Review of Economic Studies (1995) 62, 263-286 0034-6527/95/00120263$02.00

Duration to First Job and the Return to Schooling: Estimates from a Search-Matching Model

ZVI ECKSTEIN
Tel Aviv University and Boston University

and

KENNETH I. WOLPIN
New York University

First version received September 1992; final version accepted October 1994 (Eds.)

This paper investigates the properties of the joint distribution of the duration to the first post-schooling full-time job and of the accepted wage for that job within a search-matching-bargaining theoretic model. The model provides an interpretation of the observations on duration to first job and accepted wages that differentiates between behavioural influences and market fundamentals in determining the accepted wage-schooling relationship. The return to schooling is appropriately measured by differences in the wage offer distribution, which depends only on market fundamentals. We use data from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience to follow several school-leaving cohorts of young males. A model which allows for five types of heterogeneous workers within schooling/race groups fits the duration and wage data well for all such groups. Offer probabilities for all groups are estimated to be close to one. Mean offered wages are about $1000 less than mean accepted wages and the internal annual rate of return for attending college relative to graduating from high school is 32% for blacks and 17% for whites.

1. INTRODUCTION

This paper investigates the properties of the joint distribution of the duration to the first post-schooling full-time job and of the accepted wage for that job within a search-matching-bargaining model. It is a well established empirical regularity that hazard rates are decreasing with the duration of search (unemployment). It is also true for the data we use that the sample mean of the accepted wage associated with the first job is not increasing with duration. These findings hold for both black and white male youth at all levels of schooling. Conditional on schooling, however, the time between leaving school and working at the first full-time job is longer for blacks and their accepted mean wage is lower. Conditional on race, the time between leaving school and working at the first full-time job is shorter for those with higher levels of schooling and their accepted mean wage is higher. Interestingly, blacks display a similar pattern of duration and accepted wages to whites who are one "level" of schooling below, e.g. black high-school graduates and white high-school dropouts.

Observed differences in mean accepted wages between schooling levels may provide a distorted picture of the return to schooling because not all the wage offers, that is, firm-worker matches, are accepted. The return to schooling is appropriately measured by differences in the wage offer distribution which depends only on market fundamentals, such as productivity, the wage bargaining process, etc., and not on worker or firm behaviour. In our sample, mean duration to the first (full-time) job varies from one to more than six quarters depending on schooling level and race, and differentials in mean accepted wages by schooling level varies between 7% and 26%. In order to determine whether these differences over- or under-estimate the return to schooling it is necessary to estimate the wage offer distribution. This requires a model of the wage acceptance process.

We use a search-matching-bargaining model to provide an interpretation of the observations on duration to first job and accepted wages that differentiates between behavioural influences and market fundamentals in determining the accepted wage-schooling relationship. In Flinn and Heckman (1982) the wage is determined as an exogenous split of the value of the match between the firm and the worker. The value of the match is sampled from a given distribution and the probability of receiving an offer is assumed to depend on the ratio of firms to workers in the market. We modify the model to the case where the value of a sampled match is divided between the worker and the firm according to a Nash axiomatic bargaining rule, where the disagreement value is determined by the value of the option to continue to search for an alternative match.[1] Furthermore, workers and firms decide on their search effort, recognizing that effort is costly and affects the probability of a meeting between a firm and a worker. This model is described and solved, with some minor differences, by Mortensen (1982), Diamond and Maskin (1979) and Wolinsky (1987). The duration of job search and the distribution of accepted wages are jointly determined by a reservation value of the match as in all search models. The reservation match is affected by three different components of the model: the distribution of the value of the match, which reflect the workers' "true" productivity, the cost of search and a bargaining power parameter.[2]

Heterogeneity in the distribution of wages and in the duration of unemployment arises in this model because of the existence of an exogenous distribution of firm-worker match-specific productivities. Alternatively, in the equilibrium search model of Albrecht and Axel (1984) and Mortensen (1990) the main source of wage dispersion is the heterogeneity of workers with respect to their non-market productivity and the inherent heterogeneity of firms with respect to their production efficiency. Eckstein and Wolpin (1990) provide a method for empirically implementing the Albrecht and Axel model. More recently, Burdett and Mortensen (1993) have developed a model in which wage dispersion does not rely on inherent heterogeneity, and van den Berg and Ridder (1993) have implemented that model.

The econometric framework in this paper is similar to that contained in Flinn and Heckman (1982) and in Eckstein and Wolpin (1990). Because these models have constant reservation wages, in order to obtain the decreasing hazard for duration and the non-increasing mean wage with duration that are general features of unemployment job duration data, unobserved heterogeneity (or some form of state dependence) must be added to the model. Having individuals who differ in their reservation wages, but who draw from

1. However, the empirical formulation is quite close to that of Flinn and Heckman (1982).

2. We have adopted the Nash Axiomatic solution to preserve the simplicity of the model. It has been shown by Wolinsky (1987) that the efficient strategic bargaining equilibrium is nested in the axiomatic solution. Bargaining models have been used to interpret strike data (Kennan and Wilson (1989)) but, to our knowledge, there has been little effort devoted to structurally implementing bargaining models.

the same offered wage distribution, is sufficient to generate negative duration dependence in unemployment population hazard rates. However, this form of heterogeneity will also generate increasing mean accepted wages with duration. With more flexible heterogeneity, in the wage offer distribution, in search costs, and in wage offer probabilities, it is possible to obtain many different patterns of mean accepted wages with duration including non-monotonic ones. For example, to obtain decreasing accepted wages, suppose that individuals who face wage offer distributions with higher means nevertheless have lower reservation wages and thus higher hazard rates, because they also face sufficiently higher costs of search. In this case workers with, on average, high accepted wages leave the unemployment state first and the observed mean accepted wage is decreasing with duration.[3]

We use data from the 1979 youth cohort of the National Longitudinal Surveys of Labor Market Experience to follow several school-leaving cohorts of young males from the time they finally leave school until the time they accept their first full-time job (or until the last survey date). We consider white and black males within four schooling levels: high-school non-completers, high-school graduates, college non-completers and college graduates.

Given data only on the duration to the first job and accepted wages, identification of the parameters of the search–matching–bargaining model is tenuous. To reduce the number of estimated parameters, we assume that workers and firms are "symmetric" and estimate reservation wages without imposing the restrictions between the cost of search, search effort, offer probabilities and reservation wages that are implied by the model. After estimating "unrestricted" reservation wages we use the model to recover additional parameters, based on the restrictions of the model.

A model which allows for five types of heterogeneous workers within each schooling/race group fits the duration and wage data well for all such groups. Offer probabilities for all groups are estimated to be close to one. Hence, the main factor explaining the observed pattern of duration and wages is the rate at which offers are accepted. Mean offered wages are about \$1000 less than mean accepted wages and the coefficient of variation of offered wages is somewhat lower than that of observed wages. Hence, the degree of income inequality within each schooling/race group is affected by behaviour in addition to market fundamentals.

In terms of the payoff to schooling, based on the estimated schooling/race specific wage offer distributions, the internal rate of return associated with attending college (without graduating) is 32% for blacks while the return to graduating high school is 3%. For whites, the comparable rates of return are 17 and 27%. These figures differ substantially from those obtained using the accepted, rather than offered, wage distribution. Mean offered wages for blacks are always below those of whites with the same schooling level, and black–white offered wage differentials differ from black–white accepted wage differentials, at least for high-school graduates.

We use the estimated parameters from the unrestricted version of the model and the reservation wage rule obtained from the search–matching–bargaining model to calculate the implied search costs. Providing a subsidy to search of \$50 per quarter, which is about a 5 to 18% reduction in search costs, causes the mean duration of unemployment to the first full-time job to increase by 30 to 60% and mean accepted wages to increase by 3 to 30%.

The paper is organized as follows. In the next section we present the search–matching–bargaining model for the labour market. In Section 3 we discuss a simple method for

3. Berkovitch (1990) analysed a search-bargaining model where in equilibrium the wage is decreasing with the duration of unemployment due to the stigma associated with long durations of unemployment.

incorporating observed and unobserved heterogeneity of workers. The data is described in Section 4 and in Section 5 we present a particular specification for estimating the model. In Section 6 we present the estimation results and implications. Concluding remarks are presented in Section 7.

2. THE BASIC MODEL

Consider a market where a worker (w), who lives forever, meets a firm (π) with probability P which is a function of the uncoordinated effort of the worker (s^w) and of the firm (s^π). That is,

$$P = G(s^w, s^\pi; g), \tag{1}$$

where $s^j, j = w, \pi$, is the effort made by j to meet members of the other group. $G(\cdot)$ is a function of the effort of the two parties and a vector of parameters g. Furthermore, it is assumed that $G_j(\cdot) > 0$, $G_{jj}(\cdot) < 0$ and $G_{ij}(\cdot) > 0$, where $j, i = 1, 2$. The effort s^j involves incurring cost $c_j(s^j)$, which is continuously twice-differentiable with $c'(\cdot)$ and $c''(\cdot)$ positive.[4]

Once a firm and a worker meet they sample a present value for their match that is equal to $m \in [0, M]$, which is the value added by the worker to the firm's production; M is bounded away from infinity. The value of m, known immediately to both parties, is a random draw from the distribution function $F(m)$, which is assumed to be continuously differentiable.

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,

$$w(m) + \pi(m) \leq m. \tag{2}$$

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. If they do not agree, then they can search again during the next period. Time is assumed to be discrete.

The Axiomatic Nash bargaining solution of Mortensen, and Diamond and Maskin (MDM) consists of an agreement schedule $(w(m), \pi(m))$ that satisfies (2) such that;

(i) Each agent's strategy is optimal, given $w(\cdot)$ and $\pi(\cdot)$ and the search strategy of the other party.
(ii) Each agreement $(w(m), \pi(m))$ is a solution to the bilateral bargaining problem faced by the matched agents.
(iii) The system is at a steady state.

The division of the surplus m is determined by the static Nash Axiomatic solution relative to the disagreement outcome of continued search. The solution is derived as follows. Given increasing functions $w(m)$ and $\pi(m)$ and constant value s^i, let $V^j(s^j; s^i)$ denote the steady-state expected value of search by party j, $(j = w, \pi)$; it is given by the equation

$$V^j(s^j; s^i) = -c_j(s^j) + PE \max\{j(m), \delta^j V^j(\cdot)\} + (1-P)\delta^j V^j(\cdot). \tag{3}$$

The search policy is characterized by a constant reservation value m_j, that is, search

4. In some specifications of search–matching models the probability of a meeting is a complicated function of the number of workers and firms (jobs). In equation (1) we abstract from the complicated 'technology' that relates the offer probability to the endogenously determined volume of available jobs and new workers that seek employment.

until the match value m is above $m_j, j=w, \pi$. Now, the steady-state expected value of search of equation (3) can be written as

$$\begin{aligned} V^j(s^j, m_j; s^i) = & -c_j(s^j) + PE(j(m) \mid m > m_j) \Pr(m > m_j) \\ & + P\delta^j V^j(s^j, m_j; s^i) \Pr(m < m_j) \\ & + (1-P)\delta^j V^j(s^j, m_j; s^i). \end{aligned} \tag{4}$$

The optimal search strategy of party j is to maximize $V^j(s^j, m_i; s^i)$ with respect to s^j and m_j. Let $\bar{V}^j, j=w, \pi$, denote the maximized value of V^j with respect to s^j and m_j. At each date the disagreement values for the worker and the firm are the one-period discounted value of continued search, that is, $\delta^j \bar{V}^j$.

The Nash Axiomatic bargaining solution is efficient, which implies that equation (2) holds as an equality and in equilibrium it is required that only for match values $m \geqq \delta^w \bar{V}^w + \delta^\pi \bar{V}^\pi$ will the worker and the firm arrive at an agreement. Hence, it is clear that the solution is characterized by worker and firm reservation match values such that $m_w = m_\pi = m^*$, where m^* is the reservation match value satisfying

$$m^* = \delta^w \bar{V}^w + \delta^\pi \bar{V}^\pi. \tag{5}$$

The Axiomatic Nash bargaining solution for a non-symmetric case with a weight α for workers implies that the wage schedule satisfies

$$w(m) = \delta^w \bar{V}^w + \alpha(m - m^*), \tag{6}$$

and, equivalently, the profit schedule satisfies

$$\pi(m) = \delta^\pi \bar{V}^\pi + (1-\alpha)(m - m^*). \tag{7}$$

Observe that equations (5), (6) and (7) are consistent with equation (2). Furthermore, from the definition of $\bar{V}^j$, $\partial \bar{V}^j / \partial m^* = 0$, and by substituting (5) into (6) and (7) we get that $\partial w(m)/\partial m^* = 0$ and $\partial \pi(m)/\partial m^* = 0$. In order to solve for the maximum of V^j at equilibrium, we substitute equations (6) and (7) into equation (4), so that the solution for $V^j, j=w, \pi$, must satisfy

$$\bar{V}^w = \max_{s^w, m^*} [-c_w(s^w) + P\alpha E(m - m^* \mid m > m^*) \Pr(m > m^*)]/(1-\delta^w), \tag{8}$$

and

$$\bar{V}^\pi = \max_{s^\pi, m^*} [-c_\pi(s^\pi) + P(1-\alpha) E(m - m^* \mid m > m^*) \Pr(m > m^*)]/(1-\delta^\pi). \tag{9}$$

From (8) and (9) we solve for m^* satisfying,

$$\delta^j \bar{V}^j = j(m^*), \qquad j = w, \pi. \tag{10}$$

The first-order condition with respect to s^w and s^π, using (10), are

$$c'_w(s^w) = G_1(s^w, s^\pi; g) \alpha E(m - m^* \mid m > m^*) \Pr(m > m^*), \tag{11}$$

and

$$c'_\pi(s^\pi) = G_2(s^w, s^\pi; g)(1-\alpha) E(m - m^* \mid m > m^*) \Pr(m > m^*). \tag{12}$$

Equations (5), (8), (9), (11) and (12) solve for the three variables of the model; s^w, s^π and m^*. Note that the ratio $[c'_j(s^j) \div G_j]$ is increasing in s^j and it is very small for $s^j = 0$. Hence, for a given level of m^* there exists a solution for s^w and s^π. Given values for s^w

and s^{π} the solution for m^* can be verified and the existence of a Nash Bargaining solution is established for the model.[5]

The model provides a complete characterization of the joint distribution of the duration of search and accepted rents for both parties, workers and firms. This distribution is affected by the parameters, δ^w, δ^{π} and α, and the functions of the model, $G(\cdot)$, $c_w(\cdot)$, $c_{\pi}(\cdot)$ and $F(m)$.

3. WORKER HETEROGENEITY

Workers with different levels of schooling and of different race are assumed to have different fundamental parameters. In order to avoid the complications that would arise from incomplete and asymmetric information, we assume that heterogeneity among workers is fully observed by firms. Further, heterogeneity in characteristics also exists within race/schooling groups, e.g. differences in the value of non-market time for a given race/schooling type. Race and schooling are assumed to be observable by the econometrician, but differences within race/schooling groups are not.

The worker problem of Section 2 is not changed due to observable heterogeneity if it is assumed that the probability of a meeting between a worker and a firm depends only on the effort that the firm makes to meet a particular homogeneous group of workers. As long as we separate completely the groups with respect to competition for jobs the solution of the model stays exactly as it is described above.[6]

It should be noted that extending the model to consider unobserved (to us) heterogeneity may provide a potential explanation for the main observed differences between race or schooling groups. For example, as we have noted, blacks have lower mean accepted wages and longer durations of unemployment at all levels of schooling. Suppose that within each race/schooling group there are two types of individuals. One type has a low match productivity mean and the other a high mean. Because workers with a low mean will search less intensively, they will also have a lower offer probability and a lower reservation wage than the high-mean type. The resulting hazard rate of the low-mean type may be lower than that of the high-mean type, leading to a longer mean duration of unemployment. Now, if blacks are disproportionately of the low-mean type, say because they have lower quality of schooling, then they will have lower mean accepted wages and longer durations as in the data. It should be noted that the same example could explain why accepted wages decline with duration for a given race/schooling group, as noted in the introduction.[7]

4. DATA

The data for this analysis are from the National Longitudinal Survey of Labor Market Experience 1979 youth cohort (NLSY). The NLSY consists of 12,686 individuals who

5. The matching model can be solved for an equilibrium division (α) of the observed match value (m), using a version of Rubinstein's (1982) strategic bargaining specification. Wolinsky (1987) demonstrated that the strategic solution is not identical to the Nash Axiomatic solution due to the choices of effort by each of the parties.

6. The complete separation between the groups for each vacancy is an extreme assumption. A more realistic assumption could be to allow for some substitution between types for each vacancy. However, this would greatly complicate the model and the empirical analysis. Assuming, as we do, that firms have no capacity constraints and that they make a non-zero profit from each employed worker, then the solution of the model is invariant with respect to heterogeneity among workers.

7. Note that in the search equilibrium model of Eckstein and Wolpin (1990) the mean accepted wage is increasing with duration.

were 14 to 21 years old as of 1 January, 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 (1979–1986) for the white male core sample and the black male core supplemental sample.

We consider four schooling groups: high-school non-completers (HSINC), high-school graduates (HSGRD), college non-completers (COLINC) and college graduates (COLGRD). The highest grade completed is that schooling level achieved at the last school enrolment data (week) up to 31 December, 1984. We exclude individuals with inconsistent schooling information, e.g. those who report lower schooling in a subsequent survey round.

We define the duration to the first job to be the number of calendar quarters, with the first calendar quarter of school enrolment being the one after the last enrolment calendar quarter, before the individual began working at a full-time job.[8] A full-time job is defined to be a job in which the individual worked at least 30 hours per week in the entire calendar quarter (13 weeks). Employment data is available from 1 January, 1978 through 31 December, 1985. We exclude individuals who left school prior to 1 January, 1978 because we do not know the employment history prior to that time. We also exclude individuals who ever served in the military. The real weekly wage (in 1982 dollars) corresponds to the wage reported in the last calendar quarter multiplied by 13.[9]

Table 1 reports descriptive statistics for all eight race/schooling groups, although there are too few black college graduates to perform a statistically meaningful analysis. Because we observe individuals for at least a year after leaving school the great majority of spells are complete. Among all school completion levels, black males take at least one additional quarter to become employed at a full-time job relative to white males and the difference is about two quarters at the lower school completion levels.

Not only are the durations longer for black males, but once employed their mean weekly wage is also lower, with black males receiving between 83 and 95% of that received by white males at the same schooling level. The return to schooling on the first job as measured by the percentage wage difference between adjacent schooling groups is between 10% and 26% for blacks and between 7% and 25% for whites. These figures are consistent with other studies (see Donohue and Heckman (1991)) which also control for experience, sex and race. However, the return to schooling calculated using accepted wages may differ greatly from the return calculated using offered wages if reservation wages differ by schooling level. In addition, the substantial observed lower durations of unemployment for individuals with higher levels of schooling, evident in Table 1, indicates the existence of an additional potentially important return to schooling. White high-school dropouts

8. We use the last enrolment date in order to avoid jobs that are between schooling spells, e.g. working for a year before attending college. Intermittent schooling, while not unusual, does not involve a substantively large proportion of the population. Aedo (1992), reports using the same data source, that approximately 90% of white males who attend college do so in the year directly following high school graduation. Moreover, once enrolled 92% continue their schooling uninterruptedly. The comparable figures for blacks are 78 and 87%.

9. Search models that have been estimated in the literature have almost always assumed that the payoff to search is the present value of the accepted wage that is observed at the start of the spell. Particularly for young workers, using the starting wage has several problems: (1) wages tend to rise over the life cycle, and (2) job turnover is significant. There is thus an implicit assumption that the present value of the starting wage is representative of the value of taking the first job when account is taken of wage growth and turnover. Using data on the first job spell can provide some evidence on this issue. Holding race and schooling constant, estimates from a proportional hazards model (controlling also for censoring) reveals that there is a statistically positive correlation between the starting wage and the duration of the employment spell for the first job. In addition, for those who changed jobs, there is a non-negative correlation between wages on the first two jobs.

TABLE 1

*Duration to first full-time job and accepted wages by race and schooling**

	Mean duration in quarters**	Prop. complete Spell	Accepted quarterly wages ($)		
			mean	std. dev.	s.d./mean
White Males					
HSINC	4·8	0·88	2698	1348	0·50
	(184)		(150)		
HSGRD	2·5	0·94	2892	1657	0·57
	(543)		(484)		
COLINC	1·1	0·96	3625	1917	0·53
	(167)		(144)		
COLGRD	1·2	0·99	3972	1855	0·47
	(180)		(156)		
Black Males					
HSINC	6·5	0·79	2305	1006	0·44
	(174)		(128)		
HSGRD	4·6	0·87	2520	1349	0·53
	(322)		(261)		
COLINC	2·0	0·92	3018	1756	0·58
	(84)		(64)		
COLGRD	2·1	0·92	3807	1917	0·50
	(24)		(20)		

* Sample size in parentheses.
** Complete and incomplete spells.

and high-school graduates have greater wage dispersion than similarly educated blacks, where wage dispersion is measured by the coefficient of variation (s.d./mean). However, college dropouts and college graduates have more dispersion than similar whites (see also the accepted wage data distributions in Figures 3(a)–(g).[10]

Figures 1(a) to (c) present the Kaplan–Meier survivor curves for the duration to the first job. The figures show that this duration is uniformly higher for blacks at each quarter since leaving school. We used the Kaplan–Meier survivor functions to perform log-rank, Mantel–Haenszel, and Wilcoxson–Gehan tests (see Kalbfliesch and Prentice (1980)) for the equality of the survivor functions. All three tests very strongly rejected (with marginal probability of less than 2%) equality of the survivor functions for blacks and whites of the same schooling level. Interestingly, the same tests of equality of the survivor functions for black high-school graduates vs. white high-school dropouts as well as for black college dropouts vs. white high-school graduates did not lead to rejection (see Figures 2(a) and (b)).

The data also reveal that the hazard rate associated with entering full-time employment is declining with duration for all race/schooling groups. 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.[11]

10. It is also true that the duration of employment on the first job is longer at higher schooling levels for each race group. White (black) high-school dropouts spend an average of 3·6 (3·9) calendar quarters on their first job, white (black) high-school graduates 5·2 (4·4) quarters, white (black) college dropouts 6·1 (4·8) quarters and white college graduates 6·6 (4·7) quarters. The difference between schooling groups in this table is actually understated because of censored observations. For more information on this data see Wolpin (1992).

11. One way to introduce structural state dependence would be to allow for learning about the match value over the period of employment as in Jovanovic (1978) or Miller (1984). Such an assumption, however, would greatly complicate the solution of the bargaining model.

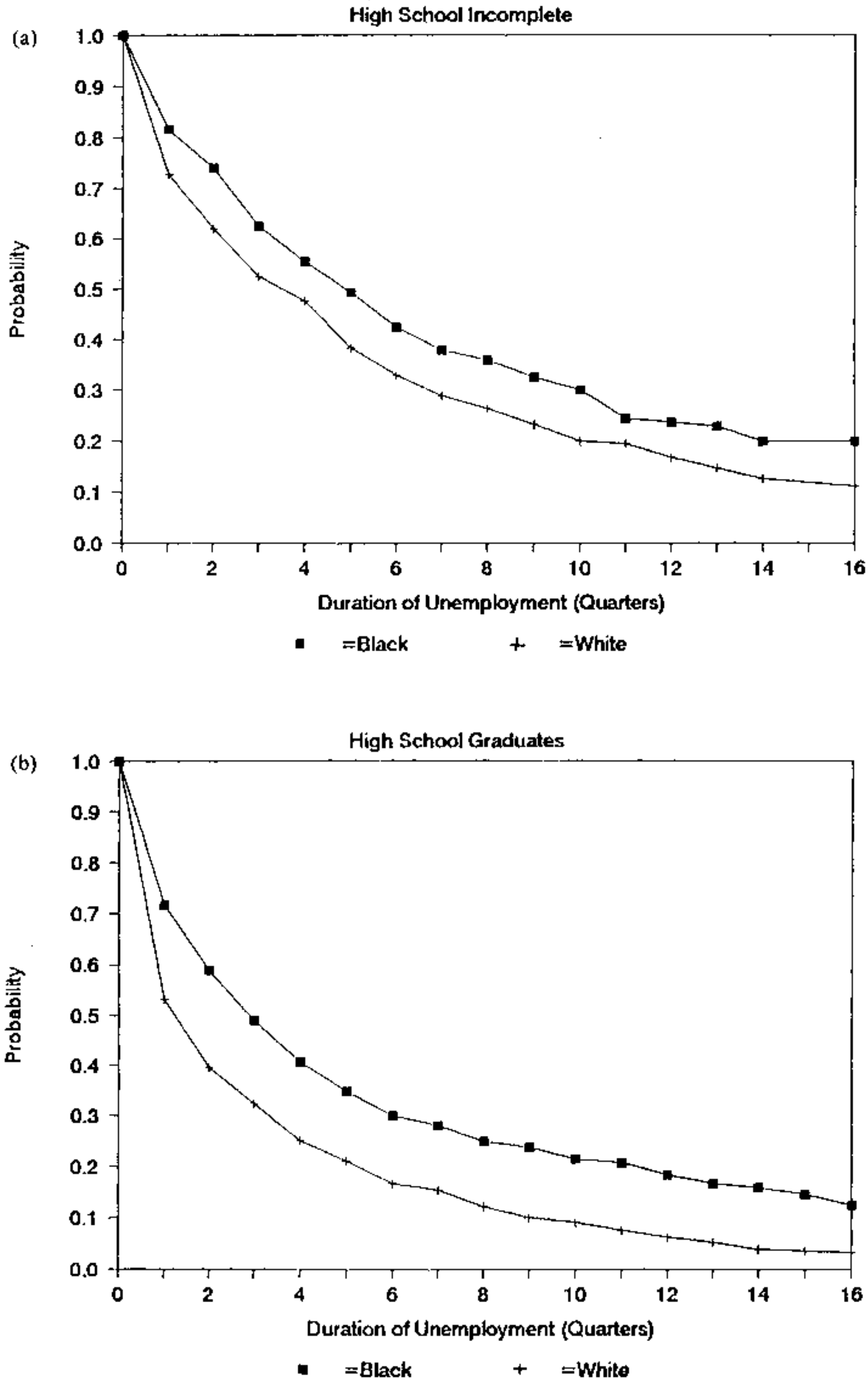

FIGURE 1
Survivor functions: Kaplan–Meier

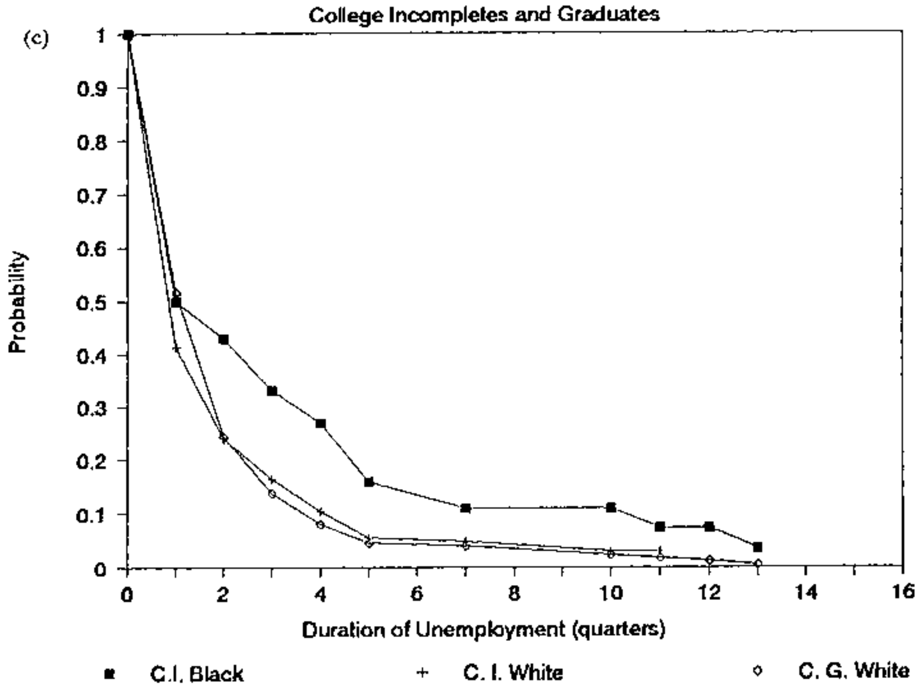

FIGURE I—continued

Table 2 presents regression results that characterize the pattern of accepted wages with the duration of unemployment. In all but one schooling/race group, and in all of the pooled groups, accepted wages decline with the duration of unemployment. While this fact is consistent with the standard finite-horizon partial-equilibrium search model, it is not consistent with a pure heterogeneity (infinite-horizon) partial-equilibrium search model in which individuals differ only in their reservation wages.

Accepted wage regressions also include a dummy variable for individuals who left school during the years 1980–1982.[12] For all race/schooling groups, accepted wages are lower in those years, although the difference is statistically significant only in the case of white high-school graduates (which is also the largest group). Accepted wages are about 10% lower during those years. Of course, the interpretation of this result is unclear. Indeed, while accepted wages may have been lower, this is not necessarily true of offered wages; accepted wages are determined both by the offered wage distribution and reservation wages.[13]

5. ESTIMATION SPECIFICATION

We estimate the model of Section 3 at the steady-state equilibrium characterization using the above data. In order to estimate the model the match value m is assumed to come

12. Those years, which correspond loosely to recession years as measured by aggregate unemployment figures, seemed to be the most different in terms of accepted wages based on regressions for most race/schooling groups, as ascertained from regressions that included individual school-leaving year dummies.

13. Interestingly, the duration of unemployment is longer for the 1980–1982 white high-school graduate school leaving cohort by almost a full calendar year. This is, however, not the case for black high-school dropouts nor for several other groups.

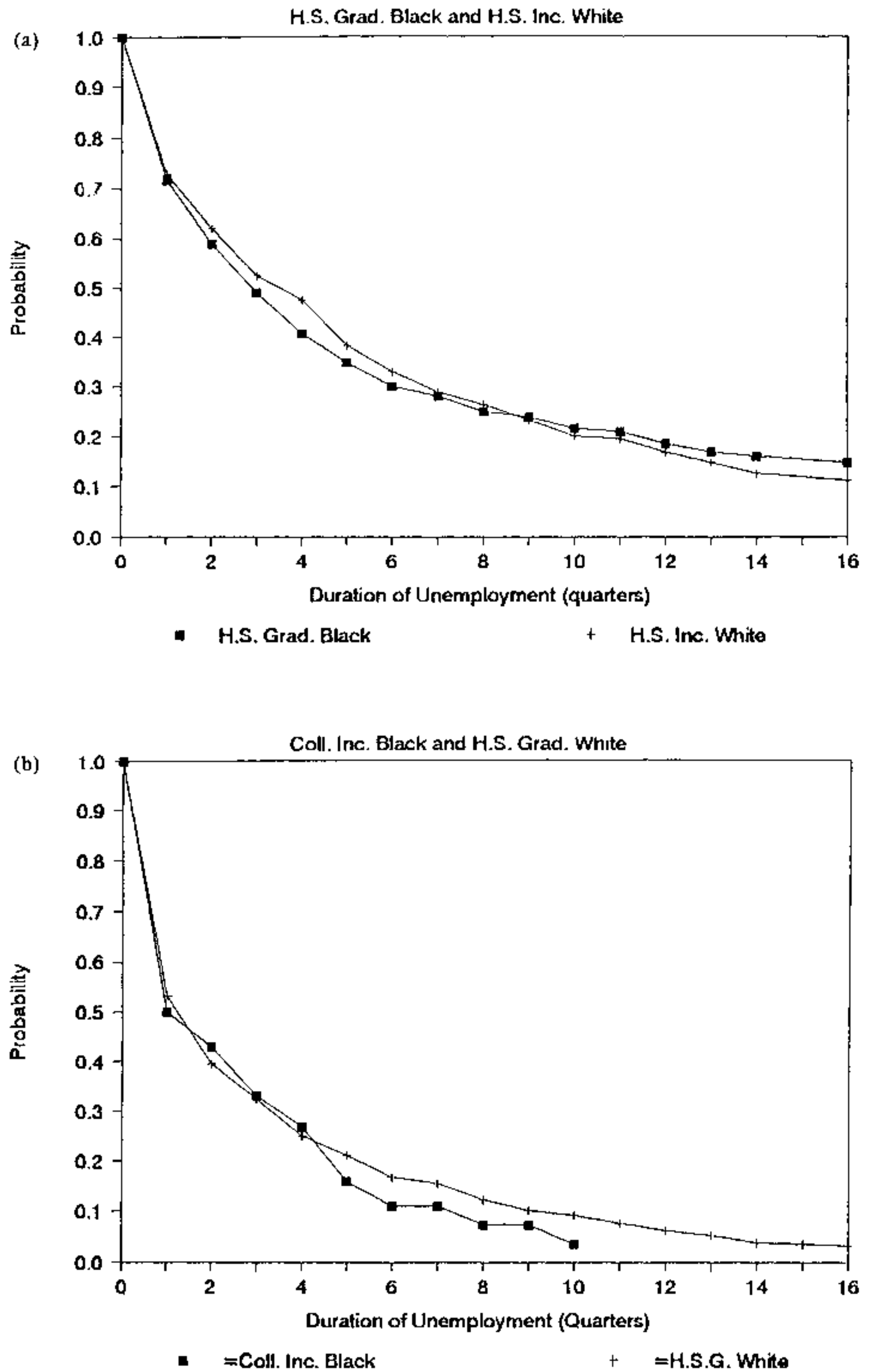

FIGURE 2
Survivor functions: Kaplan-Meier

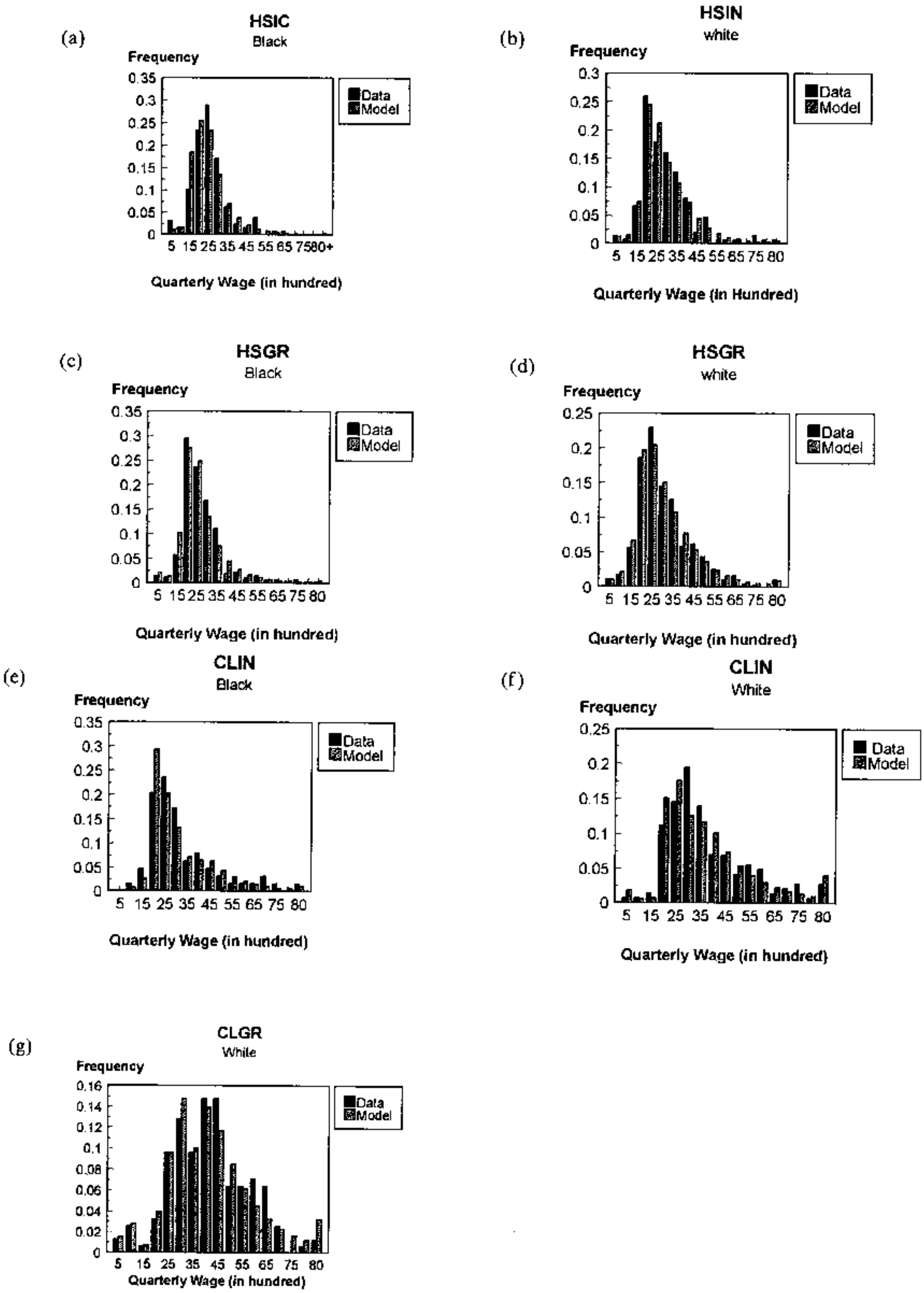
(a)
HSIC
Black
Frequency
0.35
0.3
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 7580+
Quarterly Wage (in hundred)
Data
Model
(b)
HSIN
white
Frequency
0.3
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in Hundred)
Data
Model
(c)
HSGR
Black
Frequency
0.35
0.3
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in hundred)
Data
Model
(d)
HSGR
white
Frequency
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in hundred)
Data
Model
(e)
CLIN
Black
Frequency
0.35
0.3
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in hundred)
Data
Model
(f)
CLIN
White
Frequency
0.25
0.2
0.15
0.1
0.05
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in hundred)
Data
Model
(g)
CLGR
White
Frequency
0.16
0.14
0.12
0.1
0.08
0.06
0.04
0.02
0
5 15 25 35 45 55 65 75 80
Quarterly Wage (in hundred)
Data
Model

FIGURE 3

TABLE 2

Accepted wage regressions (Hetroschedasticity corrected standard errors)[1]

Sample	Duration	Recession
White:		
HSINC	9·05	−122
	(21·2)	(221)
HSGRD	−45·3	−317
	(22·1)	(157)
COLINC	−91·0	−270
	(107)	(322)
COLGRD	−80·8	−117
	(55·4)	(292)
All schooling levels[2]	−30·5	−258
	(14·7)	(111)
Black:		
HSINC	−37·0	−28·4
	(15·6)	(173)
HSGRD	−20·5	−159
	(16·5)	(167)
COLINC	−110	327
	(51·6)	(426)
COLGRD	−466	−1000
	(139)	(773)
All schooling levels[2]	−32·4	−81·7
	(11·7)	(125)
All race and all schooling[3]	−31·4	−197
	(9·36)	(84·6)

1. Wage (dependent variable) = accepted wage for first job, Duration = number of quarters to first job, RECESSION = a dummy variable for leaving school during the years 1980–1982. Standard Errors in parenthesis.
2. Includes schooling dummies.
3. Includes race and schooling dummies.

from a log-normal distribution with density function

$$f(m)=\frac{1}{m\sigma_m(2\pi)^{1/2}}\exp\left\{-\frac{1}{2}\left(\frac{\ln m-\mu}{\sigma_m}\right)^2\right\}. \tag{13}$$

Using standard properties for the mean of the log-normal distribution (13), it can be shown that

$$E(m\mid m>m^*)\ \Pr\,(m>m^*)=\exp\,(\tfrac{1}{2}\sigma_m^2+\mu)\left[1-\Phi\left(\frac{\ln m^*-(\sigma_m^2+\mu)}{\sigma_m}\right)\right], \tag{14}$$

where Φ is the normal c.d.f. and note that σ_m affect the mean of the match distribution. The probability of accepting a job conditional on search, the hazard rate (h), is given by

$$h(s^w, s^\pi, m^*)=P(1-F(m^*)), \tag{15}$$

where P is the offer probability given optimal effort levels in equilibrium. The density function of the wage conditional on acceptance of a job is given by

$$z((w(m)-a)\mid m>m^*)$$
$$=\left[\frac{1}{(w(m)-a)\sigma_m\sqrt{2\pi}}\exp\left\{-\frac{1}{2}\left(\frac{\ln\,(w(m)-a)-(\ln\alpha+\mu)}{\sigma_m}\right)\right\}^2\right]\Big/(1-F(m^*)), \tag{16}$$

where $a = w(m^*) - \alpha m^*$ and $w(m^*) = w^*$ is the reservation wage. Thus, $\ln(w(m) - a)$ is a normal random variable. a is the linear term in the wage function, which is given by $w(m) = a + \alpha m$ for $m > m^*$.

The wage distribution parameters in (16) are μ, σ_m, α and a. Clearly α cannot be identified separately from μ. The parameter a is identified due only to the non-linear way that it is included in (16). If we assume that firms and workers in the model are symmetric, that is, that all of the fundamental parameters of the model of Section 3 are the same for workers and firms, including $\delta^w = \delta^\pi$ and $\alpha = 0{\cdot}5$, then $a = 0$. Given the tenuous identification of a, and the simplification that results in the likelihood specification (see below), we restrict our attention to the symmetric case, where $a = 0$, and thus, $w = 0{\cdot}5m$.

We extend the model by assuming that the population of workers in a given market, i.e. of a given race and schooling, consists of K different types of individuals where each type may have a different value of the cost of search (value of leisure) and a different value of the mean of the distribution of the match. The different types of workers are unobserved by us but are known to firms. Furthermore, we assume that the observed wage, w^o, is measured with a multiplicative error that is independent of the true wage. Thus, $\ln w^o = \ln w + u$, where w^o is the observed wage, w is the true wage, and u is the measurement error.[14] We assume that u is distributed $N(0, \sigma_u^2)$ and u is independent of w.

With the above assumptions, the joint probability that the wage exceeds the reservation wage and that w^o is realized is,

$$\begin{aligned}\Pr(w(m) > w^*(m), w^o) &= \Pr(w(m) > w^*(m) \mid w^o)\Pr(w^o)\\ &= \left[1 - \Phi\left(\frac{\ln w^* - (1-\rho^2)(\ln(0{\cdot}5)\mu) - \rho^2 \ln w^o}{\rho\sigma(1-\rho^2)^{1/2}}\right)\right]\\ &\quad \times \frac{1}{w^o}\cdot\frac{1}{\sigma}\phi\left(\frac{\ln w^o - (\ln(0{\cdot}5) + \mu)}{\sigma}\right), \end{aligned} \tag{17}$$

where $\sigma^2 = \sigma_m^2 + \sigma_u^2$ and $\rho^2\sigma^2 = \sigma_m^2$, and ϕ is the normal p.d.f..[15]

Given that the cost of search and the mean of the match distribution are different for the different types, the reservation wage of type j is given by equation (8). Letting the c_k be the cost of search for a worker of type k, $k = 1, \ldots, K$, and recognizing that $m_k^* = 2w_k^*$ for each type, with appropriate substitutions (8) is equal to

$$\begin{aligned} w_k^* = \frac{\delta}{1-\delta}\Bigg\{ &-c_k + P_k 0{\cdot}5 \exp\{0{\cdot}5\sigma_m^2 + \mu_k\}\\ &\times\left[1 - \Phi\left(\frac{\ln(m_k^*) - (\sigma_m^2 + \mu_k)}{\sigma_m}\right)\right] - P_k 0{\cdot}5 m_k^* P m_k\Bigg\}. \end{aligned} \tag{18}$$

14. If wages were not measured with error the existence of extremely low wages in the data would have a large impact on the estimated reservation wages, given that reservation wages are bounded from above by the lowest observed accepted wage. Moreover, without measurement error the likelihood function would be non-standard (Heckman and Flinn (1982)).

15. This expression is derived from the fact that $\ln w$ and u are independent normal random variables and, hence, w^o is log-normal. Equivalently, the joint distribution of $\ln w$ and $\ln w^o$ is bivariate normal. For the case where $a \neq 0$, recognizing that $\ln(w - a)$ is normal, and assuming that u is independent of $\ln(w - a)$ (17) becomes,

$$\Pr(w > w^*, w^o = we^u) = \int_{w^*}^{\infty} [1 - F(m^*)]z(w(m) - a) \mid m > m^*)\,\frac{1}{w^o\sigma_u}\,\phi\left(\ln\left(\frac{\ln w^o - \ln w}{\sigma_u}\right)\right)dw.$$

In addition to the variables that are as previously defined, Pm_k is the probability that there will be an agreement between the worker of type k and the firm, given that there is a match, that is,[16]

$$Pm_k = 1 - \Phi\left(\frac{\ln m_k^* - \mu_k}{\sigma_m}\right). \tag{19}$$

The likelihood function for I individuals of K types, each with a completed spell length of d_i and an observed wage of w_i^o, is

$$\begin{aligned} L(\psi) = \textstyle\prod_{i\in I} \sum_{k=1}^{K} \gamma_k & \left[1 - P_k\left(1 - \Phi\left(\frac{\ln w_k^* - (\ln (0{\cdot}5) + \mu_k)}{\rho\sigma}\right)\right)\right]^{d_i} P_k \\ & \times \left[1 - \Phi\left(\frac{\ln w_k^* - \rho^2 \ln w_i^o - (1-\rho^2)(\ln (0{\cdot}5) + \mu_k)}{\rho\sigma(1-\rho^2)^{0{\cdot}5}}\right)\right] \\ & \times \frac{1}{w_i^o}\frac{1}{\sigma}\phi\left(\frac{\ln w_i^o - (\ln (0{\cdot}5) + \mu_k)}{\sigma}\right), \end{aligned} \tag{20}$$

where γ_k is the proportion of type k in the population and ψ is a vector of parameters.[17] The likelihood function (20) can be considered as an unrestricted version of the search–matching model in which the reservation wage solution of the model (equation (18)) and the functional relation between the search cost and the offer probability are not required to be satisfied.

Given the symmetry assumption, the duration and the wage data allow identification of all the parameters of the unrestricted model, namely the K values of the μ's, the w^*'s and the γ's and σ and ρ.[18] To reduce the number of estimated parameters, we assume that the "type" proportions, γ_k's, come from a binomial distribution with a parameter γ. With this additional assumption, the parameter vector for the unrestricted model is,

$$\psi = [\gamma, P_1, \dots, P_K, \mu_1, \dots, \mu_K, w_1^*, \dots, w_K^*, \sigma, \rho]'. \tag{21}$$

The parameters of the search–matching model can be recovered only with further restrictions. One possibility is to write the offer probability and search cost functions explicitly in terms of search intensity. For example, assume that the function $G(\cdot)$, which determine the offer probabilities is given by,

$$P_k = \frac{\exp\{g_1(2s_k)^{g_2}\}}{1 + \exp\{g_1(2s_k)^{g_2}\}}, \tag{22}$$

where s_k is the effort level of the worker (and also of the firm given symmetry) to meet and g_1 and g_2 are parameters. Further, let the cost of search be given by,

$$c_k = \theta_1 + \theta_2 s_k + \theta_3 s_k^2, \tag{23}$$

16. Note from (18) that if $\delta = 0$ then $w^* = 0$.

17. For incomplete spells and completed spells without wage data the likelihood function is modified in an obvious way.

18. For the case of $a = 0$ the model becomes almost the same as that in Flinn and Heckman (1982) and the identification conditions there apply here. Replacing equation (17) with the term in footnote 15 in the likelihood function (20), a_k is identified because of the non-linearity of the model. However, implementation would require a numerical integration for each observation.

where θ_1, $\theta_2>0$, $\theta_3>0$, are parameters. Note that even though (22) and (23) are functions that are independent of type (k), s and c are type-specific because they depend on all of the other parameters of the model, some of which are different across types. Equations (22) and (23) add five additional parameters. Of course, it is possible either to increase or decrease the number of parameters by changing the functional forms of these two equations. Estimating the parameters of (22) and (23), together with the other parameters of the model, involves explicitly solving the equilibrium model numerically for optimal search intensities and reservation wages.

An alternative to this parametric approach, with full information estimation, is to estimate the parameters of the unrestricted model, and assuming a value for k, then to calculate the implied values of the cost of search (c_k) using (18). We obtain the K values of the P_k's from the unrestricted estimation and the K values of the c_k's from inverting (18). A consistency check of the model is to see whether c_k and p_k are monotonically (positively) related across types. The advantage of this approach, aside from the considerable computational savings, is that one need not impose additional parametric forms for (22) and (23). This is even more appealing given that search intensity is not measured.

6. RESULTS

Table 3 presents the estimates of the unrestricted model based on five types for each race/schooling group.[19] We arrived at five types by using likelihood ratio tests for small numbers of types relative to five. Although we did not always reject less than five types for all groups, we always rejected two types, and to allow for comparisons across groups we used five types for all groups. We order the types in Table 3 by their estimated proportions, beginning with the least represented type (the smallest γ).

As seen in Table 3, the probability of receiving an offer is very close to one for almost all individuals in each race/schooling group.[20] Except for white college graduates, less than 5% of the population within each race/schooling group are uncertain about receiving an offer, that is, have an offer probability much less than one. For white college graduates, 32% have an offer probability significantly less than one.

In all cases, not only does the least represented type have the lowest offer probability, but they also have the lowest reservation wage and the lowest mean of the offer (match) distributions. However, there is no obvious relationship between the mean wage offer and the reservation wage across types. For example, for black high-school dropouts, the type with the highest mean wage offer has the lowest reservation wage (actually zero), while for black college dropouts the type with the highest mean wage offer has the highest reservation wage. As is clear from equation (18) the search-matching model allows for one additional free parameter, the cost of search, which could account for the different correlations between offer probabilities, mean wage offers and reservation wages across groups and types.

19. The estimates in Table 3 ignore business cycle effects. We did estimate the model assuming that both offer probabilities and the wage offer mean were potentially affected by the business cycle, as measured by a 1980–1982 dummy, for the high-school graduate groups. (Recall that accepted wages were significantly different only for white high-school graduate leaving cohort of 1980–1982 (see Table 2)). Results were both quantitatively similar and not statistically different from those in Table 3.

20. This result is consistent with the findings of Eckstein and Wolpin (1990) who estimate an equilibrium search model using the same data source. However, most partial equilibrium search models tend to find the opposite, that is, very low offer probabilities (Blau (1991), Engberg (1992), Wolpin (1987)). The limited "direct" data on offer probabilities, obtained through survey questions about job offers received, are more consistent with the latter (Blau and Robins (1990)). However, such survey questions are themselves not conceptually clean.

TABLE 3

*Estimated parameters of the unrestricted model**

	Schooling/race group						
Parameter	HSINC B	HSINC W	HSGRD B	HSGRD W	COLINC B	COLINC W	COLGRD W
γ_1	0·01	0·01	0·03	0·03	0·03	0·02	0·05
P_1	0·99	0·99	0·22	0·07	0·99	0·32	0·63
μ_1	5·69	5·51	6·31	7·20	7·07	6·19	7·16
m_1^*	544·47	315·34	346·18	0·00	1870·31	523·16	310·59
w_1^o	170·35	144·00	340·22	762·11	641·54	294·38	714·44
γ_2	0·10	0·09	0·12	0·11	0·11	0·13	0·07
P_2	0·99	0·99	0·99	0·99	0·99	0·99	0·99
μ_2	8·50	8·16	6·55	8·21	7·89	8·35	7·76
m_2^*	0·00	0·00	2094·73	6301·31	4448·27	2913·03	4079·03
w_2^o	2824·89	2034·63	429·11	2076·43	1454·28	2553·09	1295·96
γ_3	0·19	0·21	0·16	0·18	0·18	0·15	0·23
P_3	0·99	0·99	0·99	0·99	0·99	0·98	0·99
μ_3	7·88	8·04	8·28	8·64	8·85	8·29	8·39
m_3^*	3909·12	5757·77	2731·44	0·00	7478·37	6441·90	4971·57
w_3^o	1529·64	1822·16	2423·29	3203·71	3815·76	2408·98	2437·20
γ_4	0·30	0·29	0·34	0·32	0·32	0·34	0·27
P_4	0·99	0·99	0·99	0·99	0·99	0·99	0·69
μ_4	6·81	7·16	7·71	7·55	7·61	8·79	8·65
m_4^*	2414·64	2991·78	3307·42	3154·94	2988·00	2910·45	3111·34
w_4^o	523·85	750·98	1371·38	1081·79	1106·66	3963·01	3156·79
γ_5	0·39	0·40	0·35	0·36	0·36	0·35	0·37
P_5	0·99	0·99	0·99	0·99	0·99	0·99	0·99
μ_5	7·59	7·80	7·31	8·24	8·47	7·95	8·85
m_5^*	3569·88	3588·69	3447·64	3827·34	3254·76	3482·44	7174·59
w_5^o	1136·47	1428·23	919·93	2157·22	2610·31	1715·33	3878·89
σ_m^2	0·28	0·31	0·42	0·25	0·18	0·37	0·21
σ_v^2	0·02	0·02	0·01	0·03	0·00	0·00	0·00
ρ	0·96	0·97	0·99	0·94	0·99	0·99	0·99
log-L	−1456·37	−1667·89	−2853·75	−5097·78	−665·08	−1468·01	−1617·32
N	174	184	322	543	84	167	180

* w_k^o = The expected offered wage for type $k = \exp\{\mu_k + 0{\cdot}5\sigma_m^2 + \ln(0{\cdot}5)\}$.

The estimated value of ρ is above 0·93 for all groups and is as high as 0·99 for four of the groups. This result, which is consistent with there being at least one type with a very low reservation wage, indicates that a very small fraction of the variance of accepted wages is due to measurement error. Therefore, almost all of the wage dispersion within schooling/race groups is due to the inherent variation in match productivity, σ_m, faced by each type and to the differences in the μ's match productivity means, across types.[21]

21. The model is estimated using Guass 3.0 and the MAXLIKE procedure to maximize equation (20) modified for complete spells with and without wage data, and incomplete spells. The parameters are restricted to their assumed domain. That is, parameters that are assumed to be strictly between zero and one (e.g. γ, P_k's, etc.) are estimated as a logistic function and positive parameters are exponentiated transforms. The programme provides standard errors for the transformed parameters. We do not report the estimated standard errors for the model's parameters to conserve on space and because our analysis of the fit of the model is based on alternative methods.

TABLE 4

*Survivor functions, goodness-of-fit tests and mean duration of unemployment**

Quarter	Schooling/race group HSINC B		HSINC W		HSGRD B		HSGRD W	
	K–M	Model	K–M	Model	K–M	Model	K–M	Model
0	1·00	1·00	1·00	1·00	1·00	1·00	1·00	1·00
1	0·82	0·79	0·73	0·76	0·72	0·75	0·53	0·57
2	0·74	0·70	0·62	0·64	0·59	0·61	0·40	0·43
3	0·62	0·62	0·53	0·55	0·49	0·51	0·32	0·33
4	0·55	0·56	0·48	0·47	0·41	0·44	0·25	0·27
5	0·49	0·50	0·38	0·41	0·35	0·39	0·21	0·22
6	0·42	0·46	0·33	0·35	0·30	0·34	0·17	0·18
7	0·38	0·42	0·29	0·31	0·28	0·30	0·15	0·16
8	0·36	0·38	0·26	0·27	0·25	0·27	0·12	0·13
9	0·33	0·35	0·23	0·24	0·24	0·24	0·10	0·11
10	0·30	0·33	0·20	0·22	0·22	0·22	0·09	0·10
11	0·24	0·30	0·19	0·20	0·21	0·20	0·08	0·08
12	0·24	0·28	0·17	0·18	0·18	0·18	0·06	0·07
13	0·24	0·26	0·15	0·16	0·17	0·17	0·05	0·06
14	0·23	0·25	0·13	0·15	0·16	0·15	0·04	0·05
15	0·20	0·23	0·13	0·13	0·15	0·14	0·04	0·05
16	0·20	0·22	0·11	0·12	0·13	0·13	0·03	0·04
17	0·18	0·21	0·10	0·11	0·11	0·12	0·02	0·04
χ^2 test	25·11		10·14		24·70		24·81	
(m.s.l.)	(0·07)		(0·43)		(0·10)		(0·10)	
Mean duration								
K–M	8·25		5·46		5·30		2·68	
Model	6·85		5·27		5·18		2·90	
Mean *Pm*	0·21		0·24		0·27		0·45	
Mean *P*	0·99		0·99		0·98		0·97	

Table 4 reports evidence on the fit of the model to the duration data. The two columns for each race/schooling groups compare the Kaplan–Meier estimate of the survivor function to the survivor function predicted by the model. The absolute difference between the two for any given quarter is no more than 0·06 and in most quarters it is between zero and 0·03. For only two of the seven groups does the chi-square goodness-of-fit test reject the model on a 5% level. The predicted mean duration differs by about 0·2 quarters from the estimate based on the Kaplan–Meier survivor function for all groups except black high-school dropouts, where the model under-predicts duration by 1·4 quarters. Overall, the unrestricted model fits the duration data quite well.

Table 4 also shows that mean offered probabilities vary only little across schooling/race groups, with the only group having an offer probability below 0·97 being white college graduates. However, because reservation wages and wage offer distributions differ across groups, the mean acceptance probabilities also differ. Specifically, the acceptance rate is always larger for whites of the same schooling level than for blacks and larger at higher schooling levels for both races. Blacks and the more educated tend to have high reservation wages relative to the mean of their wage offer distributions as compared to whites and the less educated. It is this relationship that accounts for the unemployment duration pattern by race and schooling shown in the data.[22]

22. All types are assumed to have the same match variance.

TABLE 4—continued

Quarter	COLINC B K–M	COLINC B Model	COLINC W K–M	COLINC W Model	COLGRD W K–M	COLGRD W Model
0	1·00	1·00	1·00	1·00	1·00	1·00
1	0·50	0·56	0·41	0·43	0·52	0·52
2	0·43	0·40	0·24	0·26	0·24	0·29
3	0·33	0·31	0·17	0·18	0·14	0·17
0	0·27	0·25	0·10	0·13	0·08	0·11
5	0·16	0·21	0·06	0·09	0·05	0·08
6	0·11	0·17	0·06	0·07	0·04	0·06
7	0·11	0·14	0·05	0·05	0·02	0·04
8	0·07	0·12	0·05	0·04	0·02	0·04
9	0·07	0·10	0·05	0·03	0·02	0·03
10	0·07	0·09	0·05	0·02	0·01	0·03
11	0·07	0·07	0·03	0·02	0·01	0·02
12	0·04	0·06	0·03	0·01	0·01	0·02
13	0·04	0·05	0·03	0·01	0·01	0·02
14	0·04	0·05	0·03	0·01	0·01	0·01
15	0·04	0·04	0·03	0·01	0·01	0·01
16	0·04	0·03	0·03	0·01	0·01	0·01
17	0·04	0·03	0·03	0·00	0·01	0·01
χ^2 test	29·89		16·08		5·17	
(m.s.l.)	(0·001)		(0·041)		(0·89)	
Mean duration						
K–M	2·15		1·14		1·22	
Model	2·49		1·33		1·41	
Mean Pm	0·44		0·58		0·57	
Mean P	0·99		0·97		0·90	

* K–M refers to Kaplan–Meier estimates. m.s.l. = Marginal significance level. The $\chi^2(T-1)$ statistic is defined as $\sum_{j=1}^{T}\{[f(j)_{K\text{-}M}-f(j)_{\text{model}}]^2/f(j)_{\text{model}}\}N$, where N is the number of observations, T are the number of quarters of duration in the sample and j is the index for the quarter. $f(j)_{K\text{-}M}$ is the density of the Kaplan–Meier estimate of survival function, and $f(j)_{\text{model}}=\sum_{k=1}^{K}\gamma_k(1-P_kPm_k)^jP_kPm_k$, for $j=1,\ldots,T-1$, and $f(T)=F(T)=\sum_{k=1}^{K}\gamma_k(1-P_kPm_k)^{T-1}$. The parameters values for the models are shown in Table 3. The degrees of freedom are only upper limit because we do not take into account that some of the parameters are estimated. The mean duration of the model is given by $\sum_{j=1}^{T}(j-1)f(j)$, where $T=17$ for the high-school groups and $T=12$ for the college groups. Quarter 1 refers to zero duration of unemployment.

Table 5 reports the predicted mean accepted wage conditional on duration and predicted means and variances of accepted and offered wages. Clearly, the model does well in fitting the wage data. The prediction error in the mean accepted wage is under 3% for all groups. Consistent with Table 2, predicted mean accepted wages are decreasing with duration of search, although not exactly monotonically for all groups. This pattern of decreasing mean accepted wage with duration arises because on average the acceptance rate (Pm) is lower for types with lower mean offered wages (μ). Thus, those with longer durations are lower productivity types on average. Figures 3(a)–(g) graphically depict the distributions of actual and predicted accepted wages. The results indicate that the model fits the accepted wage distribution and not only the mean of this distribution, for almost all race/schooling groups.

Mean predicted, accepted and offered quarterly wages differ by about 1000 dollars, reflecting the fact that reservation wages are significantly positive on average. More

TABLE 5

*Predicted mean accepted and offered wages**

	Schooling/race group						
Quarter	HSINC B	HSINC W	HSGRD B	HSGRD W	COLINC B	COLINC W	COLGRD W
1	2628	2464	2705	3010	3161	3740	3936
2	2415	2743	2561	2813	3205	3390	4038
3	2388	2756	2459	2752	3051	3294	4093
4	2359	2766	2398	2692	2822	3330	4091
5	2329	2773	2360	2640	2618	3390	4027
6	2299	2775	2330	2598	2462	3451	3903
7	2269	2773	2305	2568	2349	3507	3732
8	2238	2766	2283	2546	2270	3554	3533
9	2206	2756	2261	2531	2218	3592	3329
10	2175	2741	2241	2518	2186	3618	3143
11	2143	2729	2223	2508	2168	3632	2987
12	2112	2701	2205	2498	2159	3633	2868
13	2080	2677	2189	2488	2157	3621	2781
14	2049	2651	2174	2477	2160	3598	2720
15	2019	2623	2159	2466	2167	3563	2679
16	1989	2594	2145	2452	2175	3520	2651
17	1960	2564	2132	2438	2186	3467	2633
Actual accepted wage							
Mean	2305	2698	2520	2892	3018	3625	3973
(s.d.)	(1006)	(1348)	(1349)	(1657)	(1756)	(1917)	(1855)
Predicted accepted wage							
Mean	2368	2668	2471	2824	2978	3592	3971
Predicted offered wage							
Mean	1188	1358	1241	1948	2146	2667	2997
(s.d.)	(667)	(460)	(761)	(769)	(1022)	(1022)	(925)
S.D./Mean	0·56	0·34	0·61	0·40	0·48	0·38	0·31
Reservation wage							
Mean	1441	1756	1516	1525	1999	1790	2489

* The predicted mean accepted wage for an individual of type k is

$$\bar{w}_k = \frac{\exp\{\mu_k + 0\cdot 5\sigma_m^2 + \ln(0\cdot 5)\}}{Pm_k} \cdot \left[1 - \Phi\left(\frac{\ln w_k^* - \sigma_m^2 - (\ln(0\cdot 5) + \mu_k)}{\sigma_m}\right)\right].$$

The predicted mean accepted wage conditional on duration to quarter j is $\bar{w}_j^a = \sum_{k=1}^K \rho_{kj}\bar{w}_k$, where, ρ_{kj} is the proportion of workers of type k looking for a job at quarter j, given by $\rho_{kj} = [\gamma_k(1 - P_k Pm_k)^{j-1} P_k Pm_k]/\sum_{k=1}^K \gamma_k(1 - P_k Pm_k)^{j-1} P_k Pm_k$. The predicted mean accepted wage for the particular race/schooling group is $\bar{w}^a = \sum_{k=1}^K \lambda_k^d \bar{w}_k^a$, where λ_k^d is the proportion of workers of type k looking for a job for d quarters, such that $\lambda_k^d = [\sum_{j=1}^d \rho_{kj}]/[\sum_{k=1}^K \sum_{j=1}^d \rho_{kj}]$. Given the data, we let $d = 17$ for high-school dropouts and high-school graduates, and $d = 12$ for the college level groups. The predicted mean offered wage is $w^o = \sum_{k=1}^K \gamma_k w_k^o$, and $w_k^o = \exp\{\mu_k + 0\cdot 5\sigma_m^2 + \ln(0\cdot 5)\}$. The variance of the offered wage is $\sum_{k=1}^K \gamma_k \sigma_k^2 + \sum_{k=1}^K \gamma_k \bar{w}_k^{o2} - (\sum_{k=1}^K \gamma_k \bar{w}_k^o)^2$, where $\sigma_k^2 = \exp\{\sigma_m^2 - 1\}\exp\{2(\mu_k + \ln(0\cdot 5)) + \sigma_m^2\}$ is the variance of the offered wage for type k. The formula for σ_k^2 is derived from the variance of the log normal distribution. The values for the parameters are from Table 3.

importantly, the pattern of accepted wages distorts to some extent race/schooling wage differentials that exist solely due to differences in wage offer distributions which do not confound behaviour with market fundamentals (productivity, discrimination, etc.). The extent of the distortion varies. With respect to wage comparisons, among high school non-completers the black–white ratio is 0·89 using accepted wages and 0·87 using the offer wage, while for high-school graduates, the black–white wage ratio is 0·87 using accepted wages, but only 0·64 using offered wages. Interestingly, the mean wage offer for black

TABLE 6

*The annual rate of return to schooling by race**

	Offered wages		Accepted wages	
	Black	White	Black	White
High-school dropouts to High-school graduates	2·9	27·2	6·1	4·7
High-school graduates to College dropouts	31·5	17·0	9·4	11·9
College dropouts to College graduates	—	6·0	—	4·7

* The annual rate of return to schooling is calculated as follows. We assume that it takes 6 quarters to complete high-school, 8 quarters to drop out of college and additional 8 quarters to complete college. We ignore the direct cost of schooling and the possibility of working while in school. The rate of return is just the interest rate that equate the present value of the constant streams of the mean offered wage for each additional schooling level. Specifically, let Y_i = mean income at level i of schooling and let x be the number of quarters of additional schooling to move to income level Y_{i+1} $(Y_{i+1} > Y_i)$. The discount rate that solves for the equality of present value of the income streams solves the equation:

$$\textstyle\sum_{t=1}^{x} \beta^t Y_i - \beta^{x+1} \frac{Y_{i+1} - Y_i}{1-\beta} = \frac{(\beta - \beta^{x+1})Y_i - \beta^{x+1}(Y_{i+1} - Y_i)}{1-\beta} = 0.$$

As a result we can solve for β as $\beta = [Y_i/Y_{i+1}]^{1/x}$ and the annual rate of return is given by $(1/\beta)^4 - 1$.

high-school graduates is only 91% of the mean wage offer for white high-school non-completers. Also, the estimated parameters imply that blacks face a higher degree of inequality in wage offers than whites at all schooling levels.

In the model, the reservation wage of each type is proportional to their welfare as measured by expected wealth. The values reported in Table 5 indicate that welfare is increasing with schooling for both blacks and whites, except for the high-school graduate whites who have a lower average reservation wage than high-school dropout whites. Expected wealth is higher for whites relative to blacks who are high-school dropouts or graduates, but this is reversed for those who have some college education.

Table 6 compares internal annual rates of return to alternative schooling levels based on offered and accepted wages. In general, rates of return are higher when calculated with offered wages than with accepted wages. The largest differences are for white high-school graduates and for black college dropouts. Using accepted wages, the rate of return to high school graduation for whites is one-sixth as large as compared to the same calculation using offered wages. And, attending college has a 32% return for blacks based on offered wages, but only a 9% return based on accepted wages.[23]

To complete the picture, Table 7 reports the estimated (quarterly) cost of search by type and its average over types for each of the groups based on equation (18). The cost of search varies considerably within each group with there usually being one type with a very large cost of search and at least one type with a very low cost of search. Average search costs (over types) also varies by group. The lowest search costs are for high-school non-completers; $325 for blacks and $267 for whites, although it is only $270 for black high-school graduates. Costs of search generally increase with schooling level, except for the white college-graduate group whose search cost is even lower than the white high-school graduate group. Search costs also tend to be higher for whites than for blacks. As we have previously noted, it is not possible to tell whether these differential search costs across groups and types reflect different optimal search intensities or different search cost functions. Furthermore, the estimated values of c are gross of the value of non-market

23. These results are not sensitive to the number of types used in the estimation. Similar results are obtained assuming only three types rather than five.

TABLE 7

*Predicted cost of search**

Cost	Schooling/race group						
	HSINC B	HSINC W	HSGRD B	HSGRD W	COLINC B	COLINC W	COLGRD W
c_1	6	24	36	51	14	22	349
c_2	2825	2035	−5	80	16	1171	21
c_3	151	87	1147	3203	594	270	372
c_4	−14	−0	226	52	45	2481	1086
c_5	46	163	34	498	1006	366	759
Mean	325	267	270	782	470	1165	679

* The c's were calculated using equation (18).

TABLE 8

*The effect of a \$50 per quarter subsidy for search**

	Schooling/race group						
	HSINC B	HSINC W	HSGRD B	HSGRD W	COLINC B	COLINC W	COLGRD W
Mean duration	10·1	8·1	8·0	4·5	4·1	1·7	1·8
(% Δ)	(48%)	(52%)	(54%)	(55%)	(64%)	(31%)	(29%)
Mean accepted wage	2905	3120	3214	3065	3336	3761	4103
(% Δ)	(22%)	(17%)	(30%)	(8%)	(12%)	(5%)	(3%)

* %Δ Indicates the percent increase relative to the base line values of the model.

time and, therefore, it is difficult to interpret these estimates as measuring solely search costs.

To evaluate the empirical importance of search costs we use the model to perform an experiment of a \$50 per quarter subsidy to search. This subsidy should increase search effort and, therefore, should increase the probability of receiving offers. Because estimated offer probabilities (P's) are close to one for all groups and types, it is reasonable to conduct this experiment under the assumption that offer probabilities do not change as a result of this subsidy. Using equation (18) we calculated new reservation wages associated with the subsidy. The predicted mean duration of unemployment and mean accepted wages that result are reported in Table 8. The results indicate that this relatively small subsidy to search (unemployment compensation) causes a significant increase in the duration of search.[24] The mean duration of unemployment increases by 30 to 64%. The increase in mean accepted wages is 22 and 30% for black high-school dropouts and high-school graduates, respectively. It is 17% for white high-school dropouts, but less than 12% for all other race/schooling groups. The subsidy, of course, is much smaller relative to either the total cost of search or the mean wage offer for the higher schooling groups. Overall, given the estimates, unemployment compensation policies have a non-negligible impact in this model.

7. CONCLUDING REMARKS

This paper has attempted to interpret data on the duration unemployment associated with obtaining the first full-time job and the wage on that job within a search–matching–

24. A discount factor of 0·98 was assumed. Although the cost estimates are certainly affected by the choice of the discount factor, the large subsidy effect is robust to reasonable alternative values.

bargaining model. We estimated a model which, although not fully restricted by all of the conditions implied by the theory, fits the data well and provides an interesting economic interpretation for the observations. We find that differences in unemployment durations by race and schooling are primarily due to differential rates at which job offers are accepted rather than to differential job offer probabilities, which are close to one. We used the estimates of the model to calculate the internal rate of return to schooling associated with wages on the first job. We found that using accepted wages rather than offered wages distorted the estimated return to schooling significantly for some race/schooling groups. Using accepted wages also gave a misleading picture describing match productivity. We also used the estimates of the model to perform the policy experiment of providing a subsidy to search during the school-to-work transition. We found the duration to unemployment to be very sensitive to this subsidy and accepted wages to rise significantly for the less educated.

In Eckstein and Wolpin (1990) we showed that one could impose enough restrictions on a particular search equilibrium model to identify all of the parameters of the model. However, the model in that paper did not fit either the duration or the wage data well. To estimate all of the parameters of the search-matching model implemented in this paper would have required a particular parameterization that most likely would also not fit the data well. In our view, to make significant progress in estimating equilibrium models would require data on both market participants, firms and workers.

Acknowledgements. We have benefited from comments from Chris Flinn, Boyan Jovanovic, Roni Shachar, Asher Wolinsky, and three referees of the Review. Ezer Soref and Fizza Gillani provided excellent research assistance. We are grateful for the support from the National Science Foundation under grant no. SES-9109607 and the Binational US-Israel Science Foundation under grant no. 87-00121.

REFERENCES

AEDO, M. C. (1990) *The Schooling Decision: A Dynamic Model* (Unpublished Ph.D. Dissertation. University of Minnesota).

ALBRECHT, J. W. and AXEL, B. (1984), "An equilibrium model of search employment", *Journal of Political Economy*, **92**, 824–840.

BERKOVITCH, E. (1990), "A stigma theory of unemployment duration", in Y. Weiss and G. Fishelson (eds.), *Advances in The Theory and Measurement of Unemployment* (London: Macmillan).

BLAU, D. M. (1991), "Search for Non-Wage Job Characteristics: A Test of the Reservation Wage Hypothesis", *Journal of Labor Economics*, **9**, 186–205.

BLAU, D. M. and ROBINS, P. (1990), "Job Search Outcomes for the Employed and Unemployed", *Journal of Political Economy*, **98**, 637–655.

BURDETT, K. and MORTENSEN, D. (1993), "Equilibrium Wage Differentials and Employer Size" (mimeo, Northwestern University).

DIAMOND, P. A. (1982), "Wage determination and efficiency in search equilibrium", *Review of Economic Studies*, **49**, 217–227.

DIAMOND, P. A. and MASKIN, E. (1979), "An equilibrium analysis of search and breach of contracts, I: steady states", *Bell Journal of Economics*, **10**, 282–316.

DONOHUE, J. J. and HECKMAN, J. (1991), "Continuous Versus Episodic Change: The Impact of Civil Rights Policy on the Economic Status of Black", *Journal of Economic Literature*, **29**, 1603–1643.

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

ENGBERG, J. (1991), "The Impact of Unemployment Benefits on Job Search: Structural Unobserved Heterogeneity and Spurious Spikes" (mimeo, Carnegie-Mellon University).

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

JOVANOVIC, B. (1979), "Job Matching and The Theory of Turnover", *Journal of Political Economy*, **87**, 972–990.

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

KENNAN, J. and WILSON, R. (1989), "Strategic bargaining models and interpretation of strike data", *Journal of Applied Econometrics*, **4**, s87–s130.

MILLER, R. A. (1984), "Job Matching and Occupational Choice", *Journal of Political Economy*, **92**, 1086–1120.

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

MORTENSEN, D. T. (1990), "Equilibrium Wage Distributions: A Synthesis", in J. Hartog, G. Ridder and J. Theeuwes (eds.), *Panel Data and Labor Studies* (Amsterdam: Elsevier).

RUBINSTEIN, A. (1982), "Perfect equilibrium in a bargaining model", *Econometrica*, **50**, 97–100.

TOPEL, R. and WORD, M. (1988), "Job Mobility and the Careers of Young Men" (Working Paper No. 2649, NBER, Cambridge, Mass.).

VAN DEN BERG, G. J. and RIDDER, G. (1993), "Estimating an Equilibrium Search Model From Wage Data" (mimeo, Vrije Universiteit, Amsterdam).

WOLINSKY, A. (1987), "Matching, search and bargaining", *Journal of Economic Theory*, **42**, 311–333.

WOLPIN, K. I. (1987), "Estimating a structural search model: The transition from school to work", *Econometrica*, **55**, 801–817.

WOLPIN, K. I. (1992), "The determinants of black-white differences in early employment careers: search, layoffs, quits and endogenous wage growth", *Journal of Political Economy*, **100**, 535–560.