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Abstract: During the mid 1980s young, low-skilled adults in Canada were much more likely to be out of work than their U.S. counterparts. The unemployment rate gap for this cohort was 7 percentage points. At the same time wage inequality was higher in the U.S. Using panel data from the U.S. National Longitudinal Survey of Youth and the Canadian Labour Market Activity Survey, this study employs a general equilibrium search model of the labour market to identify structural differences contributing to these gaps. The results reveal both gaps are driven by a higher job destruction/separation rate in Canada and higher job offer arrival rates in the U.S. In general, the model characterizes the U.S. labour market as more 'competitive' than Canada's.

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1. Introduction

Several studies have been done on the now seemingly long-run phenomenon of the unemployment rate gap between Canada and the United States.[1] The divergence of the rates started over a decade ago with Canada experiencing an unemployment rate that was, on average, 2 percentage points higher than the U.S.'s throughout the 1980s (Card and Riddell (1993)). The fact that this divergence lasted through the recovery and expansionary period of the mid- to late 1980s has led researchers to investigate possible structural or institutional differences, rather than concentrating on demand-based, cyclical determinants.[2]

In separate, but related studies, researchers have compared how the wage distributions have changed over this same period. In particular, researchers such as Burbidge, Magee and Robb (1994), Freeman and Katz (1994), and MacPhail (1993) have investigated whether or not Canada has experienced the same increases in inequality found in the U.S.[3] The evidence so far suggests that Canada has experienced a widening of its wage distribution but not to the same extent as the U.S. This general conclusion, however, is not robust across age and education groups. Burbidge, Magee and Rob (1994) provide evidence of a decline in inequality for older, more experienced and

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[1]See Ashenfelter and Card (1986), Card and Riddell (1993), McCallum (1987), Keil and Symons (1990), and Milbourne, Purvis and Scoones (1991), Baker, Corak and Heisz (1995), Jones and Corak (1995) and Zagorsky (1996).

[2]Research has been done on immigration policies, unionization patterns and unemployment insurance in a search for possible explanations. For an interesting collection of articles on these differences, including Borjas (1993) on immigration, Riddell (1993) on unionization and Card and Riddell (1993) on UI, see Small Differences That Matter. See also Poloz (1994) for a survey of possible explanations behind the increase in Canada's unemployment rate.

[3]The increase in U.S. wage inequality and skill differentials has been well-documented by several authors including Davis and Haltiwanger (1991), Bound and Johnson (1992), Katz and Murphy (1992), Levy and Murnane (1992), Murphy and Welch (1992) and Juhn, Murphy and Pierce (1993).

educated Canadians coexisting with an increase in inequality over this time period for younger, less experienced and educated Canadians.

In this paper I study both areas within Mortensen's (1990) general equilibrium search framework, a framework that brings together several important labour market processes linking them to stocks and flows and wages. The model includes search both on- and off the job and thus imparts information on unemployment and job durations as well as competition between firms for workers. It also houses a job destruction process which relates to employment durations as well as the level of unemployment. These processes, along with productivity measures, help determine the endogenous wage offer distribution of firms and the resulting earnings distribution of workers. Thus, while the framework is simple, it combines several key structural features of the labour market in ways that tie them to both unemployment and wages. While this study is primarily focused on the search framework's interpretation of the forces behind the differences in the countries, it is also concerned with whether or not the differences are consistent with or can be captured by this framework.

The estimation of a specific model of the labour market represents a departure from the more common use of reduced form or regression analysis. Estimation of this search model leads to a second departure. Panel data, in particular panel data at the job spell level, are required to identify all of the parameters in the model. This paper documents and analyzes differences found in comparable panel data from the countries instead of using more common cross-section data. While comparable cross-section data are readily available over a substantial period of time through the U.S. Current Population Survey (CPS) and the Canadian Labour Force Survey (LFS), this is not so with panel data. In addition, estimation of the model requires panel data that provides detailed employment histories at the job spell level. This limits the choice set of data for both countries even further. In this study I use the National Longitudinal Survey of Youth (NLSY) for the U.S. and the Labour Market Activity Survey (LMAS) for Canada. The designs of these surveys lead to time and cohort restrictions. The NLSY is a study of youth only from 1979 to the present and, while the LMAS contains all ages, it covers only two years, 1986-1987.[4] Thus, the intersection of the two focuses this study on the employment histories of young people over the 1986-1987 period.

The young age group and the years 1986-1987 are not without a positive side. 1986 is well beyond the start of the divergence in unemployment rates and wages and therefore, if any structural change is to have taken place in either country, it should have happened by then. It is also a period in which the U.S. and Canada were experiencing relatively stable economic activity levels.[5] So cyclical factors are less likely to play a role. As for the age restriction, young people tend to not be insulated from recent structural changes, since they have been in the labour force only a short time and are most likely still in a job shopping stage.[6] Therefore, estimates from this age group may be more revealing than those from older samples. The cohort of interest here, white males age 20-25 with some or completed high school education, does exhibit the same general unemployment rate gap phenomenon as seen in the aggregate rates throughout the 1980s. In particular, a substantial gap exists in 1986 for this cohort with unemployment rates of 13.5% in the U.S. and

[4]There is also a second wave of the LMAS from 1988-1990. Only the first wave is used here.

[5]This could not be said of 1988-1990 period covered by the second wave of the LMAS.

[6]This job shopping helps keep the censoring rate down for the employment spells. Something that would have been a much larger problem given the short time frame with an older sample.

21% in Canada.[7] This large gap accounts for 20% of the overall male unemployment rate gap in 1986 even though this cohort houses for less than 10% of the male labour force in both countries (Bowlus (1996)). It has also played a large role with respect to the growth in wage inequality in both countries (see Burbidge, Magee and Rob (1994) and Juhn, Murphy and Pierce (1993)).

The panel data exhibits a similar gap in unemployment rates and indicates a slightly higher degree of wage inequality among young, low-skilled American males. The data also reveal three labour market patterns of importance in explaining these differences. Canadians have longer average nonemployment spells,[8] shorter average job spells, and a lower probability of transitioning from a job spell immediately into another job spell than Americans. Thus, during this all important time of job shopping for young workers, Canadians are more likely to experience a spell of nonemployment between jobs.

The search model attributes these patterns to structural differences including significantly higher job offer arrival rates and lower job destruction rates in the U.S. In reconciling the wage inequality gap the model characterizes the U.S. labour market as more 'competitive' than Canada's in terms of firm competition for workers and assigns higher monopsony power levels to Canada. With respect to the unemployment rate gap over 60% is found to stem from differences across the countries in the incidence of unemployment while the remaining 40% is due to duration differences.

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[7]These figures were calculated from the March 1986 U.S. CPS file and the April 1986 Canadian Survey of Consumer Finance (SCF) file.

[8]The division of nonemployment spells into unemployment and non-participation spells is not possible in the NLSY and LMAS. Several restrictions are imposed on the samples in order to better match the homogeneity and unemployed search assumptions of the model. However, it is still true that time spent out of work is likely composed of weeks of both unemployment and non-participation in the labour force. Thus, technically, the rates and durations from the panel data are not unemployment rates and durations but rather nonemployment, and are therefore labelled as such.

The paper is organized as follows. The model is outlined in Section 2. Descriptions of the samples taken from the NLSY and LMAS are given in Section 3. The estimation procedure is discussed in Section 4 and Section 5 contains its results. Conclusions are offered in Section 6.

2. Search Model

I use Mortensen's (1990) general equilibrium search model to investigate the features of the Canadian and U.S. labour markets underlying the unemployment rate and wage inequality gaps. Mortensen's model is a two-state (employment and unemployment) model of the labour market that features both on- and off the job search with a job destruction process. In this section I briefly describe the homogeneous version of the model.

The movement of workers in the model is driven by three exogenous processes: 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. Taking the wage offer distribution of the firms F(w) as given, workers solve the standard linear search utility maximization problem and adopt a reservation wage strategy. Following Mortensen and Neumann (1988) a worker's reservation wage while unemployed is:

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

where $\kappa_0=\lambda_0/\delta$, $\kappa_1=\lambda_1/\delta$ and b is a worker's value of non-market time. In the homogeneous version b and, therefore, r are constant across workers. Job offers arrive at rate λ_0 for unemployed workers who accept the first job that offers more than the reservation wage. Jobs are destroyed, returning workers to unemployment, at rate δ. These two processes imply a steady-state unemployment rate equal to

$$u = \frac{\delta}{\delta + \lambda_0(1 - F(r))} = \frac{1}{1 + \kappa_0(1 - F(r))} \tag{2}$$

For employed workers a job can also end because of a new job offer. Job offers arrive at rate λ_1 for employed workers and are accepted if the new wage offer is higher than the current wage. The rate at which acceptable job offers arrive depends on the current wage w and is equal to $\lambda_1(1\text{-}F(w))$.

In the homogeneous version firms are identical with productivity level P. They face a constant returns to scale production process and maximize profits by choosing a wage to pay given the wage posting strategies of other firms and the search strategy of workers. Profits are given by

$$\pi = (P - w)\, l(w) \tag{3}$$

where l(w) is the steady-state labour supply of workers to a firm offering wage w. Since all firms have the same productivity, profits in equilibrium must be equal across firms. Wage offers, however, are not equal in equilibrium. Because of on-the-job search, l(w) is increasing in w. Offering a higher wage attracts more workers and retains them longer. The balancing condition equating supply and demand is that firms offer higher wages if and only if they can expect to get an additional number of workers to cover the lower per worker profits. In equilibrium the model generates a non-degenerate wage offer distribution shown in Mortensen (1990) to be:

$$F(w) = \frac{1 + \kappa_1}{\kappa_1}\left[1 - \left(\frac{P - w}{P - w_L}\right)^{1/2}\right] \qquad w_L \leq w \leq w_H \tag{4}$$

with density

$$f(w) = [\frac{1 + \kappa_1}{2\kappa_1}]\frac{1}{\sqrt{(P - w)(P - w_L)}} \tag{5}$$

where w_L is the lowest wage offered by firms and equal to r, the reservation wage, and w_H is the highest wage offered.

Since firms never offer wages below r, F(r) equals 0 and equation (2) for the unemployment rate reduces to

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

Thus in this model an increase in the unemployment rate can be explained by either a fall in the arrival rate of job offers, an increase in the rate of job destruction, or a combination of the two. Mortensen also shows the steady-state cross-section earnings distribution, G(w), is a function of F(w).

$$G(w) = \frac{F(w)}{1+\kappa_1(1-F(w))} \tag{7}$$

Because workers move up the wage offer distribution through on-the-job search and stay longer at higher paying jobs, the earnings distribution lies to the right of the wage offer distribution. κ_1 plays a large role in determining the curvature of this distribution by controlling how fast workers advance up the wage range.

Both κ_0 and κ_1 can be thought of as measures of competition in the labour market. They are ratios of job offer arrival rates to the job destruction rate and therefore measure the relative speed of offers or firm competition for workers. This competition has an impact on workers through their reservation wage, as well as on 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. When λ_1=0, $l(w)$ is invariant with respect to w and therefore firms profit maximize by setting all wage offers equal to r. In contrast, as λ_1 moves toward infinity or perfect information, the competitive solution of all workers being paid their marginal product P emerges. For values of λ_1 between zero and infinity, the non-degenerate wage offer distribution given in (4) results. Holding r and w_H constant, an increase in κ_1 will increase the variance of wages. However, both r and w_H will respond to a change in κ_1. If λ_1 increases, the wage distribution will spread out with r falling and w_H rising causing the variance to increase. If δ decreases, the distribution will shift to the right with both r and w_H increasing having an indeterminate effect on the variance.[9] Thus the arrival and job destruction rates are important determinants of both wage inequality and unemployment.

Findings in Bowlus, Kiefer and Neumann (1995a) indicate a poor fit between the wage distribution implied by the homogeneous version of the model given above and wage data. They show that adding firm heterogeneity through variation in productivity levels greatly improves the model's fit.[10] Thus when estimating the model I assume there are Q firm types which can be ordered as $P_1 < P_2 < ... < P_Q$. Mortensen (1990) shows the solution under firm heterogeneity results in a complete segmentation of the wage offer range among firm types. Low productivity firms occupy the lower range of wages while high productivity firm types find it optimal to pay higher

[9]This comparative static result holds only when $\lambda_0 > \lambda_1$.

[10]It is also possible to add worker heterogeneity through variation in the value of non-market time. Since estimation procedures have yet to be developed for this form of heterogeneity, the assumption of worker homogeneity is maintained in this study.

wages. In this case equal profits within, but not across, firm types is required in equilibrium. The equilibrium wage offer distribution in the wage range for firm type j is now given by:

$$F_j(w) = \frac{1+\kappa_1}{\kappa_1} - \frac{1+\kappa_1(1-\gamma_{j-1})}{\kappa_1}\left[\frac{P_j-w}{P_j-w_{Lj}}\right]^{1/2}, \quad w_{Lj}\leq w\leq w_{Hj} \tag{8}$$

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 model implies the following restrictions: $w_{L1} = r$ and $F(r)=0$; $w_{Hj} = w_{Lj+1}$ and $F(w_{Hj})=\gamma_j$, $j=1,...,Q-1$; and $F(w_{HQ})=1$. All wages between r and w_{HQ} are offered, and the highest wage offered by one firm type is the lowest wage offered by the next highest productivity firm type. The relationship between G(w) and F(w) remains as in equation (7) with the wage range of the earnings distribution divided at the same wage cuts as the offer distribution's. Thus $G(w_{Hj}) = \gamma_j^G$, the fraction of workers receiving w_{Hj} or less, is equal to

$$\gamma_j^G = \frac{\gamma_j}{1+\kappa_1(1-\gamma_j)} \ . \tag{9}$$

3. Data Description & Analysis

To estimate the structural parameters of the search model panel data are needed. Specifically information on job and unemployment spells, wages, and transitions between jobs and from employment to unemployment is required. Unfortunately, many U.S. panel data sets do not collect enough information from respondents to allow for the construction of employment histories at the job spell level.[11] One notable exception, which has been used to estimate various search models including Mortensen's (1990), is the NLSY.[12] While it spans a long time frame, 1979 to the present, its main drawback is an age specific sample of youth. All respondents were between the ages of 14 and 22 in 1979. In terms of Canadian longitudinal data sets there are few choices. I use the LMAS, which does have information at the job spell level[13] and includes a nationally representative sample of all ages.[14] However, the first wave of the LMAS only covers two years, 1986-1987. Thus for this comparative study I am only able to collect estimates for a specific age group at one point in time.[15] Taking the intersection of the two surveys I study the employment histories of individuals between the ages of 20 and 25 over the years 1986 to 1987.[16]

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[11]Several data sets keep track of transitions between employment and nonemployment, but to estimate the parameters of this model job-to-job transitions are also needed.

[12]Eckstein and Wolpin (1990), Kiefer and Neumann (1993), Bowlus, Kiefer and Neumann (1995a&b), and Bowlus (1995) all make use of the NLSY.

[13]The LMAS defines a job in terms of its position and duties. Thus in this data a job-to-job transition can occur even if the respondent does not change employers. This is not the case in the NLSY which uses an employer-based job definition. Under a duty-based definition more job-to-job transitions are recorded and job spell lengths are shorter than under an employer-based definition.

[14]Alternative Canadian data sets, such as the Annual Work Patterns Survey and the Survey of 1981 Work History, are mostly one year retrospectives that do not cover enough time to be of use here primarily because of high censoring rates.

[15]This limits the focus to structural differences across the countries after the unemployment rate gap is in existence. It prevents the identification of structural changes over time that may have led to the gap itself.

[16]In 1986 the NLSY sample ranges in age from 21-29. In the LMAS age is categorized such that individuals age 20-24 and 25-34 are lumped together. To keep a close match in ages while not reducing the sample size too much I have chosen to study respondents 21-25 years old in the NLSY and 20-24 years old in the LMAS.

Under the model's assumption of homogeneous workers the sample is restricted to white[17] males who have completed at least 9 years of education but no more than a high school degree and are not enrolled in school over the sample period.[18] In addition to these restrictions, the NLSY respondents must have been interviewed in 1987 and cannot have spent any weeks in the military during 1986 or 1987.[19] There are 922 individuals in the NLSY who meet these requirements and 1498 in the LMAS. I also exclude individuals who are self-employed or work part-time in any job from which information is gathered and individuals who spend a significant fraction of either year out of the labour force,[20] since their search behaviour most likely doesn't match that described by the model.[21] The final sample sizes are 651 and 1097 for the NLSY and LMAS, respectively. Taking the restrictions as a whole results in samples of young, low-skilled males who are active in the paid, full-time labour market. While the non-participation restriction is intended to remove those individuals who are not actively looking for work while nonemployed, it does not completely eliminate the possibility of nonemployment durations containing weeks of non-participation.[22]

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[17]For the NLSY the category is non-black and non-Hispanic. For the LMAS I use respondents who are not visible minorities.

[18]Given the education categories in the LMAS it is not possible to separate high school dropouts from graduates as is the case in the NLSY. Using the NLSY Appendix A examines the effect of combining these groups on the parameter estimates.

[19]The LMAS is already restricted to civilians only.

[20]Individuals who spend more than 20 weeks out of the labour force in either 1986 or 1987 are removed from the sample. This restriction eliminates most of the respondents who never worked over the two-year period.

[21]These restrictions are similar to those used by Van den Berg and Ridder (1993) in their study of the Dutch labour market using Mortensen's (1990) search model.

[22]Limitations of the data prevent one from determining transitions between unemployment and out of the labour force during the nonemployment spells.

Thus, technically, the rates and durations calculated for the samples are nonemployment, not unemployment, rates and durations and are labelled as such.

The two surveys start to overlap the first week of January 1986. Starting the analysis at this point uses the maximum amount of common time available and thus offers the lowest censoring rates. A problem arises, however, with the existence of a large seasonal component in the nonemployment rates at the start of the year. This is especially true for Canada which has nonemployment rates of over 25% during the winter months. The search model has a difficult time reconciling these high rates with the relatively short nonemployment durations given its steady-state assumptions. Thus I have chosen to start the analysis with the first week of April 1986. This leads to nonemployment rates from both panels that are more consistent with the duration data. It does, however, increase the censoring rates.

Information is collected on the state of employment during the first week of April. Three states are distinguished: nonemployment, part-time employment (less than 35 hours per week), and full-time employment. A job must last three weeks to be considered as a job spell.[23] For those who are employed (full- or part-time) I collect information on the wage currently received, the duration (in weeks) of the job,[24] and, if the spell is not censored,[25] the state to which the respondent transits afterwards, i.e. nonemployment or another job.[26] If he transits to

[23]Unless it is censored in which case it can have a duration as short as one week.

[24]In both surveys the start dates of all jobs are available and thus left censoring of the employment spells does not occur. There are, however, right censored spells due to the finiteness of the sample period.

[25]A spell is censored if it lasts beyond the last week of 1987.

[26]If more than two weeks elapse between the end of one job and the start of another, the job spell is coded as ending in a transition to nonemployment.

nonemployment, that duration is collected as well and is determined by the time spent between job spells or until the end of the sample period (in which case the spell is censored) whichever comes first.

For those individuals who are not employed during the first week of April I collect information on the length of time from this week until they find work.[27] The wage accepted at the job is also collected, as well as the duration of the job spell and the transition following the spell, if it is not censored. Again if the transition is to nonemployment I record the duration of that spell as well. Due to the short time horizon the censoring rates are quite high, especially for the job spells at 52% for the NLSY and 48% for the LMAS. It is possible for a spell series to be inadmissible. This occurs if either the job or nonemployment spell has an erroneous start or stop date.[28]

4. Estimation Procedure

Given the structure of the available panel data, the contribution of an individual's spell series to the likelihood function depends on the state he is in during the first week of April, employment or nonemployment. I consider each case separately starting with those individuals who are employed. I have data on the durations of their job spells, D_J, their current wages, w, whether

[27] Prior to 1986 I am unable to determine under the same guidelines for both data sets the start dates of the nonemployment spells. Thus these nonemployment spells are all left censored. If the respondent does not take a job by the end of 1987, I record the nonemployment spell as containing the number of weeks between April 1986 and December 1987 and treat it as censored.

[28] A few job spells in the NLSY have no start and/or no stop dates and a few have start dates which are greater than their stop dates. These erroneous dates cause problems not only for the job spells themselves but also for neighbouring nonemployment spells.

the jobs end 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 during any week is 1 minus the unemployment rate given in equation (6) or $\lambda_0/(\lambda_0+\delta)$.[29] The observed wage for this individual is distributed as g(w), the probability density function (pdf) of the cross-section earnings distribution G(w). Since the job spells for this group are sampled from the stock rather than the flow, their durations are distributed Gamma with parameters 2 and $\delta+\lambda_1(1\text{-}F(w))$ and density

$$f(D_J|S=1) = (\delta+\lambda_1[1-F(w)])D_J\exp\{-(\delta+\lambda_1[1-F(w)])D_J\} \tag{10}$$

where S=1 if the individual is employed during the first week of April (Lancaster (1990)). The probability of a job spell ending due to another job is

$$P(C=1|w) = \frac{\lambda_1(1-F(w))}{\delta+\lambda_1(1-F(w))}\ . \tag{11}$$

The probability a job is lost, P(C=0|w), is 1-P(C=1|w) and the subsequent nonemployment spell is distributed exponential with rate λ_0. The total contribution to the likelihood function is the product of these pieces.

$$\ell(\theta) = \ell(\lambda_0,\lambda_1,\delta,P_1,...,P_Q,r,w_{H1},...,w_{HQ},\gamma_1,...\gamma_{Q-1}) \tag{12}$$

$$\ell(\theta) = \lambda_0/(\lambda_0+\delta)\ g(w)\ D_J\ \exp(-(\delta+\lambda_1[1-F(w)]D_J))\ \delta^{1-C}\ (\lambda_1[1-F(w)])^C\ [\lambda_0\ \exp(-\lambda_0 D_N)]^{1-C}. \tag{13}$$

Next, I consider the contribution of those individuals who are not employed during the first week of April (S=0). The probability of observing this state is the unemployment rate, $\delta/(\lambda_0+\delta)$. Here the nonemployment spells, D_N, seen at the start of April 1986 are sampled from the stock. I observe only the residual lives of these spells, but given the assumption of exponential nonemployment durations, these residual life spells are also exponentially distributed with parameter

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

λ_0.[30] Most of the respondents in this category accept employment sometime during 1986 or 1987. The wages they receive on these jobs are distributed as f(w), the pdf of the wage offer distribution F(w). The job spells, D_J, are sampled then from the flow and exponentially distributed with rate $\delta+\lambda_1(1-F(w))$. The probabilities of transit are the same as those discussed for the previous group. The contribution to the likelihood function is once again the product of these terms

$$\ell(\theta) = \delta/(\lambda_0+\delta)\ \lambda_0 \exp(-\lambda_0 D_N)\ f(w) \exp(-(\delta+\lambda_1[1-F(w)]D_J))\ \delta^{1-C}\ (\lambda_1[1-F(w)])^C \qquad \textbf{(14)}$$

For expositional convenience I have assumed in (13) and (14) that all spells are complete. Appropriate measures are, however, taken to deal with censored spells in the estimation process. Measures are also taken to deal with those individuals with missing wages. Their contribution to the likelihood function is their labour force status in April and any nonemployment durations.

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. The restriction $F(w_{Hj})=\gamma_j$ leads to the following relationship between the productivity levels, the wage cuts and the fraction of firms of each type:

$$P_j = \frac{w_{Hj}-B_j w_{Hj-1}}{1-B_j} \quad where \quad B_j = [\frac{1+\kappa_1(1-\gamma_j)}{1+\kappa_1(1-\gamma_{j-1})}]^2 \qquad \textbf{(15)}$$

Following Bowlus, Kiefer, and Neumann (1995b) (BKN) I replace the P_j's in the likelihood function with the expression in (15). The likelihood is now a function of the rates and the wage cutoffs with their respective γ's. There are actually kinks at the wage cuts in F(w), as well as G(w), and therefore f(w), g(w) and $\ell(\theta)$ are discontinuous at these points leaving standard maximization procedures inappropriate. BKN show the maximum likelihood estimates for r and w_{HQ} are the

[30]This result is special to the exponential case.

sample minimum and maximum, respectively, and those for $w_{H1},\ldots,w_{HQ-1}$ come from the set of observed wages. 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 according to an optimal stopping rule.[31] These estimators of the w_H's are super-convergent and therefore can be treated as known when estimating the remaining parameters using standard maximum likelihood techniques.

The estimation is actually done by 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. Unlike in BKN where all job spells are sampled from the flow and thus the observed wages all stem from the wage offer distribution F(w), here the vast majority of the wage observations are sampled from the cross-section earnings distribution G(w). To handle this difference I rewrite the likelihood function in terms of γ_j^G's instead of γ_j's using equation (9). I then conduct the 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 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 the addition of firm types. Since the exact distribution of the difference in log likelihood values is yet to be worked out, I follow BKN and compare 2 times the change to Chi-square critical values.

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

Given the estimates for r and w_{HQ} are the sample minimum and maximum, one must deal with outliers in the wage data. In the mid 1980s both the Canadian and U.S. labour markets were covered by minimum wage legislation. In the model, the lowest wage offered will be the legislated minimum (w_L=wmin) if the minimum wage is binding (r<wmin). For the low-skilled, less educated, young cohort studied here the minimum wage more than likely binds. Therefore, I treat as missing all hourly wage observations below the 1986 federal minimum wage of $3.35 in the U.S. and $4.00 in Canada.[32] The respondent is not dropped from the sample, because they can still provide information in the form of their labour force status in April and their durations. All acceptable wage responses are converted into weekly rates and left in nominal terms denominated in each country's currency. Because of problems with long right tails in the wage distribution during estimation, I also trim the data at the top.[33] The model has a difficult time reconciling the relatively few and far between 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. I trim the wage data only 1% at the top. This trim level yields more reasonable productivity values while preserving data features and estimation results.

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[32]The federal minimum wage in Canada was changed in May of 1986 from $3.50 to $4.00. It applies only to those workers in industries that fall under federal jurisdiction. Provinces can set minimum wages as well. Generally they tend to be higher than the federal minimum. For example, Ontario raised its minimum wage to $4.35 in October of 1986.

[33]Again the respondent is not dropped from the sample, but rather only his wage is treated as missing.

5. Results

Means for the samples are shown in Table 1.[34] Nonemployment rates for the first week of April are shown in the first row of the table. The 8.6 percentage point gap between the rates is similar in magnitude to the cross-section unemployment rate gap of 7.5 points given in the introduction. The levels are lower. This is to be expected given the sample restrictions, and suggests the restrictions resulted, as intended, in rates that resemble unemployment rates more than nonemployment rates.

The mean durations of both nonemployment and job spells are shown in rows 2-3 and 5-6 of Table 1, respectively. Canada exhibits longer nonemployment durations and shorter job spell durations. The censoring rates reinforce these patterns. In terms of what is driving this cohort's unemployment rate gap, row 8 of Table 1 may be the most revealing. It shows a large difference in the fraction of completed job spells that end in a transition, not to nonemployment, but to another job. In the U.S. this fraction is over 55%, whereas in Canada only 47% do not experience a period of nonemployment between jobs.[35] Mean wages, shown in rows 9 and 10, from both countries fit the model in that those for the wage offer distributions have lower values than those for the earnings distributions. After converting for the exchange rate, the average wage in the U.S. is

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[34]These means are calculated using the weights given in the data. For the NLSY the 1987 weights are used since all respondents in the sample must have been interviewed in that year.

[35]This difference is actually underestimated due to the difference in job definitions pointed out earlier in footnote 13.

higher than that in Canada.[36] In addition, the coefficients of variation in the last row reveal more earnings inequality in the U.S. for this cohort.

Turning now to the model's reconciliation and interpretation of these patterns, Table 2 shows the parameter estimates for the two samples.[37] Rows 1-3 contain the underlying rates for the three processes that drive the model. At the 5% level all three rates differ significantly across the countries. The parameter estimates indicate faster job offer arrival rates and a lower job destruction rate in the U.S. than in Canada. As shown in rows 6 and 7 this leads to higher values of both κ_0 and κ_1 for the U.S. Thus the model characterizes the U.S. labour market as being more competitive than the Canadian market. This can also be seen in the monopsony index values in row 9 where the U.S. value of .31 is smaller than the .43 value for Canada.[38] The value for P in row 4 is the average across workers working at different firm types weighted by their fractions in the market, γ^G. Applying an exchange rate of 1.38 indicates productivity is higher in the U.S.[39]

Table 3 shows the model's predictions for the unemployment rates, average durations and mean wages of the U.S. and Canada. It is useful as a guide to how well the model captures certain features of the data. Predictions regarding unemployment rates and durations are shown in the first four rows. The predicted unemployment rates are lower, more so for Canada than the U.S., than

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[36]The annual exchange rate was 1.39 in 1986 and 1.32 in 1987.

[37]As with the means, weights are used in the estimation process.

[38]This index is defined as MI=E((P-w)/P) and gives a measure of the monopsony power held by firms in a market. As w ranges between 0 and P, the index goes from 1 to 0. If all workers are paid 0, firms have complete monopsony power; whereas, if all workers are paid the competitive wage P, they have none. Thus higher values of the index indicate more power.

[39]Caution should be exercised when making claims based on the productivity values, because, while the rates tend to be robust to trimming of the wages, the productivity values are not.

those observed in the data for April 1986, because the parameter estimates must also reconcile the duration, wage and transition patterns observed throughout the sample period. The resulting gap is only 3.6 percentage points. Decomposition of this gap reveals 55-65% is due to incidence differences across the countries (the higher Canadian job destruction rate), while the remainder stems from longer unemployment durations in Canada (row 2 of Table 3). Rows 3 and 4 show the higher rate of job destruction in Canada leads to shorter job and employment spell durations, on average. The job duration mean in row 3 is influenced by two of the market processes: job offers while employed and the destruction of jobs. Here these forces oppose one another. On one hand, job offers while employed arrive faster in the U.S.: on average a job offer arrives every 70 weeks in the U.S. during employment, while in Canada the average time is 109 weeks.[40] On the other hand, job loss can be expected to occur sooner in Canada. The average time until a job is destroyed is 4.7 years in the U.S., but only 3.7 years in Canada. The fact that job spell durations are predicted to be shorter in Canada indicates the job destruction process is dominant.

The model also predicts a greater likelihood of U.S. young adults to change jobs without an intervening unemployment spell as compared to their Canadian counterparts. An examination of workers at the reservation wage in each country reveals the extent of this difference. The model predicts a reservation wage worker in the U.S. faces a 77% probability his job will end in a transition to another job rather than to unemployment, whereas for a Canadian worker the probability is only 64%.[41] Ferrall's (1994) finding, that Canadians are much more likely to

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[40]This is the length of time until any job offer arrives not until an acceptable job offer arrives which would depend upon the wage. These average durations are calculated using the exponential mean formula of 1 over the rate λ_1.

[41]These percentages are calculated using equation (11) and setting w equal to r.

experience a spell of unemployment after the completion of schooling than Americans, provides additional evidence of this tendency of young Canadians to spend time in unemployment during this important job shopping stage of their careers.

As for wages row 5 of Table 3 and Figures 1 and 2 show the model does a reasonable job of fitting the earnings distribution. The slightly higher U.S. mean is due to the model's difficulty in matching the upper right tail (see Figure 1). For the wage offer distribution the predicted means in row 6 are lower than those in Table 1.[42] This is more true for the U.S. where the higher value of λ_1 has U.S. workers moving up the wage offer distribution rather quickly. All of this results in a larger coefficient of variation for the U.S. than that seen in the data. However, the ordering between the two countries is preserved. The U.S. with its higher level of κ_1, due to a higher job offer arrival rate and a lower job destruction rate, has more variation in wages than Canada.[43] Thus the model establishes a link between offer arrival and job destruction rates in the U.S. and Canada and both the unemployment rate and wage inequality gaps.

The finding of a higher job destruction rate for Canada goes against Baldwin, Dunne and Haltiwanger's (1994) (BDH) finding of similar job destruction processes across the two countries. BDH use establishment based data for manufacturing and essentially study job flows. The use of individual survey data here gives information on worker flows. In the model these two flows are the same. However, this need not hold in the labour market. In particular, a firm may not destroy every job vacated by a worker. The discrepancy between the findings here and BDH may be due

[42]The sample means for the wage offer distributions are computed using much smaller sample sizes than those for the earnings distribution. For example, the sample size for the U.S. wage offer distribution mean is only 30.

[43]In both countries the derivative of the variance of wages with respect to δ, evaluated at the parameter estimates, is negative.

to (1) higher job destruction rates in Canada in industries outside of manufacturing or (2) more worker churning in Canada than in the U.S. The rates estimated here are consistent with two previous studies of worker flows between unemployment and employment, Blanchard and Diamond (1990) and Jones (1992).[44] Evidence for this same cohort in Bowlus (1996) indicates a higher fraction of Canadian jobs end in involuntary or firm-initiated job separations than American. Thus, if there is more worker churning, it is not due to a greater propensity of Canadians to voluntarily quit employment.

6. Conclusions

While this study is limited to the examination of the behaviour of young, low-skilled males, the results from both the panel data analysis and the search model estimations aid in understanding the U.S.-Canadian unemployment rate and wage inequality gaps of the mid 1980s. The panel data reveal Canadians are less likely than Americans to move immediately between jobs, are out of work longer, and exhibit less wage inequality.

The search model attributes these patterns to structural differences in the U.S. and Canadian labour markets. A higher Canadian job destruction/separation rate leads to a higher level of incidence of unemployment, while lower job offer arrival rates bring about longer nonemployment

[44]Blanchard and Diamond (1990) uses matched monthly CPS data for the U.S. and Jones (1992) uses matched monthly LFS data for Canada. Blanchard and Diamond find a monthly exit rate from employment to unemployment of 1.29% for the U.S. Multiplying by 4.3 to convert the weekly rates estimated here to monthly, the search model predicts a rate of 1.76%. They also find an average rate of 24.62% from unemployment to employment compared to the search model's 18.71%. For Canada Jones calculates an average flow rate from employment to unemployment of 3.08% for those 15-24 years old. The estimate from the search model is 2.28%. For the rate from unemployment to employment he finds a rate of 23.87% for young people compared to 16.21% here.

durations. The higher wage variation in the U.S. is also connected to the lower job destruction rate in the U.S. as well as a higher offer arrival rate while employed. In general, the model characterizes the U.S. labour market as more 'competitive' than the Canadian market. This competitiveness gives rise to a lower unemployment rate and more wage inequality in the U.S.

Figure 1 Empirical and Estimated CDFs of U.S. Earnings Distributic :

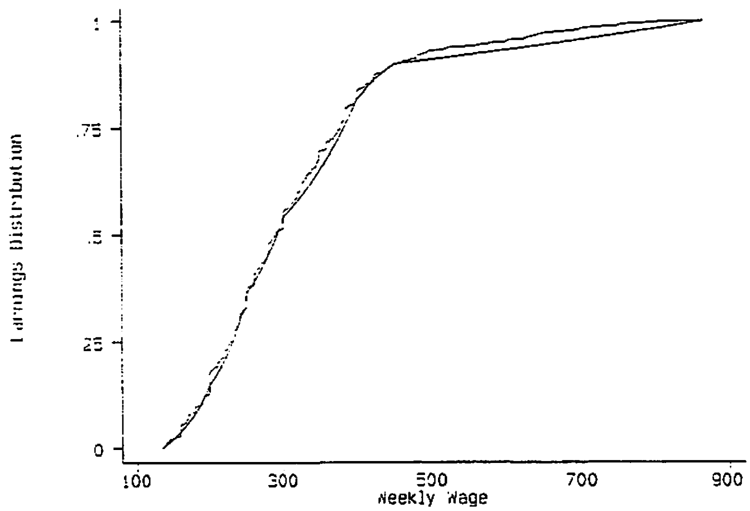

Figure 2 Empirical and Estimated CDFs of Canadian Earnings Distribution

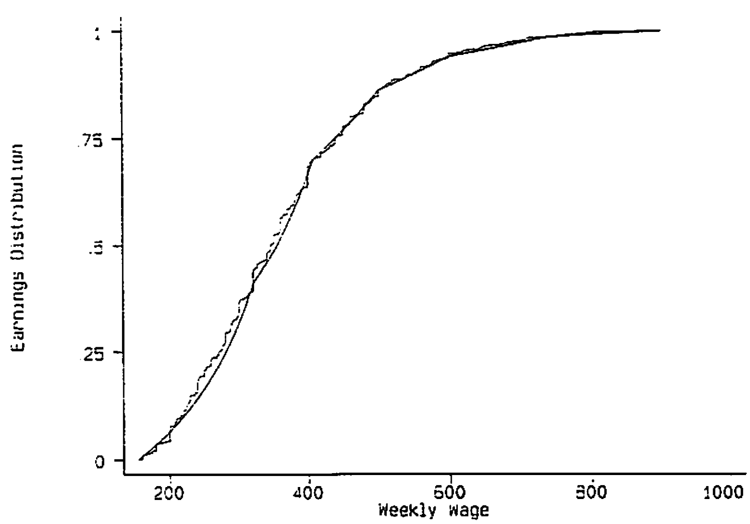

Table 1
Means for U.S. and Canadian Samples

	U.S.	Canada
Fraction nonemployed during first week of April	.0912 (.0113)	0.1771 (.0115)
Nonemployment duration (in weeks) excluding censored spells	15.42 (1.04)	18.57 (0.68)
Nonemployment duration including censored spells	19.64 (1.42)	19.55 (0.73)
Fraction of nonemployment spells censored	.1287 (.0231)	0.1429 (.0147)
Job spell duration (in weeks) - sampled from stock including censored spells	175.40 (5.43)	177.90 (4.16)
Job spell duration - sampled from flow including censored spells	45.92 (4.43)	37.81 (2.16)
Fraction of job spells censored	0.5177 (.0207)	0.4801 (.0155)
Fraction of completed job spells that end in a job-to-job transition	0.5497 (.0303)	0.4620 (.0213)
Wages sampled from earnings distribution	315.06 (5.43)	366.79 (4.58)
Wages sampled from offer distribution	298.91 (18.72)	329.98 (8.31)
Coefficient of variation for wages from earnings distribution	0.3995	0.3596
Sample size	651	1097

Note: Sample weights are used when calculating these means. Standard errors are in parentheses.

Table 2 Search Parameter Estimates		
Parameter	U.S.	Canada
λ_0	0.0435 (.0030)	0.0377 (.0017)
λ_1	0.0141 (.0013)	0.0079 (.0005)
δ	0.0041 (.0002)	0.0053 (.0002)
P (average among workers)	701.93	886.16
Q	7	6
κ_0	10.6103	7.1028
κ_1	3.4284	1.4899
r	134.62	154.00
Monopsony Index	0.3056	0.4286
-log likelihood	6627.55	12728.71

Notes: Standard errors in parentheses. Sample weights are used in the estimation.

Table 3 Model Predictions		
	U.S.	Canada
Unemployment rate	0.086	0.123
Mean unemployment duration	22.99	26.50
Mean job spell duration	149.13	132.19
Mean employment spell duration	243.96	188.22
Mean of earnings distribution	325.32	371.62
Mean of wage offer distribution	237.26	310.98
Coefficient of variation for earnings distribution	0.4394	0.3564

Appendix A: Combining High School Graduates and Dropouts

The categorization of education levels by the LMAS does not allow one to distinguish between high school graduates and high school dropouts. This distinction is possible in the NLSY. Given the homogeneity assumptions of the search model, one would, more than likely, not have chosen to group these two education categories. This appendix analyzes the effect of doing so using the NLSY data.

Table A.1 presents summary statistics for the combined sample and graduates and dropouts separately. A comparison of the groups reveals, as might be expected, dropouts have a higher nonemployment rate, longer nonemployment durations, shorter job spells and a lower job-to-job transition probability. In comparison to CPS unemployment rates the NLSY rates are again lower due to the sample restrictions. This is especially true with respect to the dropouts. The March 1986 CPS unemployment rate for graduates is 11.4% and for dropouts is 20.5%. As stated in the introduction these combine for an overall rate of 13.5%.

Table A.2 shows the search model estimates for the three samples. A comparison of log likelihoods reveals the parameters are statistically significantly different and the two groups should be estimated separately. Given the sample sizes the parameter estimates from the combined sample more closely resemble those for high school graduates.

Thus given the opportunity one would have chosen to analyze the groups separately. Since this is not possible, the U.S. results reveal the combined samples more closely describe the labour market behaviour of high school graduates. This stems from the restrictions placed on the sample and the greater fraction of high school graduates in the labour market. It is likely that the same can be said of the Canadian sample and results.

Table A.1
Means for U.S. Samples

	High School Graduates & Dropouts	High School Graduates	High School Dropouts
Fraction nonemployed during first week of April	.0912 (.0113)	.0868 (.0126)	.1117 (.0258)
Nonemployment duration (in weeks) excluding censored spells	15.42 (1.04)	13.92 (1.10)	20.72 (2.40)
Nonemployment duration including censored spells	19.64 (1.42)	17.29 (1.57)	27.36 (2.93)
Fraction of nonemployment spells censored	.1287 (.0231)	.1135 (.0267)	.1785 (.0461)
Job spell duration (in weeks) - sampled from stock including censored spells	175.40 (5.43)	180.46 (6.10)	149.79 (11.74)
Job spell duration - sampled from flow including censored spells	45.92 (4.43)	49.15 (5.16)	25.07 (6.38)
Fraction of job spells censored	.5177 (.0207)	.5309 (.0234)	.4501 (.0449)
Fraction of completed job spells that end in a job-to-job transition	0.5497 (.0303)	0.5737 (.0346)	0.4445 (.0622)
Wages sampled from earnings distribution	315.06 (5.43)	321.53 (6.27)	282.28 (9.60)
Wages sampled from offer distribution	298.91 (18.72)	314.07 (22.52)	204.67 (16.47)
Sample size	651	501	150

Note: Sample weights are used when calculating these means. Standard deviations are given in brackets. Standard errors are in parentheses.

Table A.2
Search Model Estimates

Parameter	High School Graduates & Dropouts	High School Graduates	High School Dropouts
λ_0	0.0435 (.0030)	0.0493 (.0040)	0.0318 (.0039)
λ_1	0.0141 (.0013)	0.0113 (.0012)	0.0179 (.0033)
δ	0.0041 (.0002)	0.0041 (.0002)	0.0052 (.0005)
-log likelihood	6627.55	5113.66	1467.43

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