Unemployment Insurance Eligibility and the School-to-Work Transition in Canada and the United States

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To study how the design of unemployment insurance (UI) affects people leaving school to find jobs, a model of job search in the presence of UI is developed and estimated for the United States and Canada. The level of UI benefits depends on previous earnings, which creates opposing incentives for unemployed people not receiving benefits. Which of these opposing incentives dominates the other is found to differ across demographic groups within each country. Changes in UI policy therefore can have very different effects on different individuals. The major differences found in the transition from school to work in Canada and the United States are a lower rate of job-offer arrivals and a lower rate of offer rejections in Canada. Within each country, offer-arrival rates differ across individuals much more than offer-rejection rates.

KEY WORDS: Comparative labor markets; Job search; Structural estimation.

This article studies how the design of unemployment insurance (UI) affects young people leaving school. A dynamic programming model of job search in the presence of UI is developed and estimated using comparable longitudinal datasets for Canada and the United States. To calculate the effect of UI on the currently uninsured, the theoretical job-search environment incorporates most of the rules that govern the Canadian UI system. The value of UI benefits associated with each possible job offer is taken into account when solving for reservation wages on leaving school. Estimates of the model are used to study how changing parameters of the UI system would affect the duration of nonworking spells after leaving school.

When applied to the data in either country, the maximum likelihood estimation procedure controls for observed and unobserved heterogeneity across people and jobs, measurement error in wages, and job search before leaving school. When applied to U.S. data, the model is estimated without a UI system because the incentives generated by the U.S. system are weaker and much more difficult to model than in Canada. During the 1980s, Canada spent four times as much as the United States on unemployment insurance per dollar of gross national product (Card and Riddell 1991), and young people in Canada are three times as likely to report receiving UI in the first two years after leaving school. Furthermore, UI is administered uniformly throughout Canada, whereas the rules differ considerably across U.S. states.

All of the major features of the Canadian UI system are integrated into the search environment—minimum and maximum insurable earnings, the earnings replacement ratio, the two-week waiting period, regional extended benefits, and the formula that links the length of the benefit period to the length of the previous working spell. In most Organization for Economic Cooperation and Development countries, including the United States and Canada, people leaving school to find a job are usually not eligible to receive UI compensation because one must lose a job covered by UI to qualify for benefits. Basic search theory predicts that more generous UI benefits lead the uninsured to lower their reservation wages in an effort to "get into the system." Mortensen (1977) identified this effect, which runs counter to the better-known result that higher benefits raise reservation wages for people currently receiving UI benefits.

Atkinson and Micklewright (1991) listed seven further features of most UI systems that are typically ignored in the theoretical job-search literature, including Mortensen's pioneering model. At least one of these features alters the predictions of search theory concerning uninsured people searching for jobs. In particular, when the level of benefits depends on previous earnings, the effect of UI on the uninsured becomes ambiguous. A person who accepts and then loses a better-paying job qualifies for higher UI benefits. Higher benefits encourage the uninsured to get into the system, which tends to lower reservation wages, but higher benefits also encourage the uninsured to search for higher-paying jobs, which tends to raise reservation wages.

Evidence is found that which of these opposing incentives dominates differs across demographic groups within each country. From the policy analysis carried out in Section 3, three general results arise. First, the estimated response of uninsured job searchers to some aspects of UI policy is the opposite of the predictions from the basic search model that ignores details of the UI system. Second, the differences in labor-market conditions within and between Canada and the United States lead to opposing responses to the same hypothetical change in UI policy. And third, some aspects of UI policy may have surprisingly little effect on the transition from school to work, whereas other aspects of UI may have very large effects.

Although no general statements can be made about the estimated responses to UI policies, three general patterns do arise from the estimates of the model. First, the la-

Journal of Business & Economic Statistics
April 1997, Vol. 15, No. 2

bor markets for people leaving school in the United States and Canada differ primarily in a lower rate of job-offer arrivals in Canada. Second, within countries, offer-arrival rates are lower and layoff (working-spell termination) rates are higher among people leaving high school. Third, the rate of rejecting job offers, which is determined within the job-search model, is estimated to be lower in Canada under the status quo, but within countries, job-rejection rates in different demographic groups are similar.

There is substantial evidence concerning UI's impact on employed and eligible unemployed workers (Atkinson and Micklewright 1991; Devine and Kiefer 1991). Van den Berg (1990) and Engberg (1992) estimated structural search models that take into account that the level of benefits and the length of potential benefits differ among unemployed people. Their models concern current UI benefits only and do not account for the future benefits associated with job offers. Reduced-form estimates of the incentive effects of UI were provided by Meyer (1990) for the United States and Ham and Rea (1987) and Baker and Rea (1993) for Canada (among many others).

There is much less evidence concerning how UI affects the uninsured, probably because the impact of UI on the uninsured is inherently indirect. UI affects the job search of uninsured workers through the value of losing jobs not yet held and through the decisions made by firms and other workers. Levine (1993) used variation in benefit ratios across states to provide evidence that the uninsured displace the insured when competing for jobs. He bases the analysis on a static labor-supply model, which makes the same unambiguous predictions as the basic search model.

The dynamic programming model estimated by Wolpin (1992) includes UI but assumes that new labor-market entrants become eligible for fixed UI benefits only after taking a job. Evidence is found that raising the level of UI benefits would substantially lower reservation wages for black males leaving high school in the United States. When considering the adoption of a Canadian-styled UI system, we find the opposite effect: Reservation wages and expected unemployment duration among nonwhites leaving high school would rise. The difference is due to tying UI benefit levels to past wages, which encourages the unemployed to reject low wage offers they otherwise would have accepted.

1. SEARCH FOR INSURED JOBS WHILE UNINSURED

The model of job search developed here draws on elements of several previously estimated models and adds one ingredient—the value of UI benefits associated with a job offer. In the model, previous earnings and the length of previous employment durations determine the level of future UI benefits. Workers live forever and search for jobs that are summarized by a weekly wage. In each period t, they act to maximize

$$\sum_{s=0}^{\infty} \beta^s E_t \ln w_{t+s},$$

where w_t is consumption in week t, β is a weekly discount factor, and E_t denotes expectations given information available at week t. Consumption equals the wage when working. Search models typically have been estimated assuming expected present value of wealth maximization, which is consistent with risk neutrality, perfect capital markets, or both. The insurance aspect of unemployment insurance is ruled out by wealth maximization. Identifying consumption with income, as done here, implicitly assumes that it is not possible to borrow and save. A preferred strategy would jointly model dynamic decisions concerning consumption, earnings, and job search. Because constructing continuous histories for consumption, earnings, and employment is very difficult, most empirical research has focused on only one or, as done here, two of these aspects of labor-market dynamics (earnings and unemployment).

When not working, net utility equals $\ln(e^c + b_t)$, where c is the estimated pecuniary and nonpecuniary value of not working and b_t is the amount of unemployment insurance benefits received in week t. In the estimated model, parameters such as c depend on observed and unobserved characteristics of the individual, but this is suppressed while describing the model. While searching, the person receives a job offer in each period with probability λ^o.

A working spell ends with probability λ_k^l each period after the first period, where λ_j^l takes on one of J values $(j = 1, \ldots, J)$ with corresponding probabilities π_j^l that are independent of the wage draw. At no point do workers know the type j of their working spell. The assumptions of exogenous and unknown layoff probabilities are restrictive, but they relax the assumption of permanent job offers made in most previously estimated job-search models. When a working spell ends, the person begins searching again with UI benefits if the previous wage and length of spell were such that the person qualifies for benefits.

It is convenient to express the wage-offer distribution in log or utility terms. Let $u = \ln w$ denote the utility associated with a wage of w. Log wage offers are assumed to be drawn from an exponential distribution, shifted by a lowest offer $\bar{u}$:

$$f(u) = \begin{cases} \gamma e^{-\gamma(u-\bar{u})} & \text{if } u \geq \bar{u} \\ 0 & \text{if } u < \bar{u} \end{cases}$$

$$F(u) = \begin{cases} 1 - e^{-\gamma(u-\bar{u})} & \text{if } u \geq \bar{u} \\ 0 & \text{if } u < \bar{u}, \end{cases}$$

where $f(u)$ and $F(u)$ are the density and distribution of offered utility. This assumption is equivalent to assuming that wages in levels follow the Pareto distribution. The parameter $\gamma > 0$ shifts the mean and variance of the offer distribution. The distribution $F(u)$ has the convenient property that

$$E(u|u > u^*) = u^* + 1/\gamma \tag{1}$$

for $u^* \geq \bar{u}$. The lower bound on wage offers mitigates a problem in interpreting offer probabilities. Without the shift, the mode offer is 0 and most offers are in essence non-offers. Given the truncation point $\bar{u}$, the estimated offer probability, λ^o, can be interpreted as the probability of receiving a substantial job offer.

Job search takes place in four distinct phases. In chronological order, they are (a) job search while in school, (b) job search after school but before taking a job, (c) job search after working and while receiving UI benefits, and (d) job search after UI benefits have been exhausted. Figure 1 illustrates the four phases of job search. Through backward induction, optimal search behavior can be characterized by different but related value functions and reservation wages in each phase. The model generates three endogenous variables observed in the data—the length of unemployment after leaving school denoted t_1, the length of the first working period denoted by t_2, and the log wage on the first job in the working period denoted u.

1.1 Search After UI Benefits Are Exhausted

In phase (d) of Figure 1, people expect no more unemployment benefits and enter a stationary environment of work, layoff, and search. To model the continuing buildup and exhaustion of UI benefits over time creates a dynamic programming problem of considerable analytical and computational complexity. Modeling UI benefits after only the first working period, however, is tractable, and from this initial specification one could build in benefits expected in later periods.

A worker taking a job with utility u receives u for one week for certain and thereafter stops working with probability λ_j^l in each subsequent week. The stream of utility while working has expected present value $u/(1-\beta(1-\lambda_j^l))$. Once laid off, the person returns to the state of unemployed search, which by definition has value EV^*. The expected utility of taking a job with log wage u is therefore

$$V(u) = \sum_{j=1}^{J} \pi_j^l \frac{u + \beta\lambda_j^l EV^*}{1-\beta(1-\lambda_j^l)}.$$

Defining

$$\Lambda_1 = \sum_{j=1}^{J} \frac{\pi_j^l}{1-\beta(1-\lambda_j^l)}$$

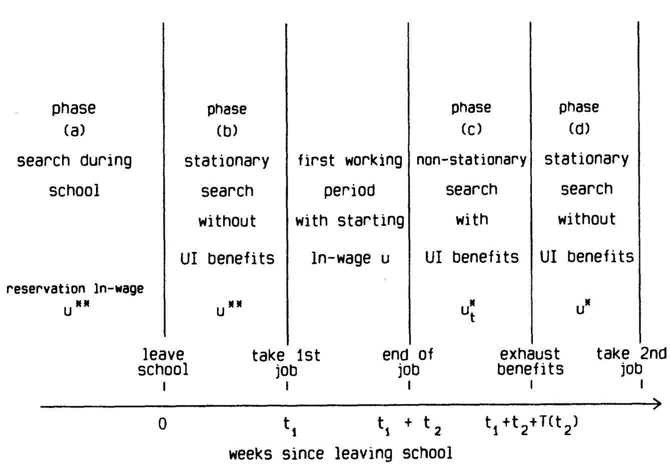

Figure 1. The Timing of Job Search in the Model.

$$\Lambda_2 = \sum_{j=1}^{J} \frac{\pi_j^l \lambda_j^l}{1-\beta(1-\lambda_j^l)},$$

then we can rewrite $V(u)$ as

$$V(u) = \Lambda_1 u + \beta\Lambda_2 EV^*. \tag{2}$$

The reservation utility in phase (d), u^*, implicitly equates the value of taking a job with that wage and the value of searching one more period: $V(u^*) = c + \beta EV^*$. Using the optimal decision rule to reject wages less than u^*, EV^* can be written

$$EV^* = (1-\lambda^o)V(u^*) + \lambda^o \left[F(u^*)V(u^*) + \int_{u^*}^{\infty} V(u)f(u)\,du\right]. \tag{3}$$

In other words, if an offer does not arrive or one arrives that is rejected, the person receives the value of continued search $V(u^*)$. Otherwise, the person accepts an offer and receives $V(u)$, which must be integrated over the distribution of acceptable offers. Using (1) and (2), the integral in (3) reduces to $(\Lambda_1 u + \Lambda_2 EV^*)/\gamma$. The reservation utility implicitly solves

$$u^* = \frac{c}{\Lambda_1} + \frac{\beta(1-\Lambda_2)}{1-\beta\Lambda_2}\left[u^* + \lambda^o \frac{1-F(u^*)}{\gamma}\right]. \tag{4}$$

Christensen and Kiefer (1991) presented the algebra for the case $J = 1$ and $\lambda_j^l = 0$ (i.e., permanent homogeneous jobs). The solution for u^* is found by iterating on (4) from an arbitrary guess of u^*, a process that is guaranteed to converge to the unique reservation wage when $0 \le \beta < 1$. For most offer distributions other than the exponential, iterating to find u^* requires numerical integration, which when combined with the UI system greatly increases the burden of estimating the model. With u^* calculated, EV^* can be computed by substituting (2) into (3):

$$EV^* = \frac{\Lambda_1}{1-\beta\Lambda_2}\left[u^* + \lambda^o \frac{1-F(u^*)}{\gamma}\right]. \tag{5}$$

If $u^* < \bar{u}$, then $\bar{u}$ replaces u^* on the right sides of (4) and (5).

1.2 Unemployment Insurance Rules and Insured Job Search

Green and Riddell (1993) provided a complete description of UI rules in Canada, but the main features are described by three equations:

$$b_t = \begin{cases} b(u) & \text{if } t > t_w \text{ and } t \le T(t_2) \\ 0 & \text{if } t \le t_w \text{ or } t > T(t_2) \end{cases}$$

$$b(u) = \begin{cases} 0 & \text{if } e^u < w_{\min} \\ \tau e^u & \text{if } w_{\min} \le e^u < w_{\max} \\ \tau w_{\max} & \text{if } w_{\max} > e^u \end{cases}$$

$$T(t_2) = \begin{cases} 0 & \text{if } t_2 < t_E \\ \min\{t_w + t_2 + t_R, 52\} & \text{if } t_e \le t_2 \le 25 \\ \min\{t_w + 25 + (t_2 - 26)/2 \\ \quad + t_R, 52\} & \text{if } 25 < t_2. \end{cases}$$

Here b_t is the level of UI benefits in week t of an unemployment spell, $b(u)$ is the level of potential weekly benefits as a function of the previous log wage u, and $T(t_2)$ is the potential number of weeks of benefits depending on the length of the previous employment spell t_2. A worker who becomes unemployed and is eligible for UI must wait t_w weeks before benefits begin. If the worker lost his job through a layoff, then $t_w = 2$. At the time of the sample, a Canadian worker who quit or was fired faced $t_w = 5$, but for simplicity people are assumed to expect all terminations to qualify for the shorter two-week waiting period. The benefit level $b(u)$ depends on the previous wage through the replacement rate τ. In 1987, the replacement ratio τ was .60, the weekly maximum insurable wage $w_{\max}$ was \$530, and the minimum insurable amount $w_{\min} = .2 \times w_{\max}$ was \$106. These values are used when estimating the model in the presence of UI.

To be eligible for UI benefits at all ($T(t_2) > 0$), a worker must have worked at least t_E weeks on insurable jobs during the 52 weeks prior to becoming unemployed (the qualifying period). In Canada, the entrance requirement t_E depends primarily on the regional unemployment rate and the individual's previous employment and UI receipt. For those who received UI in the previous year (repeaters) or who did not work in the previous two years (new entrants), $t_E = 20$. It is assumed that everyone leaving school is a new entrant. For $t_2 \ge t_E$, a person earns an extra week of benefits for every additional week worked up to 25 weeks and thereafter an additional week of benefits for every two weeks worked. $T(t_2)$ also depends implicitly on the regional unemployment rate through regional extended benefits, denoted t_R. The legislated UI regions within provinces are not identified in the Canadian data. To approximate t_R, people in the Atlantic provinces of Canada are assumed to be eligible for the full 32 weeks of extended benefits, but people in other areas are assumed to receive no extended benefits.

Now consider a worker in phase (c) of Figure 1 who is currently receiving UI benefits $b(u)$. Let the reservation utility and value of entering week t of the benefit period be denoted $u_t^*(b,T)$ and $EV_t(b,T)$. (When convenient, T and b are written without their arguments.) If the worker survives the whole benefit period without finding a job, then $u_{T+1}^*(b,T) = u^*$ in (1.1). Then

$$u_t^*(b,T) = \frac{1}{\Lambda_1}\left[\ln(b_t + e^c) + \beta EV_{t+1}(b,T)\right] - \beta\Lambda_2 EV^* \quad (6)$$

and

$$EV_t(b,T) = \Lambda_1\left[u_t^* + \lambda^o \frac{1 - F(u_t^*)}{\gamma}\right] + \beta\Lambda_2 EV^* \quad \text{for} \quad t = 1, 2, \ldots, T, \quad (7)$$

where $EV_{T+1}(b,T) = EV^*$. The last term in (6) incorporates the fact that a job taken during phase (c) will end at some point and throw the worker back into phase (d) search. Equations (6) and (7) reduce to (4) and (5) when $EV_t(b,T) = EV_{t+1}(b,T) = EV^*$ and $b_t = 0$ for all t.

1.3 Search After Leaving School

Given the model of search while receiving UI, we can now consider the problem of an uninsured worker in phase (b) searching for insured and uninsured jobs. The value of a job with utility u has a new component. As in the stationary phase, one component is the expected present utility of holding the job—$u/(1 - \beta(1 - \lambda_j^l))$. The other is again the expected present value of what occurs after the job ends, denoted $U(u)/(1 - \beta(1 - \lambda_j^l))$. $U(u)$ is the value of a job offer at the point in the future when the job ends. Without UI, $U(u)$ is simply $\beta\lambda^l EV^*$ and does not depend on u. By definition, $EV_1(b,T)$ incorporates the value of potential benefits during a spell of unemployment with a maximum of T weeks of benefits. $U(u)$ can therefore be expressed in terms of EV_1:

$$\begin{aligned} U(u) &= \beta \sum_{j=1}^{J} \pi_j^l \lambda_j^l \left[\sum_{k=1}^{\infty} (\beta(1 - \lambda_j^l))^k EV_1(b(u), T(k))\right] \\ &= \beta \sum_{j=1}^{J} \pi_j^l \lambda_j^l \left[\sum_{k=1}^{51} (\beta(1 - \lambda_j^l))^k EV_1(b(u), T(k)) \right. \\ &\qquad \left. + \frac{(\beta(1 - \lambda_j^l))^{52}}{1 - \beta(1 - \lambda_j^l)} EV_1(b(u), T(52))\right]. \end{aligned} \quad (8)$$

The summation over weeks of employment accounts for the fact that benefits begin only after the job ends. When this is expected to occur depends on the layoff probability λ_j^l. The UI system makes the value of becoming unemployed a nonlinear function of the previous spell length. $U(u)$ depends on all characteristics of a worker that affect future job-search behavior, even though the UI system does not distinguish among most worker characteristics. Because $T(k)$ becomes constant at $k = 52$, the infinite series collapses to the finite series (8).

After leaving school, people search for jobs with value $V^{**}(u) = \Lambda_1 u + \Lambda_2 U(u)$. Search is stationary because there are no UI benefits to exhaust and because eligibility in the next period of unemployment does not depend on how long it took to find a job. Let EV^{**} denote the value of entering any period after school but before holding a job. Because $V^{**}(u)$ is increasing in u, optimal decisions are still characterized by a reservation utility denoted u^{**}, which solves $V^{**}(u^{**}) = c + \beta EV^{**}$. This leads to a system of equations describing optimal search decisions:

$$u^{**} = \frac{1}{\Lambda_1}\left[c + \beta EV^{**} - U(u^{**})\right] \quad (9)$$

and

$$EV^{**} = \Lambda_1 u^{**} + [1 - \lambda^o(1 - F(u^{**}))]U(u^{**}) + \lambda^o \left[\Lambda_1 \frac{1 - F(u^{**})}{\gamma} + \int_{u^{**}}^{\infty} U(u) f(u)\, du\right]. \quad (10)$$

Iterating on (9) and (10) converges to u^{**}, although convergence is much slower than iterating on (4) to compute u^*. Unlike the reservation wage without UI, it is necessary to iterate on EV^{**} while iterating u^{**} because (9) is not written only in terms of u^{**}. If $U(u) = U(u^{**})$ for all $u \geq u^{**}$, then the integral in (10) and the term containing $U(u^{**})$ reduce to $U(u^*)$ and (10) collapses to (5). Computing EV^{**} requires the evaluation of an integral over u that contains $EV_1(b, T)$, which itself must be solved by backward induction using (7) for each value of u. Because of the cap on insurable earnings, the value of $U(u)$ is constant for $u > \ln w_{\max}$.

1.4 Search in School

Many people hold a job while going to school, and about 90% of these people are employed the first week after they leave school. The fact that search for a post-schooling job may not be sequential and may begin at different times makes it difficult to model phase (a), job search in school. Wolpin (1987) modeled search in school as sequential by extending search from the point of leaving school back for some fixed number of periods and estimating a probability of receiving an offer each week. With an estimated low probability of rejecting offers, he could match perfectly the large fraction of people having a job on leaving school. Applying the same approach here fails because estimated rejection rates are found to be high. Computing a reservation wage in each of many weeks of search during school also requires many additional evaluations of $U(u)$.

Instead, for people who hold a job during the first week of the last month they are in school, it is assumed that they receive up to two offers before the end of the leave month and choose among the offers simultaneously. The number of possible offers was set to two because of the complexity of the observed offer distribution with more than two offers (see Appendix A). The decision is made in the last week of the leave month, so the reservation wage equals u^{**}. Define $W = 0$ for people who did not work in school and $W = 1$ for those who did. For $W = 0$, the probability of accepting an offer is the probability that the maximum offer is above the reservation wage or

$$P_S(0) = 1 - F^2(u^{**}). \quad (11)$$

People who held a job in school ($W = 1$) are assumed to have already accepted a job above the reservation wage u^{**}, and the distribution of accepted wages is assumed to be the same as if they also received two offers simultaneously. No attempt is made to explain the small fraction that become unemployed between the beginning and end of the month they leave school. $P_S(1)$ was set equal to 0 for those who were not working at the end of the leave month and was set equal to 1 for those who were. In other words, their subsequent behavior is conditional on their status at the end of the leave month just as they start sequential search.

1.5 Theoretical Response to UI Policy Parameters

To predict the response of reservation wages to changes in the parameters of the UI system, the total derivative of (9) and (10) would have to be calculated. The point can be made instead using the stylized model of UI typically studied in the theoretical literature. Namely, if $T(t_2)$ and $b(u)$ are constants rather than functions of previous outcomes, then $U(u)$ is constant. If there is only one job type ($J = 1$), then the total derivative of the uninsured reservation wage is

$$\frac{du^{**}}{db} = \frac{(\beta - 1)}{1 - \beta + \lambda^o f(u^{**})/\gamma} \frac{dU}{db} \leq 0. \quad (12)$$

The response to increasing T is also negative. Behind (12) is the following argument: A job offer in hand with log wage u^{**} rises in value when b is increased. But all insured jobs increase in value by the same amount, which affects the worker's decision to accept the job in hand through βEV^{**}. To consume the value of other jobs requires that the current offer be rejected and search continued next period, which is discounted by β. To restore the equality of the reservation wage with the value of further search, the reservation wage falls. Therefore, in the stylized model of UI, uninsured people are unambiguously encouraged to take jobs they would not have taken at a lower benefit level.

When T and b depend on previous outcomes, however, their effect on uninsured job-search behavior is ambiguous. Consider an increase in $b(u)$ caused by an increase in the replacement ratio τ. From (10), the term

$$\int_{u^{**}}^{\infty} \frac{dU(u)}{d\tau} f(u)\, du \quad (13)$$

enters the expression for $du^{**}/d\tau$. For $u < \ln w_{\max}$, $\partial^2 U(u)/\partial\tau\partial u > 0$ because increasing τ increases benefits more for higher-wage jobs. The term in (13) may now outweigh the negative effect on u^{**} from the increased value of accepting the reservation job itself. Increasing the generosity of UI may lower or raise the reservation wages of unqualified job seekers. The direction depends on all of the parameters of the search environment. For instance, if $u^{**} > \ln w_{\max}$ before the change, then there is no distortion across acceptable jobs.

2. EMPIRICAL ANALYSIS

2.1 Data

The U.S. data are drawn from the National Longitudinal Survey of Youth (NLSY). The Canadian data include 17–24-year-olds in the first longitudinal file (1986–1987) of the Labour Market Activity Survey (LMAS). The two-year segment 1982–1983 of the longer running NLSY was drawn so that the length of the sample periods would be the same. These years were chosen because in 1982–1983 the age distribution in the NLSY cohort matches the 17–

Table 1. Sample Means of Variables by Country and Sex

Variable	Canada Men	Canada Women	United States Men	United States Women
Demographics				
Observations	364	362	225	214
Proportion of sample (total = 1,163)	.31	.31	.19	.18
Atlantic provinces	.25	.24	—	—
Nonwhite	—	—	.25	.18
School and work before leaving school				
Leave month (3 = March, 4 = April, 5 = May, 6 = June)	5.15	5.12	5.11	5.18
Number of months in school	5.20	5.14	5.03	5.09
Leaving post-secondary school	.42	.55	.36	.38
Working in first week of leave month	.21	.18	.40	.31
Work and unemployment after leaving school				
Positive nonworking weeks after leave month	.52	.59	.52	.61
Completed nonworking period	.96	.94	.94	.90
Weeks not working (first spell)	11.65	14.08	13.55	18.24
Completed working period	.61	.57	.61	.57
Weeks working (first spell)	43.78	43.37	41.39	35.01
ln weekly wage on first job (local $)	5.53	5.32	5.11	4.95
Received UIC in first year (86/82)	.13	.11	.04	.03
Received UIC in second year (87/83)	.27	.22	.09	.06

NOTE: See Appendix B for definitions of variables and sample criteria.

24-year-old brackets defined in the LMAS. In the United States during 1982, the economywide unemployment rate was 9.7%, whereas in Canada during 1986, the unemployment rate was 9.5%. In the second year the unemployment rate fell in both countries, although more so in Canada (to 8.8%) than in the United States (to 9.6%).

From both surveys, a sample was drawn of people leaving school for at least 18 months during the survey period. The lengths of the first working and nonworking spells after leaving school were constructed, as was the weekly wage on the first job after school. Both surveys ask about school attendance on a monthly basis, and spells are defined to begin after the last month of attendance, referred to as the leave month. Other variables used in the analysis include sex, race in the United States (white and nonwhite), region of Canada (Atlantic provinces and other areas), receipt of UI compensation in each of the two sample years, whether a job was held while in school, and two levels of education (leaving high school or leaving a post-secondary school). Appendix B describes the details of the sample selection and variable definitions.

Table 1 reports summary statistics for the sample, categorized by country and sex. The excluded reference categories are male, high school or less, Atlantic Canada, and white in the United States. The LMAS sample contains 62% of the total sample. On average, people in both countries spend over five of the first six months of the year attending school full time, and they leave school during the month of May. The women in the sample leave school with more education on average than the men. The middle panel of Table 1 indicates that patterns of work before leaving school also differ by both country and sex. For instance, young people in the United States are more likely than those in Canada to hold a full-time job in the first week of the month in which they leave school. In the United States, 40% of men hold a job in the first week of the leave month, compared to 21% in Canada. For women, these figures are 31% in the United States and 18% in Canada.

Over half the people in the sample experience one week or more of not working after the leave month, but over 90% find jobs by the end of two years. More women than men leave school without a full-time job, and for each sex the percentages are nearly identical across countries. Moreover, nonworking spells after leaving school are somewhat longer for women than men in both countries, and they are longer in the United States than in Canada. On average, U.S. women wait 18 weeks before finding a full-time job, compared to 14 weeks for Canadian women. For men, the figures are roughly 12 weeks in both countries. As for the first working spell, a little less than 60% of the people complete the first spell of working within the two-year sample period. Working spells are on average somewhat shorter in the United States than in Canada. (Survivor functions for working and nonworking spells are presented in Section 2.2.)

Canadians are roughly three times more likely than Americans to report receiving UI benefits during the two-year sample period. Less than 5% of American men and women report receiving any UI benefits in the first year, compared to 13% of Canadian men and 11% of Canadian women. In the second year, the percentage of people receiving UI roughly doubles in both countries. The relative values across countries are of the same magnitude reported by Card and Riddell (1991) based on other data sources. Although it is not possible to determine how many people received UI benefits in the job search after leaving school (the period of interest here), clearly it is a small fraction. Some who received UI in the first year did so before leaving school by working sufficient hours, and others did so after losing a job taken after leaving school. By and large, people leaving school appear to be outside the UI system until they find a job, which allows us to isolate school leavers as uninsured job searchers.

2.2 Empirical Specification

The endogenous variables for each observation are the starting log wage u and the spell durations t_1 and t_2. An indicator for working while in school, W, was defined in Section 1. Three other binary indicators were defined for each observation—S, A, and E. S equals 1 if the person left school with a full-time job, A equals 1 if the person found a job by the end of the survey, and E equals 1 if the person completed a working spell by the end of the survey.

The parameters that describe an individual's search problem are $\beta, \lambda^o, c, \gamma$, and the working-spell parameters λ^l_j and π^l_j. The number of job types J was fixed to 2. Within countries, people are distinguished by the vector of observed characteristics $\mathbf{X}$, which includes sex, region, education, race, and whether a job was held during school. Furthermore, we allow people to differ in ways observable to themselves and to employers that are not captured by $\mathbf{X}$. In particular, for the estimates reported here all individuals are of one of two types indexed by $z \in \{0, 1\}$. Let $\tilde{\mathbf{X}} = [z \quad \mathbf{X}]$ denote the vector $\mathbf{X}$ augmented with z. The worker's type is known by the worker and discovered by the worker's employers before a job is offered. Variation in z captures differences in productivity and attachment to jobs and the labor market, and it generates the negative duration dependence in working and nonworking spell durations. Engberg (1992) accounted for unobserved heterogeneity in a structural search model in a similar fashion, and Wolpin (1987) and Gönül (1989) used offer-probability functions that declined with duration to explain duration dependence. The proportion of workers of type 0 is denoted π_0, and it is allowed to depend on $\mathbf{X}$ through the two previous choices that enter as explanatory variables, level of education, and whether a job was held in school or not. This approach controls for correlation between unobserved-type and lagged-choice variables without an explicit model of the relationship.

Conditional on the person's type, the density of accepted (or actual) log wages for offers accepted with one offer in hand is

$$f^a(u^a; u^{**}) = \begin{cases} \gamma e^{-\gamma(u^a - u^{**})} & \text{if } u^a \geq u^{**} \\ 0 & \text{if } u^a < u^{**}. \end{cases} \tag{14}$$

The density for offers accepted with two offers in hand (during school) equals the density of the maximum of the two offers above u^{**}, or $2f^a(u; u^{**})F(u^{**})$. Search behavior places a lower bound on accepted wages, which implies a nonstandard maximum likelihood problem (see Flinn and Heckman 1982; Christensen and Kiefer 1991). Instead of imposing the condition that the reservation wage equals the minimum observed wage, following Wolpin (1987, 1992) and Engberg (1992), I assume that observed log wages include a normally distributed measurement error, $u = u^a + \varepsilon$, where $\varepsilon \sim N(0, \sigma^2)$, ε is distributed independently across people, and σ is the estimated standard error of the measurement error.

Let $f^0(u; W, s, u^{**})$ denote the density of log wages, taking into account measurement error, the number of possible offers, and the person's characteristics $\tilde{X}$ that jointly determine u^{**}. The density is derived in Appendix A. The distribution of observed log wages is a mixture of $f^0(u; W, S, u^{**})$ over the unobserved type z.

Flinn and Heckman (1982) showed that $\bar{u}$ is not identified from accepted wages, so $\bar{u}$ is held fixed in the estimation. In particular, $\bar{u}$ was set equal to 75% of the average log wage within country and education levels. Christensen and Kiefer (1991) proved that the vector of parameters (c, γ, λ^o) is identified from the unemployment spell length t_1 and the accepted wage offer $\ln w$ with no heterogeneity and no measurement error in wages. With unobserved heterogeneity and measurement error, the boundary condition that helps identify these parameters no longer holds. An attempt to estimate them along with β for all types resulted in unstable estimates. Instead, the discount factor for people of type 0 was fixed at $\beta_0 = .998$, and for type 1 $\beta_1 = .97$. Because the proportion of each type is estimated and is allowed to vary across observed characteristics, the average value of β varies within the bounds of the two fixed values. Using weekly data, β_1 amounts to an almost static decision process because the utility of a dollar six months in the future is discounted by $.97^{26} = .45$.

The search model parameters are related to the estimated coefficients in the following way:

$$\begin{aligned}
\beta &= \beta_z \in \{.97, .998\} \\
\frac{1}{\gamma} &= e^{\tilde{\mathbf{X}}\tilde{\gamma}} \\
c &= \tilde{\mathbf{X}}\tilde{\mathbf{c}} \\
\lambda^l_j &= \frac{e^{\bar{\lambda}^l_j + \tilde{\lambda}^l \tilde{X}}}{1 + e^{\bar{\lambda}^l_k + \tilde{\lambda}^l \tilde{X}}}, \qquad j = 1, 2, \quad \bar{\lambda}^l_2 = 0 \\
\pi^l_1 &= \pi^l_1, \quad \pi^l_2 = 1 - \pi^l_1 \\
\lambda^o &= \frac{e^{\tilde{\lambda}^o \tilde{\mathbf{X}}}}{1 + e^{\tilde{\lambda}^o \tilde{X}}} \\
\pi_0(\mathbf{X}) &= \frac{e^{\bar{\pi} + \boldsymbol{\pi}\mathbf{X}}}{1 + e^{\bar{\pi} + \boldsymbol{\pi}\mathbf{X}}}, \quad \pi_1(\mathbf{X}) = 1 - \pi_0(\mathbf{X}).
\end{aligned} \tag{15}$$

The vectors $\tilde{\gamma}, \tilde{c}, \tilde{\lambda}^l, \tilde{\lambda}^o$, and π and the scalars $\bar{\lambda}^l_0$ and $\bar{\pi}$ are estimated coefficients that convert the variables in $\tilde{X}$ into shifters of the structural parameters. The value of $\bar{\lambda}^l_2$ was normalized to 0 because λ^l_j also contains a full set of worker type-specific coefficients. The logistic form for the probabilities ensures that the values fall between 0 and 1 for all values of $\tilde{\mathbf{X}}$. To keep the number of estimated parameters as low as possible, zero restrictions were imposed on the estimated parameters in (15) so that only two observed characteristics appear in each equation. The zero restrictions were based on reduced-form results available from Ferrall (1994). The restrictions improve identification of the other parameters, and they reduce computation costs because the model must be solved for each type of worker separately. An insignificant coefficient in a reduced-form equation, however, does not guarantee that the variable should be excluded from a structural equation. The restrictions are indicated

Table 2. Maximum Likelihood Estimates

Parameter	Description	Variable	U.S. Coeff.	U.S. Std. err.	Canada Coeff.	Canada Std. err.
Lambda 1	Layoff prob.	Skill 0	−4.956*	(.33)	−5.302*	(.48)
		Skill 1	−3.024*	(.16)	−3.515*	(.22)
		Above H.S.	−.178	(.26)	−.030	(.17)
		Nonwhite	.144	(.22)		
		Atl. Cdn.			.519*	(.17)
		Job type 1	−97.742	—	−5.198	(21.27)
1/gamma	Mean wage offer	Skill 0	−.902*	(.14)	−1.054*	(.28)
		Skill 1	−1.112*	(.09)	−.650*	(.05)
		Female	−.118	(.06)	−.186*	(.05)
		Above H.S.	.042	(.08)	.070	(.07)
Lambda *o*	Offer prob.	Skill 0	14.703	—	−.473	(2.17)
		Skill 1	−1.140*	(.31)	−.842*	(.20)
		Above H.S.	.362	(.58)	.808*	(.31)
		Nonwhite	−.821*	(.38)		
		Atl. Cdn.			−.595*	(.22)
c	Value not work	Skill 0	.310	(5.65)	4.705*	(.57)
		Skill 1	4.373*	(.61)	4.633*	(.09)
		Female	−.041	(.06)	−.001	(.09)
		Above H.S.	.177	(.13)	−.277	(.15)
Pi *z*	Prob. skill type 0	Constant	−1.604*	(.48)	−1.916*	(.55)
		Education	1.770*	(.52)	.449	(.49)
		Wrk. in sch.	2.113*	(.45)	1.056*	(.45)
Pi *j*	Prob. job type 1	Constant	.064	(.07)	.293*	(.13)
Sigma	Std. dev. error		.464*	(.02)	.349*	(.02)
N. of observations			439		726	
-Ln-likelihood			2,569.7		4,528.1	

NOTE: Estimates of structural Equations (15). * denotes significance at 5% level.

in Table 2 by noting which variables enter each structural parameter equation. There are 32 distinct values of $\tilde{\mathbf{X}}$ for each country, which means that the search model must be solved 32 times to evaluate the likelihood function once.

The likelihood for an observation $(t_1, t_2, u, W, S, A, E)$ conditional on unobserved type z is

$$\begin{aligned} L_z(t_1, t_2, u, W, S, A, E) &= [P_S(W) f^0(u; W, 2, u^{**})]^S [1 - P_S(W)]^{1-S} \\ &\quad \times [1 - \lambda^o(1 - F(u^{**}))]^{t_1} \\ &\quad \times \left[\sum_{j=1}^{2} \pi_j^l [\lambda^o(1 - F(u^{**})) f^0(u; 0, 1, u^{**}) \right. \\ &\quad \left. \times (1 - \lambda_j^l)^{t_2 - 1}]^A \lambda_j^{l^E} \right]. \end{aligned} \tag{16}$$

The first line of (16) takes into account accepting or rejecting jobs in school. The second line is the probability of the individual's nonworking-spell length. The third line accounts for completed spells, the probability of the working-spell length, and the starting wage. The last line accounts for completed working spells. An observation's total contribution to the likelihood function is therefore

$$L(t_1, t_2, u, W, S, A, E) = \pi_0(\mathbf{X}) L_0(t_1, t_2, u, W, S, A, E) + \pi_1(\mathbf{X}) L_1(t_1, t_2, u, W, S, A, E).$$

The search model and the log-likelihood function for the sample, $\sum \ln L$, were programmed in Pascal. Estimation was carried out on an IBM Model 375 workstation. Quasi-Newton methods were used to converge to the maximum likelihood estimates. Estimates of the asymptotic standard errors were computed using the outer product of the gradient matrix.

2.3 Parameter Estimates

Table 2 summarizes the estimated model for the United States without a UI system and Canada with UI. Many of the parameters are estimated precisely, although two parameters are not well-behaved for the United States, the constant term $\bar{\lambda}_1^l$ and the coefficient on type 0 in λ^o. They were pushed to large negative and positive values, respectively. Through the logistic form of (16), this pushes the probabilities to the boundary values 0 and 1. This means that in the United States one type of working spell is essentially permanent and one type of worker receives a job offer every week. The estimated standard errors for these two parameters are large and unstable because the logistic transformation is flat for such large values. This did not affect the convergence of the other estimated parameters.

When looking at the patterns within countries, we see larger variation in structural parameters across unobserved types than across observed worker characteristics. Across countries, there is no consistent pattern in the coefficients on unobserved type. In both countries, type $z = 0$ people (by definition those with a higher discount factor) face lower layoff probabilities and greater job-offer probabilities. But the ranking of mean wage offers $(\bar{u} + 1/\gamma)$ and values of c differ in the United States and Canada. In both countries,

Table 3. Estimated Preferences, Job Offers, and Layoffs

Characteristics	Discount factor	ln wage offer	Offer prob.	Layoff prob.	Value of not wrk.	Working duration
U.S.						
White H.S.	.979	4.02	.50	.031	3.00	66.3
Nonwhite H.S.	.978	4.03	.38	.038	3.18	52.7
White above H.S.	.989	4.38	.78	.015	1.76	133.1
Nonwhite above H.S.	.989	4.39	.73	.018	1.80	113.9
Canada						
Non-Atlantic H.S.	.974	4.46	.31	.018	4.64	86.7
Atlantic H.S.	.974	4.47	.20	.030	4.64	49.9
Non-Atl. above H.S.	.977	4.64	.51	.016	4.37	107.4
Atlantic above H.S.	.976	4.64	.37	.027	4.37	64.0

NOTE: All values are averaged across other characteristics and unobserved type. Duration is in weeks; wages are in local currencies.

however, higher proportions of people who are leaving university and who worked in school are estimated to be of type 0 (the relationship with education is insignificant in Canada). The estimates are therefore consistent with selection into more education and into working during school by more "patient" people. We find heterogeneity in the duration of jobs as well. In the United States about 6% of working spells are estimated to be permanent. In Canada the long-lasting spells are not permanent, but the estimated proportion is 34%.

The coefficients on the observed characteristics are the same in the two countries except for the coefficient on leaving college or university in c, the value of a period spent not working. In the United States those leaving university are estimated to have a higher value than those leaving high school, all else equal. As shown in Table 3, however, selection by type 0 into higher education leads to a lower value of c on average among college leavers. In Canada, those leaving college have a lower value of c all else equal, although the difference is insignificant. People leaving a college or university face lower layoff probabilities, higher probability of job offers λ^o, and higher wage offers than those leaving high school. In both countries, women face significantly lower average wage offers than men, but they are estimated

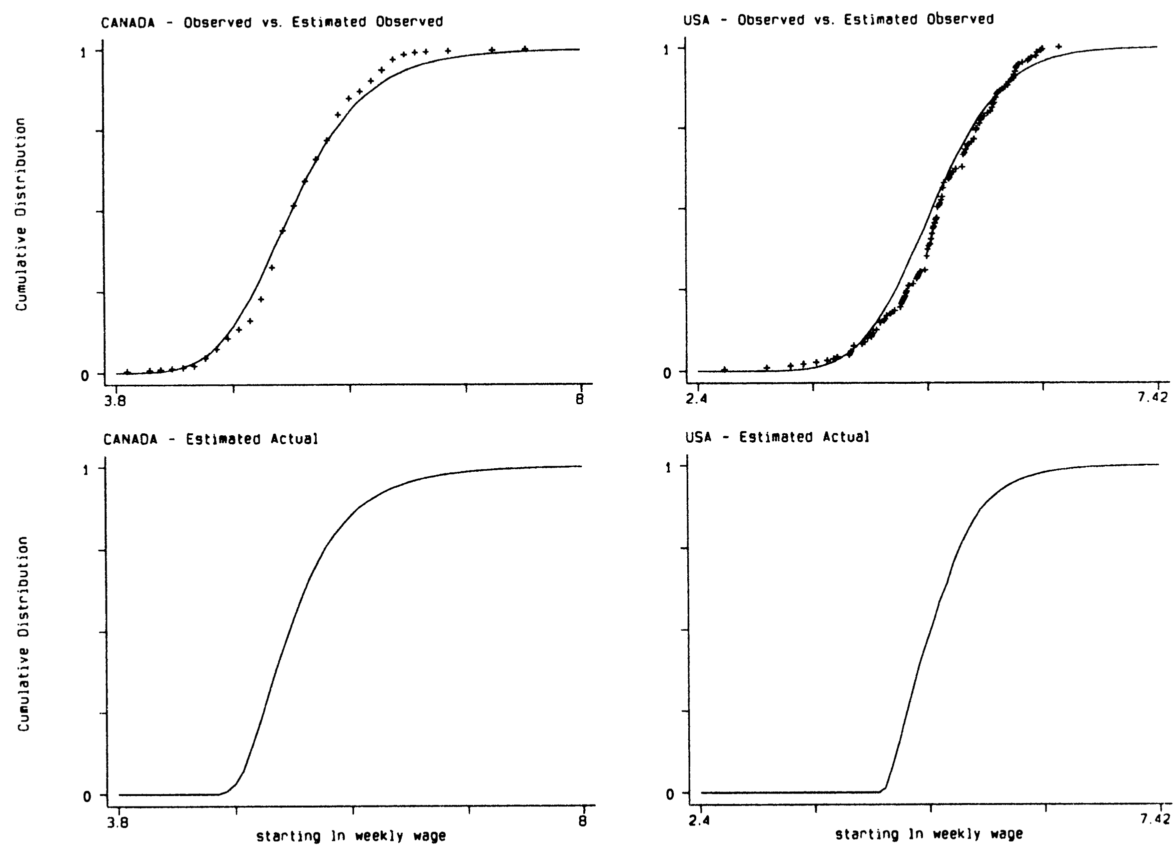

Figure 2. Actual and Predicted Wage Distributions in Canada and the United States. + indicates the empirical distribution of log wages for workers who took jobs after school. Solid lines are predicted distributions of log wages using the maximum likelihood estimates of the model averaged over unobserved type z and the sample averages of observed characteristics X. In the top panel the solid line is the predicted distribution of accepted wages after leaving school plus measurement error. In the lower panel the solid line is the predicted distribution before measurement error.

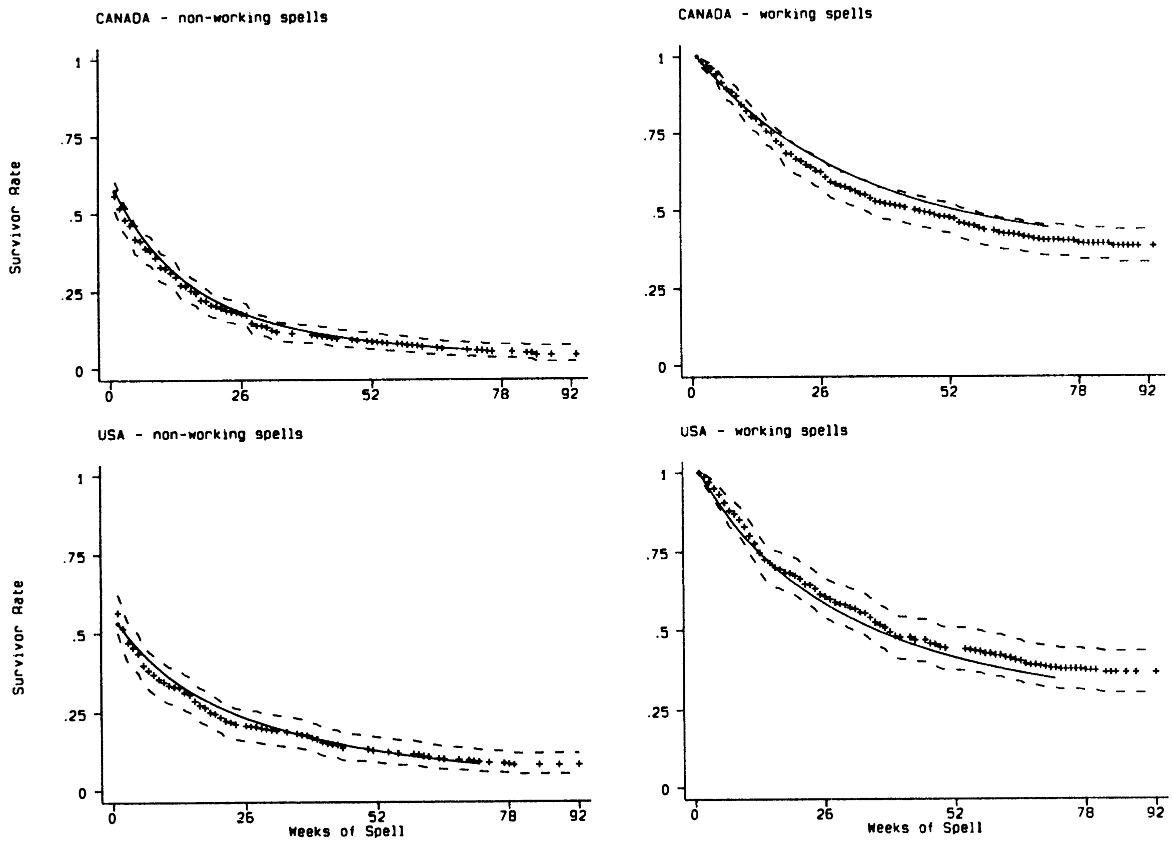

Figure 3. Actual and Predicted Distributions of First Working and Nonworking Spells After Leaving School in Canada and the United States. + indicates the Kaplan–Meier survivor function for the spell. Dashed lines indicate the 99% confidence interval for the empirical survivor function. Solid lines are the predicted survivor function at the maximum likelihood estimates averaged over observed and unobserved types. The predicted values extend to 72 weeks. The survivor function for nonworking spells includes zero length spells due to jobs taken while in school.

to have similar values of not working. Perhaps the strongest demographic effects relate to the Atlantic region of Canada and nonwhites in the United States. Both groups have significantly larger layoff probabilities and smaller job-offer probabilities than other groups in their respective countries.

Figures 2 and 3 illustrate the fit of the model to the data in each country. In all cases the estimated values are averages over observed characteristics $\mathbf{X}$ weighted by their sample proportions and over unobserved type z weighted by their estimated proportion $\pi_z(\mathbf{X})$. The top panel of Figure 2 compares the actual and predicted distributions of observed log starting wages. The predicted distribution is the accepted log wages plus measurement error, $u^a + \varepsilon$. The bottom panel is simply the distribution of u^a conditional on acceptance. In both countries the model captures the overall shape of the wage distribution. The lowest and highest wages in both countries are explained through measurement error because the distribution of accepted wages is more concentrated than in the data. The lower bound on predicted actual wages is not generated by assuming exponential wage offers but rather is an implication of optimal job search. The cumulative distribution of u^a begins to rise at the lowest estimated reservation wage.

Another measure of fit in the wage distribution is to compare the variance of actual wages with the estimated variance of the measurement error, σ^2. A regression of observed log wages on the variables in $\mathbf{X}$ results in a mean squared error of .23 in Canada and .30 in the United States. The square of σ for Canada and the United States is .12 and .22, respectively. Therefore the estimated measurement error in wages is 52% of the residual variance in Canada and 73% in the United States. The remaining variance is due to the distribution of "actual" wages generated by search behavior and the offer distribution.

Figure 3 compares the actual and predicted distributions of nonworking (t_1) and working (t_2) durations. The points are the Kaplan–Meier survivor function and the dashed lines are the point-by-point 99% confidence intervals for the empirical survivor function. The solid line is the predicted survivor function from the model for the first 72 weeks of the spell, again averaged over type and sample proportions. The survivor functions in the left panel begin below 1.0 due to the sizable proportion of people who experience no unemployment as they leave school. The predicted survivor functions stay within the confidence bounds except for a few weeks of working-spell duration in Canada. Looking at nonworking spells, which are the main focus of the article, we see that the data provide a lot of information about the survivor functions because the confidence bounds are tight. In each country, the predicted survivor function stays within the confidence intervals and tracks the data closely.

There is less information about the duration of first working spells in the data as indicated by the wider bounds in the right side of Figure 3. The data cover a two-year period minus four to six months of time in school, so the precision in estimating long-term working duration is lower than for nonworking duration. At longer durations the predicted survivor functions drift away from the empirical values in both countries, but the width of the confidence bands indicate that this is not unexpected. Throughout the duration of a spell, the model is able to explain working durations in the United States, where the UI system creates weaker incentives. Interestingly, however, the model begins to overpredict working-spell survival in Canada near the point at which new entrants become eligible for UI—namely, after about $t_E = 20$ weeks of work. This is consistent with workers and firms responding to the UI rules by raising the probability of a separation into unemployment. The simplification that a working spell ends with constant (but heterogeneous) probability captures the primary effect of UI on uninsured search.

Table 3 summarizes various computed values based on the parameter estimates in Table 2. The entries in the table are again weighted averages over the unobserved type z and the population proportions in each of the region/race and education categories. These demographic groups are highlighted because differences in job-offer and layoff probabilities are greatest among them. The average discount factor is smaller among high school leavers, and the difference is greater in the United States than in Canada because the probability of being type 0 is more strongly related to educational choices in the United States. The difference in the average offered log wages between high school and college leavers is large and similar in magnitude in both countries. The rate of job offers arriving is higher in the United States within educational categories. In the United States, white high school students receive an offer with probability .5 each week, but for nonwhites the probability is only .38. The rate is much higher for people leaving post-secondary schools and is more similar across races (.78 and .73). In Canada, high school leavers face job-offer probabilities of only .20 in Atlantic provinces and .31 elsewhere. As in the United States, the rates are higher for people leaving post-secondary schools, but they only reach values comparable to high school leavers in the United States.

Compared to job-offer rates, layoff probabilities, and through them the expected length of employment, spells follow different patterns. The expected duration for an individual with characteristics $\mathbf{X}$ equals

$$\sum_{z=0}^{1}\sum_{j=1}^{2} \frac{\pi_z(\mathbf{X})\pi_j^l}{\lambda_j^l}.$$

Working spells for high school leavers are shorter in both countries, but the difference is more pronounced in the United States, where employment spells last about one year for high school leavers and over two years for college leavers. In Canada, working spells for high school leavers in the Atlantic provinces are somewhat shorter than in the United States but much longer elsewhere. For college leavers in Canada, working durations are only 64 weeks in the Atlantic region and 107 weeks elsewhere, well below the U.S. averages. The estimated values of not working (c) are more diverse in the United States than in Canada. Average values of c are higher for high school leavers, which may reflect better opportunities to live with parents and perhaps the effect of student loans on college leavers.

3. IMPLICATIONS OF UI POLICY

3.1 The Status Quo

Table 4 summarizes values of endogenous variables under various UI policies. The status quo in the United States is the first column with no UI system in effect. The status quo in Canada is the second column with the UI system in effect during 1986–1987. The estimates imply substantial differences in average accepted wages ($u^{**} + 1/\gamma$) within countries. Given differences in both reservation wages and mean wage offers, however, these differences result in similar probabilities of rejecting offers in hand within countries. People are estimated to reject about 90% of offers in the United States and about 85% in Canada. These offer-rejection probabilities are much higher than the estimates of less than 10% found by Wolpin (1987) and Gönül (1989) for male high school leavers in the NLSY. The estimates are also higher than the directly measured rejection rate of 55% reported by Blau (1992) for the United States using a different dataset. There are numerous differences between

Table 4. Job-Search Responses to UI Policy

	No UI			CDN UI 20 Wks req., reg. ext. ben.			CDN UI 12 wks. req., no ext. ben.		
	Accepted ln wage	Prob. of reject.	Unemp. duration	Accepted ln wage	Prob. of reject.	Unemp. duration	Accepted ln wage	Prob. of reject.	Unemp. duration
U.S.									
White H.S.	4.84	.91	21.8	4.91	.93	32.1	4.92	.93	34.1
Nonwhite H.S.	4.81	.90	40.9	4.90	.93	71.3	4.92	.93	77.1
White above H.S.	5.31	.91	9.8	5.23	.89	6.9	5.25	.89	7.6
Nonwhite above H.S.	5.27	.90	12.6	5.21	.88	9.6	5.21	.88	9.6
Canada									
Non-Atlantic H.S.	5.51	.89	28.5	5.32	.84	16.5	5.43	.88	20.8
Atlantic H.S.	5.40	.86	34.7	5.39	.86	34.0	4.99	.54	14.4
Non-Atl. above H.S.	5.69	.88	17.4	5.57	.85	12.4	5.64	.87	14.0
Atlantic above H.S.	5.53	.84	16.8	5.54	.84	17.0	5.46	.82	11.5

NOTE: All values are averaged across other characteristics and unobserved type. The status quo and estimated values are no UI in the United States and Cdn UI with 20-week entrance requirement in Canada.

the other specifications and the present one, but it should be noted that the positive lower bound on wage offers $\bar{u}$ counts only plausible offers as rejections. The offer distributions used by Wolpin and Gönül included offers near 0 as rejections.

The rejection rates and job-arrival rates combine to determine the expected duration of unemployment spells after leaving school, which for an individual is

$$\sum_{z=0}^{1} \frac{\pi_z(\mathbf{X})}{1 - \lambda^o(1 - F(u^{**}))}.$$

The values in Table 4 include people who experience no unemployment as they leave school. Looking at the status quo in both countries, we see that unemployment durations are over twice as long for high school leavers than for college leavers in the United States. Nonwhites leaving high school average 41 weeks of being unemployed after leaving school, compared to 22 weeks for whites and 13 weeks for nonwhites leaving college. In Canada, the differences in unemployment are not as extreme. Outside the Atlantic provinces, the difference is only 4 weeks between high school and college leavers. The durations in the Atlantic provinces are similar to those of nonwhites in the United States.

The estimates of the search model can be used to simulate how people leaving school respond to changes in the UI system. We consider three alternatives—no UI at all, the Canadian UI system, and the same system with no extended benefits and an extrance requirement of 12 weeks worked rather than 20, which is roughly the entrance requirement for a nonnew entrant. The response to each of these hypothetical systems is measured in Canada and the United States. For each case the search model was re-solved using the structural parameters reported in Section 2. To account for differences in currencies, the minimum and maximum insurable earnings were scaled by the difference in mean wages in the NLSY and LMAS samples when applying the Canadian UI system to the United States. No regional extended benefits were given in the United States.

3.2 Switching Systems

Consider first the radical change of switching UI systems. The effects of these changes are in some respects mild in both countries, but the pattern of change underscores the importance of having structural estimates. For example, adopting the Canadian UI system in the U.S. system is estimated to raise reservation wages among students leaving high school and lower them among students leaving college. The difference comes from the details of the Canadian UI system, as was illustrated in Section 1.5. For high school leavers, a generous UI system has the overall effect of subsidizing high-wage jobs, leading uninsured job searchers to reject some offers they would have accepted without UI. For college leavers, the overall effect is to subsidize employment of any sort, leading them to accept offers they would have rejected without UI.

The change in rejection probabilities from switching systems is only about .02 in each case, which is not a drastic response to such a large change in labor-market institutions. Yet when translating that change into average unemployment duration, the effect on high school leavers is dramatic. The .02 rise in rejection probabilities among nonwhites leaving high school results in an estimated 30-week change in average unemployment spells after leaving school. The reason the effect is so dramatic is because job-offer probabilities are so low for these people. A slight change in rejection rates can translate into a large change in unemployment duration. For college leavers faced with higher job-offer rates the same absolute change of .02 in rejection probabilities only lowers unemployment durations by about three weeks.

Compared to adopting the Canadian UI system in the United States, the effect of scrapping the system in Canada is not symmetric. For people outside the Atlantic regions, the overall effect of eliminating UI is to discourage school leavers from taking jobs, and the result is an increase of unemployment duration from 16.5 to 28.5 weeks among high school leavers and 12.4 to 17.4 weeks among college leavers. The first change is the exact opposite to the same hypothetical removal of UI among high school leavers in the United States, who would stay unemployed for a shorter period. Surprisingly, the estimated response in Atlantic Canada is negligible even though the change includes the loss of extended regional benefits not available elsewhere in the estimated model. Offer rejection probabilities rise slightly among high school leavers, and the result is a slight rise in unemployment duration. A similar but opposite effect occurs among college leavers.

Card and Riddell (1991) suggested that the higher unemployment rate in Canada compared to the United States during the 1980s may have been caused by generous UI in Canada keeping more people attached to the labor market. Within a search framework, this effect should lead to lower rejection rates in Canada among school leavers attempting to get into the UI system. Among people leaving school, job-rejection rates are lower in Canada. Hypothetically, dropping the Canadian UI system moves the offer-rejection rates for high school leavers closer to those found in the United States. The result, however, is a mixture of convergence and divergence in estimated unemployment durations across demographic groups when differences in job-offer probabilities in the two countries are taken into account. In short, the effect of UI policy on the school-to-work transition is ambiguous and depends on the response of both workers and firms.

3.3 Lowering Entrance Requirements

The third panel of Table 4 illustrates that the details of UI systems matter and that assessing the impact of changing the details requires knowledge of structural parameters of the labor market. Here workers now qualify for benefits after only three months of working ($t_E = 12$). This change would have almost no additional effect in the United States. Jobless durations change less than one week among college leavers and only six weeks among nonwhite high school leavers. Compared to the change predicted with the normal 20-week system, these responses are small.

In Canada, however, the effects of lowering entrance requirements into UI are both larger and ambiguous. Outside the Atlantic region, people reject more offers and their unemployment durations increase approximately 2 to 4 weeks. In the Atlantic region, job-rejection rates fall dramatically among high school leavers. Unemployment duration falls from 34 weeks under the status quo to 14.4 weeks with a lower entrance requirement. Recall that the Atlantic region was not responsive to a complete elimination of UI. The large response to changing eligibility rules reflects the timing of qualifying for and receiving benefits. First working spells are so short in the Atlantic region that many people do not qualify for UI after their first job. Eliminating UI therefore does not cause a major change in the incentives facing uninsured workers to take jobs. After lowering the entrance requirements, many more spells qualify for UI, and the benefits are received nearer in time to when the decision to accept the job is made. The UI benefits associated with a job offer are discounted less heavily by β and therefore have a greater impact on the decision to take the job.

By its nature, this sort of experiment captures partial equilibrium responses to changes in government policy because it holds fixed the wage-offer distribution, layoff rates, and job-offer probabilities. These are held fixed because of the difficulty in capturing their endogenous formation in a structural analysis, or for that matter a reduced-form analysis. An equilibrium analysis such as that of Eckstein and Wolpin (1990) and a model of job-quitting behavior such as that of Wolpin (1992) would clearly be preferable. Incorporating the UI system into those approaches presents major theoretical and computational challenges. Baker and Rea (1993) reported evidence that employment duration in Canada lengthened in response to an unanticipated increase in the entrance requirement in 1990. Allowing for a change in working-spell duration in response to a fall in entrance requirements would enhance the effect of lower rejection rates among the unemployed.

4. CONCLUSION

This article has estimated a model of search for jobs covered by UI. The estimates were used to compare and contrast the market for workers leaving school in the United States and Canada. The major differences between the two countries are much lower job-offer probabilities and somewhat lower job-rejection probabilities in Canada. Within both countries, people leaving high school receive job offers less often and face a quicker return to unemployment after accepting a job than people leaving college. Job-rejection rates are estimated to be very similar across education levels, but expected unemployment duration is higher for high school leavers because of the lower offer-arrival rates.

For the most part, the potential changes in job-rejection rates in response to major changes in UI policy are mild, but the direction of change differs both within and across countries. Furthermore, small changes in job-rejection probabilities can translate into major changes in average unemployment durations for those groups facing low job-offer rates. The Canadian UI system accounts for some of the gap in job-rejection rates between school leavers in the United States and Canada. Narrowing the gap by eliminating UI in Canada actually widens the gap in average unemployment durations after leaving school. Lowering the entrance requirement for receiving benefits in Canada would encourage workers in the Atlantic provinces to reject fewer offers. The same change would provoke little response in the United States and in the rest of Canada, and the response is in the opposite direction. Lowering entrance requirements would raise rejection rates in these labor markets.

Although conclusions based on the estimates must be qualified in several respects, the exercise demonstrates that a realistic representation of government programs such as UI can be integrated into a rich model of labor-market behavior. The failure of the literature to account for institutional details noted by Atkinson and Micklewright (1991) need not be permanent. The exercise also demonstrates that in an empirically measurable sense the size or generosity of the UI system cannot be summarized in a single statistic such as the replacement ratio or the average benefit level. The response to a change in UI depends on whether eligibility rules, levels of benefits, or other parameters are being changed.

These results depend on the values of parameters in which we have limited confidence, including discount factors and job-offer probabilities. Estimating job-search models is still in its infancy, and continued experiments with different assumptions should eventually reveal reasonable specifications and parameter estimates with which to analyze policy changes.

ACKNOWLEDGMENTS

Research support from the Social Science and Humanities Research Council of Canada is gratefully acknowledged. Earlier versions of this article were presented at Canadian Employment Research Forums (Vancouver 1993 and Calgary 1994) and the Summer Econometric Society Meetings (Quebec 1994). I thank Michael Baker, Charles Beach, Laura Brown, Martin Browning, John Engberg, Füsun Gönül, David Green, Michael Keane, Gregor Smith, Ken Wolpin, and an anonymous referee for valuable comments. Chris Auld provided helpful research assistance.

APPENDIX A: THE DENSITY OF OBSERVED WAGES

To derive the density function of observed log wages, first the cumulative distribution function is derived for the case in which exactly one offer is in hand, denoted $F^o(u|1, \mu^o > u^{**})$. Let ϕ and Φ denote the standard normal density and distribution function. The error component of observed wages, ε, has density $(1/\sigma)\phi(\varepsilon/\sigma)$. Then

$$\begin{aligned}
&F^o(u|1, u^a > u^{**}) \\
&\quad = \Pr(u^a + \varepsilon < u | u^a > u^{**}) \\
&\quad = \frac{1}{\sigma}\int_{-\infty}^{\infty} \Pr(u^a < u - \varepsilon | u^a > u^{**})\phi(\varepsilon/\sigma)\, d\varepsilon
\end{aligned}$$

$$= \frac{1}{\sigma}\int_{-\infty}^{\infty} F(u-\varepsilon|u^a > u^{**})\phi(\varepsilon/\sigma)\,d\varepsilon = \int_{u-u^{**}}^{\infty} 0$$

$$+ \frac{1}{\sigma}\int_{-\infty}^{u-u^{**}} (1-e^{-\gamma(u-\varepsilon-u^{**})})\phi(\varepsilon/\sigma)\,d\varepsilon$$

$$= \boldsymbol{\Phi}((u-u^{**})/\sigma) - \frac{1}{\sigma}\,e^{-\gamma(u-u^{**})}$$

$$\times \int_{-\infty}^{u-u^{**}} e^{\gamma\varepsilon}\phi(\varepsilon/\sigma)\,d\varepsilon = \boldsymbol{\Phi}((u-u^{**})/\sigma)$$

$$- e^{-\gamma(u-u^{**})+\sigma^2\gamma^2/2}\boldsymbol{\Phi}((u-u^{**}-\sigma^2\gamma)/\sigma).$$

The step leading to the last line uses the fact that $\phi(x)$ is proportional to $e^{-x^2/2}$. Multiplying by $e^{\gamma\varepsilon}$ and completing the square in the exponent results in another normal density function that integrates to $\boldsymbol{\Phi}((u-u^{**}-\sigma^2\gamma)/\sigma)$.

When there are two offers in hand, the distribution of the maximum actual wage is $F^2(u)$, and it can be shown through a similar derivation that

$$F^o(u|2, u^a > u^{**}) = \boldsymbol{\Phi}((u-u^{**})/\sigma)$$

$$- 2e^{-\gamma(u-u^{**})+\sigma^2\gamma^2/2}\boldsymbol{\Phi}((u-u^{**}-\sigma^2\gamma)/\sigma)$$

$$+ e^{-2\gamma(u-u^{**})+2\sigma^2\gamma^2}\Phi((u-u^{**}-2\sigma^2\gamma)/\sigma).$$

Taking derivatives of the cumulative distributions conditional on one and two offers,

$$f^o(u;1,u^{**})$$

$$= \frac{1}{\sigma}\phi((u-u^{**})/\sigma)$$

$$+ e^{-\gamma(u-u^{**})+\sigma^2\gamma^2/2}\bigg(\gamma\boldsymbol{\Phi}((u-u^{**}-\sigma^2\gamma)/\sigma)$$

$$- \frac{1}{\sigma}\phi((u-u^{**}-\sigma^2\gamma)/\sigma)\bigg)$$

$$f^o(u;2,u^{**})$$

$$= \frac{1}{\sigma}\,\phi((u-u^{**})/\sigma)$$

$$+ 2e^{-\gamma(u-u^{**})+\sigma^2\gamma^2/2}\bigg(\gamma\Phi((u-u^{**}-\sigma^2\gamma)/\sigma)$$

$$- \frac{1}{\sigma}\phi((u-u^{**}-\sigma^2\gamma)/\sigma)\bigg)$$

$$- e^{-2\gamma(u-u^{**})+2\sigma^2\gamma^2}\bigg(2\gamma\boldsymbol{\Phi}((u-u^{**}-2\sigma^2\gamma)/\sigma)$$

$$- \frac{1}{\sigma}\,\phi((u-u^{**}-2\sigma^2\gamma)/\sigma)\bigg).$$

For jobs taken after at least one week of unemployment, the observed density of wages that enters the likelihood function (16) is

$$f^o(u|W,1,u^{**}) = \frac{f^o(u|1,u^a > u^{**})}{1-F(u^{**})}.$$

For jobs taken in school,

$$f^o(u|W,2,u^{**}) = \frac{f^o(u|2,u^a > u^{**})}{P_S(W)},$$

where $P_S(W)$ is defined in Section 1.4.

APPENDIX B: SAMPLE SELECTION AND VARIABLE DEFINITIONS

Since 1979, the NLSY has surveyed a sample of people born between 1957 and 1964. The survey includes a supplemental sample of blacks, poor people, and those serving in the military. Only the representative sample is used here. The LMAS is a large representative sample of the Canadian population between the ages of 16 and 69. Ages are coded into brackets, and men and women in the 17–19 and 20–24-year-old brackets were selected. Both surveys ask the respondent which months of the year he or she attended school full time. To identify people leaving school for work, I define a school leaver as someone who attends school full-time for three or more of the first six months of the first sample year (corresponding to calendar years) and who does not attend school in any other month during the sample. Both surveys report employment activities on a weekly basis. I define a respondent to be working in a week when he or she holds at least one job with 20 or more regular hours per week on the job. This definition captures most jobs covered by UI insurance in Canada, which covers jobs lasting 16 or more hours a week. I refer to the last month of attending school as the leave month. The working status during the first week of the leave month is used as an indicator of employment status before leaving school.

For each person leaving school in this way, I define a first nonworking and a first working spell after leaving school. The designs of the NLSY and LMAS surveys make it difficult to distinguish between workers who are unemployed and those who are out of the labor force altogether. Therefore, working and not working are the only activities after leaving school. Nonworking spells begin with the first week after the leave month, and they end the week a worker begins a full-time job. If the person emerges from the leave month holding a full-time job, then the nonworking spell is zero weeks long. Sixteen respondents in the NLSY ended nonworking spells by entering the military. These observations were dropped from the sample. The first working spell lasts from the week the nonworking spell ends until the first week the worker no longer holds a full-time job. A few observations in the NLSY had missing wages. The likelihood function accounts for this.

[Received August 1994. Revised May 1996.]

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