Bonus for High School Graduation		$25,000 Bonus Targeted to Family Income ≤ 120% of Median Income: The Reich Proposal	
	White	Black	White	Black	White	Black
Mean highest grade completed	13.0	12.0	13.6	12.8	13.5	12.8
Highest grade completed ≤ 12 (%)	26.2	37.9	4.2	8.2	7.6	10.1
Highest grade completed = 12 (%)	29.6	31.5	41.5	49.2	39.6	47.8
Highest grade completed ≥ 13 and ≤ 15 (%)	19.1	17.9	26.2	27.4	26.2	26.4
Highest grade completed ≥ 16 (%)	25.1	12.8	28.1	15.2	26.7	15.8
High school graduate receiving bonus (%)	0	0	100	100	56.4	90.3
High school dropouts receiving bonus (%)	0	0	84.0	78.4	66.6	71.7

analysis, 92% of black male youths and 62% of white male youths had a family income (in 1978) below 120% of the median.

Table 9 reports simulations of the impact on completed schooling distributions of a $25,000 bonus for high school graduation, first globally and then targeted by family income. The global graduation bonus dramatically reduces the percentage of high school dropouts both for whites and blacks, dropping from 26.2% to 4.2% for whites and from 37.9% to 8.2% for blacks. Moreover, a considerable proportion of those who would have dropped out of high school not only graduate from high school but also obtain some college, and a small proportion even graduate from college. Overall, mean schooling increases by .6 years for whites and by .8 years for blacks, reducing the black-white schooling differential from 1.0 to .8 years.

One problem with the global bonus is that all of those who would have graduated from high school without the additional incentive nevertheless receive the bonus. One way to maintain the incentive aspect of the program while minimizing its purely redistributive aspect is to target the bonus to groups who have low graduation rates. A measure of the efficiency of the targeting mechanism—family income in the Reich (1998) proposal—is to compare the percentage of (initial) high school graduates who receive the bonus to the percentage of (initial) high school dropouts who receive the bonus. In the ideal mechanism, the former would be 0%

and the latter 100%. In the untargeted program, as the last two rows in table 9 indicate, 100% of the high school graduates receive the bonus, while 84% of the white high school dropouts and 78.4% of the black high school dropouts receive the bonus.

Given the very high percentage of blacks who fall within the "low" family income group, the effect of the targeted bonus on black school completion is almost the same as that of the untargeted program, as is the efficiency of the program. However, the efficiency of the program greatly improves for whites. There is only a small drop in the proportion of high school dropouts who receive the bonus (66.6% vs. 78.4%), while a little more than one-half of high school graduates receive the bonus (vs. 100% for the untargeted program). Notice that the black-white differential in mean schooling falls only slightly more to .7 years.[37]

Suppose we could design a bonus scheme that could be targeted only to blacks and that essentially eliminated differences in the black-white schooling distribution.[38] It turns out that providing a bonus of $6,250 for high school graduation and $15,000 for college graduation would accomplish such an equalization. The last column in table 7 shows the effects of this bonus scheme on completed schooling, earnings, and welfare.[39] Perhaps surprisingly, although the bonus program equalizes schooling, its effect on later labor market success is quite small. In particular, labor market outcomes are much the same as those obtained from equalizing initial schooling distributions, the first experiment in table 7. The bottom line is that with respect to the summary welfare measures, the black-white gap is reduced by only 4.8% in the case of the PDV of earnings and by only 8.7% in the case of the PDV of utility.[40]

[37] We also experimented with two other targeting factors, whether the youth's mother was a high school dropout and whether the youth resided in a nuclear family at age 14. Both worked similarly to family income in that the targeting worked better for whites than for blacks in terms of program efficiency. Based on whether the youth's mother was a high school dropout, for blacks 46% of the high school graduates and 48% of dropouts would receive the bonus. The comparable figures for whites were 18% and 46%. Similar figures based on whether the youth resided in a nuclear family at age 14 are 27% and 29% for blacks and 13% and 33% for whites.

[38] Perhaps one could target the bonus scheme to inner-city schools, which is the way it has been implemented before on a small scale.

[39] We also experimented with an annual college tuition subsidy to blacks of $5,000 and a comparable (in present-value terms) college graduation bonus of $25,000. The yearly college bonus reduced the percentage of high school dropouts by 10 percentage points and increased the college graduation rate by 21 percentage points. The college graduation bonus had a slightly larger impact on college graduation rates.

[40] We calculated the total annual cost of this bonus program, based on a black male cohort size of 250,000, to be $2.1 billion (assuming equilibrium effects on

Given the small impact of graduation bonuses on earnings or welfare, its rationale must hinge on the existence of schooling externalities. Of course, this is not an unusual argument for school subsidies in general. However, an alternative, to which we now turn, is to subsidize wages directly.

F. Wage Subsidies

We have already reported in table 7 on a wage subsidy scheme that eliminated the potential discrimination-induced differential in skill rental prices. Recall that the implied increase in white- and blue-collar wage offers, of approximately 8% and 5% (see appendix table A1), closed about 40% of the earnings gap.[41] We do not know, however, what part of each occupation-specific wage gap is because of discrimination. Other than direct antidiscrimination programs, it is therefore more usual to see wage policies that are targeted by income group rather than by race (e.g., a minimum wage).

One proposal targeted at low-wage workers in general, suggested by Phelps (1997), is to provide an indirect subsidy to workers with wages below a fixed amount. Table 10 shows the impact on schooling, earnings, and welfare of two such subsidy schemes, one that provides a subsidy of 50% of the difference between a wage rate of $7.50 per hour and the accepted wage and a second that provides a subsidy of 33% of the difference between a wage rate of $12.00 per hour and the accepted wage.[42] The latter experiment conforms more closely to the actual subsidy level proposed by Phelps (1997).[43] Consider the impact of the smaller

---

skill rental prices to be small). We also calculated that an efficiently targeted program, one in which we could identify ex ante who would be the "new" graduates (i.e., those who would not have graduated without the bonus) based on observable characteristics, would cost $.68 billion, or approximately one-third the cost of the untargeted program. From the private perspective, because the model assumes that individuals are optimizing and that there are no financing constraints on school attendance, the (monetary value of the) utility gain to any individual black male can be no greater than the bonus itself. If total expenditures on the program ($2.1 billion) were simply divided equally among the cohort of black males, each cohort member would receive $8,464. However, the present discounted value of utility increases by only $4,516 because of the bonus scheme. Black males are induced to choose a different path, acquiring more schooling than would be optimal without the bonus, which lowers utility relative to the lump-sum transfer.

[41] It is important to recognize that this is a partial equilibrium effect. The actual market equilibrium rental prices will depend on the demand and supply elasticities of skill in each occupation.

[42] The wage is in 1993 dollars. A wage of $7.50 ($12.00) per hour is equivalent to full-time (2,000 annual hours) earnings of $15,000 ($24,000).

[43] Phelps proposes that the subsidy be paid to firms. In the simulations that follow we assume that the entire subsidy is passed on to workers. As we discuss

Table 10
The Effect of Wage Subsidies on Completed Schooling, Earnings, and Welfare

			Phelps Wage Subsidy			
	Baseline		$7.50		$12.00	
	White	Black	White	Black	White	Black
Mean highest grade completed	13.0	12.0	12.7	12.0	12.4	11.8
Highest grade completed < 12 (%)	26.2	37.9	32.4	37.4	35.8	41.2
Highest grade completed = 12 (%)	29.6	31.5	30.8	34.9	31.0	34.9
Highest grade completed ≥ 13 and ≤ 15 (%)	19.1	17.9	15.4	16.6	13.9	15.0
Highest grade completed ≥ 16 (%)	25.1	12.8	21.4	11.1	19.3	9.0
Net earnings at age:						
30	22,795	15,818	23,062	17,203	23,678	18,296
40	32,635	24,626	32,723	25,783	32,924	26,340
50	44,014	34,120	43,786	34,927	43,781	35,244
Mean present discounted value of:						
Earnings	285,400	210,258	291,972	228,971	303,797	244,680
Utility	316,898	264,886	319,890	272,274	326,870	280,582

subsidy, first on blacks. The subsidy has almost no impact on the school completion distribution, a measurable positive effect on earnings (the present discounted value of earnings increases by 9%), and a small positive impact on welfare (the present value of utility increases by 3%). For whites, the subsidy at that level produces a "perverse" impact on schooling, reducing mean schooling by .3 years. The Phelps (1997) proposal takes schooling levels as given, whereas the subsidy to low-wage workers reduces the return to schooling. In the case of whites, given their initial wage distribution, the subsidy scheme significantly increases the relative wage of high school dropouts but does so only negligibly for blacks. Although the reduced schooling offsets the overall effect of the wage subsidy on earnings, the present value of earnings and welfare increase for whites, though by a small amount (2% and 1%).[44]

below, in general equilibrium that will not necessarily be the case. Of course, it is immaterial as to whether the firm or the worker actually receives the payment.

[44] Welfare must increase, or at least cannot fall, because individuals in the model always make optimal decisions, and they are free to maintain the same levels of schooling. Note that as with a wage subsidy to firms, this effect is partial equilibrium. Market equilibrium occupation-specific wages will be changed as the subsidy draws more people into the lower paying occupations and out of the

The larger subsidy magnifies these effects. Now mean schooling falls for both whites and blacks, by .6 years for whites and by .2 years for blacks, and the present value of earnings (utility) rises by 6% (3%) for whites and by 16% (6%) for blacks. Phelps's (1997) wage subsidy scheme was not specifically designed to ameliorate racial earnings differences. However, the black-white earnings ratio (in present value terms) does increase from .74 to .81. Moreover, if the subsidy were directed only at blacks, the black-white earnings ratio would rise to .86. Thus, if closing the racial wage gap were itself a goal, even larger subsidies targeted more directly to black youths would be necessary.

V. Conclusions

We have structurally estimated a human capital investment model using longitudinal data on the school attendance, occupational choice, and wages of young white and black men. The model assumes forward-looking behavior in which the consequences for future rewards are optimally accounted for in making current school attendance and work-occupation decisions. The dynamic model is able to rationalize the lower employment and school attendance rates of young black males relative to white males as an optimal response to the combination of discrimination against blacks in wage offers and lower endowments at age 16 of labor market, school, and home production skills. In fact, we could not reject the hypothesis that all of the model's parameters (discount factors, the technology of occupation-specific skill production, nonpecuniary valuations of occupations, etc.), except for those that capture discrimination or age-16 endowments, are the same for blacks and whites.

However, the lower school attainment and work attachment of black youths is also consistent with their having less concern for future payoffs. Although (annual) discount factors for both blacks and whites when separately estimated were of similar magnitude (around .93), we also estimated the static formulation, in which only immediate payoffs are considered in decision making, in order to contrast its implications for life cycle behavior. The forward-looking model produced far more reasonable out-of-sample life cycle forecasts of choice behavior and wage realizations for both blacks and whites than did the myopic model that ignores the human capital investment components of schooling and work.

Our framework enabled us to provide an upper-bound estimate of the impact of eliminating wage discrimination on the school attainment and

---

home and as it changes the incentives for schooling. The amount of the subsidy actually going to workers will depend on the occupation-specific skill supply and demand elasticities. Given these equilibrium effects it is difficult to estimate the fiscal cost of the program or the ultimate impact on wages.

labor market success of black males. Our results indicate that eliminating wage discrimination would at most reduce the fraction of black males who drop out of high school by 14% (about 44% of the black-white gap) and increase the fraction who graduate from college by 20% (about 20% of the black-white gap). It would also increase the expected present value (at age 16) of lifetime earnings for a typical black male from 73.7% to 83.8% of that for a typical white. The large differences that remain are attributable to "endowment" differences between blacks and whites that already exist at age 16.

We also used our estimates to explore the impact on school attainment and earnings of a recent proposal by Reich (1998) to provide a $25,000 bonus to all high school graduates who come from families with income below 120% of the median. We found that such a proposal would reduce the fraction of high school dropouts by two-thirds for blacks and by one-half for whites. However, even a bonus scheme that could be directly targeted at blacks and that equalized the schooling distributions would have only a minimal impact on the racial lifetime earnings differential. The performance bonus program, although obviously creating powerful incentives for increased schooling, does not alter the fundamentals that account for the existing race differences in school attainment, namely, the age-16 endowment differences. Similarly, we found that a Phelps-type (1997) wage subsidy, even if targeted at blacks (which was not the original intention), would close less than 50% of the racial lifetime earnings gap and would actually reduce schooling levels.

In our model, individuals begin making independent decisions as of age 16. At that time, their skill endowments are given and immutable. However, these "endowments" reflect prior parental investment decisions as well as innate traits. Interventions that alter the incentives of youths as decision makers, such as eliminating wage discrimination or providing school performance bonuses or wage subsidies, may also alter the incentives of parents to invest in their children. For example, in a world with school performance bonuses, parents may be more likely to invest in their children's acquisition of "schooling" skills. In addition, there might exist important intergenerational effects. If the school attainment of blacks is increased for one cohort, then these more educated parents might invest more in their children, who form a subsequent cohort, giving these children higher age-16 endowments and reducing the bonus payments for that cohort, which is necessary to equalize schooling distributions. A complete benefit-cost assessment of such interventions therefore would require knowledge of the child investment process.

Our results indicate that equalizing age-16 endowments would by itself go a long way toward eliminating race differences in labor market success. If the proportions of blacks of each age-16 endowment type were set equal to that of whites, then the present value of lifetime earnings (utility) for a typical black would increase from 73.7% (83.6%) to 89.9% (94.2%) of that for a typical white. It would be tempting to conclude from our results that interventions that seek to increase those endowments would be more efficacious than the other interventions we have analyzed. However, such a conclusion would require cost-benefit calculations of the impact of particular policies aimed at increasing those endowments, and all such analyses would have to be sensitive to the general equilibrium considerations we have mentioned.

Appendix

The complete specification of the model is

$$\begin{aligned} R_{mk}(a) = {} & w_{mk}(a) - c_{m1} \cdot I[d_m(a-1)=0] - c_{m2} \cdot I[x_m(a)=0] + \alpha_m \\ & + \beta_1 I[g(a) \geq 12] + \beta_2 I[g(a) \geq 16] \\ & + \beta_3 I[x_3(a)=1], \quad m = 1, 2 \end{aligned}$$

$$\begin{aligned} R_{3k}(a) = {} & \exp[\alpha_3(a)] w_3(a) - c_{32} \cdot I[x_3(a)=0] \\ & + \beta_1 I[g(a) \geq 12] + \beta_2 I[g(a) \geq 16], \end{aligned}$$

$$\begin{aligned} R_{4k}(a) = {} & e_{4k}(16) - tc_1 \cdot I[12 \leq g(a)] - tc_2 \cdot I[g(a) \geq 16] \\ & - rc_1 \cdot I[d_4(a-1)=0, g(a) \leq 11] - rc_2 \cdot I[d_4(a-1) \\ & = 0, g(a) \geq 12] + \beta_1 I[g(a) \geq 12] + \beta_2 I[g(a) \geq 16] \\ & + \beta_3 I[x_3(a)=1] + \gamma_{41} a + \gamma_{42} I[16 \leq a \leq 17] + \varepsilon_4(a), \end{aligned} \tag{A1}$$

$$\begin{aligned} R_{5k}(a) = {} & e_{5k}(16) + \beta_1 I[g(a) \geq 12] + \beta_2 I[g(a) \geq 16] \\ & + \beta_3 I[x_3(a)=1]=1) + \gamma_{51} I(18 \leq a \leq 20) \\ & + \gamma_{52} I(a \geq 21) + \varepsilon_5(a), \end{aligned}$$

where $I(\cdot)$ is an indicator function equal to one if the term inside the parentheses is true and equal to zero otherwise. In equation (A1), the skill technology functions that determine the wage are

$$\begin{aligned} e_m(a) = \exp\Bigg\{ & e_m(16) + e_{m11} g(a) + e_{m12} I[g(a) \geq 12] + e_{m13} I[g(a) \geq 16] \\ & + e_{m2} x_m(a) - e_{m3} x_m^2(a) + e_{m4} I(x_m > 0) + e_{m5}(a) \\ & + e_{m6} I(a < 18) + e_{m7} d_m(a-1) \\ & + \sum_{n(\neq m)=1}^{3} e_{m,n+7} x_n(a) \Bigg\} \exp[\varepsilon_m(a)], \end{aligned}$$

$$m = 1, 2;\ a = 16, \ldots, 65. \tag{A2}$$

Table A1
Estimated Occupation-Specific Parameters

	White-Collar	Blue-Collar	Military
1. Skill functions:			
Schooling	.0709 (.0010)	.0204 (.0009)	.0656 (.0023)
High school graduate	−.0123 (.0041)	.0063 (.0032)	. . .
College graduate	.0173 (.0044)	.0207 (.0050)	. . .
White-collar experience	.0267 (.0004)	.0199 (.0004)	. . .
Blue-collar experience	.0229 (.0005)	.0463 (.0001)	. . .
Military experience	.0165 (.0011)	.0173 (.0008)	.0443 (.0022)
"Own" experience squared/100	−.0422 (.0017)	−.0760 (.0011)	−.0442 (.0063)
"Own" experience positive	.2192 (.0071)	.1551 (.0085)	.1651 (.0189)
Previous period same occupation	.2905 (.0081)	.2002 (.0085)	. . .
Age*	.0107 (.0002)	.0119 (.0001)	.0092 (.0014)
Age less than 18	−.1144 (.0409)	−.0629 (.0245)	−.1372 (.0350)
Constants:			
Type 1	8.8640 (.0060)	8.7977 (.0043)	8.420 (.0207)
Deviation of type 2 from type 1	−.0827 (.0047)	.3131 (.0059)	. . .
Deviation of type 3 from type 1	−.5827 (.0061)	−.1873 (.0061)	. . .
Deviation of type 4 from type 1	−.5285 (.0104)	−.0797 (.0061)	. . .
Differential for blacks	−.0791 (.0046)	−.0471 (.0039)	−.0727 (.0061)
True error standard deviation	.3997 (.0062)	.3904 (.0058)	.2527 (.0160)
Measurement error standard deviation	.2408 (.0096)	.1931 (.0092)	.1717 (.0168)
Error correlation:			
White-collar	1.0000	. . .	. . .
Blue-collar	.0707 (.0282)	1.0000	. . .
Military	.0020 (.0791)	.4642 (.0594)	1.000
2. Nonpecuniary values:			
Constant	−2,522 (131)	−3,101 (118)	.1388 (.0290)
Age	. . .	. . .	.0281 (.0039)
3. Entry cost:			
If positive own experience but not in occupation in previous period	1,183 (210)	1,136 (147)	. . .
Additional entry cost if no own experience	3,375 (609)	580 (474)	518 (304)
4. Exit cost:			
One-year military experience	. . .	. . .	1.362 (99)

NOTE.—Standard errors in parentheses.
* Age is defined as age minus 16.

Table A2
Estimated School and Home Parameters

	School	Home
Constants:		
Type 1	17,876 (1,405)	22,437 (396)
Deviation of type 2 from type 1	−6,920 (1,265)	−2,258 (421)
Deviation of type 3 from type 1	−12,182 (1,295)	−13,692 (383)
Deviation of type 4 from type 1	−6,656 (1,282)	−6,439 (418)
Has high school diploma	802 (94)	. . .
Has college diploma	1,821 (200)	. . .
Net tuition costs: college	5,684 (660)	. . .
Additional net tuition costs: graduate school	4,492 (1,269)	. . .
Cost to reenter high school	23,609 (1,366)	. . .
Cost to reenter college	10,179 (778)	. . .
Age*	−2,101 (151)	. . .
Age 16–17	1,993 (723)	. . .
Age 18–20	. . .	859 (288)
Age 21 and above	. . .	1,358 (301)
Error standard deviation	14,362 (840)	6,612 (255)
Discount factor	.9340 (.0013)	. . .

NOTE.—Standard errors are in parentheses.
* Age is defined as age minus 16.

Table A3
Estimated Type Proportions by Initial Schooling Level and Type-Specific Endowment Rankings

Initial Schooling	Type 1	Type 2	Type 3	Type 4
Whites:				
9 years or less	.0339 (–)	.2435 (.0316)	.4012 (.0356)	.3215 (.0363)
10 years or more	.1688 (–)	.3151 (.0216)	.3527 (.0220)	.1634 (.0197)
Black differentials:				
9 years or less	−.0262 (–)	−.1264 (.0403)	−.0349 (.0479)	.1875 (.0503)
10 years or more	−.1053 (–)	−.1712 (.0295)	.0870 (.0349)	.1895 (0.355)

NOTE.—Standard errors are in parentheses. Dashes in parentheses signify that there are no standard errors.

References

Ben Porath, Yoram. "The Production of Human Capital and the Life Cycle of Earnings." *Journal of Political Economy* 76 (August 1967): 352–65.

Cameron, Steven, and Heckman, James J. "The Nonequivalence of High School Equivalents." *Journal of Labor Economics* 11 (January 1993): 1–47.

———. "The Dynamics of Educational Attainment for Blacks, Hispanics, and Whites." Photocopied. Chicago: University of Chicago, 1998.

Donahue, John, and Heckman, James J. "Continuous vs. Episodic Change: The Impact of Civil Rights Policy on the Economic Status of Blacks." *Journal of Economic Literature* 29 (December 1991): 1603–43.

Fischer, F. "Graduation-Contingent Student Aid." *Change* (1987), pp. 40–47.

Heckman, James J. "The Incidental Parameters Problem and the Problem of Initial Conditions in Estimating Discrete Time Discrete Data Stochastic Processes and Some Monte Carlo Evidence." In *Structural Analysis of Discrete Data with Econometric Applications,* edited by Charles F. Manski and Daniel McFadden. Cambridge, MA: MIT Press, 1981.

Heckman, James J., and Sedlacek, Guilherme. "Heterogeneity, Aggregation, and Market Wage Functions: An Empirical Model of Self-Selection in the Labor Market." *Journal of Political Economy* 93 (December 1985): 1077–1125.

Keane, Michael P., and Wolpin, Kenneth I. "The Solution and Estimation of Discrete Choice Dynamic Programming Models by Simulation and Interpolation: Monte Carlo Evidence." *Review of Economics and Statistics* 76 (November 1994): 648–72.

———. "The Career Decisions of Young Men." *Journal of Political Economy* 105 (June 1997): 473–522.

Kominski, Robert, and Adams, Andrea. *Educational Attainment in the United States: March 1993 and 1992.* U.S. Department of Commerce, Bureau of the Census. Current Population Reports, Population Characteristics P20–476. Washington, DC: Government Printing Office, 1994.

Manski, Charles. "Schooling as Experimentation: A Reappraisal of the Postsecondary Dropout Phenomenon." *Economics of Education Review* 8 (1988): 305–12.

Neal, Derek A., and Johnson, William R. "The Role of Premarket Factors in Black-White Wage Differences." *Journal of Political Economy* 104 (October 1996): 869–95.

Phelps, Edmund S. *Rewarding Work: How to Restore Participation and Self-Support to Free Enterprise.* Cambridge, MA: Harvard University Press, 1997.

Reich, Robert. Op. Ed. *New York Times* (January 9, 1998).

Rosenzweig, Mark R., and Wolpin, Kenneth I. "Testing the Quality-Quantity Fertility Model: Results from a Natural Experiment Using Twins." *Econometrica* 48 (January 1980): 227–40.

Roy, A. D. "Some Thoughts on the Distribution of Earnings." *Oxford Economic Papers* 3 (June 1951): 135–46.

Smith, James P., and Welch, Finis. "Black Economic Progress after Myrdal." *Journal of Economic Literature* 27 (June 1989): 519–64.

Willis, Robert J. "Wage Determinants: A Survey and Reinterpretation of Human Capital Earnings Functions." In *Handbook of Labor Economics,* edited by Orley C. Ashenfelder and Richard Layard, 1:525–602. Amsterdam: North Holland, 1986.