Search Friction in the U.S. Labor Market:
Equilibrium Estimates from the PSID

Audra J. Bowlus
University of Western Ontario & Free University Amsterdam

and

Shannon N. Seitz*
University of Western Ontario

First draft: May 1998
Current version: August 1998

Abstract

In this paper we determine the feasibility of using data from the Panel Study of Income Dynamics to estimate the Burdett-Mortensen general equilibrium search model. The data contain sufficient information on wages, labor force states, durations, and transitions to generate estimates of the model's structural parameters. Our analysis compares the relative labor market search friction for black and white male household heads. In general we find blacks face greater search friction while unemployed than whites, but a similar level while employed. Within the model this finding implies substantial productivity differentials are needed to generate the black-white wage differentials found in the data.

---

*Department of Economics, Social Science Centre, University of Western Ontario, London, Ontario N6A 5C2, Canada. Correspond to abowlus@julian.uwo.ca and seitz@julian.uwo.ca. We thank participants at the Sandbjerg Panel Data and Structural Labour Market Models conference for helpful comments. This paper was written while Audra Bowlus was a visiting researcher at the Free University of Amsterdam. Both authors are grateful for support from the Social Sciences and Humanities Research Council of Canada. We alone are responsible for any errors.

1. Introduction

In recent years general equilibrium search models have been used to describe and explain the inter-relationships between worker flows and wage determination. The development of estimation techniques for the Burdett-Mortensen (1998) (see also Mortensen (1990)) model,[1] in particular, and the availability of panel data have allowed researchers to estimate structural search parameters for several different countries including the U.S., Canada, Denmark, the Netherlands and France. The U.S. estimates have come from the National Longitudinal Survey of Youth (NLSY). The NLSY has been used to study the school-to-work transition for black and white males (Bowlus, Kiefer and Neumann, 1998), male-female wage differentials (Bowlus, 1997) and the U.S.-Canada unemployment rate gap (Bowlus, 1998). The main drawback of the NLSY is its limited focus on youth.[2] Thus, the parameter estimates from the NLSY only reflect a small subset of the U.S. population and limit the degree to which the U.S. labor market can be characterized and cross-country comparisons can be made.

To address this deficiency of the NLSY, we examine the feasibility of using the U.S. Panel Study of Income Dynamics (PSID) for estimation. Two aspects of the PSID make it attractive for the estimation of general equilibrium search models. First, the PSID contains a representative sample of households, in direct contrast to the NLSY. As a result, estimates from the PSID apply to a broader segment of the U.S. labor market than those in previous studies. Second, detailed work history information, necessary for the estimation of the Burdett-Mortensen (1998) model, has been collected by the PSID since 1984. In 1988, the PSID switched from a position-based job spell definition to an employer-based definition.[3] To maintain a consistent job spell definition and to capture a more recent time period,[4] we use data from 1988-1992. Despite its potential advantages, we encountered some problems with the PSID while constructing the data set to be used in estimation. Unlike the NLSY, which provides codes to link job spells over interviews, job spells recorded in the PSID must be linked by matching on job characteristics. In principle this

[1]A variety of estimation techniques have been developed for different versions of the Burdett-Mortensen model. See, for example, Kiefer and Neumann (1993), Bowlus, Kiefer and Neumann (1995, 1998), Ridder and van den Berg (1998) and Bontemps, Robin and van den Berg (1997).

[2]The NLSY sample, first interviewed in 1979, was restricted to individuals between the ages of 14 and 21.

[3]The NLSY also adopts an employer-based definition.

[4]The NLSY studies tend to use first spells that occur mainly in the early 1980's.

is possible given the information collected. However, in practice we found that we were not able to successfully link job spells across the survey years for many respondents. This, in conjunction with a short panel of five years, leads to a job spell censoring rate that is quite high relative to other studies.

To estimate the Burdett-Mortensen (1998) model we use the methodology developed by Bowlus, Kiefer and Neumann (1995, 1998). We find that the data collected from the PSID are sufficient for identification of the model's parameters. As in other studies the model is able to provide a relatively good fit to the wage data. However, it has substantial difficulties fitting the job duration data in conjunction with the other labor market transitions found in the data. Because of the poor quality of the job duration data in the PSID, we also estimate the model excluding the job duration data. The model is estimated for both white and black males in the U.S. and for different age and education groups. The following conclusions can be drawn from the estimates. First, blacks tend to face a greater level of search friction while unemployed than whites, but a similar level while employed. This result holds for the overall sample, younger workers and high school graduates and implies a substantial productivity differential is needed to generate the black-white wage differentials found in the data. Second, young workers tend to have lower levels of search friction than their older counterparts. Finally, respondents with some college are found to face more search friction while employed than either those with high school or university degrees.

Our results can be compared most closely to Bowlus, Kiefer and Neumann (1998) who study the transition from school-to-work for blacks and whites using the NLSY. They find young black high school graduates face a higher job destruction rate, a lower job offer arrival rate while unemployed and a similar job offer arrival rate while employed as compared to whites. They conclude that a substantial portion of the black-white wage differential can be attributed to the higher job destruction rate for blacks. Our results, based on a sample with a broad age range, are consistent with respect to the job destruction and unemployed job offer arrival rate relationships found in the Bowlus, Keifer and Neumann study. However, we find that blacks in general have a higher arrival rate of offers while employed than whites, offsetting the effects of the higher job destruction rate for blacks. As a result, search frictions are found to play almost no role in determining the black-white wage differentials and the model is left to explain the wage differentials found in the data through productivity differences.

2. Model

The Burdett-Mortensen (1998) general equilibrium search model is a 2 state (employment and unemployment) model of the labor market with on and off the job search and a job destruction process. In the simplest version of the model, described briefly here, workers and firms are homogeneous. That is, workers have a common value of non-market time b and firms have a common productivity level P.

Three exogenous processes drive the movement of workers in the model: the arrival of job offers at rate λ_0 while unemployed, the arrival of job offers at rate λ_1 while employed, and the destruction of jobs at rate δ. Workers, taking the wage offer distribution of the firms $F(w)$ as given, solve a standard linear search utility maximization problem and adopt a state-contingent reservation wage strategy. While unemployed a worker's reservation wage r is (Mortensen and Neumann, 1988):

$$r = b + (\kappa_0 - \kappa_1)\int_r^{\infty}\left[\frac{1 - F(w)}{1 + \kappa_1(1 - F(w))}\right]dw \tag{1}$$

where $\kappa_0 = \lambda_0/\delta$ and $\kappa_1 = \lambda_1/\delta$. Since unemployed workers accept the first job that offers more than r, the exit rate from unemployment is $\lambda_0(1\text{-}F(r))$. While employed a worker's reservation wage is the current wage w. Employed workers may exit from their current jobs for two reasons. First, a job can end because it is destroyed, returning workers to unemployment at rate δ. Second, a worker can accept a new job offer at rate $\lambda_1(1\text{-}F(w))$.

The production process for firms is constant returns to scale with profits $\pi(w)$ equal to

$$\pi(w) = (P - w)\, l(w) \tag{2}$$

where $l(w)$ is the steady-state labor stock of a firm offering wage w. A firm maximizes profits by choosing w given the wage posting and search strategies of the other firms and workers, respectively. Since firms are equally productive, profits in equilibrium must be equal across firms. Wage offers, however, need not be equal in equilibrium. Because of on-the-job search, $l(w)$ is increasing in w as offering a higher wage attracts more workers and enables firms to retain them longer. Firms offer higher wages if and only if $l(w)$ increases enough to offset the lower per worker profits. In equilibrium the model generates a non-degenerate wage offer distribution shown in Burdett and Mortensen (1998) to be:

$$F(w) = \frac{1+\kappa_1}{\kappa_1}\left[1 - \left(\frac{P-w}{P-w_L}\right)^{1/2}\right] \qquad w_L \le w \le w_H \tag{3}$$

with density

$$f(w) = \left[\frac{1+\kappa_1}{2\kappa_1}\right]\frac{1}{\sqrt{(P-w)(P-w_L)}} \tag{4}$$

where the lowest wage offered by firms w_L equals r, and the highest wage offered w_H is given by $F(w_H)=1$. Since no worker will accept a job with a wage lower than the reservation wage, firms never offer wages below r and thus $F(r)$ equals 0. Workers move up the wage offer distribution through on-the-job search; therefore, the steady-state cross-section earnings distribution $G(w)$ is not equal to the wage offer distribution $F(w)$. Rather, $G(w)$ lies to the right of $F(w)$ and is given by

$$G(w) = \frac{F(w)}{1+\kappa_1(1-F(w))} \ . \tag{5}$$

Finally, the unemployment rate u is found by equating the flow of workers in and out of unemployment and is equal to

$$u = \frac{\delta}{\delta+\lambda_0} = \frac{1}{1+\kappa_0} \ . \tag{6}$$

The ratios κ_0 and κ_1 characterize the degree of search friction in the labor market. These ratios of job offer arrival rates to the job destruction rate affect workers through their reservation wage and firms through their wage offers. A value of zero for λ_1 (and therefore κ_1) results in a degenerate wage distribution with firms only offering the reservation wage r. As λ_1 increases to infinity - complete information and perfect mobility, the competitive solution where all workers are paid their marginal product P emerges. For values of λ_1 between zero and infinity, the non-degenerate wage offer distribution given in (3) results. The existence of search friction in the labor market gives firms monopsony power over workers. An index measuring monopsony power, suggested by Mortensen (1990), is defined as $MI = E((P\text{-}w)/P)$. Values of the index range between 1 and 0 as w ranges between 0 and P, where higher values of the index correspond to greater monopsony power in the labor market. If firms have complete monopsony power, workers are paid 0; firms with no monopsony power pay the competitive wage P.

Although the homogeneous version of the Burdett-Mortensen model outlined above can account for many empirical labor market regularities, it produces a convex-shaped earnings distribution that is at odds with the data. Recent studies (Bowlus, Kiefer, Neumann, 1995,1998; Bontemps, Robin and van den Berg, 1997) overcome this shortcoming by adding firm heterogeneity through variation in productivity levels, greatly improving the fit of the model to wage data. In this paper we follow Bowlus, Kiefer, and Neumann and assume a discrete distribution of Q firm types that can be ordered as $P_1 < P_2 < ... < P_Q$. Mortensen (1990) shows the solution under discrete firm heterogeneity results in a segmentation of the wage offer range among firm types: low productivity firms offer low wages, while high productivity firms pay higher wages. In equilibrium, equal profits within, but not across, firm types are required. The equilibrium wage offer distribution in the wage range for firm type j is given by:

$$F_j(w) = \frac{1+\kappa_1}{\kappa_1} - \frac{1+\kappa_1(1-\gamma_{j-1})}{\kappa_1}[\frac{P_j-w}{P_j-w_{Lj}}]^{1/2}, \quad w_{Lj} \leq w \leq w_{Hj} \tag{7}$$

where γ_j is the fraction of firms with productivity P_j or less, w_{Lj} is the lowest wage offered by a firm of type j, and w_{Hj} is the highest wage paid by a type-j firm. The following restrictions are implied by the model: $w_{L1} = r$ and $F(r) = 0$; $w_{Hj} = w_{Lj+1}$ and $F(w_{Hj}) = \gamma_j$, j=1,...,Q-1; and $F(w_{HQ})$ = 1. The wage range of the earnings distribution is divided at the same wage cuts as the offer distribution's. Thus, the fraction of workers earning w_{Hj} or less is equal to

$$G(w_{Hj}) = \gamma_j^G = \frac{\gamma_j}{1+\kappa_1(1-\gamma_j)} . \tag{8}$$

3. Data

Panel data are needed to estimate the Burdett-Mortensen (1998) model. Specifically, work history information at the job spell level is required. In the past, the NLSY was used in U.S. studies because it contains detailed employment histories. However, the NLSY is limited in focus to youths. Thus, without a more representative sample one cannot study the labor market behavior of differing age groups nor make valid cross-country comparisons. In this paper, we investigate the feasibility of using the more representative PSID to estimate the structural parameters of the search model.

The PSID originated in 1968 when a sample of 4802 households in the U.S. was interviewed. In subsequent years, the original households have been followed as well as the new households formed by their children. Detailed work history information was first collected for household heads and wives in 1984.[5] In 1988, the definition of a job changed from position- to employer-based. Currently data based on this definition are available until 1992. To maintain consistency in the job definition and to maximize the time period covered in estimation we use data from 1988-1992. This choice also allows us to compute estimates for a more recent time period as compared to the NLSY studies, which tend to cover the early to mid-1980's. The PSID records up to two main jobs and two extra jobs held by each head during the previous year.[6] If the individual held more than two jobs in either category, a supplement containing job history information for all additional jobs was collected. This supplemental information is currently not available for the 1988-1992 period. Fortunately, only 3% of the sample recorded a supplement.

To estimate the model, information on job spell durations, unemployment durations, wages and transitions between jobs and from employment to unemployment is required. Because the Burdett-Mortensen (1998) model contains only two labor market states, we focus on male household heads[7] who are active in the labor market. To do so we eliminate heads who are disabled, retired or students at the time of the 1988 interview. This restriction, however, does not remove individuals who are out of the labor force for other reasons. Thus, the rates and durations calculated for the samples are nonemployment, not unemployment, rates and durations and are labeled as such. In addition, respondents who are employed (nonemployed) in 1988 must report

[5]Prior to 1984 only information on the current or most recent job was collected.

[6]The distinction between main and extra jobs is made by the respondent.

[7]See Bowlus (1997) for a study of both males and females using a 3 state version of the Burdett-Mortensen model, where unemployment and non-participation are distinct states.

a valid current (most recent) main job.[8] There are 3,758 male household heads who meet these criteria.[9]

We attempt to follow these respondents through a complete job spell cycle, i.e. from the starting date of one job spell to the starting date of the next. Employed individuals are followed until the current job spell ends or is censored. If the job spell ends, the starting date of the next job spell is recorded. If the starting date of the next job spell is the same as or prior to the ending date of the previous spell, the individual is recorded as experiencing a job-to-job transition.[10] If not, the individual is recorded as making a transition to nonemployment and the length of the nonemployment spell is calculated. For respondents who are not employed at the 1988 interview, we find the first job reported with a starting date after the 1988 interview and calculate the nonemployment duration. As with those who are employed, we continue to follow these individuals through a complete job spell cycle.

There is no left censoring in our data, because the starting dates of all jobs and nonemployment durations are known. Right censoring, however, can occur due to attrition, the inability to match job spells over time, or because the sample period ends. Attrition only accounts for 8.3% of the censored observations; therefore most censoring occurs because the sample period ends (52.3%) or because of problems matching spells across interviews (39.4%). With respect to the former it appears that for male household heads five years is not a long enough time period to observe many transitions. This problem arises in part because of the representativeness of the PSID sample, which contains an older group of men with more stable employment than those found in the NLSY studies. The latter matching problem stems from the way the job spell information is recorded in the PSID. Unlike the NLSY which provides codes that link job spells across interviews, the PSID only records characteristics of each job reported. Therefore, to match spells over time, job characteristics must be matched across interviews.

[8]For a main job to be valid the starting (ending) month and year of the job must be recorded if the individual is currently employed (unemployed).

[9]In 1988 7,114 households were interviewed of which 5,010 had a male household head. Eliminating heads who are disabled, retired or students at the time of the 1988 interview leaves 4,209 respondents. Of those 3,758 recorded a valid main job in 1988.

[10]We know the starting and ending month and year for each job spell. Thus, if a job ends and another begins within the same month, a job-to-job transition is recorded.

To link the PSID job spells, a matching routine was created based on job starting dates and industry codes. Our matching routine attempts to maintain a balance between accurately recording each individual's work history and minimizing censoring due to mis-matches. For each individual an attempt is made to match one job spell forward from the current interview with one of four (2 main jobs and 2 extra jobs) possible jobs in the next interview. If a match can be made, we check to see whether the spell needs to be matched further with spells in the next interview or if it is complete. If a match cannot be made, we move on to the next level. If a spell remains unmatched after going through the entire matching process, it is censored at the last known date the individual held the job. We develop a set of criteria that begins by demanding a perfect match on starting dates and industry codes. If a successful match cannot be made, the matching criteria are relaxed. In order of restrictiveness the following criteria allow for a spell to make a valid match: 1) starting month, year and industry code of the spell are the same as those for its match, 2) starting year and industry codes are the same but starting month of its match is off by up to 2 months, 3) starting year and month are equal but the industry code of its match is different,[11] 4) starting year and industry codes are equal but starting month of its match is missing, and 5) for a spell starting before 1988 its match can have any starting date prior to 1988 if the industry codes are the same and any date prior to the spell's starting date if they are not.[12] Of the 3639 job spells that enter the matching routine 2698 are returned right-censored: 625 because of mismatches,[13] 438 because of missing employment information,[14] 223 due to attrition and 1412 due to the end of the sample period.

Since the supplemental job information is not available, the job spell cycle collected for respondents with supplements may also be censored, as matches for some job spells may be contained in the supplements. Supplements also cause problems when determining the starting date of the next job to complete a cycle. If the individual has a supplement in the same year he

---

[11]In most of these cases, the industry code is missing. However, it is possible that the industry codes are different across interviews. In terms of the model we need to know how long the individual is employed in the same job. Since the starting dates match, the match is deemed to be valid with regard to the information required.

[12]The reasoning for the fifth criteria is similar to that for the third criteria.

[13]Refers to cases in which spells enter the matching routine along with possible match candidates, but a match can not be made based on our criteria.

[14]Refers to cases in which no possible match candidates are recorded, i.e. no spells are given with starting dates prior to the last interview date.

ends a job, the starting date of the next job may be contained in the supplement. For this reason we stop his cycle with the end of the job spell duration. For those individuals who are nonemployed and record a supplement in the next interview we censor the nonemployment duration at the time of the last interview. Only 3% of the sample has an incomplete work history due to a supplement. Overall we are able to follow 62% of the sample through a complete job spell cycle or until the end of the survey period, which ever comes first. The other 38% experience some sort of break in their work history.

After constructing the job spell cycles we further restrict the sample to be more consistent with the model and other studies. First, to avoid issues of retirement we eliminate male heads who are older than 60 in 1988. Second, individuals who are enrolled in school at any time during the sample period are removed. Third, we restrict the sample to individuals who are employed in full-time (at least 35 hours per week), private sector jobs during the job spell cycle. Fourth, to be more consistent with a 2-state model we remove individuals who spend longer than three consecutive years out of work within their job spell cycle or prior to the start of their cycle if nonemployed in 1988. Fifth, individuals who have job durations longer than their age minus 15 are eliminated. Finally, respondents must record their age and education level in 1988, and information on race must be available. The resulting sample consists of 1480 white respondents and 618 black respondents.[15]

4. Econometric Methodology

The sample used in our analysis is similar to the sample used by Bowlus (1998) in that individuals are first sampled from the stock rather than the flow. Therefore, we use many of the modifications to the Bowlus, Kiefer, and Neumann (1998), hereafter BKN, estimation method found in Bowlus (1998). Since individuals are sampled from the stock of nonemployed and employed, the contribution of an individual's spell to the likelihood function depends on the state he is in at the time of the 1988 interview. Consider each case separately starting with those individuals who are employed. We have data on job spell durations, D_J, current wages, w, whether jobs ended because they were lost (C = 0) or left (C = 1), and if lost the subsequent durations of the nonemployment spells, D_N. In steady-state the probability of being employed is

[15]We do not study individuals who are American Indian, Aleut, Eskimo, Asian, Pacific Islander or are of Hispanic or Spanish descent.

1 minus the unemployment rate given in equation (6) or $\lambda_0/(\lambda_0+\delta)$.[16] The observed wage is a draw from the cross-section earnings distribution and is therefore distributed as $g(w)$, the probability density function (*pdf*) of $G(w)$. The job spells for employed individuals are distributed Gamma with parameters 2 and $(\delta+\lambda_1(1\text{-}F(w)))^{-1}$ and density

$$f(D_J|S=1) \;=\; (\delta+\lambda_1[1-F(w)])^2 D_J \exp\{-(\delta+\lambda_1[1-F(w)])D_J\} \tag{9}$$

where S = 1 if the individual is employed in 1988 (Lancaster 1990, p.94). The probability of a job spell ending due to an acceptable outside job offer is

$$P(C=1|w) \;=\; \frac{\lambda_1(1-F(w))}{\delta+\lambda_1(1-F(w))}\;. \tag{10}$$

The probability a job is lost, $P(C=0|w)$, is $1\text{-}P(C=1|w)$ and the subsequent nonemployment spell is distributed exponential with rate λ_0. The total contribution to the likelihood function of an employed worker $\mathcal{L}_E$ is the product of these pieces.

$$\begin{aligned}\mathcal{L}_E = \frac{\lambda_0}{\lambda_0+\delta} g(w)\,\{[\delta+\lambda_1(1-F(w))]^2 D_J\}^{d_J}\{[\delta+\lambda_1(1-F(w))]D_J+1\}^{1-d_J}\\ \exp(-[\delta+\lambda_1(1-F(w))]D_J)\,[\frac{\delta^{1-C}[\lambda_1(1-F(w))]^C}{\delta+\lambda_1(1-F(w))}]^{d_J d_T}\;[\lambda_0^{d_N}\exp(-\lambda_0 D_N)]^{(1-C)d_J d_T}\end{aligned} \tag{11}$$

where d_J (d_N) equals 1 if the job (nonemployment) spell is complete and 0 otherwise and d_T equals 1 if a transition following the job spell is observed and 0 otherwise.

Next, we consider the contribution of those individuals who are not employed at the 1988 interview (S = 0). The nonemployment spells in progress at the time of the interview, D_{N0}, are sampled from the stock and the probability of observing this state is the unemployment rate, $\delta/(\lambda_0+\delta)$. Because we know the ending date of the most recent job spell, we observe the duration of the entire nonemployment spell. As is the case for job spell durations, nonemployment spells are distributed Gamma with parameters 2 and λ_0. Most nonemployed respondents accept employment sometime before 1992. The wages received on these jobs are distributed as $f(w)$, the *pdf* of the wage offer distribution $F(w)$. The job spells, D_J, are thus sampled from the flow and

[16]In a two-state model the nonemployment rate equals the unemployment rate.

are exponentially distributed with rate $\delta + \lambda_1(1\text{-}F(w))$. The probabilities of leaving a job due to an acceptable outside offer or of losing a job are the same as for employed individuals, as is the distribution of nonemployment spells following completed job spells that are lost, D_N. The contribution to the likelihood function $\mathcal{L}_N$ is once again the product of these terms

$$\mathcal{L}_N = \frac{\delta_0}{\lambda_0+\delta}[\lambda_0^2 D_{N0}]^{d_{N0}}[\lambda_0 D_{N0}+1]^{1-d_{N0}}\exp(-\lambda_0 D_{N0})\{f(w)[\delta+\lambda_1(1-F(w))]^{d_J}$$
$$\exp(-[\delta+\lambda_1(1-F(w))]D_J)\}^{d_{N0}}[\frac{\delta^{1-C}[\lambda_1(1-F(w))]^C}{\delta+\lambda_1(1-F(w))}]^{d_{N0}d_J d_T}\ [\lambda_0^{d_N}\exp(-\lambda_0 D_N)]^{(1-C)d_{N0}d_J d_T} \tag{12}$$

where d_{N0}, d_J and d_N are completed spell indicators and d_T is defined as before.

Given the assumption of firm heterogeneity, the likelihood is not only a function of the arrival rates and productivity parameters, but of the wage cuts as well. Following BKN we reduce the number of parameters by replacing P_j in the likelihood function with the following expression derived from the restriction $F(w_{Hj}) = \gamma_j$:

$$P_j = \frac{w_{Hj}-B_j w_{Hj-1}}{1-B_j} \quad where \quad B_j = [\frac{1+\kappa_1(1-\gamma_j)}{1+\kappa_1(1-\gamma_{j-1})}]^2. \tag{13}$$

The likelihood is now a function of both arrival rates and the wage cutoffs. There are kinks at the wage cuts in $F(w)$, as well as in $G(w)$. Therefore, $f(w)$, $g(w)$ and the likelihood function are discontinuous, leaving standard maximization procedures inappropriate. BKN show the maximum likelihood estimates for r and w_{HQ} are the sample minimum and maximum, respectively, and those for $w_{H1},\ldots,w_{HQ-1}$ come from the set of observed wages. The estimators of the wage cuts are superefficient, and therefore can be treated as known when estimating the remaining parameters using standard maximum likelihood techniques.[17]

Rather than searching over all possible combinations for the wage cut estimates, BKN employ simulated annealing which randomly searches over the possible combinations and stops

[17]The presence of measurement error in wages results in the observed wage distribution being the convolution of the underlying wage distribution, which has discontinuities, and that of the measurement error. This results in a smoothing of the observed distribution that may obscure the presence of the wage cuts. While care has been taken with the data samples to remove obvious outliers, measurement error is not directly dealt with here. See Van den Berg and Ridder (1998) for an alternative specification and estimation strategy of the Burdett-Mortensen model that incorporates measurement error.

according to an optimal stopping rule.[18] Estimation involves iterating over a two-step procedure. First, holding the arrival rates fixed, w_{H1} through w_{HQ-1} are estimated from the set of observed wages using simulated annealing. In contrast to BKN, where all job spells are sampled from the flow and thus the observed wages all stem from the wage offer distribution $F(w)$, in our case most wage observations are sampled from the cross-section earnings distribution $G(w)$. To handle this difference we follow Bowlus (1998) and rewrite the likelihood function in terms of γ_j^G's instead of γ_j's using equation (8). We then search for the wage cuts over the subset of wages from $G(w)$ and estimate the γ_j^G's off the empirical cumulative density function (*cdf*) of this subset. Once a wage cut combination has been found that maximizes the likelihood function given the fixed rates, the wage cuts and their respective γ_j^G's are held fixed and standard maximization procedures are used to estimate the rates. Iterations of this procedure are conducted until convergence is achieved. Q is chosen by comparing the improvement in the log likelihood function with each additional firm type. Since the exact distribution of the difference in log likelihood values is yet to be worked out, we follow BKN and compare 2 times the change to Chi-square critical values.

Given the estimates for r and w_{HQ} are the sample minimum and maximum, respectively, one must deal with outliers in the wage data. For each job spell the PSID reports a starting, current and ending hourly wage as well as the number of hours worked per week for each period.[19] For spells ongoing in 1988 we use the current wage and for those beginning after the 1988 interview we use the starting wage. Similar to BKN's treatment of wage outliers in the NLSY we compare the hourly wages given in the PSID with upper and lower bounds collected for males of the same education from the Current Population Survey (CPS) for 1988-1992.[20] Wage observations outside of the bounds are treated as missing. Any individuals with missing wage data are not dropped from the sample: their contribution to the likelihood function includes nonemployment durations and labor force status in 1988. All acceptable wage responses are converted to real weekly wages with 1990 as the base year.

---

[18]Given the size of the data sets in this paper searching over all possible combinations becomes computationally time-intensive at Q levels beyond two.

[19]For salaried workers the PSID calculates the hourly wage rate by dividing annual (weekly) salaries by 2000 (40).

[20]We use the 5th and 95th percentiles from the set of hourly wages for paid hourly workers from the March outgoing rotation groups for each year. In this way we do not have to calculate an hourly wage based on annual earnings. The lower bounds are quite close to the minimum wage.

The model has a difficult time reconciling the relatively few and disperse high wage observations with the majority seen in the data. The best it can do is attach different firm types to each of these wages. This results in a chopping of the right tail of the wage distribution into firm types and implausible productivity values. To avoid problems with long right tails in the wage distribution during estimation, we also trim the data 1% at the top.[21] This trim level yields more reasonable productivity values while preserving the main features of the data. The combined effects of the CPS bounds and further trimming at the top results in 1120 valid wage observations for the whites and 530 for the blacks.

4. Results

In table 1 we present summary statistics for the components that enter the likelihood function for blacks and whites separately. Our sample is comprised of active labor market participants as the nonemployment rates are very close to the national unemployment rate for males of 3.2% in 1988. Comparing across blacks and whites we find whites have a lower nonemployment rate, lower mean nonemployment duration, higher mean job spell duration, and a higher job-to-job transition rate than blacks. Whites also earn on average 23% more than blacks. The *cdf*s for blacks and whites are shown in Figure 1. We note that whites are slightly older and more educated than blacks in the overall sample. The mean age for whites is 35.8 years compared to 34.1 years for blacks and the average education levels are 13.2 and 12.1 years for whites and blacks, respectively. While the overall sample sizes are relatively large, the high censoring rates and low nonemployment rates result in small sample sizes for the determination of the job-to-job transition rates and the means for nonemployment durations, job durations sampled from the flow, and wages sampled from the wage offer distribution.

Columns 1 and 2 of table 2 contain the estimates derived from our estimation procedure for whites and blacks, respectively. Our criteria for choosing Q indicates 9 firm types are needed for whites and 8 for blacks to adequately fit the wage data.[22] The parameter estimates indicate a much lower level of λ_0 for blacks than whites, a higher level of λ_1 for blacks and similar levels of δ. This results in a much lower level of κ_0 for blacks than whites but a higher level of κ_1. The

---

[21]This problem was first noted by Kiefer and Neumann (1993). Since then trimming has become a common means of generating reasonable productivity values.

[22]This level of Q is somewhat higher than that found in other studies. It is likely due to the larger sample sizes and wider wage ranges in the PSID data as well as a lower degree of homogeneity in the sample.

lower half of table 2 shows that, while the parameter estimates result in predicted unemployment rates, mean unemployment durations, mean earnings and transition probabilities that are consistent with those found in table 1 for whites, they do not for blacks. The predicted mean earnings for blacks, although a little high, is consistent with the data, but the black unemployment rate is lower and the fraction of jobs ending in a job-to-job transition is higher than those found in the data. The job-to-job transition rate for blacks is even higher than the predicted value for whites.

The inability to match the unemployment and transition rates in the data does not stem from the choice of a discrete productivity distribution, as parameter estimates from Bontemps, Robin and van den Berg's (1997) estimation method for the continuous distribution case result in a κ_1 value of 4.38 for blacks. The failure of the model lies with its difficulty in matching the job duration data, and in attempting to match the features of the duration data, it fails to match the unemployment and transition data. In the data very short job spells of less than 2 months are present in conjunction with very long spells of more than 20 years. For blacks the ratio of the 75th percentile of the job duration distribution to the 25th percentile is 6. To match such a wide range of tenure, the model requires a low job destruction rate δ to generate long spells for high wage jobs and a high arrival rate λ_1 to generate short spells for low wage jobs. Together these forces result in a high value for κ_1 for the blacks. As a result, fitting the job duration data leads to inconsistent predictions regarding the unemployment rate and job-to-job transitions which need a higher δ and a lower λ_1, respectively. For whites, the dispersion of the job duration distribution is not as great as for the blacks with a 75-25 percentile ratio of less than 4. The lower unemployment rate and higher job-to-job transition rate for the whites also helps in matching both the long and short spells found in the data.

The difficulties in fitting the job duration data within the Burdett-Mortensen (1998) model are not new. BKN also report that the fit of the duration data is poor in the NLSY sample and conclude that there are too few long spells in their sample. Consistent with the BKN findings we predict a higher mean duration than that found in the data.[23] It is impossible to directly assess the fit of the duration data as in BKN, for the vast majority of actual finished durations are unknown because of the high censoring rate. Therefore it is difficult to know whether the model has done

[23]Note that the predicted means reported in table 2 are for durations sampled from the flow, while some of the means in table 1 are derived from stock samples and thus are not directly comparable. To obtain the predicted mean from the stock sample, one must multiply the predicted means in table 2 by 2.

a reasonable job of even fitting the PSID duration distribution. In addition, considering the problems matching jobs across years in the PSID, it is unlikely that censoring is random in cases where a match could not be made, in contrast to censoring due to the sample period ending. That is, it is more likely these spells are censored exactly because a transition has taken place although we do not observe a transition in the data. As a result, the likelihood function treats these spells as having a much longer expected duration than is actually the case. Given this problem with the data and the resulting inconsistencies with respect to other moments in the data, especially for blacks, we exclude the job duration data from the likelihood function in the remainder of the paper and estimate the parameters using only the state of employment, unemployment durations, wages, and transitions. This is possible because the three search parameters, λ_0, λ_1, and δ, are all overidentified in the original estimation procedure.[24]

The parameter estimates that result from the estimation excluding job durations are presented for both whites and blacks in columns 3 and 4 of table 2, respectively. The new estimates reveal a different characterization of the black labor market. The estimate of δ is much higher now and the estimate of λ_1 lower for the blacks relative to the estimates in column 2. The resulting lower values of both κ_0 and κ_1 for the blacks lead to predictions of a higher unemployment rate and a much lower job-to-job transition rate, which are more consistent with the data. The parameter estimates for the whites do not change with the exclusion of the job duration data. For both groups the model is able to fit the earnings distribution as shown in figure 2. Compared to whites we find that blacks face more search friction in the labor market as seen through lower values of κ_0. However, the values for κ_1 are not significantly different indicating similar employed search friction levels for the two groups. This stems from a higher value of λ_1 for the blacks which counters their lower value of δ. The monopsony index is also the same for blacks and whites. Given the similarities in κ_1, substantial productivity differences are needed to generate the black-white earnings differential found in the data.[25]

To investigate the differences between blacks and whites further we divide the samples by age into two groups: those under 35 and those 35 and over. Means for the four groups are presented in table 3. We find similar nonemployment rates and durations for whites and older

[24]The labor force state in 1988 and the nonemployment durations determine λ_0. The labor force state also helps to determine δ. Wages, job spell durations and transitions all determine λ_1 and δ.

[25]The productivity average for workers (firms) is calculated by weighting the productivity values with the workers' (firms') fractions in the market γ^G (γ).

blacks, but a much higher nonemployment rate and much longer nonemployment durations for young blacks. Surprisingly older blacks have a very low nonemployment rate. As expected, young blacks and whites have shorter job durations and higher job-to-job transition rates than their older counterparts. In terms of wages there is a larger wage differential across age groups for blacks than whites. Finally, the black-white wage differential narrows with age, but remains substantial.

Table 4 presents means for blacks and whites conditional on educational attainment. For whites there are three education groups: high school, some college and university.[26] Unfortunately for blacks only the high school category has a large enough sample size for estimation. In table 4 unemployment rates decline and wages increase with education. Upon comparison of black and white high school graduates we find many of the same relationships found in other studies including a higher unemployment rate, longer mean unemployment duration, lower mean earnings[27] and a shorter mean job duration for blacks relative to whites. However, in contrast to table 1 and evidence from the NLSY in BKN, table 4 shows a higher job-to-job transition rate for high school blacks than high school whites. In addition, it appears that whites with some college have a low job-to-job transition rate as compared to the other groups.

Tables 5 and 6 contain the parameter estimates and model predictions for the age and education groups, respectively. In table 5 we find a low value of κ_0 for young blacks, consistent with their higher nonemployment rates and longer nonemployment durations. The value of κ_0 for older blacks is 5 times greater than that for young blacks primarily because of their very low job destruction rate. The difference in κ_0's across the white age groups is not as pronounced, with older workers having a lower value because of their longer nonemployment durations. In terms of κ_1 the values are higher for both young groups and very low for older blacks. Comparing productivity estimates we find a higher productivity level for young whites compared to young blacks, but the reverse for older blacks and whites. Considering the low κ_1 value for older blacks, it appears that a higher productivity level is needed to generate their mean earnings level. Older workers, especially older blacks, face higher levels of monopsony power. For young workers a similar pattern to that found in table 2 for the overall sample emerges with respect to labor market

[26]The sample size for high school dropouts was too small to be included. High school contains respondents with exactly 12 years of education completed, some college with over 12 and less than 16, and university with 16 or more.

[27]Wages were trimmed 2% for those in the high school and some college groups.

frictions: blacks face more search friction while unemployed than whites, but a similar level while employed. For older workers blacks appear to face less search friction while unemployed than whites, although the estimates of κ_0 are not significantly different, but greater search friction while employed.

With respect to education differences search frictions do not uniformly decrease as education increases. Table 6 shows that, while high school graduates face the highest level of unemployed search friction, individuals with some college have the highest level of search friction while employed. The some college market also exhibits the most monopsony power. In comparing black and white high school graduates we find that blacks have a much lower value of κ_0, but a higher value of κ_1: as a result of the latter finding, a 37% black-white productivity differential across firms is needed to generate the lower black mean wage. A high value of κ_1 and a low average productivity results in black high school graduates having the lowest level of monopsony power of all the education groups.

The results regarding high school graduates presented here differ from those found in BKN's study of the school-to-work transition for black and white high school graduates. Our values of κ_0 and κ_1 are much larger compared to their values of 6.10 and 1.82 for whites and 2.95 and 1.06 for blacks, respectively. While the relative levels of κ_0 and κ_1 are higher here, our estimates of λ_1 and δ are lower than those found in BKN.[28] This is likely due to the young age of the BKN sample. Other contributing factors may include BKN's use of a different time frame, a longer panel and therefore negligible censoring rates, and wages drawn from the wage offer distribution instead of the earnings distribution. In general, however, our conclusions regarding blacks and whites are similar to BKN's with respect to λ_0 and δ. That is, λ_0 is larger for whites and δ is larger for blacks. In contrast, our conclusions regarding λ_1 differ, as BKN find λ_1 to be the same for blacks and whites, while we find that λ_1 is higher for blacks than whites. It is this higher λ_1 level which leads in our study to a higher level of κ_1 for blacks relative to that for whites. This result is primarily driven by the differences found in the NLSY and PSID regarding the job-to-job transition rates for high school whites and blacks. BKN find a much higher job-to-job transition rate for whites than blacks while the PSID reveals a higher rate for blacks. Subdividing the PSID high school graduate samples by age reveals a higher rate for whites among

[28]Converting BKN's weekly rate estimates into monthly rates their rate estimates for whites are λ_0=0.1157, λ_1=0.0346 and δ=0.0191 and for blacks are λ_0=0.0957, λ_1=0.0340 and δ=0.0325.

those under 30, an age group similar to that in the NLSY. In fact job-to-job transition rates are higher for whites for all age groups except 30-35 where the rate is uncharacteristically low for whites at .26 and high for blacks at .64. It is this group that generates the higher job-to-job transition rate for blacks in our sample. However, in most cases the transition rate differentials are not as large as those found in the NLSY and therefore we conjecture that estimates from these samples would reveal similar κ_1 estimates for blacks and whites. Unfortunately, sample size constraints limit our ability to test this hypothesis. In sum the high school results are consistent with the overall findings that employed search frictions are similar for blacks and whites and therefore do not play a role in explaining the black-white wage differential.

6. Conclusions

This paper provides estimates of the Burdett-Mortensen model using the PSID. While the PSID contains a more representative sample of the U.S. labor market, and therefore allows more general comparisons across a wider range of groups, it is not without problems. The model appears to have difficulty fitting the job duration data in conjunction with the other data components. Our response, therefore, is to estimate the model excluding the job duration data.

Estimating the model separately for blacks and whites we find that blacks tend to face more search friction while unemployed than whites. This result is consistent with past findings using the less representative NLSY. In contrast to previous work, our results suggest that blacks and whites face similar search frictions while employed. As a result, large productivity differentials are needed to generate the observed black-white earnings differentials in the data. This pattern is found for young workers as well as high school graduates. Only older workers are found to face search friction differences that play a large role in generating the black-white wage differential.

While these results are interesting and provocative, they should be interpreted with caution for two reasons. First, the relatively short time frame within which consistent and detailed job spell information is available and the difficulties in matching spells across interviews yield a high censoring rate for the job durations and hence little information on transitions following job spells. As additional years of employment history become available the problem of a high censoring rate may be mitigated to some extent; improvements in the ability to link spells over time would certainly enhance the use of this data in the future. Second, the model has difficulties fitting the wide range of job tenure in the data. As a result, estimates from a model including the job duration data are not consistent with the unemployment and job-to-job transition rates in the data. Difficulties in fitting job duration data in the Burdett-Mortensen model are not unique to this study: understanding and overcoming this shortcoming of the model merits further investigation.

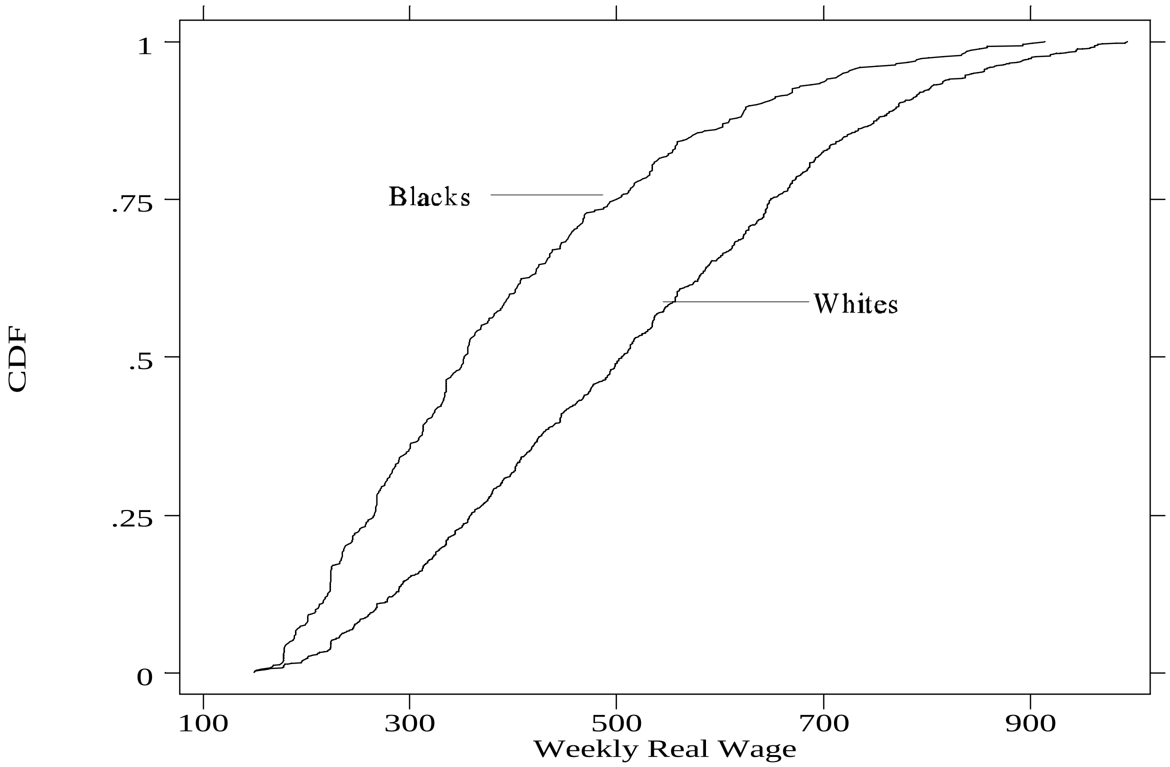

Figure 1: Comparison of empirical cumulative earnings distribution functions for blacks and whites

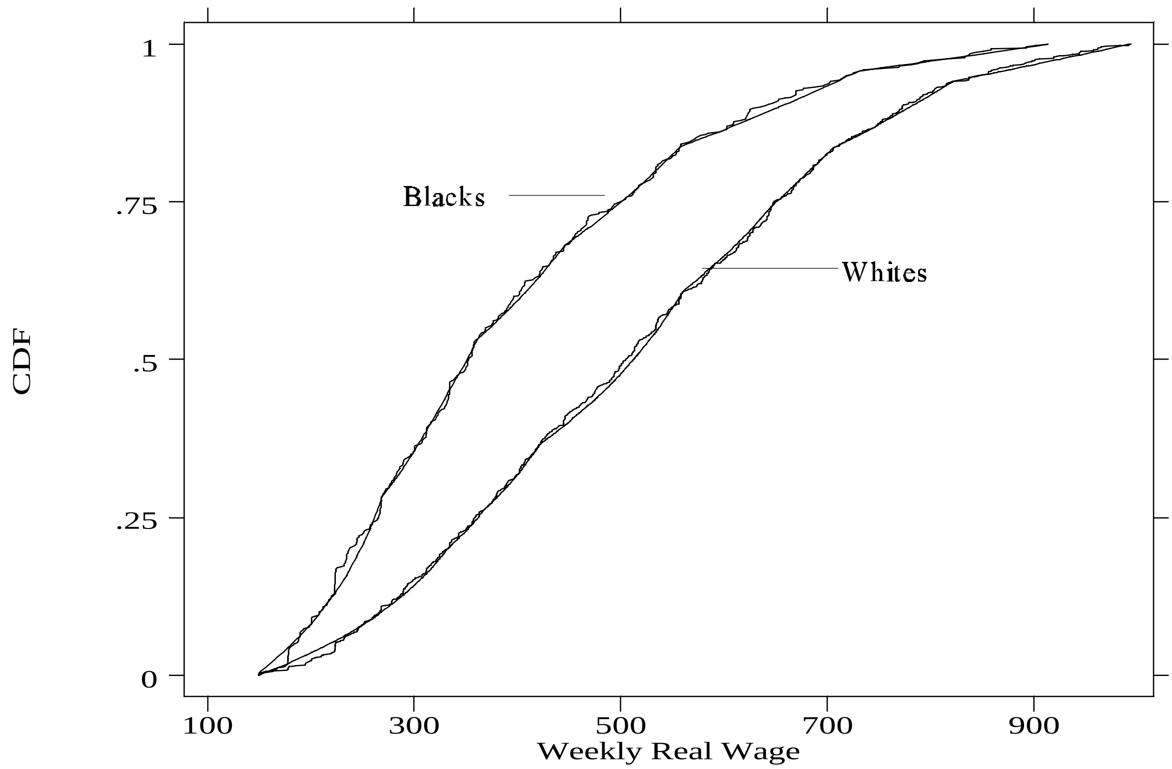

Figure 2: Comparison of empirical and estimated cumulative earnings distribution functions

Table 1
Means for White and Black Paid, Full-time Workers in the PSID

	Whites	Blacks
Fraction nonemployed at 1988 interview	0.0365 (.0049)	0.0728 (.0105)
Nonemployment duration (in months) - sampled from stock including censored spells	8.96 (1.12)	16.47 (2.05)
Nonemployment duration - sampled from flow including censored spells	6.42 (0.64)	7.06 (0.88)
Fraction of nonemployment spell censored	0.2845 (0.0297)	0.2787 (0.0408)
Job spell duration (in months) - sampled from stock including censored spells	110.09 (2.77)	100.02 (4.02)
Job spell duration - sampled from flow including censored spells	20.00 (3.13)	12.33 (2.52)
Fraction of job spells censored	0.7473 (0.0130)	0.7491 (0.0189)
Fraction of completed job spells that end in a job-to-job transition	0.4268 (.0316)	0.3898 (.0451)
Weekly real wages sampled from earnings distribution	513.84 (5.76)	392.84 (7.43)
Weekly real wages sampled from offer distribution	406.81 (33.78)	233.83 (14.70)
Sample size	1480	618

Note: Standard errors are in parentheses.

Table 2
Search Parameter Estimates and Model Predictions for White and Black Paid, Full-time Workers in the PSID

	Full Log Likelihood		Excluding Job Spells	
Parameter	Whites	Blacks	Whites	Blacks
λ_0	0.1249 (0.0080)	0.0846 (0.0074)	0.1256 (0.0089)	0.0949 (0.0088)
λ_1	0.0123 (0.0012)	0.0256 (0.0038)	0.0121 (0.0026)	0.0151 (0.0038)
δ	0.0046 (0.0002)	0.0047 (0.0003)	0.0048 (0.0007)	0.0075 (0.0013)
κ_0	26.9234 (1.9483)	18.1599 (1.7987)	26.3807 (3.6546)	12.7098 (1.9677)
κ_1	2.6516 (0.3031)	5.5026 (0.9657)	2.5358 (0.3639)	2.0184 (0.3592)
Q	9	8	9	8
r	149.05	146.13	149.05	146.13
P (average among workers)	913.78	726.96	920.70	844.50
P (average among firms)	658.06	385.51	669.45	572.59
Model Predictions				
Unemployment rate	0.0358	0.0522	0.0365	0.0729
Mean unemployment duration	8.00	11.82	7.96	10.53
Mean job duration	137.26	123.84	134.73	89.10
Mean employment duration	215.51	214.66	210.05	133.86
Fraction of completed job spells ending in a job-to-job transition	0.4393	0.5480	0.4321	0.3947
Mean of earnings distribution	515.90	400.19	515.06	394.40
Mean of wage offer distribution	382.26	259.19	385.05	303.60
Monopsony Index (*MI*)	0.4112	0.2764	0.4167	0.4110

Note: Asymptotic standard errors are in parentheses.

Table 3
Means for White and Black Paid, Full-time Workers by Age

	Whites Under 35	Whites 35 & Over	Blacks Under 35	Blacks 35 & Over
Fraction nonemployed at 1988 interview	0.0339 (.0065)	0.0393 (.0073)	0.1092 (.0165)	0.0230 (.0093)
Nonemployment duration (in months) - sampled from stock including censored spells	8.00 (1.66)	9.86 (1.54)	17.23 (2.31)	11.50 (2.86)
Nonemployment duration - sampled from flow including censored spells	4.04 (0.66)	8.45 (0.99)	7.42 (1.09)	6.56 (1.46)
Fraction of nonemployment spell censored	0.2315 (0.0408)	0.3306 (0.0424)	0.2738 (0.0489)	0.2895 (0.0746)
Job spell duration (in months) - sampled from stock including censored spells	77.42 (2.27)	152.42 (5.02)	61.27 (3.18)	151.32 (7.01)
Job spell duration - sampled from flow including censored spells	21.40 (5.35)	19.07 (3.94)	11.75 (2.77)	17.00 (5.00)
Fraction of job spells censored	0.7381 (0.0175)	0.7592 (0.0193)	0.6981 (0.0262)	0.8190 (0.0260)
Fraction of completed job spells that end in a job-to-job transition	0.4755 (.0419)	0.3592 (.0475)	0.4643 (.0547)	0.2059 (.0704)
Weekly real wages sampled from earnings distribution	484.65 (7.43)	552.11 (8.79)	345.58 (8.08)	454.42 (12.41)
Weekly real wages sampled from offer distribution	403.57 (41.37)	408.97 (50.26)	231.04 (16.45)	256.10 (6.04)
Sample size	767	713	357	261

Note: Standard errors in parentheses.

Table 4
Means for White and Black Paid, Full-time Workers by Education

	Whites High School	Whites Some College	Whites University	Blacks High School
Fraction nonemployed at 1988 interview	0.0403 (.0081)	0.0381 (.0104)	0.0275 (.0086)	0.0616 (.0141)
Nonemployment duration (in months) - sampled from stock including censored spells	10.04 (1.48)	4.85 (0.88)	11.60 (3.73)	17.28 (2.70)
Nonemployment duration - sampled from flow including censored spells	7.69 (1.14)	5.69 (1.35)	4.19 (0.90)	8.90 (1.60)
Fraction of nonemployment spell censored	0.3488 (0.0517)	0.2308 (0.0590)	0.1774 (0.0489)	0.3333 (0.0688)
Job spell duration (in months) - sampled from stock including censored spells	116.29 (4.45)	107.71 (5.53)	99.18 (5.11)	94.38 (5.64)
Job spell duration - sampled from flow including censored spells	12.85 (3.18)	31.33 (9.40)	31.33 (6.97)	17.83 (4.56)
Fraction of job spells censored	0.7705 (0.0193)	0.7702 (0.0268)	0.6829 (0.0297)	0.7773 (0.0265)
Fraction of completed job spells that end in a job-to-job transition	0.4270 (.0527)	0.3800 (.0693)	0.4800 (.0581)	0.4902 (.0707)
Weekly real wages sampled from earnings distribution	468.61 (7.20)	504.31 (10.83)	675.32 (12.79)	376.49 (9.30)
Weekly real wages sampled from offer distribution	374.58 (42.00)	284.77 (55.65)	553.98 (74.47)	223.94 (16.50)
Sample size	596	341	363	292

Note: Standard errors in parentheses.

Table 5
Search Parameter Estimates and Model Predictions for White and Black Paid, Full-time Workers by Age

Parameter	Whites Under 35	Whites 35 & Over	Blacks Under 35	Blacks 35 & Over
λ_0	0.1855 (0.0188)	0.0965 (0.0095)	0.0899 (0.0097)	0.1127 (0.0203)
λ_1	0.0201 (0.0061)	0.0064 (0.0019)	0.0373 (0.0122)	0.0016 (0.0009)
δ	0.0065 (0.0015)	0.0039 (0.0009)	0.0111 (0.0022)	0.0027 (0.0012)
κ_0	28.4575 (5.6839)	24.4328 (4.7126)	8.1283 (1.3788)	42.3121 (17.4604)
κ_1	3.0902 (0.6265)	1.6333 (0.3287)	3.3666 (0.8764)	0.6141 (0.2483)
Q	9	5	7	4
r	149.05	168.42	146.13	168.16
P (average among workers)	848.13	1040.72	695.07	1323.24
P (average among firms)	579.70	847.16	414.86	1152.72
Model Predictions				
Unemployment rate	0.0339	0.0393	0.1095	0.0231
Mean unemployment duration	5.39	10.37	11.12	8.87
Mean job duration	95.47	174.76	55.54	304.06
Mean employment duration	153.43	253.32	90.37	375.49
Fraction of completed job spells ending in a job-to-job transition	0.4636	0.3594	0.4769	0.2054
Mean of earnings distribution	486.01	554.16	351.08	456.16
Mean of wage offer distribution	349.64	448.92	257.04	406.93
Monopsony Index (*MI*)	0.3846	0.4653	0.3101	0.6226

Note: Asymptotic standard errors are in parentheses.

Table 6
Search Parameter Estimates and Model Predictions for White and Black Paid, Full-time Workers by Education

Parameter	Whites High School	Whites Some College	Whites University	Blacks High School
λ_0	0.1031 (0.0121)	0.1721 (0.0249)	0.1745 (0.0230)	0.0783 (0.0119)
λ_1	0.0094 (0.0029)	0.0120 (0.0048)	0.0159 (0.0074)	0.0210 (0.0099)
δ	0.0043 (0.0010)	0.0068 (0.0022)	0.0050 (0.0017)	0.0052 (0.0015)
κ_0	23.7921 (4.9592)	25.1511 (7.1113)	35.1937 (11.2809)	15.1680 (3.6915)
κ_1	2.1724 (0.4342)	1.7584 (0.4136)	3.1989 (0.9872)	4.0619 (1.5164)
Q	6	6	7	7
r	168.42	153.37	203.75	163.58
P (average among workers)	774.03	875.49	978.70	624.48
P (average among firms)	614.17	733.43	791.24	389.70
Model Predictions				
Unemployment rate	0.0403	0.0382	0.0276	0.0619
Mean unemployment duration	9.70	5.81	5.73	12.77
Mean job duration	151.75	99.57	124.84	116.02
Mean employment duration	230.76	146.15	201.65	193.77
Fraction of completed job spells ending in a job-to-job transition	0.4069	0.3717	0.4690	0.5052
Mean of earnings distribution	468.50	504.99	673.36	378.44
Mean of wage offer distribution	369.88	408.05	509.66	267.80
Monopsony Index (*MI*)	0.3982	0.4475	0.3705	0.2747

Note: Asymptotic standard errors are in parentheses.

References

Bontemps, Christian, Jean-Marc Robin, and Gerard J. van den Berg (1997) “Equilibrium search with productivity dispersion,” Tinbergen Institute discussion paper TI 97-081/3.

Bowlus, Audra J. (1998) “U.S.-Canadian unemployment rate and wage differences among young, low-skilled males in the 1980s.’ *Canadian Journal of Economics* 31(2), 437-64.

________ (1997) 'A search interpretation of male-female wage differentials.' *Journal of Labor Economics* 15(4), 625-657.

Bowlus, Audra J., Nicholas M. Kiefer, and George R. Neumann (1995) 'Estimation of equilibrium wage distributions with heterogeneity.' *Journal of Applied Econometrics* 10, S119-32.

________ (1998) 'Equilibrium search models and the transition from school to work.' University of Western Ontario, mimeo.

Burdett, Kenneth and Dale Mortensen (1998) “Wage differentials, employer size, and unemployment,” *International Economic Review* 39, 257-73.

Kiefer, Nicholas M. and George R. Neumann (1993) 'Wage dispersion with homogeneity: the empirical equilibrium search model.' In *Panel Data and Labour Market Dynamics*, ed. H. Bunzel, P. Jensen and N. Westergård-Nielsen (Amsterdam: North-Holland), 57-74.

Lancaster, Tony (1990) *The Econometric Analysis of Transition Data* (Cambridge: Cambridge University Press).

Mortensen, Dale (1990) 'Equilibrium wage distributions: a synthesis.' In *Panel Data and Labour Market Studies*, ed. J. Hartog, G. Ridder and J. Theeuwes (North-Holland), 279-296.

Mortensen, Dale and George R. Neumann (1988) 'Estimating structural models of unemployment and job duration in dynamic econometric modelling.' *Proceeding of the Third International Symposium in Economic Theory and Econometrics* (Cambridge University Press).

Van den Berg, G. and G. Ridder (1998) 'An empirical equilibrium search model of the labor market.' Free University of Amsterdam, *Econometrica*, forthcoming.