ESTIMATING THE EFFECT OF RACIAL DISCRIMINATION ON FIRST JOB WAGE OFFERS

Zvi Eckstein and Kenneth I. Wolpin*

Abstract—In this paper we develop and implement a method for bounding the extent to which labor market discrimination can account for racial wage differentials. The method is based on a two-sided, search-matching model that formally accounts for unobserved heterogeneity and unobserved offered wages. We find that racial differences in offered wages are proportionately twice (three times) as large as racial differences in accepted wages for high-school dropouts (high-school graduates). The results indicate that discrimination could account for the entire racial wage-offer differential for high-school dropouts and for high-school graduates, i.e., the bound on the extent of discrimination is not informative.

I. Introduction

In this paper, we develop and implement a method for bounding the extent to which labor market discrimination can account for group-based wage differentials. The vast empirical literature on estimating the determinants of black-white wage differentials has foundered on providing a compelling method for quantitatively determining the relative importance of wage discrimination and skill differentials.[1] The difficulty is in large part one of measurement. Skill bundles are inherently unobservable and an individual's wage is determined by both the skill bundle and its per-unit market valuation, which is also unobservable.[2] Measuring discrimination as the wage differential at a point in time net of the effect of a small set of observable characteristics that are a priori related to skill bundles (e.g., schooling, work experience) will misstate the extent of discrimination if measured characteristics explain only part of racial skill differentials. Although the set of presumed correlates of skills has expanded with the growth in new data sources (and race differences have persisted), it is unlikely that we will ever be able directly to measure skill bundles or to collect significantly more-convincing proxies.

A second difficulty in measuring wage discrimination is that observed wages do not correspond to offered wages if individuals engage in job search. For example, if blacks face higher job search costs, they will accept lower-wage jobs even if wage offers are not discriminatory and blacks are equally productive. Thus, wage discrimination would be inferred when it did not exist. On the other hand, observed differences in wages may understate the extent of wage discrimination. In the standard infinite-horizon job search model (Lippman & McCall, 1979), a one-dollar increase in the mean of the wage-offer distribution increases the reservation wage by less than one dollar (e.g., see Mortensen, 1986). Thus, the mean observed-wage differential will understate the mean offered-wage differential, and therefore understate the extent of discrimination. Although, it is a priori unclear as to the direction of the bias, if any, from using observed-wage differentials to measure wage discrimination, the discussion implies that specifying a model of job acceptance is necessary in testing for wage discrimination.

Although it is not feasible to resolve the fundamental identification problem, it is possible to place sufficient structure so as to identify an upper bound on the extent of discrimination. The additional structure we impose formally accounts for the nonobservability of skill bundles by positing there to be a discrete mixture of productivity types in each race group, and for the nonobservability of offered wages by positing a job acceptance strategy consistent with a two-sided search-matching-bargaining model (Diamond & Maskin, 1979; Mortensen, 1982; Wolinsky, 1987). We show that the identification of an upper bound estimate depends on the existence of a particular set of testable restrictions on the distribution of productivity types between the race groups.[3] Whether or not the upper bound estimate is in fact informative—i.e., less than 100% of the racial wage differential is an empirical issue.

We use data from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience on the duration to the first full-time job after the completion of school and the wage at that job. Individuals are classified by their completed schooling levels into three groups: high-school dropouts, high-school graduates, and college dropouts. Table 1 provides descriptive statistics on search durations (in calendar quarters) and accepted wages by race and schooling.[4] Clearly, there are large differences in the extent of search and in search outcomes by race and schooling. Mean duration is about two to three quarters longer for blacks at the lower two school completion levels than for whites, and one quarter longer for those who attended (but did not graduate) from college. The mean (accepted) wage is approximately 15% lower for blacks, with the differential being similar across schooling groups.[5]

Received for publication August 8, 1995. Revision accepted for publication November 12, 1998.

* Tel Aviv University and Boston University, and University of Pennsylvania, respectively.

We have benefited from discussions with Chris Flinn, Ed Green, Roni Shachar, and Yoram Weiss and from the comments of two anonymous referees. Oded Bizan, Fizza Gillani, and Gefan Zhu provided excellent research assistance. We are grateful for the support from the National Science Foundation under grants SES-9109607 and SBR-9309636.

[1] Cain (1986), in his survey of the discrimination literature in the *Handbook of Labor Economics* concludes that "econometric work has been useful, but in my eyes more so far its descriptive content than for testing hypotheses or for providing estimates of causal relationships" (p. 781). See also Smith and Welch (1977, 1986) and Donohue and Heckman (1991) for a discussion and evidence on alternative explanations for racial wage gaps.

[2] We abstract from nonpecuniary aspects of compensation.

[3] Given the selection that arises from systematic job-acceptance rules, identification requires the use of data on both the duration of search and accepted wages.

[4] The data are described in more detail below.

[5] Smith and Welch (1986), using dicennial census data, reports that the black/white (weekly) wage ratio in 1980 for those with a median of five

The Review of Economics and Statistics, August 1999, 81(3): 384–392

With respect to the bias in using observed wages to measure differential market wage opportunities, we find that racial differences in offered wages are proportionately twice as large as racial differences in accepted wages for high-school dropouts and proportionately three times as large for high-school graduates (who do not go to college), at least for the first postschool full-time job. Thus, there are either significantly larger skill differentials or greater discrimination—or both—than is indicated by observed wage differentials. Our results, however, indicate that, for both high-school groups, the bound on the extent of discrimination is not informative. (The overidentifying restriction was rejected in the case of college dropouts.) On the other hand, as a byproduct of the search model formulation, we find that job offer rates do not differ by race, for given schooling, and thus discrimination does not take the form of fewer employment opportunities (given wage-offer distributions).

The rest of the paper is organized as follows. In the next section, we state more concretely the fundamental identification problem and describe our method for identifying the upper bound. Section III presents the search-matching model and the estimation method; section IV describes the data, and section V presents the results. Concluding remarks are given in section VI.

II. Identifying Wage Discrimination

Tests of discrimination based on the estimation of group-specific wage functions suffer from a basic indeterminacy. In a simple spot market framework, the individual's wage is the product of a market determined (equilibrium) skill rental price (r) and embodied skill quantity (s). Skill, endowed and acquired, depends on observable and unobservable individual characteristics. In the absence of discrimination, the competitive market skill rental price is the same for all individuals regardless of their group identification. Given that skill rental prices are not observed, group-specific wage differentials can arise either from discrimination or from unobserved components of productivity. This is a fundamental identification problem.

To be concrete, ignoring the selection issue (assume we observe wage offers), consider an observably homogenous group (say white male high-school graduates with no work experience) and suppose that within that group there exists K_w distinct, but unobservable (to us but not to firms), productivity types. Assume that the skill distribution for each type is log normal with type-specific mean μ_{kw} and variance σ^2_{kw}. In this case, the ln wage density is a mixture of normal distributions and identification of the K_w means, $\ln(r_w) + \mu_{kw}$, and variances, σ^2_{kw}, as well as the K_w type proportions, γ_{kw}, has been established (Quandt & Ramsey, 1978).[6] Clearly, from the estimates of type-specific means, we can also identify differentials in type-specific mean productivities, $\mu_{kw} - \mu_{k'w}$. However, we cannot separately identify the K_w skill means and the skill rental price.[7]

Consider now a second group (say black male high-school graduates with no work experience) with an analogous wage structure, i.e., there are K_b types with type-specific means μ_{kb}, type-specific variances σ^2_{kb}, type proportions γ_{kb}, and skill rental price r_b. As in the case for the first group, type-specific mean wages are identified, $\ln(r_b) + \mu_{kb}$ for all k. But, clearly, because neither r_w nor r_b is separately identified, it is not possible to test for discrimination (i.e., for whether rental prices are equal). Suppose, however, that the following restrictions are valid: the number of types is the same for both groups ($K_w = K_b = K$), and mean skill differences for the two groups are the same for each type ($\mu_{1w} - \mu_{1b} = \mu_{2w} - \mu_{2b} = \cdots = \mu_{Kw} = \Delta\mu$).[8] Now, these restrictions do not provide an estimate of the rental price difference, i.e., each type-specific mean wage differential is equal to $(\ln(r_w) - \ln(r_b)) + \Delta\mu$. However, an upper bound can be established for the difference in (ln) rental prices by setting the common difference in type-specific productivities to zero, $\Delta\mu = 0$, assuming that $\Delta\mu$ is positive as would seem to be the implicit assumption consistent with the discrimination literature.[9]

Under these restrictions, the (ln) wage differential, $\ln w_w - \ln w_b$, can be decomposed into three components, one due to rental price differences, a second due to the common type-specific differential, $\Delta\mu$, and a third due to differences in type proportions. Specifically, we can write the ln wage differential as

$$\ln w_w - \ln w_b = (\ln(r_w) - \ln(r_b)) + \Delta\mu + \sum_{k=1}^{K} (\gamma_{wk} - \gamma_{bk})\mu_{bk}.$$

The second and third terms in this expression represent the difference in ln wages due to productivity differences that

years of potential work experience was 0.87 for high-school dropouts, 0.83 for high-school graduates, and 0.89 for college noncompleters. Although we consider only first jobs and only those that were full-time, the black/white (quarterly) wage ratios from table one are 0.85, 0.87, and 0.83, well within the range (especially given the small sample sizes).

[6] The assumption of log-normality is not critical. What is necessary is that the parameters of the underlying distributions that form the mixture be identified. We chose the log-normal because it leads to the conventional ln wage formulation.

[7] This nonidentifiability would not be affected even if types were observed.

[8] More precisely, there exists an ordering of types such that the restriction holds.

[9] It is, of course, logically possible that the type-specific difference in mean productivities, after controlling for observables such as schooling and experience, are negative, i.e., that black workers are more productive than white workers for matched types. Even in this case, the "average" black worker could be less productive than the average white if black type proportions are skewed towards the less-productive types relative to whites. If $\Delta\mu < 0$ in fact, setting it to zero would then provide a lower bound estimate of the extent of discrimination. It is more plausible in this framework that black types are less productive than white types.

arise from the two sources, and their sum is expected to be positive. However, there is no reason why it will necessarily be true that the third term—the difference between the mean productivity of blacks if they had the white type proportions rather than their own type proportions—will by itself be nonnegative. It is thus possible that the upper bound estimate of the (ln) rental price difference obtained by setting $\Delta\mu$ to zero (in which case group skill differentials arise only because the type proportions differ by group), will exceed the ln wage difference. In this case, the upper bound estimate, because it exceeds 100% would be uninformative.

We have so far considered the case in which wage offers are observed. If only accepted wages are observed, then additional assumptions are necessary to estimate the mean wage offer. In a standard job search model, which satisfies the reservation wage property, it is well known that the mean wage offer can be estimated with additional parametric assumptions when there is a single productivity type (Flinn & Heckman, 1982).[10] If there is unobserved heterogeneity in productivities, as is necessary for identification of the upper bound estimate of discrimination, then we need to specify a job acceptance model that incorporates reservation wages that are also type specific.

III. The Model

We adopt a simple search-matching model that controls for selection and also allows for discrimination in both wage offers and in job offer rates.

A. Theory

Consider a market where a worker, who lives forever, meets a firm with probability P.[11] The value added by the worker to the firm's production, m, is known immediately to both parties upon meeting and is a random draw from the distribution function $F(m)$, assumed to be log-normal. Let $w(m)$ be the wage and $\pi(m)$ be the profits of the firm from a match of value m. Then, it is required that

$$w(m) + \pi(m) \le m. \tag{1}$$

Each worker can meet at most one firm in each period. If the firm and the worker arrive at an agreement about $w(m)$ and $\pi(m)$, the game is over and the worker receives that wage forever. If they do not agree, then they can search again during the next period. Time is assumed to be discrete.

[10] In such a model, a consistent estimate of the reservation wage is the lowest observed wage. The mean offered wage can be obtained from the mean accepted wage if the wage-offer distribution satisfies a recoverability property.

[11] In a more complete model, P would be determined by optimal search intensities chosen by workers and firms (Wolinsky 1987; Eckstein & Wolpin 1995).

If the firm and the worker reach an agreement, the worker's share is a fixed proportion of the match value, that is,

$$w(m) = \alpha m, \tag{2}$$

where $\alpha \in (0, 1)$ is the marginal share of a worker in the match (the worker's bargaining power). A higher value of α represents greater bargaining power by the worker.[12] In the axiomatic Nash model, α is an exogenous parameter; it is not derived from the fundamental structure of the labor market.[13] We interpret it as an index of the level of discrimination in the labor market, as it reveals differential payments unrelated to productivity. Note that this formulation leads to an ln wage function that is observationally equivalent to that derived for the spot market model, with the bargaining power parameter α equal to the skill rental price r.

The optimal search strategy in the above model is characterized by a constant reservation wage, w^*. Search proceeds until a wage offer w is received that is above w^*. In the steady-state equilibrium, the value of w^* is equal to m^*, where m^*, the reservation match value, is a function of the underlying parameters of $F(m)$, P and the worker and firm discount factors. (For example, see Wolinsky, 1987.)

The model provides a complete characterization of the joint distribution of the duration of search and accepted wages. The probability of accepting a job offer (P_m) is given by $1 - F(m)$, and the hazard rate is given by

$$h_m = P(1 - F(m^*)). \tag{3}$$

Given the assumption that m is log normal and that $w(m)$ is proportional to m, the density function of the wage conditional on acceptance of a job is given by

$$z(w(m)|m > m^*) = \frac{1}{w(m)\,\sigma_m\,\sqrt{2\pi}} \times \exp\left(-.5\left[\frac{\ln(w(m)) - (\ln(\alpha) + \mu)}{\dfrac{\sigma_m}{(1 - F(m^*))}}\right]^2\right) \tag{4}$$

[12] In the formal model using the axiomatic Nash bargaining equilibrium concept, (e.g., Wolinsky, 1987), $w(m) = \eta + \alpha m$, where $\eta = w^* - \alpha m^*$, and m^* is the reservation match value. In a Nash bargaining model, the value of α affects the reservation wage and the reservation match. If the parameters of the model for the firm and the worker are the same, and $\alpha = 0.5$, then $\eta = 0$. Here, we approximate the wage function by assuming that $\eta = 0$. For $\alpha \neq 0.5$, the assumption that $\eta = 0$ is not necessarily consistent with the Nash bargaining solution. We make this assumption in order to simplify the calculation of the likelihood function.

[13] In a strategic bargaining model, the bargaining parameter would not be a free parameter, but would be a function of the other parameters of the model (Wolinsky, 1987).

TABLE 1.—DURATION TO FIRST FULL-TIME JOB AND ACCEPTED WAGES BY RACE AND SCHOOLING

	Mean Duration Quarters**		Proportion	Accepted Quarterly Wages (1986 Dollars)		
	Simple	K-M	Completed Spells	Mean	Std. Dev.	S.D./Mean
White Males						
H.S. Dropouts	4.8 (184)*	5.5	.88	2698 (150)	1348	0.50
H.S. Graduates	2.5 (543)	2.7	.94	2892 (484)	1657	0.57
College Dropouts	1.1 (167)	1.1	.96	3625 (144)	1917	0.53
Black Males						
H.S. Dropouts	6.5 (174)	8.2	.79	2305 (128)	1006	0.44
H.S. Graduates	4.6 (322)	5.3	.87	2520 (261)	1349	0.53
College Dropouts	2.0 (84)	2.2	.92	3018 (64)	1756	0.58

Notes: * Sample size in parentheses.
** Simple = sample mean for complete and incomplete spells.
K-M = the Kaplan-Meier estimate for the mean duration.

where μ is the mean of the match productivity distribution and σ_m its dispersion.

B. Heterogeneity

Schooling, sex, age, and race have been found to be important correlates of search (unemployment) duration and wages. We restrict our attention to young males of a single cohort, and we assume that schooling levels are predetermined state variables in the transition from school to the first (full-time) job.[14] Individuals within this cohort are observably heterogeneous (to us) only in their level of schooling and race. We assume that the labor market is completely segmented by school completion levels. Unobservables that affect search behavior are assumed to separate individuals of a given schooling-race group into K distinct types who may differ in some or all of the fundamental parameters of the model. Heterogeneity (schooling, race, and type) among workers is fully observed by firms.[15] Because we model observed heterogeneity as being discrete (schooling and race), we can apply the above model to each race-schooling group separately.

It should be noted that allowing for unobserved heterogeneity can provide an explanation for the differences among race and/or schooling groups in duration and wages observed in table 1 (for example, the observation that blacks have lower mean accepted wages and longer durations of search at all levels of schooling). To illustrate the point, suppose that, within each race-schooling group, there are two types of individuals. The first type has a low match productivity mean relative to the second type. Further, suppose that workers with the lower mean also have a lower offer probability. For both reasons, the first type will have a lower reservation wage than the second. However, because of the lower offer probability, the hazard rate for the first type may be lower than that of the second type, leading to their having a longer mean duration of unemployment. Now, suppose that blacks are disproportionately of the first type, say because they have lower quality of schooling (unobserved). Then, they will have lower mean accepted wages and longer durations of search as in the data.[16]

C. Estimation Method

Consider the model for a given race and schooling group comprising K different types of individuals. (Group subscripts are ignored for convenience.)[17] We assume that each type may have a different job offer probability, P_k, a different value of the mean of the match productivity distribution, μ_k, and a different reservation wage, w_k^*. However, as in the spot market model, the bargaining power parameter is assumed not to differ by type. The proportion of type k in the population is γ_k. Furthermore, we assume that the observed wage for an individual of type k, w_k^0, is measured with a multiplicative error that is independent of the true wage; thus,

$$\ln w_k^0 = \ln w_k + u_k \tag{5}$$

where w_k is the true wage for type k, and u_k is the measurement error. We assume that the distribution of u_k is $N(0, \sigma_u^2)$, independent of k. Because $\ln w_k$ is normal, $\ln w_k^0$ is also normal with the same mean and with variance $\sigma^2 = \sigma_m^2 + \sigma_u^2$.

[14] The exogeneity of schooling is a common assumption in estimating wage functions and unemployment duration models. Willis and Rosen (1979) and Keane and Wolpin (1997) consider the effects of self-selection on schooling on wage functions.

[15] This assumption avoids complications that would arise from incomplete and/or asymmetric information.

[16] The same example can explain why accepted wages decline with duration for a given race-schooling group.

[17] The estimation method follows the development in Eckstein and Wolpin (1995). See also Flinn and Heckman (1982).

The joint probability that the wage offer received exceeds the reservation wage and that w^0 is the observed wage, given the individual is of type k, is given by

$$\begin{aligned} &Pr(w_k > w_k^*, w_k^0) \\ &= Pr(\ln(w_k) > \ln(w_k^*)|w^0)Pr(w_k^0) \\ &= 1 - \Phi\left[\frac{\ln(w_k^*) - \rho^2 \ln(w_k^0) - (1-\rho^2)(\ln(\alpha) + \mu_k)}{\rho\sigma\sqrt{(1-\rho^2)}}\right] \\ &\cdot \frac{1}{w_k^0}\frac{1}{\sigma}\phi\left[\frac{\ln(w_k^0) - (\ln(\alpha) + \mu_k)}{\sigma}\right] \end{aligned} \tag{6}$$

where $\rho^2 = \sigma_m^2/\sigma^2$ and Φ and ϕ are the normal c.d.f. and p.d.f., respectively.[18] Given data for a sample of $i = 1, \ldots, I$ individuals on completed search spell durations, d_i, and observed wages, w_i^0, the likelihood function is

$$\begin{aligned} L(\Psi) = \prod_{i\in I}\sum_{k=1}^{K} &\gamma k \times \left[1 - P_k\left(1 - \Phi\left(\frac{\ln(w_k^*) - (\ln(\alpha) + \mu_k)}{\rho\sigma}\right)\right)\right]^{d_i} P_k \\ &\times \left[1 - \Phi\left(\frac{\ln(w_k^*) - \rho^2 \ln w_{ki}^0 - (1-\rho^2)(\ln(\alpha) + \mu_k)}{\rho\sigma\sqrt{(1-\rho^2)}}\right)\right] \\ &\times \frac{1}{w_{ki}^0}\frac{1}{\sigma}\phi\left(\frac{\ln(w_{ki}^0) - (\ln(\alpha) + \mu_k)}{\sigma}\right), \end{aligned} \tag{7}$$

where ψ is the following vector of parameters

$$\begin{aligned} \psi = [&\gamma_1, \ldots, \gamma_K, P_1, \ldots, P_K, \ln(\alpha) \\ &+ \mu_1, \ldots, \ln(\alpha) + \mu_K, w_1^*, \ldots, w_K^*, \sigma, \rho]. \end{aligned} \tag{8}$$

The likelihood function in equation (7) is specified only for complete spells with reported accepted wages. If a spell is complete but the accepted wage is missing from the data, the last term of equation (7) is dropped and the term before the last is replaced with the term $(1 - \Phi(\ln(w_k^*) - (\ln\alpha + \mu_k)/\rho\sigma))$. For incomplete spells, the only part of the likelihood is the first term in equation (7), that is, $[1 - P_k(1 - \Phi(\ln(w_k^*) - (\ln\alpha + \mu_k)/\rho\sigma))]^{d_i}$.

It is clear from equation (7) that, if we estimate the model separately for each schooling-race group, then we cannot identify the share parameter as distinct from the means of the type-specific match productivities.[19] As in the spot market example, because the wage is the product of the match (m) and the share (α), ln α enters the likelihood function only additively with the type-specific match productivity means, μ_k. However, because the share parameter in the search-matching model is isomorphic to the rental price in the spot market model, an upper bound estimate of the race difference in the ln share parameters can be identified under the restrictions developed previously, namely that, for each level of schooling, the number of types is the same for both races and the type-specific mean skill levels differ by a constant amount.

IV. Data

The data for this analysis are from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience (NLSY). The NLSY consists of 12,686 individuals who were 14 to 21 years old as of January 1, 1979. It contains a nationally representative core random sample, an oversample of blacks and of Hispanics, and a special military oversample. Respondents have been interviewed annually since 1979. We make use of the data collected in the first eight personal interviews (1979 to 1986) for the white male core sample and the black male core and supplemental samples. We consider three schooling groups: high-school dropouts, high-school graduates, and college dropouts.[20] The highest grade completed is that schooling level achieved at the last school enrollment date (week) up to December 31, 1984.

We define the duration to the first job to be the number of calendar quarters, with the first calendar quarter being the one after the last school enrollment calendar quarter, before the individual began working at a full-time job. A full-time job is defined to be a job in which the individual worked at least thirty hours per week in the entire calendar quarter (thirteen weeks). The employment data span the period from January 1, 1978, through December 31, 1985. We exclude individuals who left school prior to January 1, 1978, because we do not know the employment history prior to that time.[21]

Descriptive statistics for all six race-schooling groups are presented in table 1. Because we observe individuals for at

[18] An individual will accept a wage offer whenever the match value (m) exceeds the reservation match value (m^*). However, the axiomatic Nash bargaining model is efficient, implying that this condition is satisfied whenever the wage exceeds the reservation wage.

[19] As already noted, identification of the model's parameters when there is a single type has been established. With more than one type, the lowest observed wage (ignoring measurement error) identifies the reservation wage (and thus the job offer rate) for only one type. Identifying the reservation wages (and offer rates) of the other types can be thought of as based on matching the observed higher-order moments of the accepted wage distribution with the theoretical moments that come from the convolution of the truncated accepted-wage distributions.

[20] There are too few black college graduates in the sample to allow an informative analysis.

[21] We also exclude individuals who ever served in the military and for whom there is longitudinally inconsistent data on school completion levels.

least a year after leaving school, the great majority of spells are complete. Black males who have never attended college (high-school dropouts and high-school graduates) take approximately 2.5 additional quarters to become employed at a full-time job relative to white males, while those with some college (but who did not graduate) take only one additional quarter. Not only are the durations longer for black males, but, once employed, their mean weekly wage is also lower, with black males receiving a wage approximately 85% that of white males at the same schooling level.[22] Also, white high-school dropouts and high-school graduates have greater wage dispersion, as measured by the coefficient of variation, than similarly educated blacks.

We used the Kaplan-Meier survivor functions to perform log-rank, Mantel-Haenszel, and Wilcoxson-Gehan tests (Kalbfliesch & Prentice, 1980) for the equality of the survivor functions for the search duration data. All three tests rejected (with marginal probability of less than 2%) equality of the survivor functions for blacks and whites of the same schooling level. The data also reveal that the hazard rate associated with entering full-time employment is declining with duration for all race-schooling groups (Eckstein & Wolpin, 1995).[23] The simplest of finite-horizon search models would predict the opposite. Explaining the pattern requires either the introduction of heterogeneity within these groups or some additional form of structural duration dependence. As we have noted, we rely on unobserved heterogeneity to fit the duration dependence in the hazard rate and the wage distribution.

V. Results

The benchmark specification is the unrestricted model in which blacks and whites (of the same schooling level) have completely different sets of parameters.[24] We began with five types for each race group and compared log likelihood values as the number of types was reduced. Based on likelihood ratio tests, we concluded that no reduction in the number of types was warranted for either black or white high-school dropouts and college dropouts ($K_w = 5, K_b = 5$), but that $K_w = 4$ and $K_b = 4$ were (separately) appropriate restrictions for both black and white high-school graduates.[25] Thus, estimation of the unrestricted model led us to the same number of types for blacks and whites within each schooling group.

We next sought to reduce the number of parameters of the unrestricted model, in order to improve the precision of our estimates, by testing the assumption that the type proportions are drawn from a binomial distribution with parameter γ, that is,

$$\gamma_k = \binom{K}{k} \gamma^k (1 - \gamma)^{K-k}. \tag{9}$$

This simplification, if acceptable, would reduce the number of parameters by K-2. We were able to accept the restrictions imposed by the binomial specification only for high-school graduates.[26]

Given the estimation results from the unrestricted model, we proceeded with the test of the restricted specification in which mean skill differences for the two groups are the same for each type, $\mu_{1w} - \mu_{1b} = \mu_{2w} - \mu_{2b} = \cdots = \mu_{Kw} - \mu_{Kb} = \Delta\mu$. We did not reject the restriction for the two lower schooling groups, but did reject it for the college dropout group. (Thus, no bound on discrimination can be estimated for the college dropout group.) To gain further estimation efficiency as well as to discern whether discrimination is reflected in differential job offer rates, we tested the restriction that blacks and whites faced the same offer probabilities (P_ks equal). We did not reject the restriction for either of the remaining schooling groups, implying that discrimination does not take the form of restricting job access (conditional on wage-offer distributions).[27]

Table 2 summarizes these results. The estimated likelihood values for each race-schooling group are presented for the different specifications, from the least to the most restrictive. The first column is the unrestricted model (black

[22] The quarterly wage is the weekly wage in the last week of the quarter times thirteen and is adjusted to 1982 dollars.

[23] A significant proportion of individuals have a full-time job in the first calendar quarter upon leaving school. This is due in part to the fact that some weeks may lapse between the time an individual leaves school during the calendar quarter and the beginning of the next calendar quarter. Although it may also reflect the start of job search prior to leaving school, nevertheless, these individuals are treated as having a zero duration of search, i.e., receiving an offer and accepting it at the beginning of the first period of search.

[24] The type-specific variances were constrained to be equal (within each race group) to conserve on parameters.

[25] For high-school dropouts, the joint black and white ln likelihood value was −3108.44 for five types and −3122.35 for four types. Twice the difference in the ln likelihood is 25.8 which is statistically significant at less than the 0.01 level (eight degrees of freedom). For high school graduates, the ln likelihood for five and four types, respectively, were −7948.9 and −7955.8. Twice the difference in the ln likelihood is 13.8, which is not statistically significant at the 0.05 level (eight degrees of freedom). For college dropouts, the ln likelihood for five and four types, respectively, were −2128.9 and −2143.0. Twice the difference in the ln likelihood is 28.2, which is statistically significant at less than the 0.01 level (eight degrees of freedom). The restriction that there are only three types is rejected for all of the schooling groups. Although the likelihood ratio test is not formally valid in this application because the restrictions being tested are at the boundary of the parameter space, informal comparisons of the fit of the specifications to the duration and wage data appear to be consistent with these test results.

[26] For high-school dropouts, the binomial restriction for five types was rejected at less than the 0.01 level. For high-school graduates, the binomial restriction for four types was not rejected at the 0.05 level; and for college dropouts, the binomial restriction for five types was not rejected at the 0.22 level.

[27] We further tested the hypothesis that racial differences in wages and unemployment durations are due solely to blacks and whites differing in their type proportions, i.e., in addition to the restrictions already imposed, we also added the restrictions that $\Delta\mu = 0$, $\alpha_w = \alpha_b$, $w^*_{kw} = w^*_{kb}$. We rejected the hypothesis for both schooling groups.

TABLE 2.—TESTS OF RESTRICTED SPECIFICATIONS

	Specification*				
	(1)	(2)	(3)	(4)	(5)
H.S. Dropouts: 5 Types with Free Distribution					
no. par.	42	36	31	25	21
Ln Likelihood	−3109.15	−3112.20	−3113.73	−3127.13	−3147.21
χ 2 TEST		6.10	9.16	35.96	76.12
SIG. LEVEL		0.41	0.61	0.00	0.00
H.S. Grads: 4 Types with Binomial Distribution					
no. par.	30	25	21	16	15
Ln Likelihood	−7958.22	−7961.96	−7962.00	−7979.74	−7992.22
χ 2 TEST		7.48	7.56	43.04	68.00
SIG. LEVEL		0.19	0.56	0.00	0.00
Coll. Dropouts: 5 Types with Binomial Distribution					
no. par.	36	30	25	19	18
Ln Likelihood	−2133.08	−2143.20	−2146.90	−2154.28	−2160.46
χ 2 TEST		20.24	27.64	42.40	54.76
SIG. LEVEL		0.00	0.00	0.00	0.00

Notes: * The specifications are defined as follows:
(1) Unrestricted—all the parameters in equation (8) are different for blacks and whites.
(2) The differences in the μs of blacks and whites are restricted to be equal.
(3) In addition to (2) above, the Ps of blacks and whites are restricted to be equal.
(4) Only the type proportions are different between blacks and whites.
(5) All the parameters of blacks and whites are the same.

and white parameters estimated separately). The second column restricts the mean productivities of each type to differ by the same amount, and the third further column restricts the offer probabilities to be the same for blacks and whites. The chi-squared test statistics for likelihood ratio tests comparing the restricted specifications to the unrestricted specification and the associated significance levels are reported in the last two rows for each group.

A. *Model Fit*

Chi-square fit tests were conducted for the unrestricted and the final restricted specifications by comparing predicted and actual duration distributions. In all cases, the unrestricted specifications were not rejected (at the 5% level). Similarly, restricted specifications were not rejected for black and white high-school dropouts and for black high-school graduates, although the restricted specification fits somewhat less well for white high-school graduates.[28]

B. *Parameter Values*

Table A.1 presents the estimated parameters (and standard errors) for the restricted specification for each schooling group.[29] Although those parameters are of some interest in their own right, we concentrate on summary measures that average over types, which are presented in table 3. Mean (over types) job offer rates are predicted by the model to be close to one for all schooling-race groups.[30] However, job acceptance rates, conditional on receiving an offer, are higher for whites than for blacks of the same schooling, but are always less than 0.5. In the case of high-school dropouts, the mean reservation wage exceeds the mean wage offer for both race groups, while the opposite is true for white high-school graduates.

The importance of modeling job search in quantifying wage opportunity differentials is illustrated by comparing the race differences in mean accepted and mean offered wages. As noted, white high-school dropouts (graduates) receive wages that are 15% (16%) above blacks with comparable schooling. However, the differential based on accepted wages considerably understates the wage-offer

TABLE 3.—PREDICTED CHARACTERISTICS OF THE DURATION AND WAGE DISTRIBUTIONS

Schooling Level	Specification (3) High-School Dropout		Specification (3) High-School Graduate	
Race	B	W	B	W
Predicted Mean duration	6.90	5.33	5.17	2.84
Predicted mean acceptance rate − P_m^1	0.18	0.25	0.27	0.44
Predicted mean offer rate − P^2	0.96	0.98	0.99	0.98
Predicted mean accepted wage[3]	2321.32	2666.49	2440.74	2838.48
Predicted mean offered wage[4]	981.46	1294.18	1303.48	1886.75
Predicted S.D. of offered wage[5]	919.32	1139.36	1108.34	1646.93
Predicted mean reservation wage[6]	1483.32	1719.46	1611.73	1667.29

Notes: [1] $P_m = \Sigma_k \gamma_k P_{mk}$.
[2] $P = \Sigma_k \gamma_k P_k$.
[3] The predicted Mean Accepted Wage for the particular race/schooling group is $w^A = \Sigma_{k=1}^K \lambda_k w_k^A$, where λ_k is the proportion of workers of type k looking for a job for d quarters, such that $\lambda_k^d = [\Sigma_{j=1}^d \rho_{kj}]/[\Sigma_{k=1}^K \Sigma_{j=1}^d \rho_{kj}]$ and where $w_k^A = (1/Pm_k)\exp\{\mu_k + 0.5\sigma_m^2 + \ln(\alpha)\}[1 - \Phi((\ln(w_k^*) - \sigma_m^2 - (\ln(\alpha) + \mu_k))/\sigma_m)]$.
[4] The Predicted Mean Offered Wage is $w^o = \Sigma_{k=1}^K \gamma_k w_k^o$, where $w_k^o = \exp\{\mu_k + 0.5\sigma_m^2 + \ln(\alpha)\}$.
[5] The variance of the offered wage is $\Sigma_{k=1}^K \gamma_k \sigma_k^2 + \Sigma_{k=1}^K \gamma_k w_k^o - (\Sigma_{k=1}^K \gamma_k w_k^o)^2$, where $\sigma_k^2 = \exp(\{\sigma_m^2 - 1)\exp\{2(\mu_k + \ln(\alpha)) + \sigma_m^2\}$ is the variance of the offered wage for type k.
[6] Mean reservation wage $= \Sigma_{k=1}^K \gamma_k w_k^*$.

[28] The source of the rejection is that the specification underpredicts entry into full-time employment in the first period. The test results are available from the authors and are similar to those in Eckstein and Wolpin (1995).

[29] As we previously explained, only the black/white ratio of shares is identified. The black share was set to 0.5 as a normalization.

[30] The fact that these estimates approach the boundary of the parameter space may be indicative of a lack of identification of type-specific offer probabilities and reservation wages in this data.

differential.[31] The latter, reflecting labor market opportunities, is 32% for high-school dropouts and 45% for high-school graduates.[32]

C. *The Upper Bound on the Extent of Discrimination*

We obtain an upper bound estimate of the proportion of the wage-offer differential by equating the type-specific mean productivities, $\Delta\mu = 0$. The estimated ratio of the bargaining power parameters (see table A.1) under that restriction, α^w/α^b, is 1.40 for high-school dropouts and 1.56 for high-school graduates.[33] Thus, as an upper bound, if blacks and whites were of equal productivity, white high-school dropouts would receive a wage offer that is 40% higher than black dropouts and white high-school graduates an offer that is 56% higher than black graduates. Unfortunately, these estimates imply that the upper bound estimate of the extent to which discrimination accounts for the wage offer differential is above 100% for both schooling groups ($1.40/1.32 = 1.06$ for high-school dropouts, and $1.56/1.45 = 1.08$ for high-school graduates), which is not informative.

VII. Concluding Remarks

There are at least three separate reasons for racial disparities in observed wages (for otherwise observationally identical individuals):

(i.) There exists racial wage discrimination.
(ii.) There exists unobserved skill differentials.
(iii.) There exist race differences in reservation wages (due to differences in job offer rates, search costs, or nonmarket opportunities).

Although the literature has recognized the existence of a fundamental identification problem that arises from the inherent nonobservability of skill levels and market skill rewards, no methods have been proposed to circumvent the problem, and the literature has ignored almost universally the endogeneity of wage outcomes.[34]

In this paper, we have proposed and implemented a methodology that potentially provides an informative upper bound estimate of the proportion of the market wage-offer differential due to discrimination, accounting for the fundamental identification problem and for job search behavior. The procedure relies on the nonrejection of a testable overidentifying restriction. We did not reject the restriction for two of the three schooling groups we considered. Our estimates revealed that the black/white wage-offer ratio is considerably less than the black/white observed wage differential for high-school dropouts and for high-school graduates (who did not attend college). Unfortunately, however, we also found for this sample that the upper bound estimate of the extent to which wage-offer differentials were due to discrimination was not less than 100% for both schooling groups. There is nothing inherent in the procedure that would lead to this result, and possibly other samples will provide more-informative estimates.

[31] Although this may seem to be a problem confined to the first postschool job, to the extent that initial accepted wages also determine future job search behavior (e.g., on-the-job search), observed wages at later life-cycle stages would also differ from offered wages. (See Wolpin (1992) for evidence on this point.)

[32] Note, however, that in absolute terms accepted wages actually overstate the differential for high-school dropouts (by $30), but understate the differential for high school graduates (by $185).

[33] The standard errors of these estimates are 0.10 and 0.12, respectively. (See table A.1.)

[34] But see the recent paper by Bowlus and Eckstein (1997).

REFERENCES

Bowlus, A., and Z. Eckstein. "Discrimination and Skill Differences in an Equilibrium Search Model," mimeograph, Tel Aviv University (1997).

Cain, G., "Labor market discrimination." In O. Ashenfelter and R. Layard (eds.), *Handbook of Labor Economics* (Amsterdam: North Holland, 1986), 693–785.

Diamond, P. A., and E. Maskin, "An equilibrium analysis of search and breach of contracts, I, steady states," *Bell Journal of Economics* 10 (1979).

Donohue, J. J., and J. Heckman, "Continuous versus episodic change: The impact of civil rights policy on the economic status of blacks," *Journal of Economic Literature* XXIX, 4 (1991), 1603–1643.

Eckstein, Z., and K. I. Wolpin, "Estimating an equilibrium search model from panel data on individuals," *Econometrica* 58 (1990), 783–808.

——— "Duration to First Job and the Return to Schooling: Estimates from a Search-Matching Model," *Review of Economic Studies* 62 (1995), 263–286.

Flinn, C., and J. Heckman, "New methods for analyzing structural models of labor force dynamics," *Journal of Econometrics* 18 (1982), 115–168.

Kalbfleisch, J. D., and R. L. Prentice, *The Statistical Analysis of Failure Time Data* (New York: John Willey & Sons, 1980).

Keane, M., and K. I. Wolpin, "The career decisions of young men," *Journal of Political Economy* 105 (1997), 473–572.

Lippman, S., and J. McCall, "The Economics of Job Search: A Survey," *Economic Inquiry* 14 (1976), 155–189.

Mortensen, D. T. "Job Search and Labor Market Analysis." In O. C. Ashenfelter and R. Layard (eds.), *Handbook of Labor Economics, Vol. II* (Amsterdam: North Holland, 1986), 849–919.

Mortensen, D. T., "The matching process as a noncooperative bargaining game." In J. J. McCall, (ed.) *The Economics of Information and Uncertainty* (Chicago: University of Chicago Press, 1982).

Mortensen, D. T., "Equilibrium Wage Distributions: A Synthesis," In J. Hartog, G. Ridder, and J. Theeuwes (eds.), *Panel Data and Labor Studies* (North-Holland: Elseveir Science Publishers B. V., 1990).

Quandt, R., and J. Ramsey, "Estimating Mixtures of Normal Distributions and Switching Regressions," *Journal of the American Statistical Society* 73 (1978), 730–738.

Rubinstein, A., "Perfect equilibrium in a bargaining model," *Econometrica* 50 (1982), 97–110.

Smith, J., and F. Welch, "Black-white wage ratios: 1960–1970," *American Economic Review* 67 (1977), 323–338.

——— "Closing the Gap: Forty Years of Economic Progress for Blacks," Rand mimeograph R-3330 (1986).

Willis, Robert J., and Sherwin Rosen, "Education and Self-Selection," *J. P. E.* 87, no. 5 pt. 2 (1979), S7–S36.

Wolinsky, A., "Matching, Search and Bargaining," *Journal of Economic Theory* 42 (1987), 311–333.

Wolpin, K. I., "Estimating a Structural Search Model: The Transition from School to Work," *Econometrica* 55 (1987), 801–817.

——— "The Determinants of Black-White Differences in Early Employment Careers: Search, Layoffs, Quits and Endogenous Wage Growth, *Journal of Political Economy* 100 (1987), 535–560.

TABLE A.1.—ESTIMATED PARAMETERS

	H.S. Dropouts, Specification 3, Five Types, Free Proportions				H.S. Graduates, Specification 3, Four Types, Binomial Proportions			
	Whites		Blacks		Whites		Blacks	
γ_1	0.02	(0.01)	0.04	(0.02)	0.03	(**)	0.02	(**)
P_1	0.15	(0.06)	0.15	(0.06)	0.22	(0.06)	0.22	(0.06)
μ_1	6.57	(0.24)	6.57	(0.24)	6.75	(0.18)	6.75	(0.18)
w_1^*	35.55	(261.0)	25.52	(115.1)	0.00	(*)	0.00	(*)
w_1^o	605.51		431.91		775.82		499.89	
γ_2	0.09	(0.04)	0.19	(0.05)	0.19	(**)	0.14	(**)
P_2	0.99	(*)	0.99	(*)	0.99	(*)	0.99	(*)
μ_2	6.35	(0.13)	6.35	(0.13)	6.79	(0.15)	6.79	(0.15)
w_2^*	1144.01	(90.1)	1006.52	(11.05)	1408.92	(153.94)	1177.69	(129.8)
w_2^o	487.05		347.42		806.11		519.41	
γ_3	0.14	(0.04)	0.35	(0.21)	0.20	(**)	0.42	(**)
P_3	0.99	(*)	0.99	(*)	0.99	(*)	0.99	(*)
μ_3	7.83	(0.14)	7.58	(0.14)	8.06	(0.10)	8.06	(0.10)
w_3^*	1000.64	(80.50)	1799.30	(18.40)	1467.30	(188.5)	1625.31	(73.05)
w_3^o	2139.15		1188.82		2866.01		1846.69	
γ_4	0.30	(0.06)	0.31	(0.22)	0.44	(**)	0.42	(**)
P_4	0.99	(*)	0.99	(*)	0.99	(*)	0.99	(*)
μ_4	7.58	(0.13)	7.83	(0.13)	7.50	(0.08)	7.50	(0.08)
w_4^*	2469.31	(194.15)	1842.45	(22.72)	2034.29	(199.17)	1805.16	(77.49)
w_4^o	1666.64		1525.87		1644.38		1059.54	
γ_5	0.45		0.25					
P_5	0.99	(*)	0.99	(*)				
μ_5	7.06	(0.11)	7.06	(0.11)				
w_5^*	1639.09	(128.27)	1380.43	(13.10)				
w_5^o	986.11		703.40					
σ_m	0.62	(0.04)	0.62	(0.04)	0.56	(0.04)	0.56	(0.04)
σ_e	0.000246	(*)	0.000246	(*)	0.18	(0.03)	0.18	(0.03)
α	0.70	(0.05)	0.50		0.78	(0.06)	0.50	

Notes: Standard Errors in parentheses. w_k^o is a calculated number.
* The parameter is on the boundary of the likelihood function and no S.E. is available.
** The binomial coefficient is 0.702 (s.e. = 0.023) for whites and 0.748(s.e. = 0.033) for blacks.