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Transition from School to First Job: the influence of educational attainment

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TRANSITION FROM SCHOOL TO FIRST JOB: THE INFLUENCE OF EDUCATIONAL ATTAINMENT

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ABSTRACT

This paper investigates the transition from high school to first job using data from the *National Education Longitudinal Study 1988-2000*. A proportional hazards model is estimated to identify the determinants of time-to-first-job. In contrast to earlier studies, there is strong evidence of positive duration dependence after controlling for unobserved heterogeneity. Time-to-first-job is correlated with educational attainment and type of school program attended. Attending a vocational program reduces time-to-first-job, but dropouts who obtain the General Educational Development qualification do not improve their chances of getting a job more quickly. Family background is insignificant.

JEL: J24, J64

Key words: School-to-work transition, duration dependence, GED, school program

I. INTRODUCTION

This paper is concerned entirely with those youths proceeding directly to the labour market after leaving high school. About one in four youths in the USA leave school each year and do not proceed to post secondary education. The majority of these enter the labour market and begin their search for a job. The experience of these new entrants into the labour market differs substantially, however, between individuals. Some take longer to get a job than others. The reasons for these differences are worth exploring since the early labour market experience of school leavers can have long-lasting effects on subsequent lifetime outcomes. Previous research has shown that those school leavers who take longer to get a job have a higher future probability of being unemployed and are more likely to have lower future earnings (OECD 1998, Margolis *et al*. 1999).

There are several reasons why the early experience of youths in the labour market may affect subsequent outcomes. First, human capital theory suggests that substantial investment in human capital should be made during the early years of work. Joblessness during the early years of entry into the labour market may consequently be particularly costly since the non-accumulation of new, and the deterioration of existing, human capital adversely affects future employment and earnings. Second, sorting models imply that employers may use employment records as a signal of potential productivity in screening job applicants. Early employment experiences are observable and they could play an important role in determining the future employability and future earnings of youths. Third, dual labour market theory suggests that early non-employment might lead to poor work habits, weak labour force attachment, human capital deterioration and alienation from society. Experience of joblessness may alter the attitudes of unemployed youths if they become increasingly discouraged about their chances of finding a job, and this may result in less intensive job search.

This paper has three objectives. The first is to show that the time it takes to get a job after leaving high school has longer-term effects on labour market outcomes. The second objective is to investigate the potential determinants of the time it takes to get a job after entering the labour market for the first time. We are particularly concerned with the impact of an individual's educational attainment and type of educational program attended on time-to-first-job. Educational attainment is defined in terms of whether an individual has a high school diploma, a General Educational Development (GED) qualification, or no qualification at all.

Type of school program distinguishes between academic, vocational and other programs.[1] The third objective is to investigate whether the probability of getting a job after leaving school changes as the duration of job search lengthens. Specifically, we wish to know whether the probability of getting a first job is affected by the length of time a school leaver has been in a state of non-employment.

This paper therefore aims to provide a better understanding of the hazard of getting a job for those youths that decide not to proceed to college but to enter the labour market directly after high school. These issues are investigated by using data obtained from several sweeps of the National Education Longitudinal Study (NELS).[2] The remainder of the paper is as follows. Section II provides some new evidence that time-to-first-job is correlated with labour market outcomes several years after the end of high school. Section III reviews the empirical literature on factors influencing the time taken for school leavers to enter their first job. Section IV describes the method used to model time-to-first-job. Section V discusses the data obtained from the NELS and the construction of variables used in the empirical analysis. Section VI presents the results. Section VII concludes.

II TIME-TO-FIRST-JOB AND SUBSEQUENT LABOUR MARKET OUTCOMES

Previous research has shown that both incidence and duration of unemployment adversely affect the future probability of being in a job (Narendranathan and Elias 1993, Omori 1997, Mroz *et al*. 1999, Arulampalam *et al*. 2000). More importantly for school leavers, the length of time-to-first-job adversely affects subsequent employment experience (OECD 1998, Margolis *et al*. 1999). The OECD reports that 'irrespective of education or gender, getting a job in the first year after school is associated with a greatly increased likelihood of being employed at the moment of each subsequent annual interview compared with starting off without a job' (pp.105-106). There is also evidence that the time-to-first-job is associated with lower future earnings (Mroz *et al*. 1999, Margolis *et al*. 1999) and may even scar an individual's subjective well-being (Clark and Oswald 1994, Korpi 1997, and Clark *et al*. 2001). Further evidence that time-to-first-job affects labour market outcomes is available from the recently published fourth sweep of the NELS, which allows us to investigate the

[1] The proportion of students enrolled on these three programs in the National Education Longitudinal Survey was 0.679 (academic), 0.064 (vocational) and 0.257 (other).
[2] Source of survey: National Centre for Educational Statistics, US Department of Education, Office of Educational Research and Improvement (http://www.nces.ed.gov/surveys).

relationship between time-to-first-job and labour market outcomes for up to eight years after the end of high school.

Table 1 provides some indication of the differences in labour market outcomes (eight years after leaving school) between those who got a job quickly after leaving school and those who took much longer to get a job. Males who did not proceed to further education and who got a job within three months of leaving school had an unemployment rate, for example, of 3.7%. This compared to an unemployment rate of 9.5% for those who took over twelve months to get a job. The gap was even larger for females (7.4% compared to 17.2%). A very similar result is obtained for the *per cent* obtaining a full-time job: around 90% of males who got a job within three months of leaving school were in full-time employment compared to 77% for those who took over twelve months to get a job. Again, a similar gap is observed for females. Finally, the annual earnings of males who got a job within three months of leaving school is 20% greater than those who took over twelve months to get a job. The corresponding differential for females is over 40%.

These data suggest that getting a job soon after completing compulsory education is potentially advantageous in terms of its effect on labour market outcomes in later years. It is consequently of policy relevance to investigate the determinants of the time-to-first-job. In particular, we need to know how the probability of getting a first job changes as the search duration lengthens. This is the aim of the remainder of the paper.

III. DETERMINANTS OF TIME-TO-FIRST-JOB: A REVIEW OF THE LITERATURE[3]

Previous studies have attempted to discover whether the time-to-first-job is influenced by the duration of a school leaver's time spent in the state of non-employment. According to earlier studies (Chuang 1999, Nielsen *et al*. 2001), the longer that school leavers stay unemployed, the less likely they are to become employed. This negative duration dependence, however, may be a result of unobserved heterogeneity, for which previous studies have failed to control (Franz *et al*. 2000 and Nielsen *et al*. 2001). Andrews *et al*. (2002), for example, find no evidence of negative duration dependence after controlling for unobserved heterogeneity. In particular, they find that after the initial period there is no duration dependence at all.

[3] The transition to first job fits naturally into the literature on the transition into employment, which focuses on the hazard of re-employment (Devine *et al*. 1990). The transition from school to first job, however, has its own characteristics. For example, most of the labour market entrants are first-timers in the labour market, have no previous labour market experience, and are therefore not entitled to unemployment benefit.

In addition to considering how the probability of getting a job changes over time (i.e. hazard to first job), previous studies have identified factors that influence the time it takes to get a job after leaving school. A factor found to have a substantial influence on the time-to-first-job is educational attainment. Ferral (1997) finds that people with lower levels of educational attainment receive job offers less often than those with higher levels of attainment, leading to longer expected unemployment duration for those with lower levels of attainment. Similarly, findings from empirical studies based upon reduced-form models indicate that individuals with higher levels of educational attainment have a higher employment probability, and consequently a shorter period of unemployment (Nielsen *et al*. 2001, Dolton *et al*. 1994, Upward 1999, and Andrews *et al*. 2002).[4]

Gender and ethnicity are also related to unemployment duration. Several studies indicate that females take longer to find their first job than males (Eyland *et al*. 1989 Lassibille *et al*. 2001, Genda and Kurosawa 2000).[5] This earlier work therefore suggests that females and males should be analysed separately. The effect of ethnicity on the transition to first job have been investigated by Wolpin (1992), who finds that blacks have a higher probability of receiving a job offer than whites. In a later study, longer unemployment duration and lower accepted mean wage for blacks are explained as being due primarily to differential rates at which job offers are accepted rather than to differential job offer probabilities (Eckstein and Wolpin 1995). However, Bowlus *et al*. (2001) find that while black males have a similar reservation wage to white males, they face a significantly lower offer arrival rate while unemployed. This leads to the conclusion that the longer unemployment duration for blacks stems from the lower arrival rate of job offers while they are unemployed.

Several family background factors such as parents' occupation, employment status and education have also been investigated in previous studies as potential determinants influencing the hazard of exiting non-employment to a job (Dolton *et al*. 1994, Lassibille *et al*. 2001, Andrews *et al*. 2002, Franz *et al*. 2000, Nielsen *et al*. 2000, Betts *et al*. 2000). No consistent findings have been established, however.

[4] Although higher education may theoretically improve one's chance of receiving a job offer, it is difficult to determine the effect of education on employment probability *a priori* in reduced-form hazard models. This is because individuals with higher levels of education may have a higher reservation wage, rendering the net effect of higher education levels on the hazard of exiting to a job ambiguous. In the NELS, the mean time-to-first-job is 5.3 for high school graduates, 11.5 for those with a GED qualification and 13.9 for dropouts with no equivalent qualification.

[5] In the NELS, the mean time-to-first-job is 5.7 months for males and 8.6 months for females.

Schools might influence the transition of school leavers to jobs since the links between schools and employers may be closer for some schools than for others. High schools and employers are in an interdependent relationship in which employers depend on schools to supply educated workers, and schools depend on employers to hire their graduates. One potentially important determinant of time-to-first-job that has not been investigated in previous studies is the influence of the type of school program attended. This omission is rectified in the present study by distinguishing between three broad types of school program: a general program, an academic program and a vocational program.[6] We also distinguish between public and private (including Catholic) schools.

IV A FRAMEWORK FOR ANALYSIS

Standard job search theory provides a useful framework for the analysis of duration to first job. Under the job search model, individuals start their job search process immediately on leaving school. The conditional probability that an individual school leaver will be employed in a given interval (t, t + dt) is the product of two probabilities – the probability of receiving a job offer, and the probability of then accepting it. The latter is the probability that the wage offer exceeds the reservation wage.

The hazard of exiting non-employment to first job can vary over the spell either due to changes in the offer probability or due to changes in the reservation wage. The offer probability, for example, may change with the length of the unemployment spell. The individual's search intensity may change or employers may be wary of employing those who have been unemployed for a long time. The probability of receiving a job offer will be determined by factors that make an individual school leaver more attractive to a potential employer such as education, training, personal characteristics and local labour market conditions. After having received a job offer, the probability of then accepting it depends on the individual's reservation wage. The determination of the reservation wage will depend upon the offer probability, leisure preferences, the expected wage offer distribution and the cost of continuing the search. The reservation wage may also change due to changes in the behaviour of the searcher. When the offer probability is very low, the individual may lower the reservation wage to the extent that the first offer that comes along will be accepted.

[6] In the NELS, the mean time-to-first-job (in months) is 5.1 for those attending a vocational program, 5.3 for those attending an academic program and 9.8 for those attending a normal program.

Since duration is measured in months, the discrete time hazard model is appropriate (Prentice and Gloecker 1978, Meyer 1990, 1995). We organise the data into 'sequential binary form' in which individual i contributes T_{i} observations, where T_{i} is the number of periods individual i searches for a job before entering the first job. The probability of a spell being completed by t+1, given that it was still continuing at t, is given by the discrete time or grouped hazard:

$$h_j(t) = 1 - \exp\{-\exp[(\mathbf{x}_i'\boldsymbol{\beta}) + \gamma_j(t)]\} \tag{1}$$

where $\gamma_j(t) = \int_{t_{j-1}}^{t_j} h_0(u)\,du$ is the baseline hazard.

In our analysis, we estimate the baseline hazard nonparametrically to avoid possible misspecification (Meyer 1990; Han and Hausman 1990).[7] The likelihood is then given by

$$L(\beta,\gamma) = \prod_{i=1}^{n}\prod_{j=1}^{J} h_j(t\,|\,\mathbf{x})^{y_{ij}}\,[(1-h_j(t\,|\,\mathbf{x})]^{1-y_{ij}} \tag{2}$$

It is well known that failure to control for unobserved heterogeneity may bias the baseline hazard as well as the estimated effects of the covariates (Heckman and Singer 1984, Lancaster 1990). Following standard practices, we allow for unobserved heterogeneity with a positive-valued random variable v. The hazard function will become

$$\begin{aligned} h_{ij}(t) &= 1 - \exp[-v_i\ \exp(\mathbf{x}_i'\boldsymbol{\beta} + \gamma_j(t))] \\ &= 1 - \exp[-\exp(\mathbf{x}_i'\boldsymbol{\beta} + \gamma_j(t) + u_i)] \end{aligned} \tag{3}$$

where $u = \log v$ with density $f_u(u)$. The amended likelihood can then be written as

$$L(\boldsymbol{\beta},\gamma) = \prod_{i=1}^{n}\int_{-\infty}^{+\infty}\left[\prod_{j=1}^{t} h_j(\mathbf{x}_i', u)^{y_{it}}[1-h_j(\mathbf{x}_i', u)]^{1-y_{it}}\right] f_u(u)\,du \tag{4}$$

where $h_j(\mathbf{x}_i', u_i) = 1 - \exp[-\exp(\mathbf{x}_i'\boldsymbol{\beta} + \gamma_j + u_i)]$.

[7] The transition from school-to-work, particularly the time-to-first-job, has been analysed in the context of a structural search model (Bowlus *et al*. 2001, Wolpin 1987, Eckstein *et al*. 1995, Ferrall 1997). The method of estimation for structural job search models is still in its infancy (Ferral 1997). Estimation of structural search models also has very heavy data demands. Not only are data on the duration of the search required to estimate such models but also data on the accepted wage (Wolpin 1987, 1992), the subsequent transitions (Bowlus *et al*. 2001), and for both firms and workers (Eckstein *et al*. 1995).

Two approaches have been used to model the unobserved heterogeneity. The first is to assume a particular parametric distribution for the heterogeneity term, such as the frequently used Gamma or Gaussian distributions. Once a particular distribution function is assumed, the unobserved heterogeneity term can be integrated out of the likelihood function. The advantage of the Gamma distribution is its analytical tractability since it yields a closed form solution (Lancaster 1979, Meyer 1990). The Gaussian distribution is justified on the grounds that the heterogeneity term might capture the large number of unobserved characteristics (Narendrenathan and Stewart 1993; Stewart 1996).[8] The problem with specifying a parametric distribution for the heterogeneity term is that the estimated parameters may be sensitive to the particular distribution adopted, especially where the baseline hazard is not sufficiently flexible. An alternative approach suggested by Heckman and Singer (1984) is to use the mass point technique, which approximates a continuous distribution by a finite discrete distribution of unrestricted form. The u_i and $f_u(u_i)$ are then approximated nonparametrically by a discrete distribution whose mass point locations and associated probabilities are to be estimated together with other parameters in the model.[9] Since economic theory is not informative as to the functional form of the heterogeneity term, we adopt both approaches and estimate three mixing models, the Gamma, the Gaussian and the nonparametric.

V. DATA AND VARIABLES

The analysis of the transition from school is based on data from the National Education Longitudinal Study (NELS), from the base year (1988) through the third follow-up (1994). The NELS is a nationally representative sample of 14,915 students in the USA that provides a detailed history of the transition to first job for a sample of 3687 youths. Since those pursuing a college education delay their entry into the labour market, these individuals are not included in our analysis. Our dependent variable is the duration of the initial spell of non-employment after leaving school, or more simply time-to-first-job.

The hazard of exiting non-employment to first job is determined by the joint probability of receiving and accepting a job offer. Economic theory suggests a variety of factors that may influence either the offer probability or the reservation wage. Personal characteristics such as

[8] Although the Gaussian mixing distribution does not yield a closed form solution, it can be easily approximated by the widely available Gaussian-Hermite quadrature procedure. See Stewart (1996) for derivation of the likelihood function.

age, gender, ethnicity and education may influence the hazard of exiting non-employment. In our analysis, gender differences are explored by estimating separate models for males and females; and five ethnic groups are distinguished (whites, blacks, Asians, Hispanics and American Indians).

The potential impact of educational attainment on the duration of non-employment spells is uncertain since education and training can affect the probability of receiving a job offer as well as the reservation wage. Those with higher levels of attainment may get more job offers but are also likely to have a higher reservation wage. Education level is measured here by including two indicators of educational qualifications. These are the traditional high school diploma and the GED qualification, which can be obtained by those who fail to graduate from high school. We know that those individuals who obtain the GED are essentially high school dropouts, and it is therefore of interest to see if the GED has any beneficial effect on improving the chance of getting a job.[10] This is potentially interesting since employers may want more training and skills than is indicated by the GED.[11]

The potential effect of schooling on the transition from school to work implies the need to include some school level variables in the model. The available data allow us to test for the influence of two school characteristics. First, schools can be classified as public or private (including Catholic schools). Do students from private schools have an advantage over those from public schools in the search for a job? Second, and more importantly, students attend three distinct types of education program: academic, vocational and other. We expect students who have been enrolled on a vocational program and who go straight into the labour market on leaving school to have an advantage over students enrolled on other programs. Schools offering vocational programs, for example, are more likely to have close links with local employers, thus giving students on such programs an advantage over students a on non-vocational program. Furthermore, students on a vocational program are likely to have a

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[9] Since the maximum likelihood estimation of this discrete mass point method is non-standard and computationally demanding, in the empirical analysis reported below we follow Heckman *et al*. (1990) to fix the number of mass points at two.

[10] Cameron and Heckman (1993) have found that GED earners are not distinguishable from high school dropouts in terms of earnings.

[11] Sorting models imply that potential employers may use educational qualifications to screen job applicants. Some high school dropouts may send a signal to potential employers by obtaining a GED qualification to distinguish them from their fellow dropouts.

clearer idea of what they want to do on leaving school and may even have had some experience in working with a local employer as part of their course.[12]

Family background may influence the hazard of exiting non-employment through its effects on both the reservation wage and the probability of receiving job offers. School leavers whose parents are unemployed or have low educational attainment, for example, may have a lower probability of receiving a job offer because of a lack of information and poor networking within the job market. This is a typical problem for low-income youths (Holzer 1987, Rosenbaum 1999). We also control for family size and family structure. The number of siblings, for example, can influence the intensity of job search through competition for scarce parental time.

A further group of factors that may be expected to influence the time-to-first-job relate to the geographical labour market in which a school leaver is located. Youths living in urban areas are expected to have a higher hazard rate of leaving non-employment than those in rural areas because job opportunities are likely to be more readily available, though there may be specific problems in inner city areas (Ihlanfeldt *et al*. 1998). Finally, macroeconomic conditions may also influence the employment probability. We therefore control for recruitment cycles and the business cycle by including dummy variables for the quarter and the year that school leavers exit to first job. The appendix provides a list and description of the variables used in the estimation of the models.

VI. RESULTS AND DISCUSSION

5.1 Baseline hazard

Four models have been estimated. The first is a homogeneous proportional hazards model with a nonparametric baseline hazard for males and females separately. We then estimated three mixed proportional hazards models with different distributional assumptions (i.e. Gaussian, Gamma and HS nonparametric mixing) for the unobserved heterogeneity. Due to data thinning, time intervals are grouped into two-month periods for intervals from 25 to 34, and all later periods are grouped into a single final period. The baseline hazard therefore consists of (i) the first 24 monthly periods, (ii) the next 5 two-monthly periods, and (iii) the

[12] Stratified educational programmes are less common in the USA compared with countries such as Germany, Switzerland and the Netherlands, where students are separated early on into academic and vocational tracks. In the USA, tracking begins at a later age, the curricula of various tracks are similar, and there is mobility between tracks (Muller and Shavit 1998).

final period. Since the shape of the baseline hazard for the three heterogeneous models is very similar, and since the HS nonparametric mixing model gives the largest log-likelihood values, we discuss only the HS model here. The baseline hazard for the first 24 periods are plotted for the homogeneous model and the HS mixing model in Figures 1-3.[13]

The shape of the homogeneous baseline hazard indicates some evidence of negative duration dependence for the first twelve months after leaving school, particularly for males (Figures 1 and 2). This is in common with findings reported by other studies that do not control for unobserved heterogeneity (Upward 1999, Nielsen 1999). After the first twelve months, no particular pattern can be determined for the baseline hazard. We also detect a spike in the second month after leaving school for both males and females. This is consistent with findings in the literature that the hazard of leaving non-employment is very high during the first few periods after leaving school.

After controlling for unobserved heterogeneity, however, distinctly different results are obtained. The baseline hazard now indicates positive duration dependence for the first twelve to fourteen months for both males and females. After that, the hazard varies from one period to the next and does not exhibit an obvious trend. This result contrasts with evidence of negative duration dependence reported in Nielsen *et al*. (2001) and Franz *et al*. (2000). Our results here are consistent with findings by Andrews *et al*. (2002), who report that after controlling for unobserved heterogeneity there is no negative duration dependence.

A further finding is that the hazard increases considerably after the first 14 months. This is consistent with the view that after a long period of searching unsuccessfully, the searcher becomes more willing to accept a job offer, as indicated by a reduction in the reservation wage as unemployment duration increases (Mortensen 1986). In the extreme, the individual may accept the first job offer that comes along when the likelihood of receiving an offer is very low (Narendranathan and Stewart 1993). The decline in the probability of receiving a job offer may occur due to the scarring effect of unemployment or due to the depreciation of human capital resulting from an increase in unemployment duration. Furthermore, the extra utility obtained from being unemployed (leisure) may become negligible and possibly negative because of the disutility arising from the social stigma attached to being unemployed and because of the debilitating effects of being unemployed for long periods (Nickell 1979).

[13] Estimates of the baseline hazards for all model specifications are available from the authors on request.

The evidence of positive duration dependence observed in Figure 3 therefore seems plausible in light of these arguments.

5.2. Effects of covariates

The estimated results for the homogeneous model and the nonparametric mixing model are provided in Tables 2 and 3 for males and females respectively. The results obtained from the homogeneous model are included to highlight the fact that the estimated risk ratios differ substantially in some cases from those estimated by the HS mixing model. We focus here on the results that are likely to have the greatest policy relevance.

Schooling

The results in Tables 2 and 3 suggest that schooling may affect time-to-first-job in two ways. First, attendance at a Catholic or private school reduces the risk of exiting to a job for males by over 30 *per cent* points, but has little impact on females. The lower risk of exiting to first job for males may be a consequence of less pressure to find a job quickly for those who can afford to finance private education.

The type of training students received during their high school years is highly significant in improving their chances of getting a job. Using attendance in the 'other' program as the base group, we found that those attending a vocational program had a much greater chance of getting a first job (46 *per cent* higher for males and 65 *per cent* higher for females). These results are consistent with previous findings that attending a training scheme or vocational program is beneficial for the participants in their transition to the labour market. The effect of attending an academic program is somewhat lower than that of attending a vocational program (38 *per cent* higher for males compared to 31 *per cent* higher for females).

Educational qualifications

Individuals having a traditional high school diploma are found to have a far higher hazard of exiting non-employment to first job than those without any qualification. The estimated parameters are large in magnitude and highly significant. For males, having obtained a traditional high school diploma improves a person's chance of exiting to a first job by nearly fourteen times compared to dropouts with no qualification. For females with a high school diploma, the probability is somewhat lower (eight times) but is still very high compared to dropouts.

A more striking result, however, is the effect of the GED. For females, having a GED reduces the hazard of exiting non-employment by twenty-four *per cent* but the estimated coefficient is not statistically significant. For males, a GED reduces the hazard of exiting non-employment by nearly 30 *per cent* compared to those with no qualification. This finding is pertinent because the proportion of young adults who complete secondary education by passing the GED examination has increased markedly over the last 30 years (Murnane *et al*. 2000). GED is the one educational credential that many school dropouts obtain (approximately half a million a year). One of the purposes of the GED is to give high school dropouts a second chance of completing their secondary education with the expectation that the qualification will help in their transition from school to work. Given the results obtained here, the policy of promoting the GED appears questionable, at least in terms of its effect on time-to-first-job.

The implication of these findings is that policy designed to improve the transition from school to work needs to address the roots of the problem. This means reducing the probability of dropping out of high school and providing students with the skills and knowledge necessary in the work place rather than simply providing dropouts with a post-school qualification. The current practice of lowering the standards adopted to make it easier to complete high school education may not help high school dropouts in their transition to work. Murnane *et al*. (2000) show that the GED benefits only a small group of dropouts who left school with very low academic skills. They also point out that the existence of the GED may induce some high school students to drop out and acquire the alternative credential. This is certainly a potential disadvantage of the GED.

VII. CONCLUSION

Getting a job soon after completing secondary education has been shown to be highly significant in its potential effect on a person's future labour market outcomes. Specifically, those school leavers entering the labour market who get a job quickly after leaving high school are less likely to be unemployed eight years later. They are also more likely to be in a full-time job and to be earning more than those who take longer to get a job after leaving school.

These adverse effects on labour market outcomes over the longer term indicate a need to gain a better understanding of the factors determining the time-to-first-job. This paper has investigated the transition from school by using data from the National Education

Longitudinal Study (NELS). A proportional hazards model with flexible baseline hazards has been estimated. We have also controlled for unobserved heterogeneity. The main findings are as follows.

First, after controlling for unobserved heterogeneity we do not find the initial spike for the first period in the baseline hazard reported in previous studies. Neither do we find evidence of negative duration dependence. On the contrary, we find evidence of positive duration dependence. The underlying trend in the hazard of exiting non-employment to a job increases during the first eighteen months after leaving high school.

Second, we confirm the importance of education and training on the hazard of exiting non-employment to first job for school leavers. School leavers with a traditional high school diploma are found to have a substantially higher hazard of exiting to a first job than is the case for dropouts. GED recipients, on the other hand, are found to be even worse off than dropouts who do not acquire the GED. This finding calls into question the value of the GED, at least in terms of its value in helping school leavers to get their first job. Indeed, failure to get a job may induce dropouts to study for the GED.

Third, we find strong evidence that those enrolling in a vocational program at school have a much higher probability of exiting to first job on leaving school than those enrolled in other programs. There is also strong evidence that enrolling in an academic program at school substantially improves the chance of getting a job on leaving school. Students enrolled in 'other' school programs (37% in our sample) are at a disadvantage in terms of their chances of exiting to first job compared to those in vocational and academic programs.

Finally, we do not find any evidence that ethnicity or family background have any effect on the hazard of exiting non-employment to first job.

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TABLE 1. Time-to-first-job: earnings and employment status eight years after leaving high school

Time-to-first-job (in months)		Sample size	Annual earnings (dollars)	In full-time job (%)	Unemployed and wanting a job (%)
Males	0 -3 months	1299	30070	89.5	3.7
	4 - 12 months	279	26920	84.2	7.2
	Over 12 months	304	25150	77.3	9.5
Females	0 -3 months	1015	16810	66.5	7.4
	4 - 12 months	301	14770	59.5	9.6
	Over 12 months	453	11700	53.4	17.2

Note: The sample includes only those who left high school but did not proceed to further education.
Source: National Educational Longitudinal Study.

Table 2 Homogeneous model versus HS heterogeneity model: males

	Homogeneous model			HS heterogeneity model		
	Coef.	Risk Ratio	P value	Coef.	Risk Ratio	P value
Qualifications						
Diploma	1.297	3.659	< 0.000	2.626	13.814	< 0.000
GED	-0.136	0.873	0.250	-0.330	0.719	0.026
School						
Catholic/private school	-0.282	0.754	0.018	-0.372	0.690	0.004
Academic program	0.611	1.842	<0.000	0.320	1.376	< 0.000
Vocational program	0.651	1.917	<0.000	0.381	1.463	< 0.000
Variance				0.493		
Mass point 1- location				-3.343		
Mass point 1- probability				0.042		
Mass point 2- location				0.147		
Mass point 1- probability				0.957		
No. of observations	1917			1917		
No of person-period observations	11864			11864		
Log-likelihood	-3559			-3415		

Note: The variables used as controls are given in the appendix. These include: age, ethnicity, number of siblings, family structure, parents' educational attainment, parents' employment status, parents' occupation, region of residence, stage of recruitment cycle. Full results are available on request to the authors.

Table 3 Homogeneous model versus HS heterogeneity model: females

	Homogeneous model			HS heterogeneity model		
	Coef.	Risk Ratio	P value	Coef.	Risk Ratio	P value
Qualifications						
Diploma	1.211	3.358	< 0.000	2.125	8.372	< 0.000
GED	-0.116	0.890	0.368	-0.273	0.761	0.150
School						
Catholic/private school	0.119	1.126	0.343	-0.066	0.936	0.616
Academic program	0.525	1.691	< 0.000	0.271	1.311	< 0.000
Vocational program	0.640	1.896	< 0.000	0.502	1.653	< 0.000
Variance				0.407		
Mass point 1- location				-2.649		
Mass point 1- probability				0.055		
Mass point 2- location				0.154		
Mass point 1- probability				0.945		
No. of observations	1770			1770		
No of person-period observations	15628			15628		
Log-likelihood	-3515			-3432		

Note: The variables used as controls are given in the appendix. These include: age, ethnicity, number of siblings, family structure, parents' educational attainment, parents' employment status, parents' occupation, region of residence, stage of recruitment cycle. Full results are available on request to the authors.

Appendix: Variables used in the estimation of the hazard functions

Variable	Description	Males Mean	Males Standard Deviation	Females Mean	Females Standard Deviation
White*	Racial origin = white	0.684	0.465	0.660	0.474
Black	Racial origin = black	0.109	0.312	0.112	0.315
Asian	Racial origin = Asian	0.046	0.211	0.040	0.196
Hispanic	Racial origin = Hispanic	0.143	0.350	0.167	0.373
American Indian	Racial origin = American. Indian	0.017	0.130	0.022	0.145
Born in 1972	Born in 1992 or earlier	0.112	0.316	0.067	0.251
Born in 1973	Born in 1973	0.386	0.487	0.316	0.465
Born in 1974/75*	Born in 1974 or 1975	0.502	0.500	0.616	0.486
Sibling 0*	No sibling	0.051	0.219	0.044	0.205
Sibling 1	One sibling	0.278	0.448	0.254	0.435
Sibling 2	Two siblings	0.255	0.436	0.246	0.431
Sibling 3	Three siblings	0.186	0.389	0.176	0.381
Sibling 4	Four siblings	0.231	0.421	0.280	0.449
Two -parent family*	Two parents family	0.627	0.484	0.605	0.489
Partner-parent family	Mother-partner or father-partner family	0.192	0.394	0.203	0.403
Single parent family	Single parent family	0.182	0.386	0.192	0.394
Parent not high school graduate*	Parents not graduated from high school	0.162	0.368	0.202	0.402
Parent high school graduate	Parent school graduate	0.313	0.464	0.318	0.466
Parent some college	Parent some college education	0.405	0.491	0.376	0.485
Parent college education	Parent college graduate	0.082	0.275	0.066	0.248
Parent masters/PhD	Parent higher degree	0.038	0.191	0.038	0.192
Parents unemployed	Both parents unemployed	0.058	0.234	0.050	0.217
Mother professional / manager	Mother professional or manager	0.093	0.291	0.098	0.298
Mother non-manual worker	Mother occupation in non-manual group	0.476	0.500	0.471	0.499
Mother manual worker	Mother occupation in manual group	0.028	0.164	0.034	0.181
Mother other*	Mother unskilled, semi-skilled, at home	0.403	0.491	0.397	0.490
Father professional / manager	Father professional or manager	0.124	0.330	0.128	0.334
Father non-manual worker	Father occupation in non-manual group	0.168	0.374	0.164	0.371
Father manual worker	Father occupation in manual group	0.191	0.393	0.185	0.389
Father other*	Father unskilled, semi-skilled, at home	0.517	0.500	0.523	0.500
Catholic / private school	Attended Catholic or private school	0.053	0.225	0.054	0.227
Public school*	Attended public school	0.947	0.225	0.946	0.227
Academic program	Attended academic program	0.493	0.500	0.492	0.500
Vocational program	Attended vocation program	0.141	0.349	0.107	0.309
Normal program*	Attended normal program	0.366	0.482	0.402	0.490
Respondent has HS diploma	Normal high school diploma	0.803	0.398	0.788	0.409
Respondent has GED	General Educational Development	0.073	0.260	0.076	0.266
Respondent has no qualification*	No qualification	0.124	0.330	0.136	0.343
Respondent lives in urban area	Urban area	0.241	0.428	0.243	0.429
Respondent lives in suburbs	Suburb area	0.374	0.484	0.373	0.484
Respondent lives in rural area *	Rural area	0.386	0.487	0.384	0.487
Respondent lives in South	Southern state	0.374	0.484	0.387	0.487
Year 1990	Got job in 1990	0.038	0.191	0.032	0.175
Year 1991	Got job in 1991	0.047	0.212	0.041	0.199
Year 1992*	Got job in 1992	0.780	0.414	0.679	0.467
Year 1993/94	Got job in 1993/1994	0.135	0.341	0.248	0.432
Quarter 1	Got job in first quarter	0.054	0.226	0.080	0.271
Quarter 2*	Got job in second quarter	0.703	0.457	0.599	0.490
Quarter 3	Got job in third quarter	0.111	0.314	0.109	0.312
Quarter 4	Got job in fourth quarter	0.077	0.266	0.102	0.303
No. of respondents		1917		1770	
No. of person-months		11864		15628	

Note: * indicates the excluded categories (i.e. the base group) in the proportional hazards models reported in Tables 2 and 3.

Figure 1. Baseline hazard, females

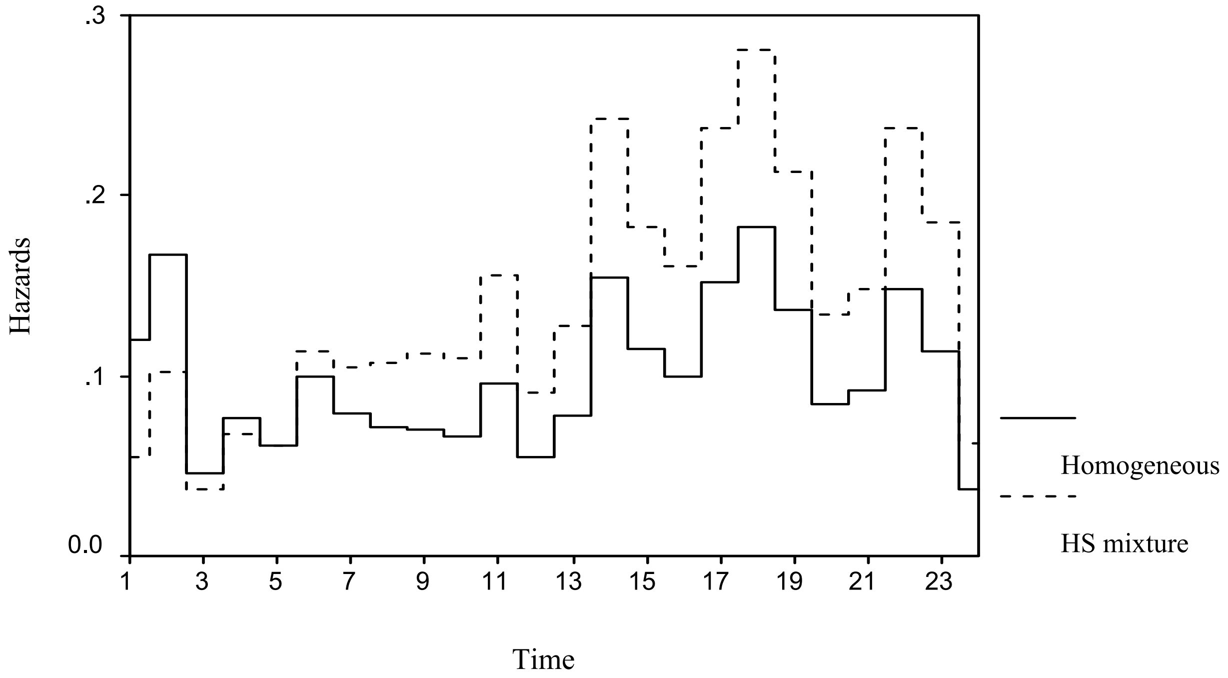

Figure 2 Baseline hazard, males

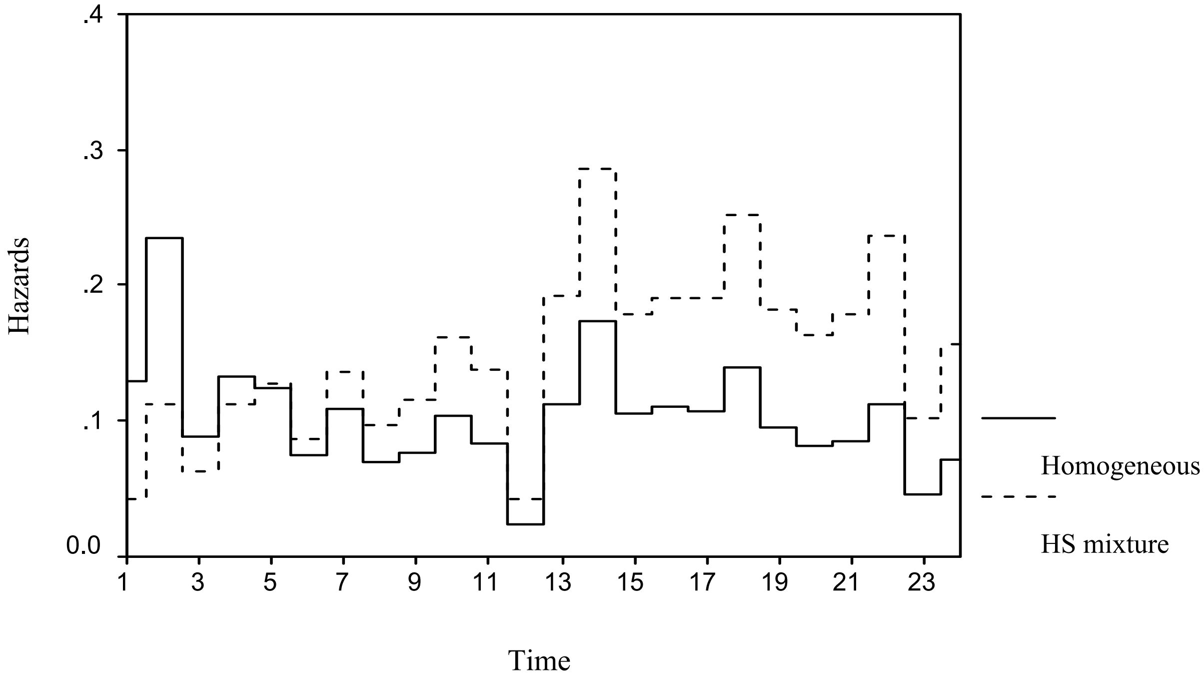

Figure 3 HS mixing models, baseline hazards