The Transition to Work for Canadian University Graduates: Time to First Job, 1982-1990

by

Julian Betts*, Christopher Ferrall and Ross Finnie*****

No. 141

11F0019MPE No. 141
ISSN: 1200-5223
ISBN: 0-660-18313-7

Price: $5.00 per issue, $25.00 annually

Business and Labour Market Analysis
25-F, R.H. Coats Building, Ottawa, K1A 0T6
*University of California-San Diego
**Queen's University
***Queen's University and Statistics Canada: (613) 951-3962
Facsimile Number: (613) 951-5403

October 2000

We acknowledge research support from Statistics Canada and the Social Sciences and Humanities Research Council of Canada. Mr. Finnie acknowledges support from the Applied Research Branch of Human Resources Development Canada. Together we thank Garnett Picot for his support and Michel Villeneuve for exceptional assistance. The authors are solely responsible for the views expressed in the paper.

Aussi disponible en français

ELECTRONIC PUBLICATIONS AVAILABLE AT

www.statcan.ca

Table of Contents

ELECTRONIC PUBLICATIONS AVAILABLE AT

www.statcan.ca

Abstract

Using three waves (1982, 1986, 1990) of the National Graduate Survey (NGS) we analyze the time it takes graduates of Canadian universities to start a full time job that lasts six months or more. We analyze duration to first job using the Cox proportional hazards model. Our results suggest large differences in the speed of the transition to work both within and between cohorts. They also suggest that the differences in duration to first job across NGS cohorts are not just driven by differences in business cycle conditions at the time of graduation. Over certain segments of duration the patterns of job-starting are similar across cohorts. Within cohorts the differences in the school-to-work transition across certain demographic groups are small, and for some the differences remain stable across cohorts.

Keywords: job search; school-to-work transition; youth in the labour market

ELECTRONIC PUBLICATIONS AVAILABLE AT

www.statcan.ca

I. Introduction

Using three waves (1982, 1986, 1990) of the National Graduate Survey (NGS) we analyze the time it takes graduates of Canadian universities to start a job that lasts six months or more. The size of the NGS samples and the amount of detail available in them enable us to focus attention on graduates with little or no prior labour market experience. This narrow focus sheds new light on how university graduates make the transition from school to the world of work.

Understanding this phase of labour market careers is important for several reasons. First, university education is a long and costly investment. Changes in the time it takes to gain steady employment afterwards have immediate effects on the returns to education and may therefore have long-term consequences for skill accumulation. Second, since young people have more fluid employment statuses, their experiences can also amplify larger trends affecting all workers. The 1980s are thought to have been a period of rapid change in the labour market, and our analysis provides new evidence on this view. Finally, since completion of a university degree appears to fundamentally alter a person's labour market opportunities, unemployment rates and other aggregate statistics that average over new and older graduates do not adequately convey the post-graduation experience. Duration analysis is therefore particularly useful for understanding the school-to-work transition.

The National Graduates Surveys (and Follow-Ups) comprise large, representative sample of those who successfully completed programmes at Canadian colleges and universities over three one-year time periods (1982, 1986, 1990). The databases include information on the educational experiences, early labour market experiences, current job characteristics, and basic demographic information of graduates.[1] The present analysis using the NGS has two advantages and one disadvantage relative to previous work (which is discussed in the next section). First, the three NGS surveys cover three cohorts of the graduating population spanning eight years. Other analyses based on single longitudinal surveys are not able to make such comparisons. Second, since the NGS is a large survey of the graduating population, the number of useful observations is up to ten times that of work using non-specialized surveys. One disadvantage of the NGS is that it does not provide post-graduation labour force status week-by-week, which limits our ability to match previous studies based on other data sources. Changes in the 1990 survey questionnaire also make it difficult to definitively compare duration to first job with the earlier cohorts.

We analyze duration to first job using the Cox proportional hazards model which is introduced in Section III. The Cox framework has complete flexibility in the baseline hazard function, which allows us to isolate differences in the survey questionnaires to some extent. The model lets us focus on two related issues. First, how do patterns in duration to first job differ across cohorts? Second, within cohorts how do the relative chances of starting a real job differ across graduates? In Section IV we present our results in detail. In general our results suggest large differences in the speed of the transition to work both within and between cohorts. They also suggest that the differences in duration to first job across NGS cohorts are not just driven by differences in business cycle conditions at the time of graduation. Our analysis does point to some areas of stability and similarity. Over certain segments of duration the patterns of job-starting are similar

---

1 For more information on the NGS and a portrait of the 1990 cohort, see Statistics Canada (1996).

across cohorts. Within cohorts the differences in the school-to-work transition across certain demographic groups are small, and for some the differences remain stable across cohorts.

II. Existing Evidence

There is little existing evidence concerning the school-to-work transition in Canada and little evidence anywhere about how it differs across cohorts of graduates, particularly for post-secondary graduates. Ferrall (1997) compares the school-to-work transition in Canada and the U.S. using the 1986-87 Labour Market Activity Survey (LMAS) and the U.S. National Longitudinal Survey of Youth (NLSY). The average non-working duration for college and university leavers in Canada is 12 weeks and nearly half experience no spell of non-employment. Non-working spells are on average roughly 33% longer for Canadians than Americans.[2]

Eckstein and Wolpin (1995) analyze the school-to-work transition by estimating a search/matching model using the NLSY. The model provides a joint model of accepted wages and search duration. Eckstein and Wolpin find an average duration to full-time job for white university graduates of approximately one month. Using estimates of the search model they find that a job-search subsidy would have a significant effect on mean jobless duration. Ferrall (1997) also finds that the Canadian unemployment insurance (UI) system has an impact on the school-to-work transition even though most people are not covered by UI during the transition. The impact comes through the incentives generated by the UI system to take jobs. Eliminating UI would increase jobless duration for college/university leavers.

These results are based on much looser definitions of a first `real' job than used in the present analysis based on the NGS.[3] A real job is defined as one that is full-time (usually 30 hours or more per week) and lasts six months or more. Eckstein and Wolpin define a real job to be one in which the person worked in the job every week in a quarter for at least 30 hours each week. Ferrall (1997) defines it to be any job lasting 20 hours or more in one week. In the 1986-87 LMAS, roughly 40% of first full-time working durations so defined end before six months. Most of these transitions to employment would therefore be excluded in the present analysis and would be added to our jobless durations. The consequence is that average duration of the school-to-work transition for the 1986 NGS cohort (most comparable in time to the 1986-87 LMAS) is nearly two years rather than 12 weeks in Ferrall (1997) without the six month requirement. Furthermore, only 14% of our NGS sample hold a real job upon graduation compared to the nearly 50% figure in Ferrall (1997) based on the definition of the first job that does not require it to be long term. These comparisons suggest that the school-to-work transition is complicated and that applying different definitions and methods can lead to further understanding of this important period.

[2] Using the same data sets to study youth unemployment durations (not specifically people leaving school), Bowlus (1998) estimates unemployment spells are 15% longer for young Canadians without post-secondary schooling than their American counterparts.

[3] Another difference between our analysis and previous work is that our sample includes people who go back to school after graduation. Most other studies use an *ex-post* definition that the person is not observed to return to school.

III. Data and Empirical Framework

Sample Selection

The National Graduates Surveys and Follow-Ups have been developed by Statistics Canada in conjunction with Human Resources Development Canada. These surveys are representative samples of Canadian post-secondary graduates who completed their programmes in 1982, 1986, and 1990. Respondents are interviewed two and five years after graduation, and the response rates are approximately 80 percent for each of the first interviews and about 90 percent for the second interviews (conditional upon the first interview). However, the 1990 follow-up survey was not available in time for this study.

We selected individuals in the NGS who met the following criteria. First, the degree they achieve in the cohort year is a bachelor's degree or higher, and the degree is higher than their previous degree. To focus on new labour market entrants we included only individuals who report less than one year of work experience and hold no job lasting longer than six months prior to entering the programme. Thus, the typical person in our sample has gone through their post-secondary schooling with at most summer jobs. We included graduates who are age fifty or younger at graduation.

The outcome variable we study is the number of months before a graduate takes a `real' job. For our purposes, a real job is defined as a full-time job (usually 30 or more hours per week) that lasts six months or longer. The 1982 and 1986 questionnaires ask about this job directly, but the 1990 questionnaire does not. A number of possible job start dates that appear in the 1990 survey can be the correct one to use, but in some cases it is impossible to be certain that a given job qualifies as a real job under our definition. We therefore exercise caution in interpreting differences in duration across cohorts. The details of how we inferred duration to the first real job for the 1990 cohort is described in the appendix.

Empirical Framework

Define:

$M^{*} \equiv$Month first real job started-month programme completed (1)

M^{*} is the duration of a person's school-to-work transition. We excluded people who held a real job before starting their degree programme. However, a sizeable fraction (14%) of graduates hold real jobs after graduation that they started during their degree programme. Thus, as defined in (1) M^{*} is often a negative number, implying the graduate experienced no period of transition after graduation. Since students can respond to their expected post-graduation situation before graduation, it would be inaccurate to study only positive post-graduate durations. All cases of $M^{*}<0$ are treated as minimum durations. On a technical note, the computer programme used to carry out the duration analysis (Stata) requires that duration variables be coded as positive values. These two considerations lead to the outcome variable we study:

$M \equiv \max\{M^{*}+1,1\}$ (2)

M=1 implies the person experiences no gap between holding a real job and finishing their degree programme.[4]

Our duration analysis is based on the *proportional hazard framework*:[5]
$h(m;z,X,\beta) = Prob(M=m \mid M \geq m) = h_z(m)\exp(X\beta)$. (3)

Here $h(m;z,X\beta)$ is the *hazard rate* out of the transition period and into a real job. That is, it is the probability that a person with characteristics (z,X) who is still in transition at month m will end the transition that period by starting a real job. In the proportional hazards framework, the hazard rate is composed of two factors. The first is the *baseline hazard function* $h_z(m)$ which is a different function of duration for each value of the *stratifying variable z*. In the proportional hazards framework the baseline hazard function is estimated non-parametrically. That is, $h_z(m)$ is estimated as a separate parameter for each value of m at which a spell is completed. The second factor is $\exp(X\beta)$, which depends upon other individual characteristics X and estimated coefficients β. We report estimates based on four different specifications of $h_z(m)$ and a variety of specifications of X. In all cases we report estimated standard errors for β that are robust to heteroscedasticity.[6]

While we define the model in terms of the hazard function $h(m;z,X,\beta)$, it helps to display the results graphically using the *survivor function*:

$S(m;z,X,\beta) = Prob(M > m) = \Pi_{s=1}^{m}\ (1-h(s;z,X,\beta))$. (4)

This is simply the probability that a spell lasts more than m months.[7] For example, $S(1;z,X\beta)$ is simply the probability that a person does not have a real job at the time of graduation ($M>1$).

The simplest version of (3) is the case that no regressors are included in the X vector. In this case, the hazard function is then identical to the baseline hazard function:

$h_z(m;z,0)=h_z(m)$ (5).

Without regressors the baseline hazard can be estimated using the *Kaplan-Meier estimator*, which provides the maximum likelihood estimator of $h_z(m)$ for each value of m. The Kaplan-Meier estimate of the survivor (or, equivalently, the hazard) function takes into account the right censoring of spells, cases in which we do not see a graduate begin a real job.

4 Adding a constant value to the duration variable has no effect on the models estimated in this paper.

5 See Kiefer (1988) for a general introduction to duration analysis.

The nature of the NGS data limits our attention to models where the covariates are constant throughout the duration of the school-to-work transition.

6 Stata's `robust' option uses the method of Lin and Wei (1989) to compute robust standard errors.

7 The positive connotation of 'survival' of a jobless spell and the negative connotation of the 'hazard' of starting a job are holdovers from the early use of duration analysis to study life expectancy and failure time of equipment. This also explains the use of 'failure' to denote the completion of a spell.

The NGS is based on a stratified sample across province, degree level, and field of study. Our estimates use the sample weights from the first interview. It is not necessary for us to use the second interview weights which take into account lack of response. When the person failed to be interviewed the second time there are two possibilities. Either the transition was complete by the first interview, so the missing information would not be used, or the transition spell was incomplete and it can simply be marked as incomplete. If the person starts their first full-time job less than six months before the last interview with them, their spell is also considered censored at the start of that job. This is because after that date it is not possible to determine when they began a real first job.

Exogenous Variables in X

All the explanatory variables we use to explain the distribution of *M* are time invariant for an individual (see Table 1 for a list). The NGS does provide some covariates that change after graduation, including region of residence and marital status as of each interview. However, since it is not known when these variables changed values between interviews, it is not possible to include them as time-varying covariates in the hazard analysis. Furthermore, in the case of region of residence, graduates presumably base the decision to move to a new region partly on their current labour market status. The graduate's location later in the spell cannot therefore be considered exogenous to the graduate's job finding process. Nonetheless, marital status at the first interview and the number of children as of the first interview are included as constant covariates. By treating these variables as exogenous and fixed over duration, we essentially assume that any changes in their values after graduation (and before the first interview) were expected ahead of time and affected duration starting from the date of graduation.

We augment the variables in the NGS with an unemployment rate constructed from series in the Labour Force Survey and merged with the NGS data. We use the unemployment rate for people aged 20-24 for the respondent's sex and region of institution (listed in Table 1) in the graduation month.[8] The usefulness of the available unemployment rate is somewhat limited, because we can already control for institutional region and sex. However, the unemployment rate does pick up extra variation through differences in the date of graduation, and it provides an interaction between region, sex, and graduation date without requiring interaction terms.

Summary Statistics

Table 1 provides summaries of all the variables used in the duration analysis by cohort. Unlike the later tables and estimation results, the statistics in Table 1 are based on unweighted data so as to summarize the information available in the sample for the hazard analysis. Since the sampling weights are not used to create Table 1, cross-cohort trends or population averages should not be inferred from it. The sample contains over 5,000 observations for each cohort, providing a great deal of information about the transition from school to work among `first time' university graduates. In the 1982 and 1986 cohorts, for which five years of post-graduation information is

[8] It is possible to refine this unemployment rate to be specific to the person's actual age at graduation as well, but this would likely be misleading since our sample is restricted to `new' labour market entrants. Unemployment rates control for neither prior labour market experience nor years of education. The unemployment rate for older people would therefore be dominated by people with greater experience who left full-time schooling several years earlier than our respondents.

available, roughly 75% of the sample is observed to start a real job. The 52% figure in the 1990 cohort reflects the limit of only eighteen months or so of post-graduate information. For the first two cohorts the average observed duration to first job (not controlling for censoring) is nearly two years. In the 1990 cohort, surveyed about two years after graduation, it is over one year.

Roughly three-quarters of our sample graduated from a four-year bachelor's programme. About one third of the sample is married as of the first interview; very few have children. About 75% were raised in an English-speaking home and 20% in a French-speaking home, leaving 5% raised in other languages. Not many (about 10%) attend a school in a different region than they resided in prior to starting the programme. The average age at graduation is between 23 and 26. Mother's years of education is missing for a fairly large fraction (9%) of respondents.[9] We estimate duration models including and excluding this variable to check whether this missing information may have a major effect on the estimates.

Demographic Trends Across Cohorts

Tables 2 and 3 provide some evidence on how the demographic composition of the graduates in our sample changed across cohorts. These tables use the sampling weights and therefore represent the reference population. The greatest difference across cohorts is the increased proportion of female graduates from 50% in 1982 to 58% in 1990. The other major difference between cohorts is that the average person graduating in 1982 faced an unemployment rate of nearly 16% versus an 11% rate for someone graduating in 1990. Average age at graduation increased only slightly to 24 years of age in 1990. This average age does not account for workers going back to school to learn new skills, since our sampling criteria eliminate those graduates with substantial work experience. Other than the increase in female graduates and the drop in unemployment rates, there is little trend in demographics across cohorts.

Table 3 looks more closely at regional migration. The proportion of people moving before entering their programme and the regional source of graduates barely changed from 1982 to 1990. Only in British Columbia do fewer than 90% of students remain within the province to attend school. A striking 97% of Ontario graduates were in the province before their programme began. Ontario receives about 5% of the students from each of the other regions and 47% of foreign students (of those who remain in Canada to be included in the NGS sample).

IV. Analysis of Time to First Job

Kaplan-Meier Estimates

We begin the analysis by summarizing the distribution of M across cohorts using the Kaplan-Meier survivor function for M, based on the simple proportional hazard model (5). The survivor functions are displayed as curves in Figure 1. The shorter curve for 1990 indicates the end of information coming from the first interview. The relative position of the curves reflects both the different characteristics of graduates across cohorts and any differences in experiences within demographic groups. The population survivor functions for 1982 and 1986 during the first 18 months after graduation are remarkably similar. After this point the 1986 cohort is more likely to start a real job.

---

9 Mother's years of education is constructed from several survey questions.

Another way to illustrate this effect is to look at median duration, the amount of time by which half the population has completed the transition to a real job. The median can be seen in Figure 1 where the survivor function reaches the value of 0.50. Median duration is almost identical across cohorts, a little over fifteen months. But for those whose spells go beyond that point the patterns begin to diverge. The median *extra* duration for those lasting past fifteen months can be determined from Figure 1 by looking at where the survivor function reaches 0.25. For the 1986 cohort this figure is an additional 28 months, but for the 1982 cohort it is an additional 37 months.

The 1990 survivor curve in Figure 1 cannot be compared too closely to the earlier cohorts' curves due to the change in the survey instruments. With this caveat in mind, the survival rates are still quite similar from three months to roughly twenty months. This is remarkable considering the differences in unemployment rates that confronted the three cohorts (as shown in Table 2).

Basic Model with Covariates

Our study of the effect other variables have on the duration to the first real job begins with estimates of the basic proportional hazard model:

$h(m;0,X,\beta) = h_0(m)\exp(X\beta)$. (6)

That is, a common baseline hazard function $h_0(m)$ is estimated for everyone by setting the stratifying variable z equal to 0 for everyone. We allow for cross-cohort differences by including in X dummy variables for the 1982 and 1986 cohorts. Thus a person's cohort affects the hazard rate simply as an element of X. Equation (6) therefore assumes that the relative log-hazard rate between identical people in different cohorts is constant across the transition period. Formally,

$\ln h_{1982}(m) - \ln h_{1990}(m) = \beta_1$
$\ln h_{1986}(m) - \ln h_{1990}(m) = \beta_2$

Table 4 summarizes the results from estimates of equation (6). We focus only briefly on this model, as it is nested by later specifications. However, some patterns in these results are interesting to discuss and useful to compare with the more flexible specifications. First, in Table 4 we see that the hazard rates for 1986 are statistically significantly higher than the 1990 rates, even when controlling for differences in unemployment rates: people in the 1986 cohort got real jobs sooner than the 1990 cohort. The rates are also greater for the 1982 cohort except for column 2, but in all cases they are not significant. Second, facing a higher unemployment rate at graduation leads to longer duration to the first real job. This is not surprising. However, the unemployment rate (based on sex and institution's region) is not significant when the regional dummies are also included in the model (columns 2, 4 and 5). At the same time, the estimated size of its effect does not drop substantially (from -0.012 in column 3 to -0.008 in column 5).

The estimated coefficients are, for the most part, similar in sign and magnitude across the different specifications in Table 4. Indeed, the only coefficients that change sign are for the unemployment rate and the indicator for the 1982 cohort when regional indicators are included without other demographic variables (column 2). It is therefore straightforward to summarize the patterns in these results. Bachelor's and master's graduates start real jobs at a slower rate (lower hazard rate due to the negative coefficient) than PhD graduates. Female graduates have lower hazard rates than males, and older graduates have lower rates than younger ones. Married graduates are faster to start real jobs than single ones, but those having children start more slowly than those without. The language the graduate was brought up in does not appear to have a strong effect on the hazard rate, although this result changes in the later specfications. Mother's years of education has a negative effect on the hazard rate. Graduates from outside of Canada and those that moved regions before beginning the program have slower transition speeds than those that attended school in their region.

Stratified Model

In Table 5 we report estimates of a stratified model:

$h(m;\text{cohort},X\beta) = h_{\text{cohort}}(m)\exp(X\beta)$. (7)

Now z equals the person's cohort. This model allows the baseline hazard function to vary arbitrarily across cohorts as well as across duration.[10] It obviously contains (6) as a special case, where the baseline hazard functions are proportional to each other. However, comparison of the likelihood values is not standard because stratifying the baseline hazard curve introduces a (potentially) infinite number of parameters[11], % which rules out use of the χ^2 distribution for the likelihood ratio statistic. Nonetheless, twice the difference in likelihood values (27864.8) between the two full specifications is large enough to conclude that the data are much better described by stratifying across cohort. Below we carry out a formal test of the stratified model itself.

The improvement in fit between the basic model (6) and the stratified model (7) may be an artifact of the change in the survey instrument in 1990 and the difficulty in identifying the first real job after graduation for that cohort. However, (unreported) estimates of both equations using only the 1982 and 1986 cohorts produce a likelihood ratio statistic of 14817.3, which is also extremely large. (Again, it's not accurate to use the χ^2 distribution to create a p-value for this statistic.)

Figure 1 suggested that durations followed different patterns in 1982 and 1986, but Figure 1 does not account for changes in the demographic composition of the cohorts as Table 5 does. Now, we can confidently conclude that identical individuals in different graduation cohorts did not experience the same duration pattern. Furthermore, the difference between cohorts is not merely proportional, as the basic model (6) assumes. It appears that this statement holds true for 1990 as well, although limitations of the data force this conclusion to be tenuous.

Mother's years of education is included in the last specification of both Tables 4 and 5. It has a strong negative effect on duration, but its value is missing for a number of respondents. However, when mother's education is included none of the other coefficients change sign and

[10] We do not control for unobserved heterogeneity in the hazard rate. This has the well-known effect of creating a bias in hazard rates towards negative duration dependence. However, our focus is not on testing for duration dependence *per se*.

[11] Each value that duration m takes on introduces a new parameter, and m is theoretically unbounded. Given the data, the actual number of restrictions imposed by the basic model is finite. In our case the stratified model adds 95 parameters in net.

only the indicator for moving before the programme changes its statistical significance. Marital status and number of children, the two variables known only at the first interview, are also significant. The other coefficients are not sensitive to their inclusion in the model. Furthermore, the patterns across specifications are very similar in Tables 4 and 5. The pattern of coefficients across specifications in Tables 4 and 5 suggest that the full specification in column 5 is satisfactory, and therefore it will be used for the most general specification of the hazard function.

Figure 2 graphs the baseline survivor curves from the stratified model in column 5 of Table 5. Comparing the three cohort's baseline curves must be done with caution. The baseline curve is the survival curve for an observation with an X vector containing all zeros. Since we include variables such as age and mother's education, the level of the curves is itself not empirically relevant. However, since the stratified model restricts the coefficient vector β to be the same across cohorts, the relative position of the three curves does determine the relative survival curve for otherwise identical people in different cohorts.

The baseline survival curves are remarkably similar in shape across the three cohorts, particularly given the major change in the likelihood value when stratifying the model. The one major difference is the much lower survivor curve early in the spell for the 1990 cohort. This appeared in the overall hazards in Figure 1 as well, and it is likely due to the difference in the survey instruments. The survivor curves begin to diverge about six months into the spell, which is earlier than in the overall curves in Figure 1.

The General Model

In Table 6 we report estimates of our most general model:

$h(m;\text{cohort},X\beta_{\text{cohort}} = h_{\text{cohort}}(m)\exp(X\beta_{\text{cohort}})$, (8)

which allows both the baseline and the effect of other characteristics to differ across cohorts. With three separate cohorts equation (8) introduces two new coefficients for each element of X, or a total of 40 extra degrees of freedom compared to (7). Since the baseline survivor curve was already stratified by cohort, a comparison between likelihood values for (8) and (7) is standard. The total likelihood value of -113864 for all cohorts in Table 6 leads to an extreme likelihood ratio of 1956 with the comparable model reported in column 5 of Table 5. We can therefore reject the hypothesis of parameter stability across cohorts. The same test based only on the 1982 and 1986 cohorts also leads to clear rejection of the stratified model, so the result is not simply an artifact of the difference in variable construction between 1990 and the earlier cohorts.

Since both the baseline survivor curve and the coefficients differ across cohort in the general model it is difficult to interpret differences in the baseline survivor curves. The level of the baseline curves does not indicate the relative survivor probability for identical individuals in different cohorts since their probabilities are also determined by different coefficient vectors. Therefore, when discussing these results we will focus on the relative position of different demographic groups within cohorts by comparing coefficient estimates.

Comparison with the Stratified Models

Significant differences in hazard rates exist among the three degree levels (PhD is the excluded category). The coefficients for bachelor's and master's graduates in Table 6 are negative and similar to each other in all three cohorts, which indicates that PhDs are the quickest to start real jobs. The gap between PhDs and the other two degrees is much larger in 1986 and 1990 than in 1982. This change across cohorts was masked by the stratified models reported in Tables 4 and 5 which force variables to have the same relative effect on duration in all three cohorts.

A second difference between the general model and the earlier ones is the size of the gender differences in duration. In the stratified models the indicator for females had a negative and significant coefficient in all cases. In the general model the coefficient is smaller in magnitude relative to the other coefficients in the model and is only significant in the 1986 cohort. One possible explanation for this is that the stratified models use the changing proportion of female graduates across cohorts (see Table 2) to help explain more general differences in duration across cohorts. When both the coefficients and the baseline survivor curve are allowed to vary freely the importance of a male-female difference in transition speeds is reduced. Age at graduation also loses its statistical significance in Table 6, but both married and number of children retain theirs.

Some variables gain strength in the general model in explaining the school-to-work transition. The language indicators do not appear very important in Tables 4 and 5, but in Table 6 we see that graduates raised in French households have larger hazard rates than both English-raised graduates and those from other languages.

Regional Differences in Duration to First Job

The model estimated in Table 6 includes controls for region of institution and region of residence before starting the programme, although only the coefficient for whether the person was living abroad before the programme is reported. (Of the nine regional variables, it has the only significant coefficient, including the first two specifications of the model in Tables 4 and 5.) Furthermore, note that in Table 6 the indicator for whether the person moved regions to attend school is insignificant. This variable is not related to any specific regional effect, but rather captures selection by people into regions, for example, if the best students tend to move out of province to attend university. Since its coefficient is insignificant it appears that this type of selection is not important for determining the school-to-work transition.

A formal test for the existence of regional differences (within Canada) in the duration to first job involves a comparison of the estimated model in Table 6 with one in which the eight Canadian regional indicators (for programme and previous residence) are not included in the model. Estimates of that model (not reported here) yield a total likelihood value across cohorts equal to -13880. The χ^2 likelihood ratio statistic with comparison to the model with the regional effects equals 32.62, with 24 degrees of freedom (eight for each of the three cohorts), and the p-value from the distribution is 0.112. Thus, the likelihood ratio is insignificant at both the 5% and 10% levels. We fail to reject the hypothesis that, after controlling for the regional unemployment rate, there are not differences in the duration to first job across regions of the country, in terms of both where the graduates come from and where they study. Furthermore, the local unemployment rate is itself relatively small and insignificant in Table 6 and changes sign in the 1990 cohort. (In the restricted model without regional indicators the unemployment rate is only significant in 1986.)

The indicator for whether the person moved regions is also not significant in any of the cohorts in Table 6.

Overall, there is no compelling evidence that the market for university graduates differs greatly across regions. This is somewhat surprising given the size and economic diversity of the regions of Canada, as well as the differences in the universities within those regions. What appear more important than location are the measured characteristics of the individual, particularly their degree level and family background (which may measure other aspects of the person's ability besides degree) and other personal characteristics that would affect the graduate's willingness to work. The only exception with regards to the lack of regional effects is the significantly slower transition rate for foreign graduates (that is, those outside the country before the programme started). Since people who leave the country before the first interview are not included in this sample, the transition rate among foreign graduates may be subject to stronger sample-selection problems than among domestic graduates. Which direction this selection problem would bias the estimated coefficient on being outside Canada before the programme is not obvious. Discouraged people may choose to return home, tending to leave the most motivated or employable in the sample. On the other hand, foreign students may be searching for jobs inside and outside Canada simultaneously.

Hazard Ratios

The estimated coefficients in the proportional hazard model are not easy to interpret since they appear in the exponent in (3). It is more straightforward to consider the hazard ratio,

$\mathrm{HR}_j = \exp\{\beta_j\}$,

associated with a coefficient β_j. HR_j is the relative hazard rate in all periods when variable j changes by one unit. A value of HR_j of 2 implies the hazard rate doubles with a one unit change in j. A value of 0.5 implies the hazard rate is cut by half. For example, if j is the indicator variable for female then HR_j=2.0 means that females are twice as likely as males to start a real job in every period. A hazard ratio of 1.0 is equivalent to β_j=0 and would imply no difference between males and females in their rates of starting jobs.

Table 7 reports the hazard ratios for selected variables from the general model in Table 6. The variables are listed in descending order according to the average absolute deviation of the hazard ratio from 1.0. For instance, the indicator variable with the greatest relative effect on the hazard ratio is that for master's graduates, relative to PhD graduates. The effect is negative, in the sense that master's degrees have lower hazard ratios and, thus, longer durations to first job. For the 1986 and 1990 the hazard rate for master's graduates is less than half of that for PhD graduates. The next largest impact is between bachelor's and PhDs. These ratios are closer to 1 than for master's which implies that master's students have the longest durations among the three degree levels (holding all other variables constant).

The next largest relative impacts are for foreign graduates. The hazard rate for graduates from outside of Canada are three quarters that of residents. Students from French-speaking households have larger hazard rates that those from other languages. The estimated coefficients for graduates from English-speaking households in Table 6 are significantly negative but numerically small. This implies their hazard rates are very close to 100% and that the French-speaking hazard is also the relative effect between French- and English-graduates.

Married graduates have quicker transitions than non-married graduates, but those with children have somewhat longer transitions than those without children. Children increase the value of non-market time and would lead to higher reservation wages and consequently longer transitions. While female graduates have lower hazard ratios than their (otherwise identical) male counterparts, the effect is statistically significant only in the 1986 cohort.

Stability of the coefficients across cohorts is clearly rejected by the data in the likelihood ratio tests reported earlier. However, the hazard ratios in Table 7 demonstrate a great deal of qualitative stability. For example, the hazard ratio for graduates from outside of Canada barely changes across cohorts. The only hazard ratio that crosses over 100% is mother's years of education. In the 1982 and 1986 cohorts a more educated mother is associated with faster transitions but a slower transition in 1990. Mother's education may pick up characteristics of the graduate that employers care about, or it may possibly pick up the effect of parent's household income on search costs. The switch in its relative effect may indicate a change in the weighing of these two factors across cohorts.

V. Conclusion

How long does it take university graduates to make the transition from school to work? The answer depends on when you think the transition is complete. The definition employed in this paper is, partly by necessity and partly by design, tighter than what is used in most studies. A person is defined to have made the transition when they have held a full-time job for six months or more. Based on this definition, the transition takes a fairly long time. Although a large fraction of graduates get jobs within a few months, the median duration to the start of the first job is over fifteen months in each graduating cohort.

Did the transition from school to work change from the early 1980s to the early 1990s? Our answer is that it did change in a number of important ways, but not necessarily in a steady trend. In some ways the overall experiences of the 1990 cohort of graduates appear more similar to those in the 1982 cohort than the 1986 cohort. The differences are more pronounced in longer durations, that is for those graduates taking six or more months to make the transition. Our results also indicate that differences in the transition between cohorts are not summarized well by a single number. Instead, to fit the data well we find it necessary to allow the whole pattern of the school-to-work transition to differ across cohorts.

Perhaps more importantly, our results also indicate that the relative duration patterns among demographic groups differed across cohorts. Here, the most striking effect concerns degree level. While PhD graduates experience shorter durations in all three cohorts, the difference is much greater in the 1986 and 1990 cohorts. This suggests a marked change in the market for PhD students since the early 1980s, relative to other graduates.

Our results also suggest some areas of stability across cohorts and similarities across demographic groups. Bachelor's and master's graduates have similar transition rates in all three cohorts. The relative speed of transition across languages and between foreign and resident graduates are also quite similar across cohorts. Furthermore, differences across regions of the country in the transition to a real job are statistically insignificant in all three cohorts. The hazard rates for female graduates are somewhat lower (i.e. longer durations) than for males but statistically significant only in the 1986 cohort.

These results are consistent with something like a national market for university graduates at each degree level, with possible differences only across language. In general, the major predictors for differences in duration to first job appear to be related more to variables affecting the graduates' costs of job search than the external factors related to a local labour market.

At least three directions for further research are clearly suggested. First, our analysis focuses on the ultimate effect that all possible strategies have on the time spent before starting a real job. Yet during the time between graduation and the start of a real job many people are not looking for work full time. Many people may work part time, return to school, or for some other reason may not be available for work. These outcomes are `competing' risks with regular employment all during the post-graduation period. The interplay between steady, full time work and other uses of time is one obvious step to take next. The second next step is to study other aspects of post-graduate labour market experience, most notably income and the characteristics of jobs. Betts et. al (1998) and Finnie (1998a and 1998b) analyze these broader aspects of post-graduate experiences for the three recent NGS cohorts. Finally, the NGS also contains graduates of vocational and technical colleges. Comparing their transitions to work with university graduates would be of interest. One would expect that the market for college graduates to exhibit more local differences than appear to operate for university graduates.

Appendix: Identifying First Real Job in the 1990 Cohort

For those not working in June 1992 (at the first interview and the only interview of the 1990 cohort used in this study), the following set of questions was asked:

qC6 In what month and year did you start working at the first job you had after completed your programme?
qC7 Was this a job at which you usually worked 30 or more hours a week?
qC8 Was this a permanent or a temporary position?

If qC7 and qC8 were both answered yes, then this job was considered a potential real job. Although it is not guaranteed that it lasted six months or more it was characterised as a `permanent' position by the respondent. The duration between its start date and graduation was defined as *D1* and $M^*=D1$ (see equation 1). The spell was considered complete. If either of these questions was answered no, or there was no first job, then duration was set as June 1992 minus the graduation date, and duration was considered incomplete. However, it is conceivable that there was another job (whether the second or later) that started after graduation, was a real job, but ended before June, 1992. Unfortunately, information about such a job is unavailable for the 1990 cohort. So for some people the duration variable may be too *long* (because another real job began and ended between the first job and June, 1992), or it may be too *short* (because the first job was permanent and full time but did not actually last six months). We rely on the possibility that the number of respondents to whom these special circumstances apply is small and that the under- and over-estimation of duration tend to balance out.

For those respondents holding a job in June 1992 it was possible to determine its start date (call it *D2*) and whether it was full time job. If this job is a `real' job it can be identified as such, but it is not necessarily the *first* real job. If the job had no breaks associated with it, then qC41 (``Was that job the first job after graduation'') was asked. If it was answered yes and if the current job had lasted six months or more then $M^*=D2$. If the current job started less than six months earlier than June 1992, the spell was considered censored as of the start of the job. That is, for current jobs that started within six months of the interview it can only be determined that the first real job began after the start of the current job.

If qC41 was answered no, then a set of questions similar to qC6-qC8 above was asked about the first job after school. If it was permanent and full time, then its duration-to-start was defined as *D3* and $M^*=D3$. The same caveat applies here as above: for some respondents *D3* is too short a duration and for others it is too long.

A similar procedure was followed for current jobs that had breaks or if the person was temporarily absent from the current job. In both cases, the person was asked if this was the first job after graduation and if not whether the first job was permanent and full time.

In summary, the variable used in the analysis as duration must be constructed from a set of derived durations. For any particular respondent only one of two durations would apply based on whether the person was employed in June 1992.

Table 1. Summary of Sample by Cohort (Unweighted)

	1982 Cohort: 7090 obs.		1986 Cohort: 9098 obs.		1990 Cohort: 5632 obs.	
	Mean/ Pct	**Std. Dev.**	**Mean/ Pct**	**Std. Dev.**	**Mean/ Pct**	**Std. Dev.**
Variable / Values		.		.		
M=Months to First Job	23.16	22.11	22.76	21.12	12.72	11.24
Completed Spell	0.72	0.45	0.76	0.43	0.52	0.5
Unemp. Rate at Graduation	16.27	4.72	14.91	4.52	11.58	3.39
Degree Program	3.28	0.53	3.25	0.46	3.45	0.66
Bachelor's %	76		77		64	
Master's %	20		22		26	
PhD %	4		1		9	
Female	0.47	0.5	0.48	0.5	0.53	0.5
Married	0.32	0.47	0.31	0.46	0.36	0.48
Number of Children	0.12	0.44	0.1	0.46	0.19	0.62
Raised in English Household	0.75	0.43	0.78	0.41	0.96	0.18
Raised in French Household	0.2	0.4	0.18	0.39	0.41	0.49
Moved Region before Pro.	0.09	0.29	0.08	0.27	0.12	0.32
Region of Institution (1-5)	2.74	1.21	2.89	1.22	2.71	1.16
Atlantic %	21		18		17	
Quebec %	20		18		26	
Ontario %	32		27		32	
Prairies %	19		28		18	
British Columbia %	8		8		7	
Region before Programme (1-6)	2.8	1.3	2.94	1.27	2.84	1.31
Atlantic %	20		17		17	
Quebec %	22		19		26	
Ontario %	30		27		30	
Prairies %	18		28		17	
British Columbia %	8		8		7	
Outside Canada %	2		2		4	
Age at Graduation	24.16	4.84	23.89	3.99	25.47	5.7
Mother's Years of Education	11.86	3.15	12.18	3.1	12.11	3.45

Note: Atlantic region includes: Newfoundland, New Brunswick, Nova Scotia, and Prince Edward Island.
Prairies region includes: Manitoba, Saskatchewan, and Alberta.

Table 2. Demographic Trends Across Graduate Cohorts

Variable/Values	1982	1986	1990	Overall
Degree Program (pct)				
Bachelor's	91	92	90	91
Master's	7	7	9	8
PhD	1	0	1	1
Female (pct)	50	51	58	53
Mover before Programme (pct)	7	6	8	7
Region of Institution (pct)				
Atlantic	9	9	9	9
Quebec	28	26	27	27
Ontario	44	43	45	44
Prairies	14	15	13	14
British Columbia	6	7	7	6
Mean Age at Graduation	23.5	23.3	24	23.5
Mean Unem. Rate at Graduation	15.9	13.9	10.7	13.8

Note: Data weighted by first interview sampling weight.
Percentages may not total 100 because of rounding.

Table 3. Regional Mobility to Attend University (All Cohorts)

	Region of Institution					
Region before Pro. %	**Atlantic**	**Quebec**	**Ontario**	**Prairies**	**B.C.**	**Total**
Atlantic	92	2	5	1	0	100
Quebec	1	94	5	0	0	100
Ontario	1	1	97	1	0	100
Prairies	0	1	4	93	2	100
British Columbia	1	2	6	6	85	100
Outside Canada	7	24	47	15	8	100
Overall %	9	27	44	14	6	100

Note: Data weighted by first interview sampling weights.

Table 4. Proportional Hazard Models with Common Baseline Hazard Across Cohorts

Variable	1	2	3	4	5
Cohort82	0.021	-0.049	0.024	0.003	0.034
	(0.03)	(0.04)	(0.04)	(0.04)	(0.04)
Cohort86	0.106 *	0.066 *	0.105 *	0.091 *	0.09 *
	(0.03)	(0.03)	(0.04)	(0.04)	(0.04)
Unemp. Rate at Graduation	-0.009 *	0.0057	-0.012 *	-0.008	-0.008
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Bachelor's	-0.395 *	-0.401 *	-0.433 *	-0.435 *	-0.417 *
	(0.06)	(0.07)	(0.07)	(0.07)	(0.08)
Master's	-0.471 *	-0.455 *	-0.458 *	-0.462 *	-0.457 *
	(0.07)	(0.07)	(0.07)	(0.07)	(0.08)
Female			-0.148 *	-0.135 *	-0.138 *
			(0.02)	(0.03)	(0.03)
Age at Graduation			-0.014 *	-0.014 *	-0.014 *
			(0.00)	(0.00)	(0.00)
Married			0.174 *	0.18 *	0.181 *
			(0.03)	(0.03)	(0.03)
Number of Children			-0.158 *	-0.154 *	-0.144 *
			(0.03)	(0.03)	(0.03)
Raised in English Household			0.081 *	0.067	0.073
			(0.04)	(0.04)	(0.04)
Raised in French Household			-0.062	-0.048	-0.073
			(0.04)	(0.05)	(0.05)
Mother's Years Education					-0.018 *
					(0.00)
Moved Region before Pro.			-0.169 *	-0.092 *	-0.072
			(0.04)	(0.05)	(0.05)
Outside Canada		-0.41 *		-0.283 *	-0.298 *
		(0.12)		(0.14)	(0.14)
Observations	21712	21452	21395	21395	19584
ln Likelihood	-142149	-140556	-140071	-140052	-128775

Note: Robust standard errors in parentheses. '*' indicates significance at the 5% level.

Excluded category: cohort90, PhD, male, single, household other language, Atlantic Canada, non-mover.

Columns including 'Outside Canada' also include indicators for other pre-program regions and the institutions' region.

Table 5. Proportional Hazard Models Stratified by Cohort

Variable	1	2	3	4	5
Unemp. Rate at Graduation	-0.009 *	0.0055	-0.013 *	-0.008	-0.008
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Bachelor's	-0.383 *	-0.388 *	-0.419 *	-0.421 *	-0.400 *
	(0.06)	(0.06)	(0.07)	(0.07)	(0.07)
Master's	-0.458 *	-0.442 *	-0.444 *	-0.448 *	-0.439 *
	(0.06)	(0.07)	(0.07)	(0.07)	(0.07)
Female			-0.152 *	-0.139 *	-0.143 *
			(0.02)	(0.03)	(0.03)
Age at Graduation			-0.014 *	-0.014 *	-0.013 *
			(0.00)	(0.00)	(0.00)
Married			0.174 *	0.179 *	0.181 *
			(0.03)	(0.03)	(0.03)
Number of Children			-0.156 *	-0.153 *	-0.143 *
			(0.03)	(0.03)	(0.03)
Raised in English Household			0.082 *	0.067	0.074
			(0.04)	(0.04)	(0.04)
Raised in French Household			-0.062	-0.048	-0.073
			(0.04)	(0.05)	(0.05)
Mother's Years of Education					-0.018 *
					(0.00)
Moved Region before Pro.			-0.171 *	-0.095 *	-0.075
			(0.04)	(0.05)	(0.05)
Outside Canada		-0.414 *		-0.284 *	-0.300 *
		(0.12)		(0.14)	(0.14)
Observations	21712	21452	21395	21395	19584
ln likelihood	-126935	-125469	-125020	-125001	-114842

Note: See Notes to Table 4.

Table 6. Proportional Hazard Models by Cohort

Variable	1982	1986	1990
Unemp. Rate at Graduation	-0.003	-0.001	0.010
	(0.01)	(0.01)	(0.01)
Bachelor's	-0.284 *	-0.699 *	-0.624 *
	(0.10)	(0.12)	(0.07)
Master's	-0.209 *	-0.790 *	-0.780 *
	(0.10)	(0.12)	(0.06)
Female	-0.061	-0.122 *	-0.021
	(0.05)	(0.03)	(0.04)
Age at Graduation	-0.012	-0.009	-0.007
	(0.01)	(0.00)	(0.01)
Married	0.130 *	0.156 *	0.218 *
	(0.04)	(0.03)	(0.04)
Number of Children	-0.092 *	-0.108 *	-0.198 *
	(0.05)	(0.04)	(0.04)
Raised in English Household	-0.010 *	-0.011 *	-0.021 *
	(0.00)	(0.00)	(0.01)
Raised in French Household	0.282 *	0.173 *	0.179
	(0.08)	(0.06)	(0.11)
Mother's Years of Education	0.123	0.164 *	-0.111 *
	(0.09)	(0.07)	(0.05)
Moved Region before Pro.	-0.119	-0.064	0.084
	(0.07)	(0.05)	(0.07)
Outside Canada	-0.292	-0.303 *	-0.347 *
	(0.17)	(0.13)	(0.16)
Observations	5786	8418	5380
ln likelihood	-36175	-54870	-22818

Note: See notes to Table 4.

Table 7. Selected Hazard Ratios for Table 6

Variable	1982	1986	1990
Master's	81% *	45% *	46% *
Bachelor's	75% *	50% *	54% *
Outside Canada	75%	74% *	71% *
Raised in French Household	133% *	119% *	120%
Married	114% *	117% *	124% *
Mother's Years of Education	113%	118% *	89% *
Number of Children	91% *	90% *	82% *
Female	94%	89% *	98%

Note: Hazard Ratio = exp(coefficient in Table 6).
* indicates coefficient in Table 6 is significant at 5% level
Variables are sorted by average of abs(HR-1) across cohorts

Figure 1. Cohort Survival Curves, population weighted

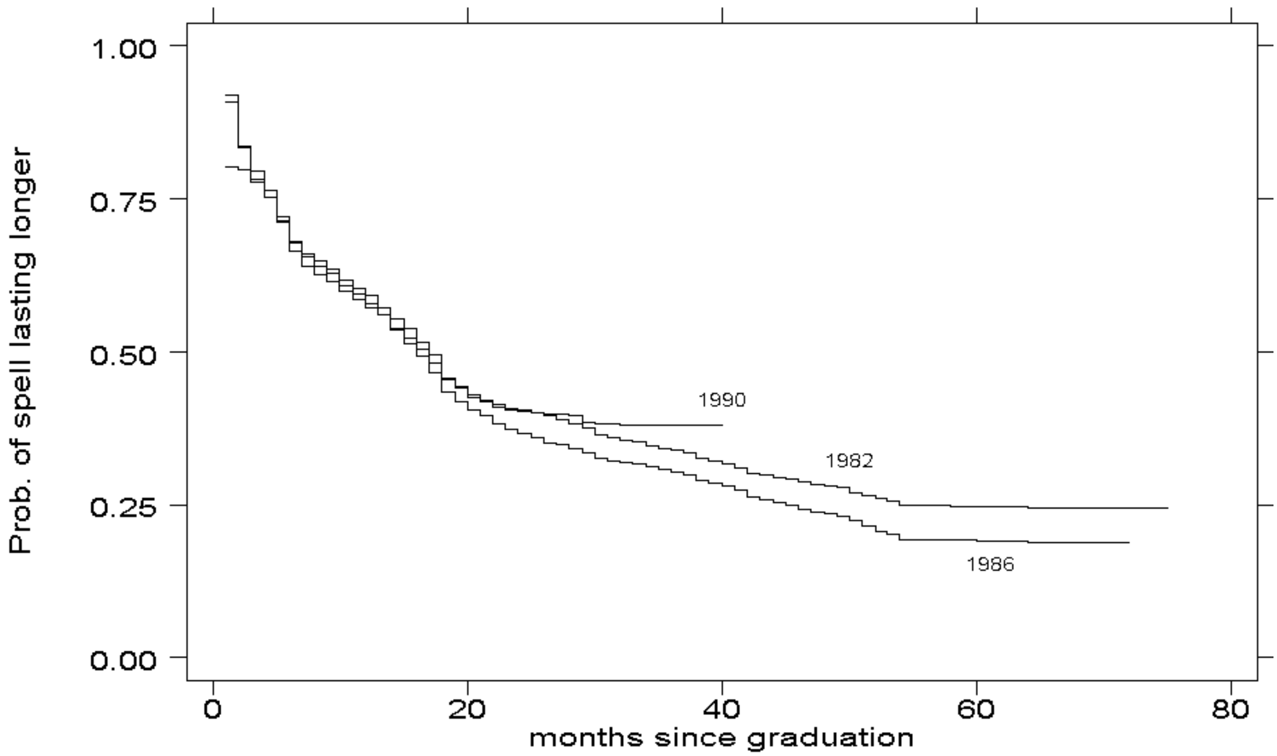
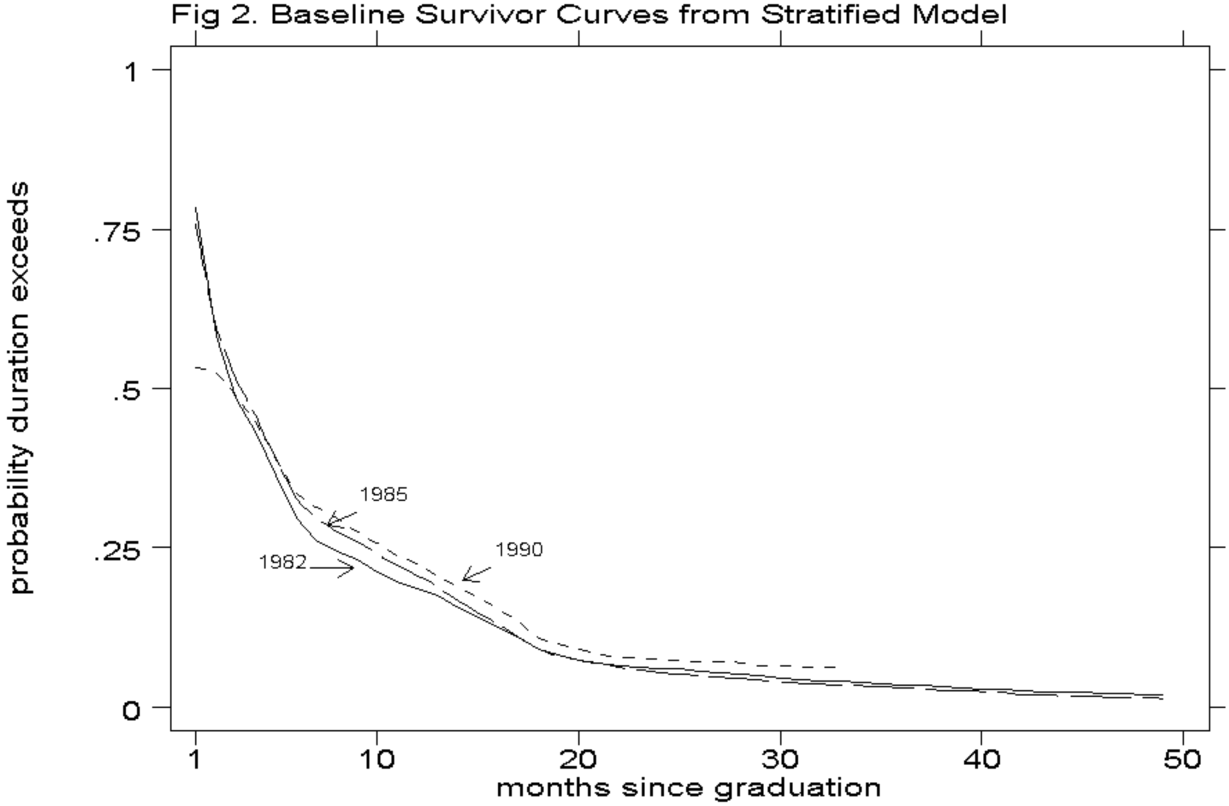


Fig 2. Baseline Survivor Curves from Stratified Model



References

Betts, Julian, Christopher Ferrall and Ross Finnie. 1998. ``The Role of University Quality in Determining Post-Graduate Outcomes: Panel Evidence from Three Recent Canadian Cohorts," work in progress.

Bowlus, Audra. 1998. ``U.S. Canadian Unemployment Rate and Wage Differences Among Young, Low Skilled Workers in the 1980s," forthcoming, *Canadian Journal of Economics*.

Eckstein, Zvi and Kenneth I. Wolpin. 1995. ``Duration to First Job and the Return to Schooling: Estimates from a Search-Matching Model," *Review of Economic and Statistics* 62, 263-286.

Ferrall, Christopher. 1997. ``Unemployment Insurance Eligibility and the Transition from School to Work in Canada and the United States," *Journal of Business and Economic Statistics 15*, 2, April, 115-129.

Finnie, Ross. 1998a. ``Process Not Event: A Longitudinal Analysis of the School-to-Work Transition of Canadian Post-Secondary Graduates," work in progress.

Finnie, Ross. 1998b. ``Holding Their Own: The Employment Patterns and Earnings of Canadian Post-Secondary Graduates in the 1980s and 1990s", work in progress.

Kiefer, N. 1988. "Economic Duration Data and Hazard Functions," *Journal of Economic Literature* 26, 649-79.

Lin, D. Y. and L. J. Wei. 1989. ``The Robust Inference for the Cox Proportional Hazards Model", *Journal of the American Statistical Association*, 84, 1074-78.

Statistics Canada 1996. *The Class Of 90: A compendium of findings from the 1992 National Graduates Survey of 1990 Graduates*, Catalogue No. 81-577-XPE.