

---

**WORKING PAPER SERIES**

---

**NO. 66**

**DURATION DEPENDENCE AND ROUTES OUT OF  
JOBLESSNESS FOR YOUNG PEOPLE**

**DUNCAN McVICAR and JAN PODIVINSKY**



**NORTHERN IRELAND  
ECONOMIC RESEARCH CENTRE**

# **Duration Dependence and Routes Out of Joblessness for Young People**

**Duncan McVicar\* and Jan M. Podivinsky\*\***

\*Northern Ireland Economic Research Centre, Queens University Belfast  
46-48 University Road, Belfast BT7 1NJ  
Tel: +44 (0) 28 9026 1807  
Fax: +44 (0) 28 9043 9435  
d.mcvicar@qub.ac.uk

\*\*Department of Economics, University of Southampton  
Southampton SO17 1BJ  
j.m.podivinsky@soton.ac.uk

September 2001

The support of the Northern Ireland Departments of Enterprise, Trade and Investment, and of Finance and Personnel, through NIERC's Human Resources and Economic Development Research Programme, is gratefully acknowledged. Thanks also to the Northern Ireland Training and Employment Agency for supporting the collection of the original survey data on which this paper is based. Finally, thanks to Aidan McMahon for excellent research assistance and Stephen Jenkins for helpful advice. All views are those of the authors.

## **Abstract**

This paper examines young people's exits from joblessness using recent survey data for Northern Ireland. A reduced form search model is estimated, allowing for a fully flexible specification of the baseline hazard function, with young people exiting joblessness into either employment or education and training. There is evidence of negative duration dependence in the conditional probability of exiting joblessness. The relationship is not monotonic, however, there being a small 'spike' in the hazard rate around 5-6 months consistent with an effect of the introduction of the New Deal for Young People in 1998.

The paper argues that, given negative duration dependence, policy makers would do better to increase targeted interventions on those facing the toughest barriers to employment in the first month of joblessness, where possible, rather than wait for six months as is the current case with New Deal. The main barriers to exit from joblessness observable in the data include caring for a young child, having a family background of unemployment and low levels of job search. Having poor qualifications also acts as a barrier to exit from joblessness, particularly into education. The introduction of Job Seekers Allowance in 1996 is correlated with shorter jobless spells. Further research is suggested to examine these issues in more detail, and in particular to examine in greater depth the effects of the introduction of the New Deal on duration dependence in joblessness among young people.

## 1. Introduction

Throughout the 1980s and early 1990s unemployment amongst young people in the UK stood at record high levels. In recent years, in the context of a new social inclusion policy agenda, policy makers have come to recognise unemployment, and particularly long-term unemployment, as a key mechanism by which young people may become socially excluded.<sup>1</sup> The need to arrest the drift into long-term unemployment for many young people lies behind the UK government's recent introduction of a major active labour market policy for young people (the New Deal). This follows earlier interventions such as the removal of unemployment benefits for under-18s, the guarantee of a training place for all school leavers, and the tightening of benefit eligibility for all unemployed persons following the introduction of the Job Seeker's Allowance (JSA). These policy concerns and interventions have all extended to Northern Ireland (NI) also.

A major question still not adequately resolved in the applied labour economics literature is whether young people become less likely to find employment the longer they find themselves in unemployment. In other words, is there negative duration dependence in youth unemployment? Such duration dependence may result from skills deterioration or dwindling motivation for job search (e.g. Phelps, 1972; Layard *et al.*, 1991), or because potential employers see unemployment as a signal of poor productivity (e.g. Blanchard and Diamond, 1994). The lack of a clear consensus is partly the result of difficulties in distinguishing 'true' duration dependence from 'spurious' duration dependence. In short, 'true' duration dependence (unemployment causes unemployment) may be observationally equivalent to unobserved individual heterogeneity that increases individuals' propensity to be unemployed at whatever point in time. People may be long-term unemployed because they have poor job prospects rather than people having poor job prospects because they are long-term unemployed (see Heckman and Borjas, 1980; Machin and Manning, 1999).

---

<sup>1</sup> "The best defence against social exclusion is having a job..." (Tony Blair, Social Exclusion Unit, 1999).

The existence or otherwise of duration dependence in unemployment for young people has important policy implications, particularly for the New Deal. New Deal is geared towards intervention after six months of unemployment and claiming benefits, offering individually tailored guidance and support in job search; education, training or subsidised employment placements; and voluntary work and ‘environmental taskforce’ options (see Van Reenen, 2001, for a detailed account). If there is no duration dependence, however, then there is no deterioration in young peoples’ employability during unemployment, and scarce resources might be better spent *ex ante* on making sure young people leave school with marketable skills and qualifications or on regeneration of disadvantaged communities than on *ex post* intervention. Equally, if there is negative duration dependence, why wait six months to intervene? Skills and motivation will have already been lost that may have been maintained given earlier intervention. Presumably intervention has to be rationed for resource reasons or because policy makers are not confident they can identify those most at risk *ex ante*. A better policy might strengthen intervention in the first month for all those thought to be at risk of long-term joblessness and maintain a New Deal type intervention after six months as a safety net. Positive duration dependence would reduce the need for intervention, since the unemployed would become more likely to obtain employment the longer their unemployment spell. Finally, matters may be complicated further if duration dependence is non-monotonic. For example, if duration dependence is positive for the initial months before becoming negative, it may be optimal to intervene earlier or later than six months, depending on where the turning point is (or how long it takes an unemployed young person to begin to lose ‘employability’).

Following a recent trend in the literature, this paper examines duration dependence, in our case in joblessness among young people, using the Prentice-Gloeckler non-parametric methodology allowing for non-monotonicity in the duration relationship (Prentice and Gloeckler, 1978). In other words, we do not restrict duration dependence to be either positive or negative, but allow for different signs at different durations. This method is less prone to bias from spurious duration dependence due to unobserved heterogeneity than traditional parametric approaches (Narendranathan and Stewart, 1993). Given the key role of post-16 education and training in the choice set available to many young jobless, and in policy interventions aimed at helping young

jobless, exits from joblessness into employment and into education and training are identified separately in a competing risks framework, alongside a ‘pooled exits’ model. By using these techniques the paper aims to improve understanding of the nature of duration dependence in joblessness for young people and to enable policy interventions to be made more effectively.

The paper uses recent longitudinal survey data that follow a cohort of young people in NI from when they are first eligible to leave school in the summer of 1993 until the summer of 1999. This survey data allows us to focus on the wider definition of joblessness (all those young people not in employment, education or training) in preference to the narrower definitions of unemployment given by either ILO or the claimant count. The distinction between time spent ‘other jobless’ and time spent unemployed may be small for young people (Lynch, 1989). Other studies argue that joblessness and not unemployment *should* be the focus of research on young people (e.g. Rees, 1986). Hannan (1998) and Portugal and Addison (2000) are recent examples of duration analyses examining joblessness in preference to unemployment.

Our interest in the NI data in particular stems from three observations. First, NI displays relatively high levels of joblessness, particularly long-term joblessness, among young people, compared to most other UK regions (McVicar, 2000). Second, the Good Friday Agreement has brought with it devolved government, and this implies a need for research specific to NI if policy is to be evidence-based. Despite a considerable unemployment duration analysis literature in GB, the US, and elsewhere, we know of no existing study for NI. Third, our sample period spans the introduction of two major policy initiatives aimed at improving the employment prospects of young jobless – New Deal and JSA. As far as the data allow, we examine the effects of these policy interventions on joblessness duration.

We find evidence of negative duration dependence in the conditional probability of exiting joblessness into employment and in exiting joblessness overall. The relationship is not monotonic, however, there being a small ‘spike’ in the hazard rate around 5-6 months. We argue that this is consistent with a New Deal effect, where both entry to the New Deal and *expected* entry to the New Deal stimulate additional efforts to exit joblessness on the part of young people. Data limitations prevent a

definitive examination of the *overall* effects of the introduction of New Deal on young peoples' joblessness duration, however. JSA is unambiguously correlated with shorter joblessness spells for young people. A number of other factors are identified that act as barriers to employment for young people, including a family background of unemployment, having young children (for young women), and low levels of job search. Poor qualifications also act as a barrier to escaping joblessness, particularly into education or training.

The remainder of this paper is set out as follows. The following section discusses the framework behind the study (a simple model of job search) and discusses the approach to its estimation in the light of the existing empirical literature on unemployment duration. Section 3 outlines the data and presents descriptive statistics. Section 4 presents and discusses the results from estimation of our empirical model. Section 5 concludes with policy implications and suggestions for further research.

## 2. The Framework

Like most of the existing empirical literature, search theory provides the starting point for our model of joblessness duration. Jobless young people search for jobs, which entails costs, at varying rates. Given our broad definition of joblessness, for some, e.g. those caring for young children or those taking a ‘year out’ on holiday before continuing education, this search rate may be close to zero. The probability of a ‘match’ at a given point in time (exiting joblessness into employment) is the product of the probability of getting a job offer and the probability of accepting that job offer (Mortensen, 1987). The probability of getting an offer is likely to depend on factors such as search intensity, individual and local labour market characteristics. The probability of accepting an offer is likely to depend on the level of the wage offer relative to the individual’s reservation wage, which in turn may depend on factors such as income during unemployment, past and expected earnings and other individual characteristics.

We examine exits from joblessness into employment. Education and government supported training, however, given often extended and complex transition patterns from school to work, offer another possible route out of joblessness for young people. Our search framework must therefore be extended to allow young people to search for education and training places alongside the search for employment places. The probability of getting an offer of an education or training place may depend on existing qualifications, amongst other things. The probability of accepting such an offer is likely to depend on financial considerations, the suitability of the education or training place (e.g. where it is, the level of the qualification on offer) and some optimising behaviour over expected present value utility of accepting the place or continuing to search. Equation 1 shows the probability of exit from joblessness for individual  $i$  at time  $t$  implied by this framework. We go on to estimate a reduced form of this model.

$$P_{it}(\text{exit}) = \sum_{j=1,2} \phi_{itj} \eta_{itj}, \quad (1)$$

where  $j = 1$  (employment) or  $2$  (education/training) and  $\phi$  and  $\eta$  denote the probability of receiving and accepting an offer respectively, defined below.

$$\phi_{itj} = f_j(X_{it}, E_{it}, \varphi_{ij}), \quad \eta_{it1} = g_1(w_{it}(X_{it}, b_{it})), \quad \eta_{it2} = g_2(q_{it}(X_{it}, b_{it})), \quad (2)$$

where  $X_{it}$  is a vector of individual characteristics,  $E_{it}$  are environmental characteristics (e.g. local unemployment rate),  $\varphi$  denotes search intensity,  $w_{it}$  and  $q_{it}$  denote the wage offer relative to the reservation wage and the ‘course offer’ relative to some ‘reservation course’ respectively, and  $b_{it}$  denotes income during the jobless spell.

Standard search theory implies that the probability of getting a job offer is independent of joblessness duration. Various extensions, however, can predict either positive or negative duration dependence. If reservation wages fall with unemployment duration, perhaps because of benefit exhaustion, then job offers will be accepted more frequently and we might see positive duration dependence (e.g. Mortensen, 1977). Conversely, Layard *et al.* (1991) suggest that as the unemployment spell goes on, the unemployed worker might lose motivation for job search and this might reduce the probability of getting a job offer implying negative duration dependence. Phelps (1972) suggests that human capital might depreciate during an unemployment spell, also reducing the chances of getting a job offer. Blanchard and Diamond (1994) suggest that employers might see unemployment as a negative productivity signal and be reluctant to hire previously unemployed workers. These arguments may, at least to an extent, also apply to exits to education/training.

Where theory has been put forward that can argue negative or positive duration dependence in unemployment, no duration dependence at all, or a combination of all three, empirical research takes on added importance if we are to understand the true relationships between unemployment duration and exit. Unfortunately, despite a considerable empirical literature, both internationally and in GB, evidence is mixed. Winter-Ebmer (1998), for example, finds evidence of positive duration dependence in Austria. Lancaster (1979), Nickell (1979) and van den Berg and van Ours (1994) find

evidence of negative duration dependence in GB. Narendranathan *et al.* (1985) and Narendranathan (1993) find little evidence of duration dependence of either sign for GB as do van den Berg and van Ours (1994) for France. Arulampalam and Stewart (1995) and Boheim and Taylor (2000) find a non-linear (inverted u) relationship between duration and exit probability.<sup>2</sup> Finally, some studies find evidence of ‘spikes’ around the time of benefit exhaustion (e.g. Meyer, 1990; Carling *et al.*, 1996) or other policy interventions (e.g. Dolton and O’Neill, 1996a, b), although others do not (e.g. Stancanelli, 1999). Reviews of this literature are provided by Devine and Kiefer (1991) and Machin and Manning (1999). Studies examining duration dependence in unemployment for young people specifically include Heckman and Borjas (1980), Lynch (1989), Korpi (1995), and Russell and O’Connell (2001). Between them these studies find evidence of no duration dependence, negative duration dependence and non-linear duration dependence.

As argued above, one reason for this ambiguity is the observational equivalence of duration dependence and unobserved heterogeneity. Failure to consider unobserved heterogeneity can lead to a false conclusion of negative duration dependence because of negative bias (Heckman and Borjas, 1980). Unobserved heterogeneity can also bias the coefficients on the covariates. The standard approach to this unobserved heterogeneity problem is to specify some mixing distribution for the hazard function including a (parametric) ‘random effects’ error term element (e.g. Portugal and Addison, 2000). A related alternative is to specify a non-parametric mass point distribution for the unobserved heterogeneity (e.g. Heckman and Singer, 1984;

---

<sup>2</sup> The curve illustrating the relationship between duration and exit probability at a given point in time (conditional upon exit not having occurred already), other things being equal, is called the baseline hazard function. If there is no duration dependence, such a hazard function will be horizontal. Negative duration dependence corresponds to a downward sloping hazard function and positive duration dependence to an upward sloping hazard function. Observed predetermined (time-invariant) factors that might influence the probability of a match between vacancy and job seeker, such as qualification level, essentially have the effect of shifting the baseline hazard function up or down (Portugal and Addison, 2000). A tightening of the labour market, for example, might be expected to shift the hazard function upwards, so increasing the chance of exit into employment at a given point in time.

Kettunen, 1997; Stancanelli, 1999).<sup>3</sup> Narendranathan and Stewart (1993) argue, however, that there is no reason to believe that the resulting distortion from introducing such specifications for unobserved heterogeneity is any less severe than any distortion that would arise from ignoring the unobserved heterogeneity in the first place. Including more observed covariates will help, but is not always possible given data restrictions.

Arulampalam and Stewart (1995) argue that the duration dependence bias and the bias on the coefficients of the covariates can be reduced by using a flexible specification for the baseline hazard function (see also Ridder, 1987; Portugal and Addison, 2000). In other words, given a sufficiently flexible baseline hazard specification, it becomes unnecessary to model unobserved heterogeneity (Boheim and Taylor, 2000). In any case a restrictive specification of the baseline hazard function (e.g. the relatively common, and monotonic, Weibull) can in itself bias the results (Narendranathan and Stewart, 1993). Given our interest in the possibility of a non-linear baseline hazard, such a restrictive approach is not attractive. Many recent studies adopt a piecewise constant baseline hazard that separates T (total time at risk) into different segments with different hazard rates (e.g. Arulampalam and Stewart (1995); Narendranathan and Stewart (1993); Bratberg and Vaage, 2000; Portugal and Addison, 2000; Boheim and Taylor, 2000). This methodology was originally introduced by Prentice and Gloeckler (1978). To allow flexibility in the shape of the baseline hazard function, and to reduce the likelihood of bias resulting from unobserved heterogeneity, we follow the Prentice-Gloeckler (PG) approach here.

The discrete time hazard rate for individual  $i$  in time interval  $t$  to exit joblessness into activity  $j$  (employment or education/training) is given by Equation 3, where  $\theta_j(t)$  is a function relating the hazard rate with spell duration. The PG approach assumes there is a specific parameter that is constant over each period identified. This parameter can

---

<sup>3</sup> Despite its widespread use in the empirical literature and its apparent attractiveness, the mass point approach to modelling heterogeneity faces two principle difficulties. Firstly, it is not uncommon to encounter computational difficulties in implementing the non-parametric maximum likelihood or mass point estimator (see e.g. Baker and Melino, 2000). Secondly, although Heckman and Singer provide conditions under which this non-parametric maximum likelihood estimator is consistent, there is as yet no derivation of the sampling distribution of this estimator. Hence standard errors calculated by conventional formulae (e.g. from the estimated information matrix) have no formal justification and might well be misleading. For these reasons in this paper we adopt parametric specifications for heterogeneity.

be interpreted as the logarithm of the integral of the baseline hazard over the particular time interval (Boheim and Taylor, 2000).

$$h_{jt}(X_{it}) = 1 - \exp\{-\exp[X_{it}'\beta_j + \theta_j(t)]\} \quad (3)$$

The extension of the standard single risk model to two exit destinations is referred to as the competing risks model (e.g. Boheim and Taylor, 2000). The two exits are treated as independent, so the probability of exit into employment is assumed not to depend on the probability of exit into education and *vice versa*. Kollmann (1994) argues that negative duration dependence in terms of chances of getting a job might lead to increasing exit from the labour market into economic inactivity, suggesting that independent risks may not be a suitable assumption. However, we treat economically inactive spells (outside of education) as unemployment spells in our model, so this argument is less relevant here. Nevertheless, our independence assumption rests on the belief that duration dependence in exits to employment does not affect the likelihood of exits to education and training. Further research might relax this assumption.

In practical terms, the independent competing risks framework treats other exits as right censored (see Lancaster, 1990). Boheim and Taylor (2000) argue that the distributional assumptions for including unobserved heterogeneity are even stronger in the competing risks framework than in the standard single risk case, and are therefore reluctant to assume any particular specification for such heterogeneity in their model. Roed *et al.* (1999) add that the standard negative bias on duration dependence of unobserved heterogeneity does not necessarily hold in a competing risks framework. We estimate the competing risks model both with and without specifying unobserved heterogeneity, following the extension to the PG methodology proposed by Meyer (1990) and allowing for either a gamma distributed random effect or a normally distributed random effect. The results are compared to the single risk model where all exits are treated as homogenous and a single hazard function estimated.

### 3. The Data

The data are taken from the 1999 Status Zero Survey of young people first eligible to leave school in NI in July 1993. A representative sample of 712 young people is followed for 72 months from July 1993 until June 1999 (from age 16 to age 22).<sup>4</sup> For each individual we have detailed monthly activity information, including details of job-search activities for spells of joblessness, along with considerable background information including qualifications achieved in compulsory education and family background details. This level of detail is generally not available from register data. However, the benefits from using the survey data come at a cost mainly in terms of sample size, but also in the potential for recall error such as ‘heaping’, where respondents might ‘round-off’ past unemployment spells to the nearest six months, for example (e.g. Torelli and Trivellato, 1993).

Around 35% of the sample from the Status Zero Survey have experienced at least one spell of unemployment or other joblessness during the sample period. 24% have experienced a spell of at least six months and 17% a spell of at least 12 months. The figures for unemployment (using the ILO definition) are a little smaller but roughly in proportion (see Table 1). Most young people who experience joblessness experience only one spell, but a significant minority experience two or more spells (see Figure 1). Our final sample for this paper is all spells of joblessness experienced by the Status Zero Survey sample over the 72 months of the survey,  $n=357$ . Following Boheim and Taylor (2000) we treat separate spells for the same individual as separate observations.<sup>5</sup>

The focus on joblessness rather than narrower definitions of unemployment has a cost in terms of including those individuals who describe themselves as being ‘on holiday’ – treated here as jobless spells. The numbers in this category are small, however, and are identified by a binary dummy in the empirical model. This cost is offset by

---

<sup>4</sup> No jobless spells are possible *before* the first observation in the data set since all sample members are in school, by definition, therefore there are no left-censored joblessness spells.

<sup>5</sup> Boheim and Taylor (2000) and Russell and O’Connell (2001) argue that by including previous unemployment experience in the explanatory variables, any correlation of the error terms for two spells experienced by the same individual is unlikely to be a problem. We follow this approach largely for reasons of sample size, and in any case we find no evidence of significant unobserved heterogeneity.

additional practical benefits of the focus on joblessness, including the fact that no spells end for ‘artificial’ reasons, such as failure to sign on the unemployment register in a given month.

For our sample of 357 jobless spells, the mean duration is 12 months (median 6 months – see Table 2). Given the inclusion of censored spells, these statistics will be biased downwards (Boheim and Taylor, 2000). Around 26% of jobless spells in the sample are right-censored. 68% of uncensored spells end in employment and the remaining 32% end in education or training. The mean duration for spells ending in employment is 9 months (median 5.5 months) compared to 8 months for spells ending in education/training (median 3 months).

Figure 2 shows the distribution of spell duration. This too will be biased downwards because of right censoring. Nevertheless, two points in particular stand out. Firstly, many jobless spells for young people are short – between one and three months. Part of the explanation for this lies in the summer periods between academic years, particularly the initial summer period in 1993, where some young people describe themselves as being on holiday, some claim to have looked unsuccessfully for a summer job and others describe themselves as waiting for a job or education/training place already obtained to begin. This partly explains the temporary peak in jobless rates in the summer of 1993 and the smaller peak in the summer of 1994, shown in Figure 3. Secondly, there is some suggestion of ‘heaping’ measurement error, particularly at 24 months and perhaps at 48 months. We discuss the robustness of our analysis to the possibility of such recall error in the following section.

Table 2 gives a life table for joblessness spells with all exits pooled and separately for exits to employment and exits to education and training.<sup>6</sup> A third of all spells do not survive past three months. For exits to education and training, the proportion of spells surviving into the fourth month is less than half. In general, joblessness spells ending in education/training are shorter than those ending in employment.

---

Ondrich and Rhody (1999) discuss a generalisation of the likelihood to account for multiple spells but make clear that it is far from straightforward. We leave this generalisation for later research.

Explanatory variables for the empirical model are included according to two considerations. First, as far as our data allow, we include those factors generally found to influence unemployment or joblessness duration in existing studies. Second, we include those factors particular to NI previously found to influence labour market outcomes and education/training participation (e.g. community/religion and sub-regional geography). Table 3 lists the explanatory variables for the (inclusive) model, along with sample means and standard deviations.

Three policy dummy variables are included in the empirical model (and Table 3) to capture the effects of the introduction of JSA and New Deal. The JSA dummy takes the value one for all months in joblessness following the introduction of JSA, for all individuals. A jobless spell spanning the introduction of JSA will correspond with zero values for JSA before October 1996 and one thereafter. Two New Deal dummies are included in the model. The first, NEWDEAL1, is defined in the same way as the JSA dummy, with zeroes before April 1998 (the date of the introduction of New Deal) and ones thereafter. This dummy is intended to capture the overall effect of New Deal – on the labour market, on employer incentives and on incentives for jobless young people in general. The second New Deal dummy, NEWDEAL2, is intended to capture the effects of New Deal on *eligible individuals* only during their period of eligibility. The variable takes the value zero for all months of joblessness before April 1998, and for the first six months of any spells of jobless beginning after April 1998, and one otherwise.<sup>7</sup> Where jobless spells span the introduction of New Deal, the dummy takes the value zero until the six-month threshold is reached or until April 1998, whichever is later.

Table 3 also gives simple pairwise correlation coefficients for the explanatory variables with the duration variable. Factors correlated with shorter spells of joblessness (from the raw data) include having good qualifications at age 16 (5+ GCSEs at grades A-C), more job search activity, the New Deal and JSA dummies, age at start of spell and the local unemployment rate. Factors correlated with longer

---

<sup>6</sup> The employment category includes part-time employment where it is described as the primary economic activity during the month.

<sup>7</sup> Because we focus on joblessness and not unemployment, New Deal will not be compulsory for all young people in the sample jobless for more than six months. New Deal does extend to single mothers and disabled young people, however, but on a voluntary basis.

spells of joblessness include having an unemployed father, coming from the west of NI (a region of comparative deprivation), women with children, and coming from the Catholic community. These patterns in the raw data generally have the expected sign, with the exceptions of the local unemployment rate, previous joblessness and previous wage.

#### 4. Results and Discussion

Below we discuss the results of estimation of the PG model with and without unobserved heterogeneity. In each case, we first treat all exits as pooled and then identify exits to employment and exits to education/training separately. Estimation is carried out on the whole sample, due to sample size.<sup>8</sup> Estimation of the PG and Meyer models uses the Jenkins (1997) procedure for Stata.

The PG model set out in Equation 3 (repeated below) assumes a piecewise constant specification for the baseline hazard function. An extension of the model introduced by Meyer (1990) explicitly models unobserved heterogeneity using a gamma distribution, as shown in Equation 3a.

$$h_{jt}(X_{it}) = 1 - \exp\{-\exp[X_{it}'\beta_j + \theta_j(t)]\} \quad (3)$$

$$h_{jt}(X_{it}) = 1 - \exp\{-\exp[X_{it}'\beta_j + \theta_j(t) + \log(e)]\} \quad (3a)$$

where  $e$  is a gamma distributed random variable with unit mean and variance  $\sigma^2$ . An alternative extension replaces the gamma distributed error term with a normally distributed error term in order to check robustness.

The interval-specific baseline hazard can only be identified for those duration intervals in which exits occur (Jenkins, 1997). In addition, the adequacy of the asymptotic normal approximations may be suspect if some time intervals include very few failures (Prentice and Gloeckler, 1978). Jenkins therefore suggests additional grouping of duration intervals for which the number of events is small in order to improve estimation precision. In our application, given the gaps in the distribution of durations shown in Figure 3, and our small sample size overall, we need to group duration intervals.<sup>9</sup> We therefore specify the following groups of duration intervals: separate groups for months 1-12, 13-15 months, 16-21 months, 22-24 months, 25-36 months and 37+ months. At least one of these duration interval dummies must be

---

<sup>8</sup> Results from estimation of the model for pooled exits, separately by gender, are available from the authors on request and are briefly discussed later in this section.

omitted for the equation to be identified – we omit the one-month and 37+ months interval dummies.

The results of the PG model for pooled exits and for exits to employment and education/training separately are presented in Table 4. While the PG model appears well specified for the pooled exits and exits to employment, it is less so for exits to education treated separately. This is unsurprising given the small number of exits to education/training in the data set. Our discussion concentrates on the pooled exits and the exits to employment, although the estimates for the exits to education/training model are reported for completeness.<sup>10</sup> The Meyer (1990) extension introducing a gamma distribution modelling unobserved heterogeneity is an insignificant addition to the PG model for pooled exits and exits to employment, as is the alternative normal specification for heterogeneity. The Meyer extension cannot be estimated for exits to education and training because of the small number of such exits in the sample. The insignificance of unobserved heterogeneity is a common finding in studies that adopt the PG model (e.g. Carling *et al.*, 1996; Boheim and Taylor, 2000).<sup>11</sup> We therefore follow Boheim and Taylor (2000) in reporting only the results from the standard PG model without unobserved heterogeneity.

### **Pooled Exits**

Not all duration dummies are significant in the pooled model and between seven and fifteen months it is not possible to discern any strong pattern of duration dependence (although the general trend appears to be negative but not strongly). Nevertheless a clear downward sloping pattern emerges up to seven months duration and beyond fifteen months duration, with a slight non-linearity at five months duration. In other words, there is clear negative duration dependence, at least initially, which suggests a loss of employability as jobless spells continue beyond the first month. The local peak, or ‘spike’, in the baseline hazard at five months is within a standard error of the

---

<sup>9</sup> Inference on  $\beta$  and the survivor function estimator should nevertheless be relatively robust to the small number of failures in some specific time intervals (Prentice and Gloeckler, 1978).

<sup>10</sup> In any case, as Boheim and Taylor (2000) point out, the estimated coefficients for the pooled model are essentially a weighted average of the impact of the covariates on the probability of exit into different states, in our case into employment or into education and training.

<sup>11</sup> Carling *et al.* (1996), for example, find the addition of Meyer-type unobserved heterogeneity to be irrelevant, with the variance of the gamma mixing distribution approaching zero.

baseline hazard at four and six months, so we cannot strongly reject negative duration dependence over the first six months. Figure 4 shows the baseline hazard function from the parsimonious PG model for pooled exits, calculated for the average young person.<sup>12</sup>

Despite its statistical insignificance, the spike in the baseline hazard function at five months is of potential interest. Spikes are sometimes found around the time of benefit exhaustion or reduction (e.g. Carling *et al.*, 1996). Young jobless aged 16 and 17 are in most cases not entitled to benefits. Above this age, the JSA has brought in a benefit regime in which the level of benefits can be reduced after six months of unemployment under certain conditions.<sup>13</sup> It is unlikely, however, that many young people will be subject to this benefit threshold effect of JSA. A more likely possibility is that the spike is capturing an effect of New Deal. After six months of claiming unemployment benefits, most young people are required to enter the New Deal or lose entitlement to benefits. This may be enough to stimulate additional job search (or the search for a training or education place) at five months. If we consider the spike to extend to six months (as it does for exits to employment), it may be that the baseline hazard function is also capturing an effect where some young people are helped rapidly out of joblessness on entry to New Deal.<sup>14</sup>

If the spike at 5-6 months is a New Deal effect, then we would expect to see a contrast in the shapes of the baseline hazard functions before and after the introduction of New Deal. Our sample period contains New Deal's introduction, so in principle we can examine this question. In practice, however, the sample is too small to examine the post-New Deal period in isolation. We can, however, estimate the model for pooled exits for that part of the sample period that falls before the introduction of New Deal.

---

<sup>12</sup> The average young person has all dummies set to 0 (the sample mean of all dummies is below 0.5) and continuous variables set to their sample means. She is female, Protestant, 18.6 years old, with less than five GCSEs at grades A-C. She has nine months of work experience (including training placements) with a previous (take-home) wage of £50-£100 per week and two months previous experience of joblessness. Her father is in employment but not in a managerial, professional or related occupation, and she lives in the south, north or east of NI, facing an unemployment rate of 20.5% in her area. She does not have children and began her jobless spell before JSA or New Deal were introduced, applying for one job per month during her jobless spell.

<sup>13</sup> If sufficient contributions have been made during previous employment and if means-testing would result in lower benefit payments after the contributions-based benefit had been exhausted (six months).

<sup>14</sup> The spike for the exits to employment competing risks model is *not* within a standard error of the baseline hazard at seven months.

We do not report the parameter estimates for this restricted sample period model, but Figure 5 gives the resulting baseline hazard function for the average pre-New Deal young person. There is no spike for the pre-New Deal sample, which appears to support our speculation that the spike in the full-sample baseline hazard may be a New Deal effect. If we could estimate post New Deal in isolation, our guess is that this spike would be more pronounced.

Torelli and Trivelato (1993) present evidence of substantial measurement errors in retrospective survey data. In particular, they show how recall error of unemployment spells can lead to spikes in the duration distribution at certain points, such as six months, 12 months, 24 months etc., or ‘heaping’. At first glance, this seems an alternative potential explanation of the spike at six months. There *are* spikes in the duration distribution at 24 months and 48 months in the Status Zero Survey data (see Figure 3). These spikes are small, however, and there is little evidence of a similar effect at six months or 12 months, so we conclude our spike is not a result of measurement error.<sup>15</sup> In general, given the relative absence of heaping in our data, and given that we have in any case grouped durations, we take no explicit account of heaping in the model. In further support, Kraus and Steiner (1998) find parameter estimates to be insensitive to explicit specification of heaping.

The model for the full sample period contains two step dummies for New Deal (see Section 3 for details) which are intended to capture the effects of New Deal on the hazard rate. Our prior for the second dummy (NEWDEAL2), which takes the value one for all eligible individual-months, is that it will have a positive sign (increase the conditional probability of exit). At least for those claiming JSA benefits, continued joblessness becomes more difficult on eligibility for New Deal. Young people are instead offered guidance and assistance in moving out of joblessness into education, training or employment (during the Gateway stage), and then a number of options including subsidised employment should the initial Gateway stage prove unsuccessful. The expected sign of the other New Deal dummy (NEWDEAL1) is less clear. It is negatively correlated with duration in the raw data, which suggests durations are shorter since New Deal’s introduction. But if the NEWDEAL2 dummy

---

<sup>15</sup> The Status Zero Survey is a face-to-face interview based survey, which may reduce such recall error compared to postal or telephone surveys.

controls for the beneficial effects of New Deal on those young people eligible for its assistance, it is possible that NEWDEAL1 might capture negative substitution-type effects. Jobless young people existing in a world with New Deal that are nevertheless outside New Deal may receive less job offers than competitors within New Deal, or than they might have received before the introduction of New Deal, for a variety of reasons. First, young people outside of New Deal may receive less assistance with job search if scarce Employment Service resources are concentrated on those in New Deal. Second, young people outside New Deal may be less attractive to employers given that they can hire subsidised employees through New Deal (see e.g. Riley and Young, 2001). Third, young people outside New Deal may feel less incentive to job search if they know they are shortly to enter New Deal and receive the help and guidance that entails.

The estimated sign on the NEWDEAL2 dummy is indeed positive in the pooled model, although insignificant (see Table 4). The weakness of this effect might be explained by its voluntary nature for some jobless young people, notably single parents, for whom it may have made little practical difference. When the sample is restricted to spells of ILO unemployment only, NEWDEAL2 becomes slightly more significant (at 20%) with the same sign. A purely statistical explanation of the weakness of this effect lies in the possibility of the dummy being endogenous. In other words, the value of the dummy may be influenced by the hazard rate up to six months – those more likely to exit joblessness early are less likely to become eligible for New Deal. If this is the case, however, we know the sign of the bias on the NEWDEAL2 coefficient will be negative. The estimated coefficient in Table 4 should therefore be interpreted as a lower bound on the New Deal eligibility effect – it may be stronger.

The estimated sign for the NEWDEAL1 dummy suggests the opposite – jobless spells *in general* have increased in duration following the introduction of New Deal, other things being equal, although again the effect is not statistically significant. Such a result is not without precedent. Carling *et al.* (1996) find little effect of Active Labour Market Policies on the hazard rate to employment in Sweden, for example. Van Ours *et al.* (2000) argue that active labour market policies in Switzerland may actually have a ‘prolonging’ effect on unemployment duration, at least for men. Riley and Young

(2001) find some weak evidence of a negative New Deal effect on some of those outside the programme. Despite its weakness, does this result suggest there may be some marginal level of negative substitution-type effects following the introduction of New Deal?

Similar to the NEWDEAL1 dummy, a dummy is included for JSA (zero for all individual-months before October 1996 and one thereafter). Our prior for this is a positive sign, given that JSA was intended to intensify job search activities for the unemployed. This is indeed what the results suggest, so the introduction of JSA does seem to be significantly correlated with shorter joblessness spells. The effect is quite large, with hazard rates around a third higher for jobless individuals following the introduction of JSA.<sup>16</sup>

Turning to the other covariates, the pooled model suggests that joblessness duration is increasing with age. The results suggest that hazard rates for 17 year olds may be as much as twice those for otherwise identical 21 year olds. With a single cohort it is possible that the negative coefficient on the age variable is also capturing some other time-related effect, however, so we are reluctant to interpret it as entirely an age effect. In the unemployment duration literature as a whole, age is widely accepted as a significant determinant of the probability of exiting unemployment and getting a job, with older workers less likely to get re-employed (e.g. Narendranathan and Stewart, 1993; Stancanelli, 1999; Firth *et al.*, 1999; Kettunen, 1997; Bloeman, 1997). Studies of young people specifically have drawn mixed conclusions of the age/duration relationship. Russell and O'Connell (2001), for example, find a significant negative effect of age on duration of unemployment spell in several European countries. Heckman and Borjas (1980) and Lynch (1989) find no significant relationship in the US.

Boheim and Taylor (2000) find mothers have reduced exit rates to employment compared to childless women. Bratberg and Vaage (2000) find similar reduced exit rates for mothers with children under five years of age. Firth *et al.* (1999) find the opposite for women with children over five years of age. In our model the interactive

---

<sup>16</sup> We do not examine the *quality* of these faster job matches in this paper.

dummy for women with children (although we do not have information on the age of children, there are likely to be few women with children over five years old in the sample) is negative and significant. In other words, joblessness duration is higher for young women with children than for others in the sample. Given our broad definition of joblessness, some young women with children in the sample will be largely disengaged from the labour market, so this result is unsurprising. The scale of this effect suggests that having a child roughly halves hazard rates for otherwise identical females. Gender is otherwise insignificant in the pooled model, as is religion/community.<sup>17</sup> Boheim and Taylor (2000) also find hazard rates to be similar for males and females in GB.

When the (pooled) model is estimated separately for males and females, covariate estimates are broadly similar, as are the estimated baseline hazard functions, albeit less precise. There are a number of differences, however, between the genders. The effects of having an unemployed father and the effects of number of job applications are not significant for females whereas they are for males. On the other hand, males do not display a significant JSA effect. Perhaps most interesting is that the NEWDEAL2 dummy is significant (and positive) for males when estimated separately, but insignificant for females. In other words, there is a positive effect of New Deal eligibility on the hazard rate for young males. Again this should be considered as a lower bound due to the fact that we cannot rule out some negative bias on this coefficient. This is consistent with the suggestion that jobless females caring for children (and therefore outside of the compulsory New Deal) are pulling down the size of the effect in the full sample. Figure 6 gives baseline hazard functions for pooled exits for the average female and average male. Hazard rates are higher for females than for males at short durations and the spike at five months is more pronounced for females, though this will be partly offset by the significance of the NEWDEAL2 dummy for males.<sup>18</sup> Otherwise the functions are broadly similar.

---

<sup>17</sup> Crouchley *et al.* (1991) find unemployment duration to be longer for Catholics than for Protestants in NI, other things being equal, but do not examine young people separately.

<sup>18</sup> The effect of this dummy is not shown on the hazard function for males since dummies are set to zero for the 'average' young male. For those eligible, however, the hazard rate shown in Figure 6 would essentially shift upwards from six months onwards.

Our two general measures of social background are dummies for father unemployed and dummies for father managerial, professional or related. Neither is strongly significant in the pooled model, although father unemployed increases joblessness duration with marginal significance and to a small extent. This latter result is consistent with existing studies of joblessness and disadvantage (e.g. Dearden *et al.*, 1997), although we have found no existing duration study which controls for father's employment status. The negative effect of having an unemployed father may capture both attitudes and aspirations, and the importance of informal job search through family contacts (see, e.g. Hannan, 1998). The insignificant effect of father managerial, professional or related may reflect offsetting effects of attitudes and aspirations and wealth effects, which might on their own prolong jobless spells (e.g. Bloeman and Stancaelli, 2001).

Numerous studies have found education level or years of schooling to reduce the duration of unemployment spells (e.g. Bradley and Taylor, 1991; Narendranathan and Stewart, 1993; Carling *et al.*, 1996; Bratberg and Nilsen, 2000; Firth *et al.*, 1999; Wolbers, 2000). We include a binary dummy for achieving five or more GCSE passes at grades A\*-C by the end of compulsory education. The results for the pooled model suggest, although the effect of having good qualifications on probability of exit from joblessness is positive, it is not significant. Kettunen (1997) finds a non-linear relationship between education and duration, argued to be the result of high reservation wages counteracting the higher arrival rate of job offers for the highly educated. It may be that these two effects counteract one another in our model, where education is captured by a single binary dummy. Many young people have no previous employment experience and therefore no previous wage data on which to base a measure of reservation wages, so it is unlikely that our previous wage variable, itself insignificant, entirely controls for this reservation wage effect.

Two other personal characteristics are included in the model. First, previous employment experience is included (number of months in employment prior to the jobless spell), but is not significant. Information on work experience during school would be advantageous here, but is not present in the data set. Second, to capture any unemployment scarring effects, a variable is included for months of joblessness experienced prior to the current jobless spell. Heckman and Borjas (1980) refer to this

as occurrence dependence, and evidence of unemployment scarring for GB has recently been found by Narendranathan and Elias (1993), Narendranathan and Stewart (1993), Arulampalam *et al.* (2000), and Boheim and Taylor (2000). In our case however, perhaps given that most jobless spells in our sample are first spells, this variable is insignificant. Binary dummies for previous employment and joblessness are similarly insignificant.

Two variables are included in the model to capture aspects of the jobless spell itself. Firstly, we include a variable for the average number of job applications made by the individual per month of the jobless spell.<sup>19</sup> We expect young people that search more intensely for jobs to experience shorter jobless durations, as suggested by the search model. This is indeed the case as shown in Table 4, although the magnitude of the effect is not large – applying for one more job per month adds around 2% to the hazard rate. Information on job search intensity has not been available for many of the data sets used to analyse unemployment duration, which may explain the omission of such measures from many existing studies (Chuang, 1999, is a recent exception). Secondly, a small number of the jobless spells in our sample can be categorised as holidays. Such spells are likely to differ in nature to other jobless spells, with little or no job search, for example, and perhaps a pre-arranged entry into employment, education or training in some cases.<sup>20</sup> In the pooled model the holiday variable is, however, insignificant.

Economic environmental factors can influence unemployment duration. Lynch (1989) and Arulampalam and Stewart (1995) find local unemployment rates to be a significant determinant of unemployment duration (Stancanelli, 1999, finds local unemployment rates to be insignificant). Bloeman (1997) finds significant area dummies. Two local economic environment measures are included in the model – the travel to work area (TTWA) adult unemployment rate (in the absence of a youth unemployment rate by local area) and a binary dummy for living in the West of NI (a region of relative social and economic deprivation). Neither the local unemployment

---

<sup>19</sup> Month-specific information on job search is not available.

<sup>20</sup> These will include summer holidays between educational institutions as well as ‘gap years’, for example.

rate nor the West dummy is significantly related to jobless durations for young people in our model, other things being equal.

### **Exits to Employment and Education/Training**

The estimates of the parameters for the independent competing risks model for exit to employment are also shown in Table 4. Again many of the duration group dummies are insignificant, suggesting no strong relationship between duration and the hazard rate beyond eight months. However, similarly to the pooled model, there is evidence of negative duration dependence over the first four months of duration, followed by a small rise in the baseline hazard in months five and six. This spike, centred on six months rather than five months, is the main driver of the pooled spike interpreted earlier as a New Deal effect.<sup>21</sup> The baseline hazard for exits to employment is shown in Figure 4.

Many of the covariates that drop out of the pooled model also drop out of the exits to employment model. The exception is the holiday dummy, which here is significant and negative. In other words, those young people describing themselves as on holiday have lower exit probabilities to employment than others. This, coupled with the insignificance of the term in the pooled regression, suggests that holidays are mostly taken between education spells (e.g. gap years or summer breaks). Unsurprisingly there is a positive coefficient on the holiday dummy in the exits to education/training model.

All variables significant in both the pooled and the exits to employment model have the same sign in both models. Older young people, young women with children, and young people with unemployed fathers all experience longer joblessness duration. Young people engaging in more job search experience shorter joblessness duration. This effect is stronger in the exits to employment model than in the pooled exits model, and is insignificant in exits to education/training. The JSA dummy also has a stronger positive effect on exit probability than in the pooled model, also

---

<sup>21</sup> The spike at five months in the pooled exits model reflects some weighted average of the spike at six months for exits to employment and a spike at five months for exits to education/training (see Figure 4).

corresponding with its insignificance in the exits to education/training model. The introduction of stricter rules on job search for those claiming unemployment benefits seems to have paid off in terms of reducing joblessness duration among young people in NI. It has been argued that, rather than leading to increased employment, JSA may have led many previously unemployed people, particularly from older age groups, to move out of the labour market altogether on to alternative benefits not covered by the new job search rules (e.g. see Beatty and Fothergill, 1999). Our results, given our broad definition of joblessness, suggest this is not the case for young people.

There are few exits to education/training in the sample so we do not place too much emphasis on the parameter estimates in columns six and seven of Table 4. There is little discernible pattern to the baseline hazard function (most duration group dummies are insignificant), as shown in Figure 4. Older young people and young women with children appear less likely to exit joblessness into education or training. Those describing themselves as on holiday are more likely to exit to education/training. Perhaps most interestingly, two variables are found to be significant which are not in the other models. Firstly, young people with good qualifications at the end of compulsory education have increased exit probability to education/training. This is consistent with those studies finding education level or years of schooling to reduce the duration of unemployment, and also the literature on post-compulsory participation in education (e.g. Rice, 1999). Secondly, those young people with more previous employment experience are less likely to exit joblessness into education/training. Once the decision to enter the labour market has been made, and young people have been employed, they appear to be unlikely to return to education/training, even if jobless.

Finally, we re-estimate the pooled exits model restricting the sample to ILO unemployment spells only, to test the robustness of our results to our broad definition of joblessness.<sup>22</sup> We have already suggested that the effects of the New Deal dummies, though still insignificant at conventional levels, are slightly stronger in the ILO model. Otherwise the results are very similar to those for the pooled model estimated on the full jobless sample, both in terms of the shape of the baseline hazard

---

<sup>22</sup> Results available from the authors on request.

and the effects of the covariates on the hazard rate. The main difference is that the mother with child variable, as we might expect given the more restrictive definition of unemployment, becomes insignificant.

## 6. Concluding Remarks

The paper examines the duration of jobless spells for a sample of young people in NI between 1993 and 1999. Separate routes out of joblessness are identified – into employment or into education and training – in contrast to much of the existing literature, which treats education together with joblessness other than unemployment as a single ‘out of the labour force’ state. Duration dependence is shown to be generally negative, i.e. it gets harder to exit joblessness the longer one has been jobless. This suggests that targeted policy intervention during the first month of joblessness may be a useful policy addition to current intervention, which focuses on young people that have been jobless for six months or more (during which time employability has been deteriorating). Where it is possible, prevention is likely to be better than cure both in financial terms and in reducing psychological costs to jobless young people themselves.

Although generally downward sloping, the hazard function also displays a spike around 5-6 month’s duration for both pooled exits and exits to employment. We argue that this spike corresponds with an effect of the introduction New Deal. The model suggests a further New Deal effect, captured by a dummy variable, but only significant for males: there is a positive shift in the hazard function from six months duration onwards following the introduction of New Deal. In other words, exit probabilities increase for male jobless on eligibility for New Deal.

There is a clear need for more detailed research on the effects of New Deal on the duration of joblessness for young people, not only in NI but also elsewhere in the UK. This paper provides a framework for such further research, and provides some exploratory analysis, but with a small sample and no suitable control group cannot claim to provide definitive evidence of the effects of New Deal. Such evidence is likely to be available from register data, however, covering large numbers of individuals and allowing use of the 25-29 year old age group (not covered by New Deal) as a control group. This is the next stage of our research. Of course, the benefits of register data will come at a cost in terms of detail on covariates and exclusion of those jobless young people not covered by claimant counts. Taken together, however, the present study based on survey data, and the suggested study of register data, are

likely to provide a strong evidence base for policy makers to ensure New Deal is having its intended, and widely supported, effects.

The introduction of JSA is correlated with shorter jobless spells for young people through increased exit rates to employment. In other words, JSA has been successful in turning greater pressure for job search into faster job matches, at least for young people in NI over our sample period. In general, greater job search activity is shown to reduce joblessness duration. The main observable barriers to exit from joblessness include caring for a young child and having a family background of unemployment. Having poor qualifications also acts as a barrier, particularly to exit into education. Overall, where significant, the estimated effects of the covariates are generally consistent with those suggested by search theory.

## **Tables and Figures**

**Table 1: Cohort Proportions with Unemployment and Joblessness Spells**

<b>Proportion of Cohort</b>	<b>Jobless Spell, %</b>	<b>ILO Unemployment Spell, %</b>
At least one spell	35	27
At least one spell $\geq$ 3 months	29	22
At least one spell $\geq$ 6 months	24	17
At least one spell $\geq$ 12 months	17	12

**Source:** McVicar *et al.* (2000). Note: Weighted sample proportions.

**Table 2: Joblessness Duration by Destination State (% Survival)**

	<b>Duration (months)</b>										
<b>Exit</b>	<b>1</b>	<b>2</b>	<b>3</b>	<b>4</b>	<b>5</b>	<b>6</b>	<b>9</b>	<b>12</b>	<b>18</b>	<b>24</b>	<b>36</b>
<b>All</b>	92	75	64	59	54	50	40	30	19	11	08
<b>Emp</b>	94	72	63	58	53	47	39	27	21	08	04
<b>Ed/tr</b>	80	66	45	37	31	27	19	14	06	03	02

**Notes:** Figures show (i) % of all jobless spells with duration of n+ months (including censored spells), (ii) % of all jobless spells ending in employment with duration of n+ months, (iii) % of all jobless spells ending in education/training with duration of n+ months. Unweighted.

**Table 3: Sample Means of Explanatory Variables and Correlation with Duration**

<b>Explanatory Variable</b>	<b>Sample Mean (Standard Deviation)</b>	<b>Correlation with Duration</b>
GCSE5 (1=yes)	.236 (.425)	-.153
Father unemployed (1=yes)	.244 (.430)	.110
Male (1=yes)	.483 (.500)	-.073
Catholic (1=yes)	.483 (.500)	.193
Previous employment (mts)	8.81 (14.0)	-.128
Previous wage*	1.09 (1.38)	-.121
Previous joblessness (mts)	2.24 (5.67)	-.140
Father SOC Group 1 (1=yes)	.188 (.391)	.020
JSA	.466 (.499)	-.215
New Deal 1	.177 (.382)	-.299
New Deal 2	.108 (.311)	-.089
Age at start of spell (yrs)	18.6 (1.84)	-.273
West of NI (1=yes)	.199 (.400)	.102
On holiday (1=yes)	.065 (.246)	-.121
Mother (have child * female)	.101 (.302)	.314
Job applications (monthly mean)	1.00 (2.03)	-.274
TTWA Unemployment Rate	20.5 (6.41)	-.320

**Notes:** \*Previous wage data is recorded in £50 bands, 1=0-£50, 6=£250+. Observations are unweighted.

**Table 4: PG Model, Pooled Exits and Competing Risks, Males and Females**

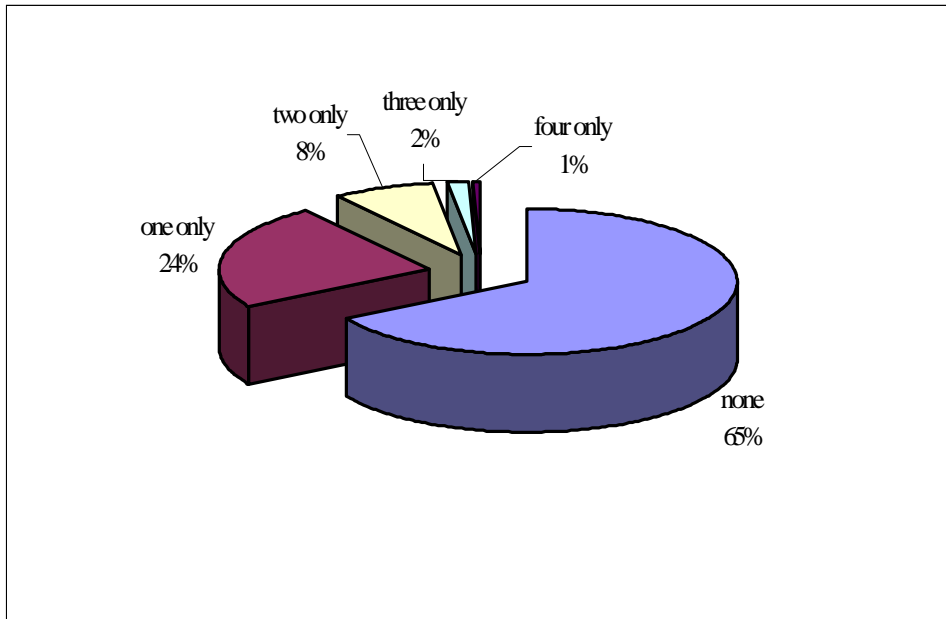
	Pooled Exits		Exits to Employment		Exits to Education/Training	
	Inclusive Model	Parsimonious Model	Inclusive Model	Parsimonious Model	Inclusive Model	Parsimonious Model
Constant	3.88* (1.53)	4.85* (1.03)	3.08 (1.83)	4.33* (1.21)	2.45 (2.77)	3.54* (1.59)
Month 2	1.73* (.23)	1.75* (.22)	2.33* (.31)	2.32* (.29)	.42 (.40)	.43 (.38)
Month 3	1.30* (.26)	1.38* (.24)	1.54* (.36)	1.52* (.34)	.89* (.37)	.93* (.35)
Month 4	.69* (.37)	.81* (.35)	1.09* (.51)	1.20* (.49)	.10 (.57)	.25 (.53)
Month 5	.88* (.34)	.92* (.32)	1.25* (.44)	1.26* (.42)	.16 (.65)	.39 (.53)
Month 6	.67* (.34)	.72* (.33)	1.33* (.40)	1.37* (.38)	-1.05 (1.06)	-1.01 (1.02)
Month 7	.02 (.44)	.09 (.42)	.79 (.48)	.77 (.45)	-1.20 (1.09)	-.85 (1.03)
Month 8	.13 (.49)	.04 (.47)	.78 (.63)	.70 (.62)	-.68 (.80)	-.83 (.77)
Month 9	-.01 (.40)	.07 (.40)	.47 (.49)	.49 (.48)	-.84 (.78)	0.64 (.75)
Month 10	-.28 (.35)	-.21 (.35)	.42 (.42)	.47 (.41)	-1.56* (.77)	-1.55* (.76)
Month 11	-.41 (.78)	-.25 (.73)	1.13 (1.10)	1.15 (1.07)	-1.66 (1.10)	-1.88 (1.09)
Month 12	-.73 (.43)	-.62 (.41)	-.24 (.52)	-.24 (.50)	-1.27 (.78)	-1.38 (.77)
Month 13-15	-.43 (.30)	-.37 (.29)	-.21 (.42)	-.14 (.41)	-.63 (.45)	-.69 (.44)
Month 16-21	-.62 (.34)	-.64 (.33)	-.01 (.40)	-.01 (.38)	-2.01 (1.04)	-2.07* (1.02)
Month 22-24	-.85* (.39)	-.84* (.39)	-.34 (.45)	-.35 (.44)	-1.76 (1.04)	-1.62 (1.02)
Month 25-36	-1.32* (.45)	-1.35* (.44)	-.59 (.48)	-.56 (.47)	-17.91 (1622.01)	-17.88 (1622.44)
Age	-.36* (.08)	-.4* (.06)	-.38* (.10)	-.43* (.07)	-.29* (.14)	-.36* (.09)
Male	-.06 (.14)		-.01 (.16)		-.19 (.25)	
Catholic	-.13 (.15)		-.13 (.18)		-.13 (.29)	
Child	-1.13* (.30)	-1.01* (.28)	-.87* (.34)	-.74* (.32)	-1.83* (.65)	-1.75* (.62)
Father U	-.26 (.17)	-.30 (.16)	-.35 (.20)	-.35 (.19)	.07 (.29)	
Father SOC1	-.01 (.17)		-.14 (.22)		.32 (.30)	
GCSE5	.22 (.16)		.05 (.20)		.59* (.27)	.62* (.26)
Previous U	.01 (.01)		.00 (.02)		.06 (.04)	
Previous E	-.00 (.01)		.01 (.01)		-.10* (.04)	-.10* (.04)
Previous wage	.01 (.08)		.03 (.09)		-.01 (.20)	

**Table 4: Model, Pooled Exits and Competing Risks, Males and Females cont.**

	Pooled Exits		Exits to Employment		Exits to Education/Training	
	Inclusive Model	Parsimonious Model	Inclusive Model	Parsimonious Model	Inclusive Model	Parsimonious Model
No. Job applications	.05 (.03)	.05 (.03)	.07* (.03)	.07* (.03)	.05 (.07)	
Holiday	-.01 (.26)		-1.68* (.63)	-1.72* (.62)	.91* (.34)	.98* (.31)
Local U rate	.01 (.02)		.02 (.02)		-.00 (.03)	
West of NI	-.25 (.21)		-.36 (.25)		.02 (.38)	
JSA	.72* (.23)	.74* (.21)	1.23* (.27)	1.22* (.26)	-.58 (.52)	
New Deal 1	-.31 (.29)		-.26 (.32)		-1.42 (1.01)	
New Deal 2	.48 (.35)		.44 (.38)		1.83 (1.02)	
$\chi^2$	311.06* (32)	303.73* (20)	214.72* (32)	209.32* (21)	192.17* (32)	185.93* (20)
Log L	-702.69	-706.35	-548.61	-551.31	-276.85	-280.32

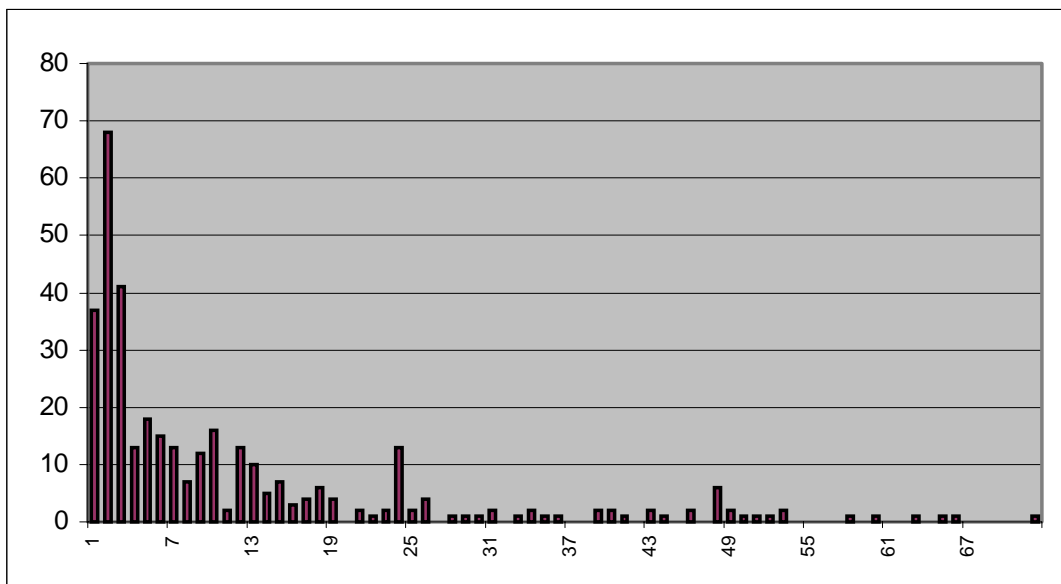
**Notes:** \* denotes significant at 5%. The  $\chi^2$  statistic tests the explanatory power of the model compared to the intercept-only model (degrees of freedom in parentheses). Parameters reported for explanatory variables are coefficients (standard errors in parentheses).

**Figure 1: Number of Spells of Joblessness, October 1993-March 1999**



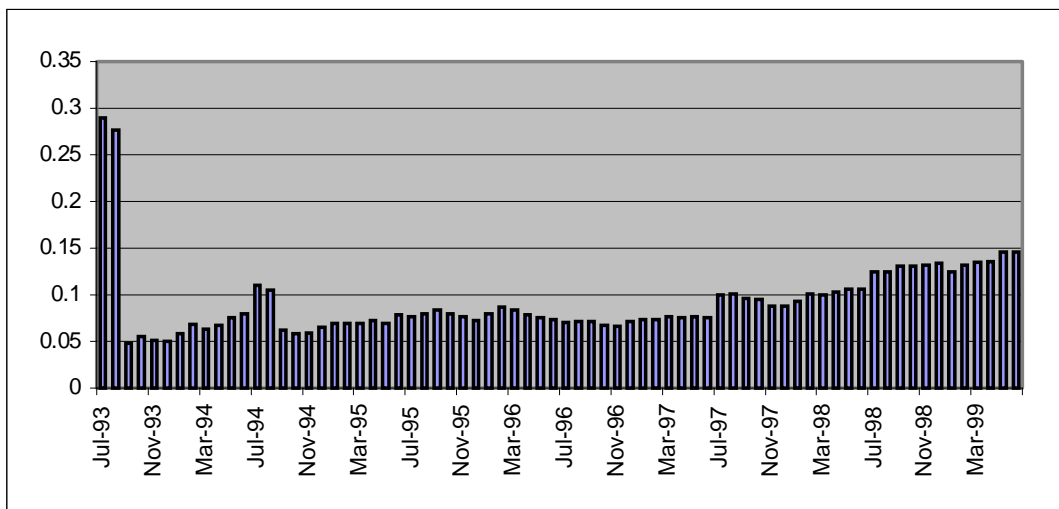
**Source:** 'Young People and Social Exclusion in Northern Ireland: Status 0 Four Years On'. Department of Higher and Further Education, Training and Employment. Belfast, 2000. Notes: No sample members experienced more than four spells of non-participation. Figures are rounded, based on weighted sample proportions.

**Figure 2: Frequency Distribution of Duration, Months**



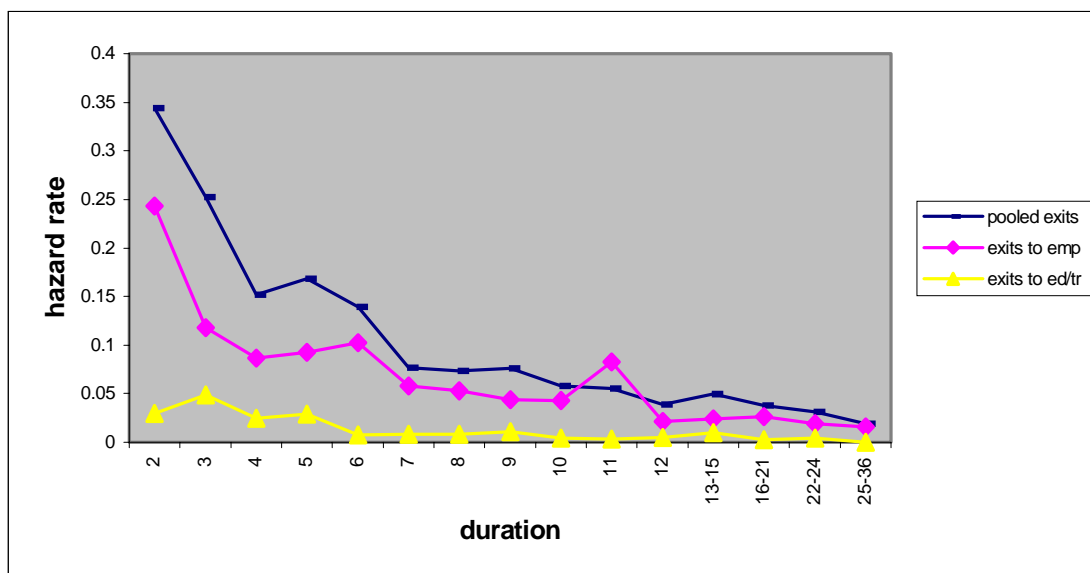
**Notes:** Includes censored spells. Unweighted.

**Figure 3: Proportion of Cohort Jobless, 1993-1999**



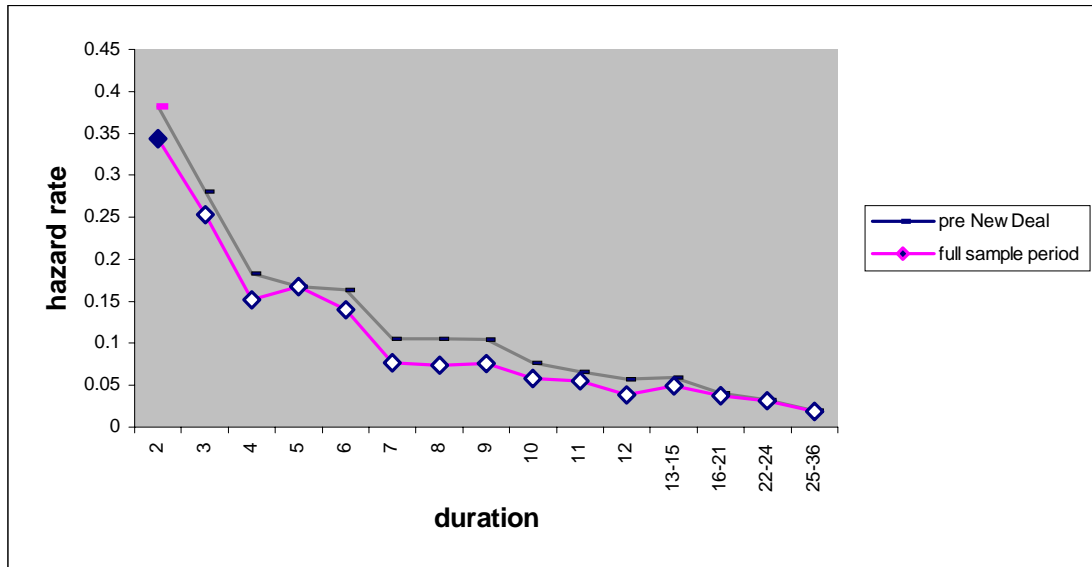
**Notes:** Figures give number of jobless as a proportion of the cohort. Weighted.

**Figure 4: Baseline Hazard Rates, PG Model, Pooled Exits and Competing Risks, 'Average Young Person'**



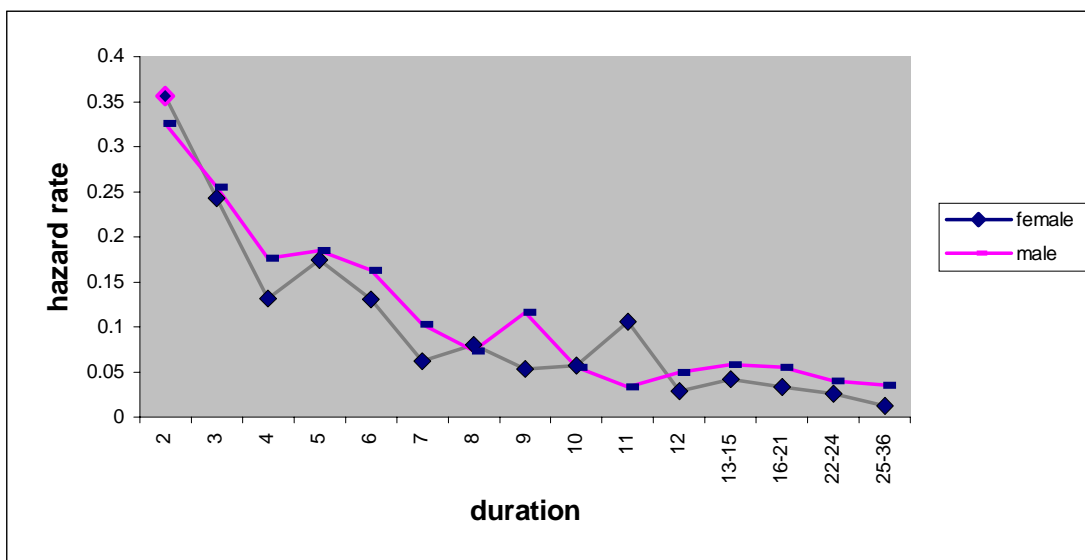
**Note:** All dummies are set to zero, continuous variables are set to their sample means.

**Figure 5: Baseline Hazards, Pooled Exits,  
Full Sample Period and Pre-New Deal Sample Period**



**Note:** Pre-New Deal sample period = spells starting July 1993 – March 1998, inclusive. Spells ending after March 1998 are treated as right-censored. All dummies set to zero, continuous variables set to sample means.

**Figure 6: Baseline Hazards, 'Average' Male, 'Average' Female, Pooled Exits**



**Note:** All dummies set to zero, continuous variables set to (gender-specific) sample means.

## References

- Arulampalam W, Booth A and Taylor M (2000) 'Unemployment persistence.' *Oxford Economic Papers*, 52, 24-50.
- Arulampalam W and Stewart M (1995) 'The determinants of individual unemployment durations in an era of high unemployment.' *Economic Journal*, 105, 321-332.
- Baker M and Melino A (2000) 'Duration dependence and non-parametric heterogeneity: A Monte Carlo study.' *Journal of Econometrics*, 96, 357-393.
- Beatty C and Fothergill S (1999) 'Incapacity benefit and unemployment.' Mimeo, Sheffield Hallam University.
- Blanchard O and Diamond P (1994) 'Ranking, unemployment duration and wages.' *Review of Economic Studies*, 61, 417-434.
- Bloemen H (1997) 'Job search theory, labour supply and unemployment duration.' *Journal of Econometrics*, 79, 305-325.
- Bloeman H and Stancanelli E (2001) 'Individual wealth, reservation wages and transitions into employment.' *Journal of Labour Economics*, 19, 2, 400-439.
- Boheim R and Taylor M (2000) 'Unemployment duration and exit states in Britain.' Mimeo, University of Essex.
- Bradley S and Taylor J (1991) 'An empirical analysis of the unemployment duration of school leavers.' *Applied Economics*, 24, 89-101.
- Bratberg E and Vaage K (2000) 'Spell durations with long unemployment insurance periods.' *Labour Economics*, 7, 153-180.
- Bratberg E and Nilsen O A (2000) 'Transitions from school to work and the early labour market experience.' *Oxford Bulletin of Economics and Statistics*, 62, 909-929.
- Carling K, Edin P A, Harkman A and Holmlund B (1996) 'Unemployment duration, unemployment benefits and labor market programs in Sweden.' *Journal of Public Economics*, 59, 313-334.
- Chuang Hwei-Lin (1999) 'Estimating the determinants of the unemployment duration for college graduates in Taiwan.' *Applied Economics Letters*, 6, 677-681.
- Crouchley R, Pickles A R and O'Farrell P N (1991) 'Duration of unemployment among men in Northern Ireland: regional inequality or denominational discrimination?' *Environment and Planning A*, 23, 903-920.
- Dearden L, Machin S and Reed H (1997) 'Intergenerational mobility in Britain.' *Economic Journal*, 107, 440, 47-66.

Devine T and Kiefer N (1991) *Empirical Labour Economics*. Oxford University Press, Oxford.

Dolton P and O'Neill D (1996)a 'Unemployment duration and the Restart effect: Some experimental evidence.' *Economic Journal*, 106, 387-400.

Dolton P and O'Neill D (1996)b 'The Restart effect and the return to full-time stable employment.' *Journal of the Royal Statistical Society Series A*, 159, 2, 275-288.

Firth D, Payne C and Payne J (1999) 'Efficacy of programmes for the unemployed: Discrete time modelling of duration data from a matched comparison study.' *Journal of the Royal Statistical Society Series A*, 162, 111-120.

Hannan C (1998) 'Beyond networks: Social cohesion and unemployment exit rates.' Mimeo, University of Essex.

Heckman J and Borjas G (1980) 'Does unemployment cause future unemployment? Definitions, questions and answers from a continuous time model of heterogeneity and state dependence.' *Economica*, 47, 247-283.

Heckman J and Singer B (1984) 'A method for minimizing the distributional assumptions in econometric models for duration data.' *Econometrica*, 52, 271-320.

Jenkins S P (1997) 'Estimation of discrete time proportional hazards models: pgmhaz.' Mimeo, ESRC Research Centre on Micro-Social Change, University of Essex.

Kettunen J (1997) 'Education and unemployment duration.' *Economics of Education Review*, 16, 2, 163-170.

Kollmann R (1994) 'The duration of unemployment as a signal.' *Economics Letters*, 45, 373-377.

Korpi T (1995) 'Effects of manpower policies on duration dependence in re-employment rates: the example of Sweden.' *Economica*, 62, 353-371.

Kraus F and Steiner V (1998) 'Modelling heaping effects in unemployment duration models – with an application to retrospective event data in the German Socio-Economic Panel.' *Jahrbucher f. Nationalokonomie u. Statistik*, 217, 5, 550-573.

Lancaster T (1990) *The econometric analysis of transition data*. Cambridge University Press, Cambridge.

Lancaster T (1979) 'Econometric methods for the duration of unemployment.' *Econometrica*, 47, 4, 939-956.

Layard R, Nickell S and Jackman R (1991) *Unemployment, Macroeconomic Performance and the Labour Market*. Oxford University Press, Oxford.

Lynch L (1989) 'The youth labour market in the eighties: determinants of re-employment probabilities for young men and women.' *Review of Economics and Statistics*, 71, 37-45.

Machin S and Manning A (1999) 'Longterm unemployment in Europe,' in Ashenfelter, O. and Card, D. (eds) *Handbook of Labour Economics*, Volume 3C. Amsterdam: North-Holland.

McVicar D (2000) 'Marginlised young people and social inclusion policy in Northern Ireland.' *Regional Studies*, 34, 9, 883-888.

Meyer B D (1990) 'Unemployment insurance and unemployment spells.' *Econometrica*, 58, 4, 757-782.

Mortensen D (1977) 'Unemployment insurance and job search decisions.' *Industrial and Labor Relations Review*, 30, 4, 505-517.

Mortensen D (1987) 'Job search and labour market analysis.' In *Handbook of Labour Economics*, Volume 2 (ed O Ashenfelter and R Layard). Amsterdam: North Holland.

Narendranathan W, Nickell S and Stern J (1985) 'Unemployment benefits revisited.' *Economic Journal*, 95, 307-29.

Narendranathan W (1993) 'Job search in a dynamic environment: An empirical analysis.' *Oxford Economic Papers*, 45, 1-22.

Narendranathan W and Elias P (1993) 'Influences of past history on the incidence of youth unemployment: Empirical findings for the UK.' *Oxford Bulletin of Economics and Statistics*, 55, 2, 161-185.

Narendranathan W and Stewart M (1993) 'Modelling the probability of leaving unemployment: Competing risks models with flexible baseline hazards.' *Applied Statistics*, 42, 1, 63-83.

Nickell S (1979) 'Estimating the probability of leaving unemployment.' *Econometrica*, 47, 1249-1266.

Ondrich J and Rhody S E (1999) 'Multiple spells in the Prentice-Gloeckler-Meyer likelihood with unobserved heterogeneity.' *Economics Letters*, 63(2), 139-144.

Phelps E (1972) *Inflation Policy and Unemployment Theory: The Cost Benefit Approach to Monetary Planning*, MacMillan, London.

Portugal P and Addison J (2000) 'Short and long term unemployment.' *Economics Letters*, 66, 107-112.

Prentice R and Gloeckler L (1978) 'Regression analysis of grouped survival data with application to breast cancer data.' *Biometrics*, 34, 57-67.

Rees A (1986) 'An essay on youth joblessness.' *Journal of Economic Literature*, 24, 2, 613-618.

Rice P (1999) 'The impact of local labour markets on investment in further education: evidence from the England and Wales Youth Cohort Studies.' *Journal of Population Economics*, 12, 287-312.

Ridder G (1987) 'The sensitivity of duration models to misspecified unobserved heterogeneity and duration dependence.' Mimeo, University of Amsterdam.

Riley R and Young G (2001) 'Does welfare-to-work policy increase employment? Evidence from the UK New Deal for Young People.' Mimeo, National Institute of Economic and Social Research, London.

Roed K, Raaum O and Goldstein H (1999) 'Does unemployment cause unemployment? Micro evidence from Norway.' *Applied Economics*, 31, 1207-1218.

Russel H and O'Connell P (2001) 'Getting a job in Europe: The transition from unemployment to employment among young people in nine European countries.' *Work, Employment and Society*, 15, 1, 1-24.

Social Exclusion Unit (1999) *Bridging the Gap: New Opportunities for 16-18 year olds not in Education, Employment or Training*. HMSO, London.

Stancanelli E (1999) 'Unemployment duration and the duration of entitlement to unemployment benefits: An empirical study for Britain.' *Applied Economics*, 31, 1043-1051.

Torelli N and Trivellato U (1993) 'Modelling inaccuracies in job search duration data.' *Journal of Econometrics*, 59, 187-211.

Van den Berg G and van Ours J C (1994) 'Unemployment dynamics and duration dependence in France, the Netherlands and the United Kingdom.' *Economic Journal*, 104, 432-443.

Van Ours J C, Zweimuller J and Lalive R (2000) 'The impact of Active Labour Market Policies and benefit entitlement rules on the duration of unemployment.' Centre for Economic Policy Research, Discussion Paper 2451.

Van Reenen J (2001) 'No more skivvy schemes? Active Labour Market Policies and the British New Deal for the Young Unemployed in context.' Mimeo, University College London.

Winter-Ebmer R (1998) 'Potential unemployment benefit duration and spell length: lessons from a quasi-experiment in Austria.' *Oxford Bulletin of Economics and Statistics*, 60, 3-46.

Wolbers M (2000) 'The effects of level of education on mobility between employment and unemployment in the Netherlands.' *European Sociological Review*, 16, 2, 185-200.