Young People, Unemployment Duration and the New Deal in Northern Ireland

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Executive Summary

In September 2001 the Department for Employment and Learning (DEL) commissioned a research team from the Northern Ireland Economic Research Centre (NIERC) and the University of Southampton (UoS) to examine the effects of the introduction of the New Deal for 18-24 Year Olds (ND18-24) on the duration of unemployment spells for the 18-24 year old age group in Northern Ireland (NI). This revised report sets out the findings of the NIERC/UoS study.

Young people between the ages of 18 and 24 years old must enter ND18-24 after being unemployed and claiming Jobseeker's Allowance (JSA) for six months. Some, mostly those with special needs, can choose to enter before this. Participants initially enter a Gateway stage involving one-to-one advice and assistance with job search. If still unemployed after four months on Gateway, participants join one of four options – subsidised employment, full time education and training (FTET), the environmental taskforce (ETF) or voluntary sector work. Voluntary sector work has been the most common option in NI. By November 2001 there had been 35,671 ND18-24 episodes in NI, roughly two thirds male and one third female. Around half of all ND18-24 episodes end with the young person finding employment, most of these 'sustained' employment (i.e. lasting three or more months).

Little quantitative analysis of the effects of ND18-24 on employment and unemployment for young people has been published to date in NI. Studies of the effects of ND18-24 in Britain (GB) suggest it may have boosted employment chances for young people by between 5% and 40%. Positive ND18-24 effects have previously been more easily identified for young men than for young women.

We investigate the effects of ND18-24 on the probability of the 18-24 age group leaving unemployment (i.e. leaving the JSA register) after different lengths of time spent unemployed by estimating *hazard functions* (which show this probability) before and after the introduction of ND18-24. Similar hazard functions for the 25-29 age group are used to estimate the *counterfactual* (i.e. what would have happened to the 18-24 age group had ND18-24 *not* been introduced). Exits to employment, exits to

education and training, exits to other benefits and other exits (mostly to unidentified destinations) are identified separately.

The data used in the study are a 20% random sample taken from the computerised unemployment register for 16-30 year olds in NI. The data span January 1995-July 2001 and contain details of 86,965 unemployment spells. Individual characteristics of sample members contained in the data are included in our study as covariates (regressors). To this we add a binary JSA dummy to control for the introduction of the JSA in October 1996, and the Noble Multiple Deprivation Index (MDI) score for the electoral ward in which sample members live.

Our analysis suggests that ND18-24 has increased the male hazard rate for exits to employment (the probability of getting a job) by a factor of almost 40% over unemployment durations between six and twelve months. In terms of magnitude, this figure is in line (and towards the top end) with existing estimates for GB. In terms of timing, our findings are also consistent with existing findings for GB – it is a direct programme participation effect that we observe. We find little evidence of an ND18-24 effect at durations shorter than six months, i.e. there is no evidence of ND18-24 'anticipation' or 'avoidance' effects of ND18-24 on young men before they enter the programme.

ND18-24 has increased the female hazard rate for exits to employment by a factor of around 20%, again primarily over unemployment durations between six and twelve months. This effect is smaller than that for males, but stronger than that found by some existing GB studies. The timing again suggests we are picking up a direct programme participation effect. There is also a small negative effect on the female hazard rate for exits to employment at short unemployment durations, suggesting a possible weak effect whereby some young women wait for entry to ND18-24 rather than take the first available job offer on entering unemployment.

The probability of exits to education and training for both genders has increased dramatically following the introduction of ND18-24, by a factor of around 300% at unemployment durations longer than six months. Although this effect is very strong, actual probabilities of exits to education and training are generally quite small,

so the 20-40% increase in exits to employment represents roughly the same *number* of additional exits from unemployment as the 300% increase in exits to education and training. There is a small negative effect of ND18-24 on exits to education and training at short unemployment durations, suggesting that some young people may wait to enter ND18-24 rather than taking the first available education or training place.

The probability of exits to other benefits for both genders has also increased following the introduction of ND18-24, by a factor of around 100% after six months of unemployment. There is no effect in earlier months of unemployment. The suggestion here is that on eligibility for ND18-24 an increased number of young people move off JSA onto alternative forms of benefits for which participation in ND18-24 is not compulsory. Again, although the effects are strong it must be remembered that the actual probabilities of exit to other benefits are small. Overall, the employment and education and training effects of ND18-24 far outweigh this effect on exits to other benefits.

Finally, the probability of other exits (primarily exits to unknown destinations) has also increased following the introduction of ND18-24, by a factor of around 70% at unemployment durations between six and twelve months. It is likely that some of these exits (generally believed to be around half) are exits to employment, although not recorded as such.

In short, ND18-24 has shifted significant numbers of young people in NI out of unemployment and into employment, education and training, onto other benefits and to other (unknown) destinations.

When we put the increased probabilities of exits to all the above destinations together, we can calculate the effects of ND18-24 on the chances of remaining unemployed for different lengths of time. **Our estimates correspond to a 45-50% reduction in the chances of a young male remaining unemployed for a full year since the introduction of ND18-24. For females, the corresponding figure is 40-45%.** These estimates control econometrically for other labour market trends and background

factors that might obscure the effects of ND18-24, i.e. they isolate the effects of ND18-24 itself.

Although the introduction of ND18-24 in NI has significantly reduced the chances of young people aged 18-24 experiencing spells of unemployment lasting one year or more, it has not *completely eradicated* such long-term unemployment. The guidelines of ND18-24 imply that very few young people in the 18-24 age group should be unemployed for longer than ten months[1] (after this time, in theory, they must enter a ND18-24 option). It appears, however, that ND18-24 is not being implemented strictly according to these guidelines – in particular, *Gateway* often lasts longer than four months. This has also been found to be the case in GB. Of course, there may be good reasons for these extended stays in Gateway, e.g. particular severe barriers to employment for a small number of young people.

An additional conclusion relates to the introduction of JSA in NI in October 1996. We find JSA to have had a significant positive effect on hazard rates for exits to employment for both genders and a small negative effect on exits to education and training. Overall, the introduction of JSA has coincided with a reduction in unemployment duration for young people in NI.

The study raises a number of questions for further research. We make five recommendations for future evaluations of ND18-24:

1. To merge the New Deal database held by DEL with the JSA-register database held by DETI.

2. Using this merged database, to evaluate the effects of ND18-24 also on inflows to unemployment in NI.

3. To evaluate what happens to young people that have passed through ND18-24 in the longer term, i.e. beyond their first destination on leaving the programme.

[1] The exceptions are those that complete three months of follow-through followed by six months of unemployment and a further four months on Gateway, i.e. 13 months.

4. To compare the effects of ND18-24 in NI with those in other UK regions.

5. To investigate the implementation of ND18-24 in NI, in particular the causes and consequences of overstaying on Gateway.

1. Introduction

In September 2001 the Department for Employment and Learning (DEL) commissioned a research team from the Northern Ireland Economic Research Centre (NIERC) and the University of Southampton (UoS) to examine the effects of the introduction of the New Deal for 18-24 Year Olds (ND18-24) on the duration of unemployment spells for the 18-24 year old age group in Northern Ireland (NI). This report sets out the findings of the NIERC/UoS study, first published in June 2002 and revised in March 2003.

The research set out to answer the following questions:

1. **How has ND18-24 affected the *duration* of unemployment spells for young people aged 18-24?**

2. **How large are the effects of ND18-24 on the probability of exit from unemployment – the hazard rate – after different unemployment durations and to different destinations?**

3. **After what duration of unemployment are the effects of ND18-24 on the hazard rate strongest?**

4. **Are there any unemployment durations over which ND18-24 has reduced the hazard rate for exit from unemployment?**

5. **Have the effects of ND18-24 on unemployment duration been uniform** across gender, age within the 18-24 age group, and location, or have they been heterogeneous?

The research uses individual level data that cover all spells of claimant count unemployment in NI beginning between January 1995 and July 2001. Access to these data was provided by NI's Department of Enterprise, Trade and Investment (DETI). Additional data were provided by DEL giving details of all individual episodes on

ND18-24, and all episodes on New Deal for 25+ (ND25+) for those aged 25-29 years in NI. The two databases are, unfortunately, as yet not matched. We cannot therefore identify the effect of *individual experiences* of ND18-24 in this study – we do not know from the JSA register which individuals have been on or are currently on ND18-24. Nevertheless, given that ND18-24 is compulsory for all 18-24 year olds that experience long-term unemployment, the effects of the *introduction* of ND18-24 on the target age group can be identified using the JSA register data. It will be necessary to merge the two databases in the future, however, if New Deal is to be more comprehensively evaluated in NI.

The research uses hazard functions – graphs showing the probability of leaving unemployment for a particular activity after having been unemployed for different lengths of time – to assess the effects of ND18-24. Hazard functions are presented for different exit destinations from unemployment – employment, education and training, other benefits and other exits – both before the introduction of ND18-24 and after the introduction of ND18-24. After controlling for other relevant factors, and using 25-29 year olds as a comparison group to control for general economic and labour market trends, the difference between the pre and post ND18-24 hazard functions measures the actual effects of ND18-24 on outflows from unemployment and unemployment duration for the target age group. The primary purpose of this report is to present and discuss these hazard functions in order to improve our understanding of the impact ND18-24 has had on outflows from unemployment and unemployment duration among young people in NI since its introduction.

The proposed research is the first to examine the effects of ND18-24 on unemployment duration explicitly using the hazard function. Policy makers, in NI particularly but also elsewhere in the United Kingdom (UK), will be able to draw on the analysis to improve understanding of the effects of ND18-24. It is only from such a knowledge-based position that policy makers can ensure the format and delivery of ND18-24 in NI and elsewhere is having its maximum potential net benefit.

Before continuing, let us briefly mention what this study does *not* examine. We do *not* examine (at least not in detail) the effects of ND18-24 on *inflow* rates to unemployment for young people. Therefore we do not draw any conclusions about the

effects of ND18-24 on the *unemployment rate* amongst young people in NI. Although it is certainly the case that *outflow* rates from unemployment have increased for the target age group following the introduction of ND18-24, if inflow rates have similarly increased there may have been little resulting change in the unemployment rate. This issue is left for further analysis.

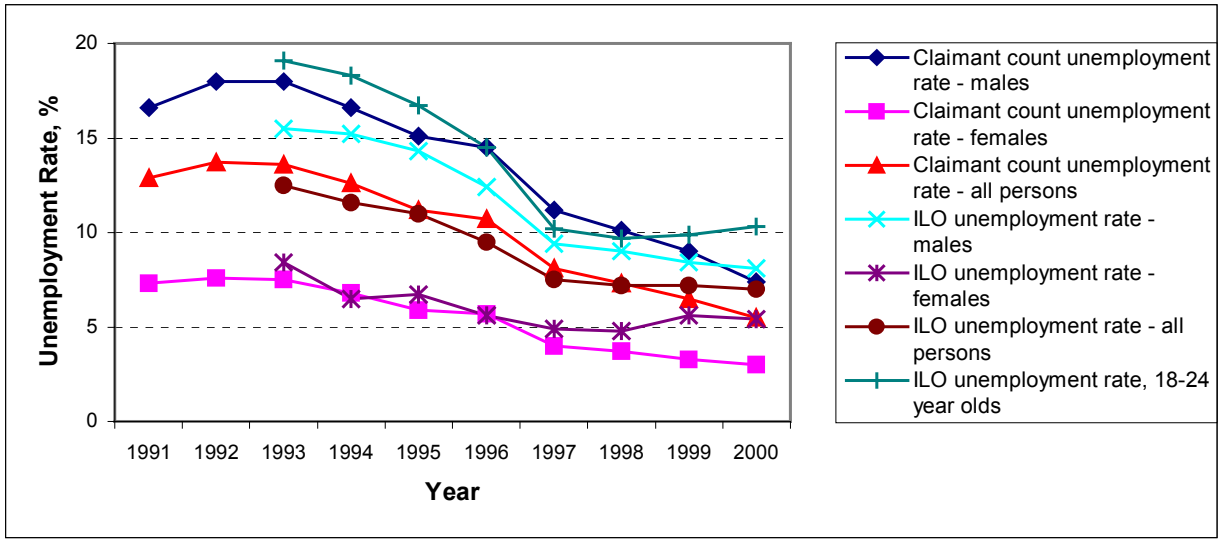
The following section presents the background to the study – discussing the New Deal policy and presenting some summary statistics for ND18-24 in NI, reviewing existing evaluations of ND18-24 in GB and NI, and setting out the economic context in NI over the last ten years. Section 3 provides a brief introduction to job search theory and duration analysis of unemployment (from where these hazard functions are derived) and sets out the model that we go on to estimate in this study. Section 4 introduces the data and presents some descriptive statistics for the unemployment register in NI (and for a 20% random sample used in estimation), for 16-30 year olds, over the period January 1995-July 2001. Section 5 presents and discusses the empirical analysis, drawing out the effects of ND18-24 on outflows from unemployment and unemployment duration. Section 6 concludes with a brief summary and a handful of recommendations. Appendices provide further detail on the data and estimation techniques used.

2. Background

2.1. The 1990s Labour Market and Policy Context in NI

The 1990s have seen substantial falls in unemployment in NI, as measured by both the claimant count (those unemployed, seeking work and claiming unemployment-related benefits) and the government's preferred International Labour Organisation (ILO) definition (those unemployed and seeking work). This fall has slowed somewhat in recent years, at least according to the ILO definition of unemployment. Figure 1 shows claimant count and ILO unemployment for NI, for all persons, and separately by gender, for 1991-2000. Unemployment rates for males are everywhere higher than for females. The ILO unemployment rate for 18-24 year olds is also shown. In all years this age group has higher unemployment rates than those for all adults. ILO unemployment for 18-24 year olds fell steeply between 1993 and 1997 and has remained roughly constant, at around 10%, since.

Figure 1: Unemployment Rates in NI 1991-2000



Source: NI Annual Abstract of Statistics.

The currently low levels of unemployment for NI mask significant geographical differences within the region, with claimant counts (all persons) in June 2000 of 9.4%

in Strabane Travel-to-work-area (TTWA) and 3.6% in Dungannon TTWA, for example. This compares to rates of 23.6% and 17.9% respectively in 1992.[2]

Long-term unemployment has also been falling. Around 36% of all persons defined as ILO unemployed in 2000 had been unemployed for at least one year, compared to around 50% in 1991.[3]

Employment rates have increased alongside falling unemployment rates over the 1990s. Over the same period, however, the number of people classed as economically inactive (neither employed nor seeking work) has also increased slightly. It is likely that most of those leaving unemployment have moved into employment. But a minority may have moved from unemployment benefits onto alternative benefits – a move off the unemployment register but not necessarily out of joblessness.

Prior to the introduction of ND18-24 and ND25+, the most significant policy change with direct implications for unemployment over the latter half of the 1990s was the introduction of the Jobseeker's Allowance (JSA) in October 1996. This was intended to tighten the relationship between claiming unemployment-related benefits and actively searching for work through regular interviews and monitoring of job search progress for all claimants. Benefit sanctions were possible for those not showing evidence of sufficient job search effort.

By introducing tougher job search requirements, JSA is likely to have reduced unemployment levels, at least according to the claimant count definition of those unemployed and claiming unemployment-related benefits. It is possible that the effects of JSA on ILO unemployment have been weaker, however, since those unemployed but not claiming unemployment-related benefits are not directly affected by the policy change. This may be one possible factor in the divergence of the claimant count and ILO unemployment rates since 1997. Beatty and Fothergill (1999) remark that some of those previously on the unemployment register may have moved from unemployment-related benefits onto other (alternative) benefits, such as Disability Living Allowance (DLA), for example, following the introduction of JSA.

[2] Source: Northern Ireland Annual Abstract of Statistics.
[3] Source: Northern Ireland Annual Abstract of Statistics.

2.2. The New Deal in NI

The need to tackle long-term unemployment among young people lies behind the government's introduction of ND18-24 – a key part of the government's Welfare to Work strategy – in April 1998, both in GB and in NI.[4] ND18-24 – a major active labour market policy for young people aged 18-24 – is geared towards intervention after a young person has been unemployed and claiming JSA benefits for six months. It offers individually tailored guidance and support, particularly in job search, during an initial *Gateway* period that is intended to last up to four months.[5] If at the end of that time the young person is still unemployed, a compulsory *option* must be taken up (or benefits can be withdrawn). These options include full time education or training (FTET), subsidised employment placements, or voluntary or environmental work (ETF). The subsidised employment option pays a wage and the voluntary or environmental options may also pay a wage or attract a small supplement in addition to the continuation of benefits. Options usually last for up to six months with FTET usually lasting for up to one year. Young people on a New Deal option are counted as having left the unemployment register. If, after completing a New Deal option, a young person is still without a job, they enter a *follow-through* stage (and go back on the unemployment register), with three months of further guidance and assistance in job search. At any time, a young person may leave ND18-24 to take up an unsubsidised job. Figure 2 shows a timeline for an ND18-24 episode. Van Reenen (2001) provides further information on the ND18-24 framework.

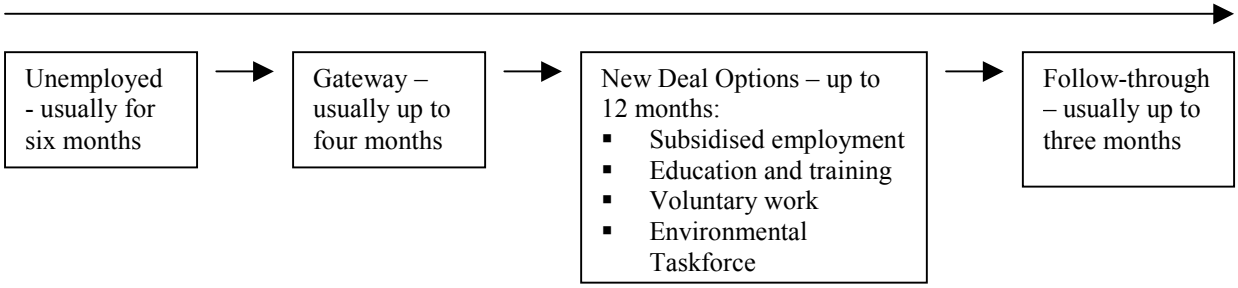
Some young people are eligible to join ND18-24 early, i.e. before they have been unemployed for six months, if they choose to do so. These include young people who would otherwise be unemployed for six months but for short breaks in claiming JSA (amounting to not more than 28 days), young people with a health condition or disability, young people needing help with basic skills, young people whose first language is not English, lone parents, and returners who have been out of the labour

[4] Some areas in GB (Pathfinder Areas) introduced ND18-24 (called New Deal for Young People (NDYP) in GB) in January 1998 to act as pilots. There were no such pilot areas in NI.

[5] In practice, Gateway sometimes lasts considerably longer than four months. 20% of all Gateway episodes fall into this category, with some episodes apparently lasting a year or more (Source: NIERC calculation from New Deal database).

market for two years or more. Also, those who leave unemployment at the Gateway stage, but who return to claimant unemployment within 13 weeks, automatically re-enter Gateway and do not have to wait an additional six months to re-qualify for ND18-24 (Riley and Young, 2001a).

Figure 2: ND18-24 Timeline



By November 2001 there had been 35,671 ND18-24 episodes, 25,676 male and 9,995 female in NI. This represents 29,396 people, 20,687 male and 8,709 female.[6] There were 4,876 episodes with early entry (before six months of unemployment), of which 3,159 were males and 1,717 females. There were 2,938 episodes in follow-through, of which 2,060 were males and 878 females. Notice that for all these headline figures, the gender split is roughly two thirds male and one third female, reflecting the claimant count unemployment rate differential between the genders (there are more males on the unemployment register). Numbers in ND18-24 were fairly evenly spread over its first four years, with the exception of higher numbers in 1998 as the 'stock' of existing young long-term unemployed entered the program (8,138 in 1998, 7,523 in 1999, 7,491 in 2000 and 6,115 in 2001 up to November).

Table 1 shows number of episodes in NI, by gender, on the various ND18-24 options up to November 2001. The most common is the voluntary sector option, followed by FTET, subsidised employment and ETF. Females are more likely than males to choose the FTET option and are very unlikely to choose the ETF option.

[6] Some young people enter ND18-24 more than once. An 'episode' ends when an individual has left an ND18-24 activity and not returned to another ND18-24 activity within 13 weeks. In Northern Ireland,

Table 1: Numbers of Young People Participating in ND18-24 Options

Option	Males	Females	All persons
Subsidised Employment	1,936	964	2,900
FTET	1,744	1,171	2,915
ETF	1,240	64	1,304
Voluntary Sector	2,697	1,530	4,227

Source: NIERC calculation from New Deal Database. Figures up to November 2001.

Of all the ND18-24 episodes, 17,409 (around half) ended with the young person finding work. Of these, 13,205 are described as 'found sustained employment', i.e. employment that lasted for at least three months. Over three-quarters of those episodes ending with finding work were for males.

ND25+ was introduced in June 1998 in NI and GB. Those aged 25 or over that have been in receipt of JSA benefits for 18 months or more are required to participate. As for ND18-24, job seekers are assigned a personal advisor during an initial *Gateway* period, where assistance in job search is offered. Those remaining unemployed after Gateway enter either subsidised employment, FTET or a 13 week Intensive Activity Period (IAP) providing work experience placements. An enhanced ND25+ was introduced in April 2001 offering a longer Gateway period and a longer Preparation for Employment Programme (PEP) to replace IAP (NI Assembly Oral Question AQO/334/01). Since its introduction, there have been a total of 29,392 ND25+ episodes in NI (25,198 male and 4,193 female).[7]

2.3. Existing Evaluations of the New Deal for 18-24 Year Olds

Despite some criticisms, ND18-24 in NI has been generally well received both among professionals working with young people and eligible young people themselves (see, e.g. McVicar et al., 2000; DEL, 2001). Quantitative evaluation in NI is at an early stage, however, and there are as yet no published estimates of the *net effect* of ND18-

23,906 young people have had one ND18-24 episode, 4,689 have had two, 717 have had three, 54 have had four and four have had five episodes. (Source: NIERC calculation from New Deal Database).

[7] Source: NIERC calculation from New Deal Database.

24 on employment and unemployment for young people, although some preliminary analysis is given in DEL (2001) and McVicar and Podivinsky (2001).

DEL (2001) outlines the findings of the first stage of a participant survey-based evaluation of ND18-24 in NI, based on interviews with 950 young people around nine months after first entering ND18-24 in summer 1999. As such it provides baseline information on the characteristics of ND18-24 participants and on their experiences to date of the delivery and implementation of ND18-24. So far, of the 950 responses to the survey, 170 are recorded as leaving ND18-24 because they have found work and 251 are recorded as leaving ND18-24 for 'other' reasons. 167 of the respondents to the survey had been on or were currently on the subsidised employment option, 127 were or are on the FTET option and a further 235 were or are on the voluntary sector option or ETF. Continued tracking of these individuals is likely to offer valuable insights about the longer term effects of ND18-24 in NI.

As part of a more general study on youth joblessness, McVicar and Podivinsky (2001) use longitudinal survey data for a cohort of young people first eligible to leave school in June 1993 (The Status Zero Survey) and followed until June 1999 to examine the effects of ND18-24 on joblessness duration in NI. They use hazard functions to estimate the probability of leaving joblessness pre and post ND18-24 after different durations of joblessness.[8] They find a 'spike' in the hazard function for exit from joblessness after 5-6 months duration following the introduction of ND18-24 which is not present pre-ND18-24, and argue that this is consistent with ND18-24 (and the anticipation of ND18-24) stimulating additional job search activity on entry and just before entry to Gateway. They also find some evidence that the probability of exit from joblessness in general has increased for males following the introduction of ND18-24, other things being equal. The study's sample period, however, only continues for 14 months following the introduction of ND18-24, and the sample size is too small for the authors to place particular emphasis on these results.

Evaluation in GB has been more extensive than in NI. In particular, two national studies – by the Institute of Fiscal Studies (IFS) and the National Institute of

[8] Joblessness is any activity outside employment, education or training, i.e. unemployment and other forms of economic inactivity. It is sometimes referred to as NEET or non-participation.

Economic and Social Research (NIESR) – have been carried out and are presented in papers by Van Reenen (2001), Blundell et al. (2001) and Riley and Young (2001a, b). Van Reenen (2001) also gives a good overview of previous evaluation studies for other labour market policies in the UK, EU and US.

Van Reenen (2001) uses administrative (claimant count) data up to the end of 1999 (the JUVOS database) to examine the effect of ND18-24 on the employment chances of young men in GB.[9] He uses both pilot areas where ND18-24 was introduced three months earlier than in other areas, and the 25-29 year old age group as comparison groups in order to identify ND18-24 effects with a difference-in-differences approach to estimation. Two critical assumptions are firstly that the 18-24 year old age group and the 25-29 year old age group react to macroeconomic trends in the same way, and secondly that there are no 'substitution' effects where 18-24 year old ND18-24 participants take jobs that would otherwise have been taken by 25-29 year old non-participants. Given these assumptions, Van Reenen estimates that ND18-24 has increased the chances of young men getting jobs by between 20% and 40%. This increased chance of getting a job includes subsidised jobs that are part of ND18-24 itself (the subsidised employment option).

Blundell et al. (2001) use the same data and adopt a similar approach to that of Van Reenen (2001) focussing on the effect of ND18-24 on outflows to employment during the Gateway stage of ND18-24 for males and females. They also find a ND18-24 boost to employment chances of between 20% and 40% for young men, but find no significant ND18-24 effect on employment chances for young women, partly because differential trends for the 18-24 age group and the 25-29 age group make identification of such effects problematical. They also find a 'program introduction effect' where the first three months of ND18-24 boosted employment chances for many young people that had been unemployed for well over six months (i.e. the 'stock' of existing young long-term unemployed).

Riley and Young (2001a) estimate matching functions between job seekers and job vacancies including ND18-24 as an explanatory (dummy) variable. As for the IFS

[9] ND18-24 is known as the New Deal for Young People (NDYP) in GB. For simplicity, we refer to it as ND18-24, however.

study, ND18-24 effects are identified by using older age groups and pilot/non-pilot areas as comparison groups. They find ND18-24 to have reduced measured unemployment among the target group partly by shifting them into non-work but also by raising employment. Outflow rates to employment for those 18-24 year olds unemployed for six months or more are estimated to have increased by between 5% and 15%.

Interestingly, Riley and Young also find tentative evidence of a small *negative* effect of ND18-24 on outflow rates for 18-24 year olds unemployed *less* than six months. This may represent an anticipation or substitution effect where those shortly to enter ND18-24 engage in less job search or where ND18-24 participants take jobs previously taken by young people that had been unemployed for a shorter duration. Riley and Young argue that it may also reflect higher *inflow* rates to short-term unemployment due to ND18-24 participants returning to the unemployment register following completion of ND18-24 options. Those still without a job following ND18-24 options are likely to be among the least 'employable' of the age group.

Riley and Young (2001b) examine the effects of ND18-24 on the wider economy, and finds some evidence of a beneficial effect on wage pressure.

To summarise Sections 2.1-2.3:

- Unemployment has fallen in NI over the 1990s.
- According to the ILO definition, unemployment has levelled off both for all adults and for 18-24 year olds, since 1997. Claimant count unemployment rates have continued to fall since 1997.
- Unemployment rates for young people are higher than for other age groups.
- Unemployment rates for males are higher than those for females.
- There are significant geographical differences in unemployment rates across NI.
- The number of adult long-term unemployed has fallen as a proportion of all unemployed, but rates remain high.
- A major policy change – JSA – was introduced in October 1996 and is likely to have reduced unemployment levels, especially as recorded by the claimant count.

- Young people must enter ND18-24 after being unemployed and claiming JSA benefits for six months. Some, mostly those with special needs, can choose to enter before this.
- Participants initially enter a Gateway stage involving one-to-one advice and assistance with job search.
- If still unemployed (officially) after four months on Gateway, participants join one of four options – subsidised employment, FTET, ETF or voluntary sector work. Voluntary sector work has been the most common option in NI.
- By November 2001 there had been 35,671 ND18-24 episodes in NI, roughly two thirds male and one third female. 4,876 of these started early, i.e. before six months of unemployment.
- Around half of all ND18-24 episodes end with the young person finding employment, most of these 'stable' employment.
- Little quantitative analysis of the effects of ND18-24 on employment and unemployment for young people has been published to date in NI.
- Studies of the effects of ND18-24 in GB suggest it may have boosted employment chances for young people by between 5% and 40%. Positive ND18-24 effects have been more easily identified for young men than for young women.

3. The Model

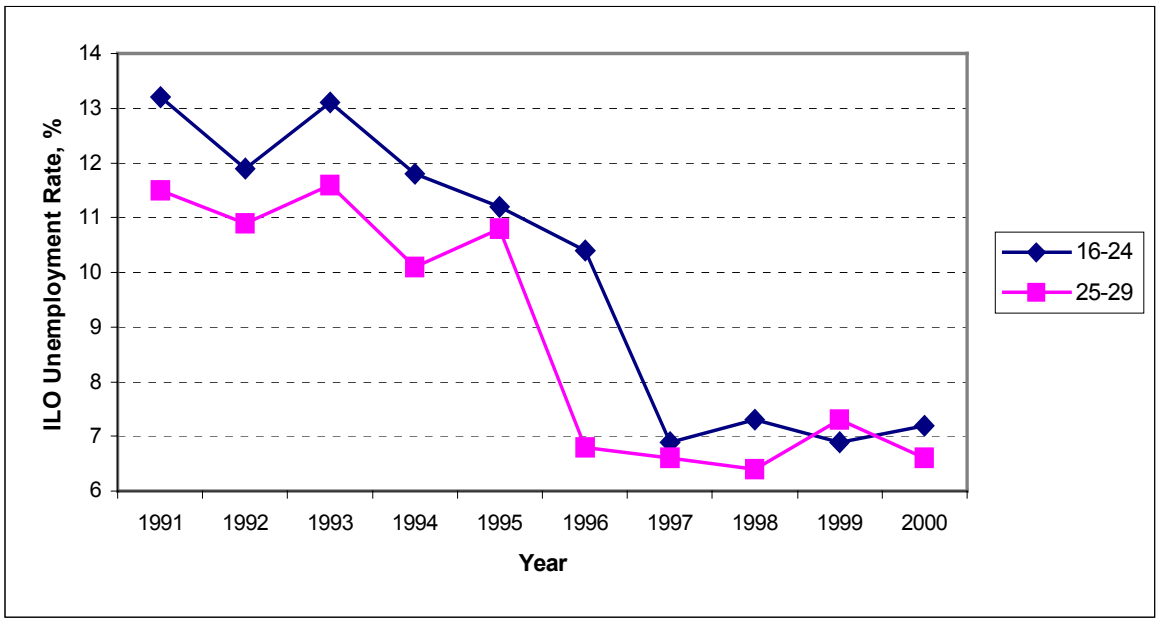
This research uses hazard functions – graphs showing the probability of leaving unemployment for a particular activity after having been unemployed for different lengths of time – to assess the effects of ND18-24. Hazard functions are presented for different exit destinations from unemployment – employment, education and training, other benefits and other exits – both before the introduction of ND18-24 and after the introduction of ND18-24. After controlling for other relevant factors, and using 25-29 year olds as a comparison group to control for economic and labour market trends, the difference between the pre and post ND18-24 hazard functions essentially measures the effects of ND18-24 on outflows from unemployment and unemployment duration.

This methodology relies on the same two critical assumptions behind the Van Reenen (2001) study outlined in the previous section: (i) that economic trends affect the 18-24 and 25-29 age groups in the same way and (ii) that there are no substitution effects between age groups as a result of ND18-24.

The first assumption can be defended by an examination of unemployment rates over time for the two age groups, as shown in Figure 3 below. Bearing in mind the statistical error involved with the Labour Force Survey (LFS) in NI for narrow age bands, the unemployment rates for 16-24 year olds and 25-29 year olds have followed broadly similar paths over the 1990s, certainly up to the introduction of ND18-24 in 1998.

Assumption (ii) is more a matter of faith, however, although no evidence has yet been presented to suggest such inter-age group ND18-24 substitution effects do exist to any significant degree. What we do know is that if such effects do exist, and ND18-24 has led to the younger age group taking jobs that would have been taken by the older age group had it not been for ND18-24, then our estimated ND18-24 effects on the 18-24 age group may be biased upwards. In other words, they will represent an *upper bound* to the true ND18-24 effects. It is less likely that any substitution effects work in the opposite direction, i.e. that ND18-24 has resulted in 25-29 year olds taking jobs that would previously have been taken by 18-24 year olds had it not been for ND18-24.

Figure 3: ILO Unemployment Rate by Age Group, 1991-2000, %.



Source: DETI (LFS).

One further assumption (assumption (iii)) is implicit in our methodology, namely that unemployed young people before the introduction of ND18-24 have similar characteristics as unemployed young people after the introduction of ND18-24, or at least that characteristics change in similar ways for both age groups. We compare observed characteristics for unemployed 18-24 year olds pre and post ND18-24 in Section 4, finding very little contrast. Nevertheless, if post ND18-24 unemployed young people are in some *unobserved* way less employable than their counterparts before the introduction of ND18-24, and the same is not true of 25-29 year olds, we may be under-estimating the true effects of ND18-24 on the hazard rate.

Finally, hazard rates beyond 18 months unemployment duration for the 25-29 age group are likely to have been affected by the ND25+. Since we are primarily interested in effects over the first 18 months of unemployment (and more particularly the first 12 months of unemployment), however, the ND25+ is unlikely to have any significant implications for our interpretation of the effects of ND18-24.

The following sections briefly discuss the economics and statistics behind hazard functions for exit from unemployment and set out the particular model we use to estimate the hazard functions presented in Section 5.

3.1. Duration Analysis of Unemployment and Job Search Theory

Search theory states that unemployed (young) people search for jobs at varying rates. The probability of a 'match' between a job seeker and a job vacancy at a particular point of time is the product of the probability of getting an offer and the probability of accepting that offer (see, e.g. Mortensen, 1987). These probabilities in turn will depend on factors such as search intensity, individual and local labour market characteristics. McVicar and Podivinsky (2001) argue that search theory can be extended for young people to include matches between education and training vacancies and unemployed young people seeking education and training places as alternatives to jobs.

If reservation wages (the minimum level of wages at which a young person would be willing to work) fall with unemployment duration then job offers will be accepted more readily the longer a young person remains unemployed. This is called *positive duration dependence* (see, e.g. Mortensen, 1977). In other words, if we have positive duration dependence, the probability of a young person leaving unemployment for a job at any particular time would increase with time spent unemployed. If, however, an unemployed young person loses motivation for job search the longer they have been unemployed, this might reduce the probability of getting a job offer, implying *negative duration dependence* (see, e.g. Layard et al., 1991). Negative duration dependence might also be caused by depreciation of human capital (skills) during an unemployment spell (e.g. Phelps, 1972), or because employers might see unemployment as a negative productivity signal and be reluctant to hire previously unemployed workers (e.g. Blanchard and Diamond, 1994). These arguments may, at least to an extent, also apply to exits to education/training (McVicar and Podivinsky, 2001).

The nature of duration dependence is reflected in the slope of the hazard function. Downward sloping hazard functions correspond to negative duration dependence and

upward sloping hazard functions to positive duration dependence. If there is no duration dependence hazard functions are horizontal. Non-linear duration dependence (e.g. some combination of positive and negative duration dependence) is also possible, and this would correspond to a non-linear hazard function (e.g. an inverted U-shape).

Empirical evidence on the shape of hazard functions for exits from unemployment has been mixed, both across and within countries. Reviews of this literature are provided by Devine and Kiefer (1991) and Machin and Manning (1999). Some studies for GB have found downward sloping hazard functions, i.e. negative duration dependence (e.g. Lancaster, 1979; Nickell, 1979; van den Berg and van Ours, 1994). Others find more or less horizontal hazard functions, i.e. no duration dependence (e.g. Narendranathan *et al*., 1985; Narendranathan, 1993). Others still find non-linear and non-monotonic hazard functions, such as the 'inverted u-shape' relationships found by Arulampalam and Stewart (1995) and Boheim and Taylor (2000). Finally, some studies find evidence of 'spikes' in the hazard function (temporary jumps in the probability of exit from unemployment) 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).

Studies examining duration dependence in unemployment for young people specifically are not common, but internationally Heckman and Borjas (1980), Lynch (1989), Korpi (1995), and Russell and O'Connell (2001) between them find evidence of no duration dependence, negative duration dependence and non-linear duration dependence. McVicar and Podivinsky (2001) find evidence of generally downward sloping hazard functions for young people in NI, although they also find some evidence that there may be a spike in the hazard function after 5-6 months of joblessness.

Although mixed empirical results on the shape of hazard functions across countries, regions and age groups may well reflect actual differences in the nature of duration dependence, another factor likely to be contributing to this ambiguity is the near observational equivalence of duration dependence and unobserved heterogeneity (differences between individuals that we do not generally observe). In other words,

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 (McVicar and Podivinsky, 2001). Failure to consider unobserved heterogeneity can lead to a false conclusion of negative duration dependence because of negative bias (Heckman and Borjas, 1980).

There are different ways of approaching this problem of unobserved heterogeneity, including specifying a 'random effects' type of error term for each individual, either parametrically (according to some standard statistical distribution) or non-parametrically (essentially assigning people to a number of different discrete groups). These methods are discussed in more detail in Appendix B. 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.

A better way to overcome this problem may be to adopt a more flexible specification for the hazard function than those commonly used in earlier studies (Arulampalam and Stewart, 1995). Boheim and Taylor (2000) go further, arguing that given a sufficiently flexible specification for the hazard function it becomes unnecessary to model unobserved heterogeneity at all. Many recent studies adopt such a flexible specification for the hazard function, based on a model first introduced by Prentice and Gloeckler (1978). Examples of these studies include Arulampalam and Stewart (1995); Narendranathan and Stewart (1993); Bratberg and Vaage, (2000); Portugal and Addison, (2000); Boheim and Taylor (2000) and McVicar and Podivinsky (2001). This study adopts such a flexible specification – the now widely used Prentice-Gloeckler (PG) methodology, discussed in the following section.

3.2. Prentice-Gloeckler Estimation of Hazard Functions

What follows is a brief description of PG estimation of hazard functions as used in this study. Further detail is given is Appendix B.

The hazard rate (probability of exit, denoted h_{ijt}) for individual i in time interval t to exit unemployment to destination j (e.g. employment, education/training or other benefits) is given by Equation 1, where $\theta_j(t)$ is a function relating the hazard rate with the duration of the unemployment spell. The PG approach assumes there is a specific parameter that is constant over each period identified, i.e. the function $\theta_j(t)$ is piecewise constant.

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

X_{it} is a set of observed explanatory variables, or *covariates*, that are also likely to affect the probability of exit from unemployment, as suggested by search theory and previous empirical studies. These covariates are listed in Section 4 and defined in Appendix A. β_j is a set of coefficients corresponding to the covariates which measure the relationship between each covariate and the hazard rate. The covariates can essentially be thought of as shifting the hazard function up or down, depending on the sign and size of β_j. A young person from Dungannon, for example, might be expected to have a hazard function parallel to but above that of a young person from Strabane because of Dungannon's relatively strong labour market.

Because we identify several destinations after leaving unemployment we adopt the competing risks PG model (see, e.g. Boheim and Taylor, 2000). This model treats exits to different destinations as independent, i.e. the probability of exit to one destination (e.g. employment) is assumed not to depend on the probability of exit to a different destination (e.g. education/training). The following section gives details of the separate destinations identified in this study.

The PG model controls for those unemployment spells that have not been completed by July 2001 (the last point at which we observe the data) by treating them as right-censored (see Appendices A and B). When estimating pre ND18-24 hazard functions, we also treat any unemployment spells that start before April 1998 but do not end before April 1998 as right-censored.

To summarise Section 3:

- We identify the effects of ND18-24 on the probability of the 18-24 age group leaving unemployment to various destinations after different lengths of time spent unemployed by estimating hazard functions pre and post ND18-24's introduction.
- Our approach rests on two critical assumptions: (i) that economic trends affect the 18-24 and 25-29 age groups in the same way and (ii) that there are no substitution effects between age groups as a result of ND18-24.
- A third assumption is that the unemployed pre ND18-24 have the same characteristics as the unemployed post ND18-24.
- Duration analysis of unemployment stems from the economic theory of job search.
- Negative duration dependence, i.e. that the probability of leaving unemployment falls the longer you are unemployed, arises if motivation or skills are lost during unemployment, or if employers see unemployment as a signal of poor productivity.
- Hazard functions show the duration dependence relationship. Negative duration dependence corresponds with a downward sloping hazard function.
- Existing empirical studies find mixed results including upward sloping hazard functions, downward sloping hazard function, flat hazard functions and non-linear hazard functions. The only existing study of duration dependence for young people in NI finds a generally downward sloping hazard function.
- Unobserved differences between individuals may partly explain mixed empirical results in the past, although there are various ways of 'dealing' with this problem. The 'cutting-edge' methodology is currently PG techniques, which we adopt in this study.
- The PG approach assumes the hazard function is piecewise constant, i.e. that the hazard rate at different durations of unemployment is unrestricted.
- Observed covariates shift the hazard rate up or down depending on their estimated sign.
- We identify different types of exits from unemployment separately (exits to employment, education and training, to other benefits and other exits) in a competing risks framework.

4. The Data

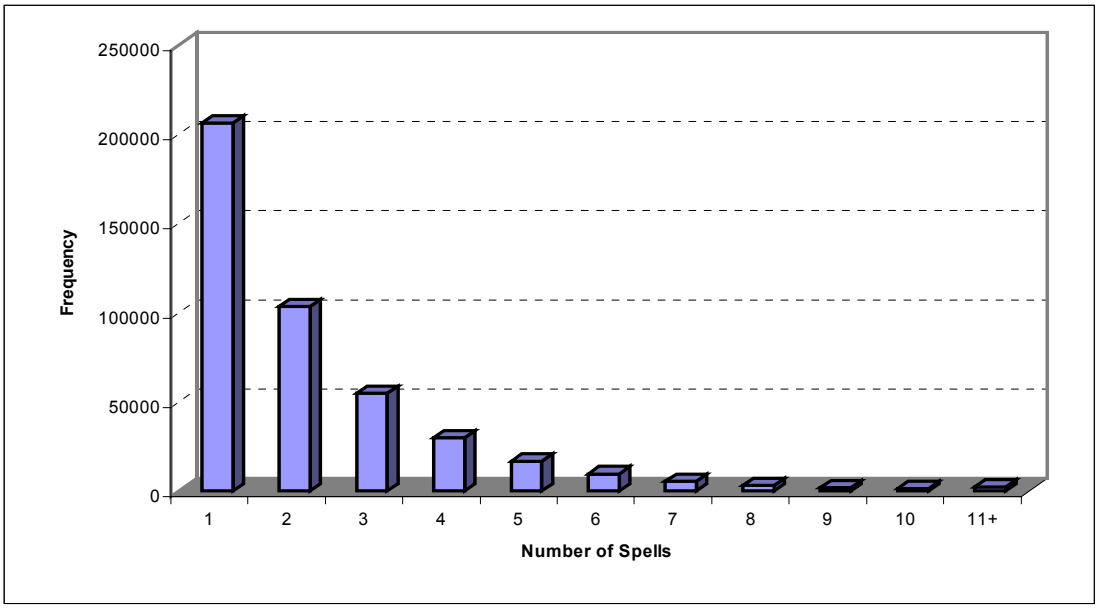
The data we analyse in this study are computerised unemployment register records for all those unemployed and claiming JSA benefits (claimant count) in NI aged between 16 and 30 on entry to unemployment. All spells on the register since January 1995 are recorded (we have excluded those spells that start before January 1995, where coverage of the data is not universal) and the last point of observation is July 2001. The data therefore span six and a half years, or 335 weeks, and are partly longitudinal (i.e. we can track individuals who return to the register through time). There are 434,826 unemployment spells recorded in the data set covering a total of 206,317 individuals. Section 4.1 presents descriptive statistics for the entire data set (the population of JSA claimants in NI aged 16-30). Section 4.2 presents data for a random 20% sample of the population for the 18-24 and 25-29 age groups separately. Because of the size of the data set, sampling was necessary in order to carry out estimation.

4.1. Some Descriptive Statistics for the Population of JSA Claimants Aged 16-30 in NI

Figure 4 shows the number of spells of unemployment experienced by individuals in the data set over the period January 1995 – July 2001. Just under half (47.4%) experience just one spell, with the others experiencing multiple spells. The maximum number of spells experienced by an individual over the period is 35. The mean number of spells is 2.2.

Although duration of spells on the register is expressed in days, for our purposes we round the length of spells to the nearest week. Figure 5 shows the frequency distribution of the duration of all spells between January 1995 and July 2001. The mean duration of spells is 20.5 weeks with standard deviation 31.2 weeks. The minimum duration is 0 weeks (three days or less) and the maximum 335 weeks.

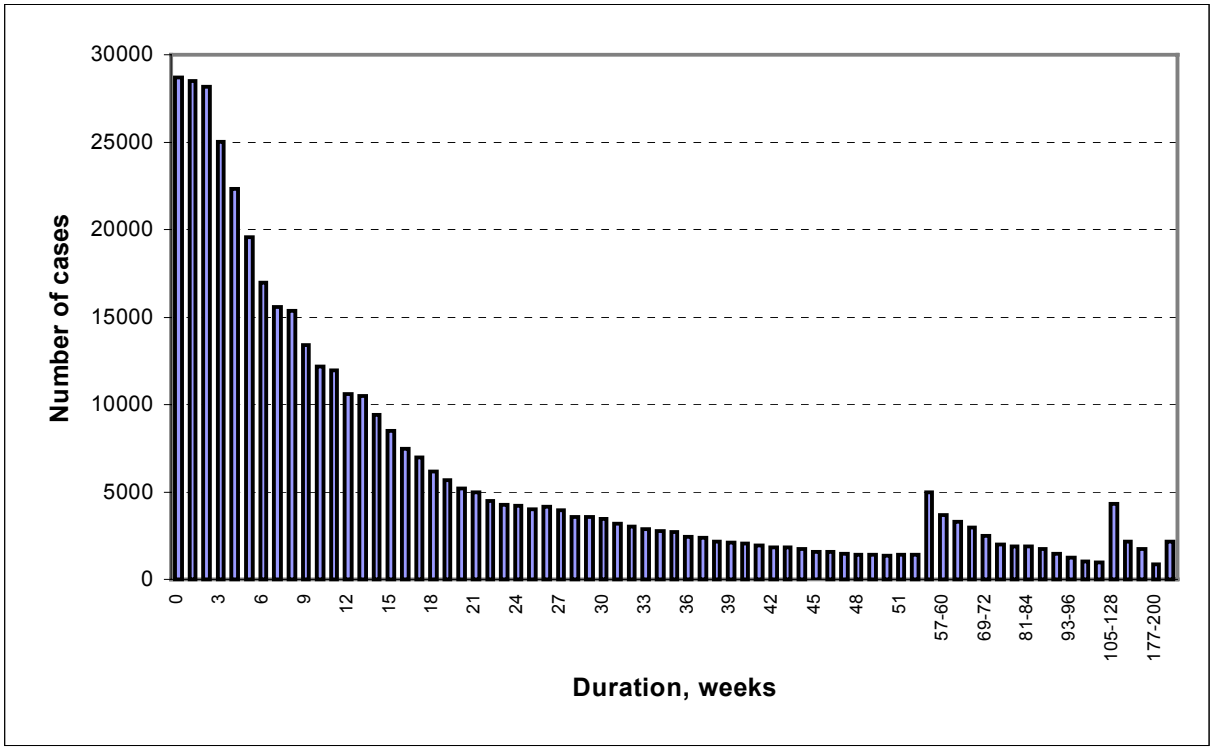
Figure 4: Frequency of Number of Spells of Unemployment Experienced by Individuals, Ages 16-30, January 1995 – July 2001



Source: DETI.

We identify four separate categories of exits from the unemployment register – exits to employment, exits to education and training, exits to other benefits, and other exits. The latter category includes (in order of importance) failure to sign, not known, emigration, claim withdrawn, court/prison and death. Table 2 shows the frequencies of each category of exit for all spells covered by the data, by broad age group. The most common type of exit, for both age groups, is into employment. For the 18-24 age group, exits to education/training and exits to other benefits are roughly equal in frequency. For the 25-29 age group, however, exits to education/training are considerably less frequent compared to exits to other benefits. Overall, exits to employment and education and training account for 60% and 58% of all exits from the register for 18-24s and 25-29s respectively.

Figure 5: Frequency of Spell Duration, All Spells, Ages 16-30, January 1995 – July 2001, Weeks



Source: DETI. Notes: Duration expressed to the nearest whole week. Includes right-censored spells (i.e. incomplete spells at last point of observation).

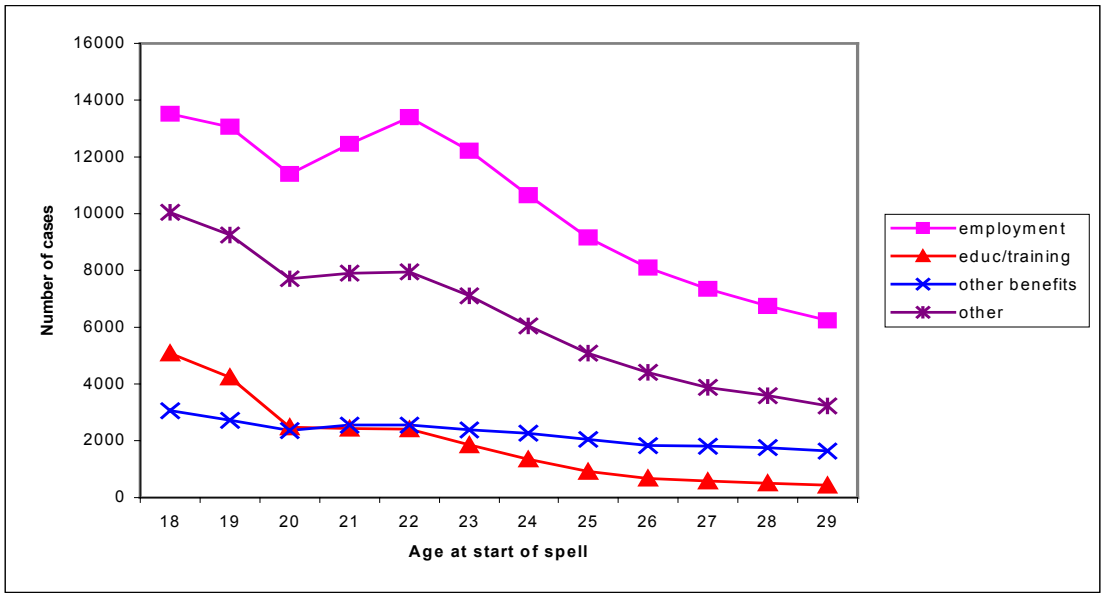
Table 2: Exits from Unemployment – Alternative Destinations, All Spells, Ages 18-24 and 25-29, January 1995 – July 2001

Destination	18-24 No. of Exits	25-29 No. of Exits
Employment	145,728	56,477
Education/Training	34,447	4,434
Other Benefits	32,454	13,920
Other	87,651	30,019
Total	**300,280**	**104,850**

Figures 6 and 7 give the frequency of exits to the various destinations by age, separately for males and females. The trend with age for exits to other benefits to gradually replace exits to education and training as the third most common destination is clear in both cases. The number of exits to employment fall with age until age 20

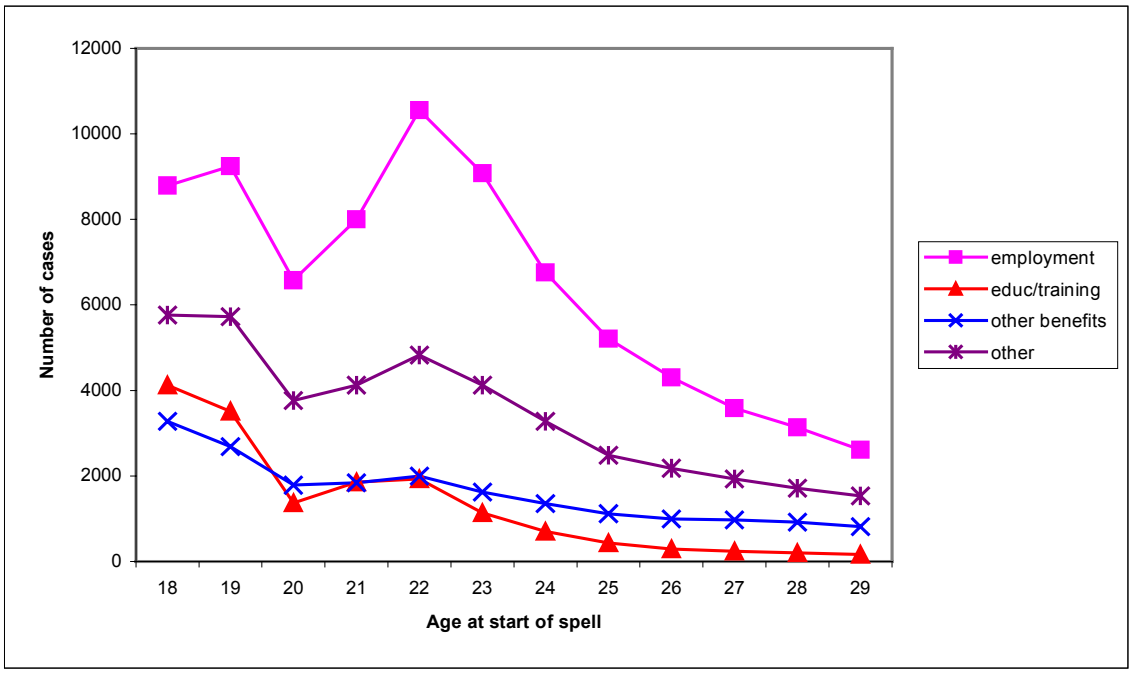
then rise until age 22 and then fall off once more, for both males and females. Exits to employment as a proportion of all exits does not change markedly with age, with the exception of ages 18-20 where relatively less exits to employment occur.

Figure 6: Destinations on Leaving Unemployment by Age on Entry, Males



Source: DETI.

Figure 7: Destinations on Leaving Unemployment by Age on Entry, Females



Source: DETI.

Of the total number of spells observed in the data, 318,079 (73.15%) are observed for the 18-24 age group and 112,883 (25.96%) are observed for the 25-29 age group. 62% of spells observed in the data set are for males and 38% for females. This broadly corresponds to the two-thirds/one-third gender split observed for claimant count unemployment rates (see Figure 1).

4.2. Descriptive Statistics for the 20% Sample

The PG model could not be estimated on the full unemployment register population with available computing resources. The analysis is therefore conducted on a random 20% sample of all JSA claimants aged 16-30 between January 1995 and July 2001. This 20% sample, containing 86,965 unemployment spells, is very closely representative of the full population.

To set up the PG model, we need to specify the duration intervals for which interval-specific parameters giving the shape of the hazard function are to be estimated. Although we can pinpoint the duration of an unemployment spell to the nearest day, some degree of aggregation is necessary for two reasons. Firstly, aggregating the duration intervals significantly cuts down on the computing power needed to estimate the model.[10] Secondly, the interval-specific hazard parameters can only be identified for those duration intervals in which exits occur, and can be imprecise where only few exits occur, therefore Jenkins suggests additional grouping of duration intervals in which the number of exits is small (Jenkins, 1997). In this case, the number of exits from unemployment is smaller at longer durations, therefore we group unemployment durations according to Table 3 below.

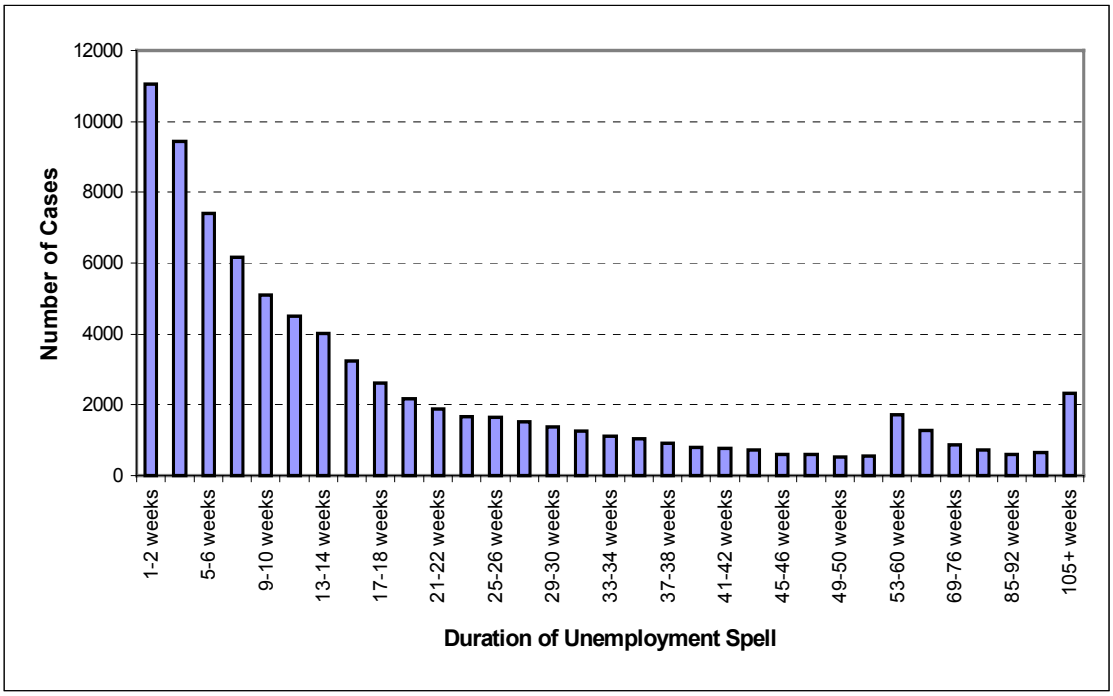
The distribution of spell durations based on this grouping is shown in Figure 8 below. As for the full population, durations are distributed as smooth curve (see Figure 5).

[10] Computing power is a major constraint with a data set this large when conducting PG estimation – several rows of data are required for each spell of unemployment, with the number of rows proportional to the length of the unemployment spell (see Section 5).

Table 3: Grouping of Unemployment Durations

Duration Group	Duration of Unemployment Spell
1-26	Fortnightly, i.e. Group1 = 1-2 weeks, Group2 = 3-4 weeks etc.
27-31	8-weekly groups, i.e. Group27 = 53-60 weeks
32	12-weekly group, weeks 93-104
33	105+ weeks

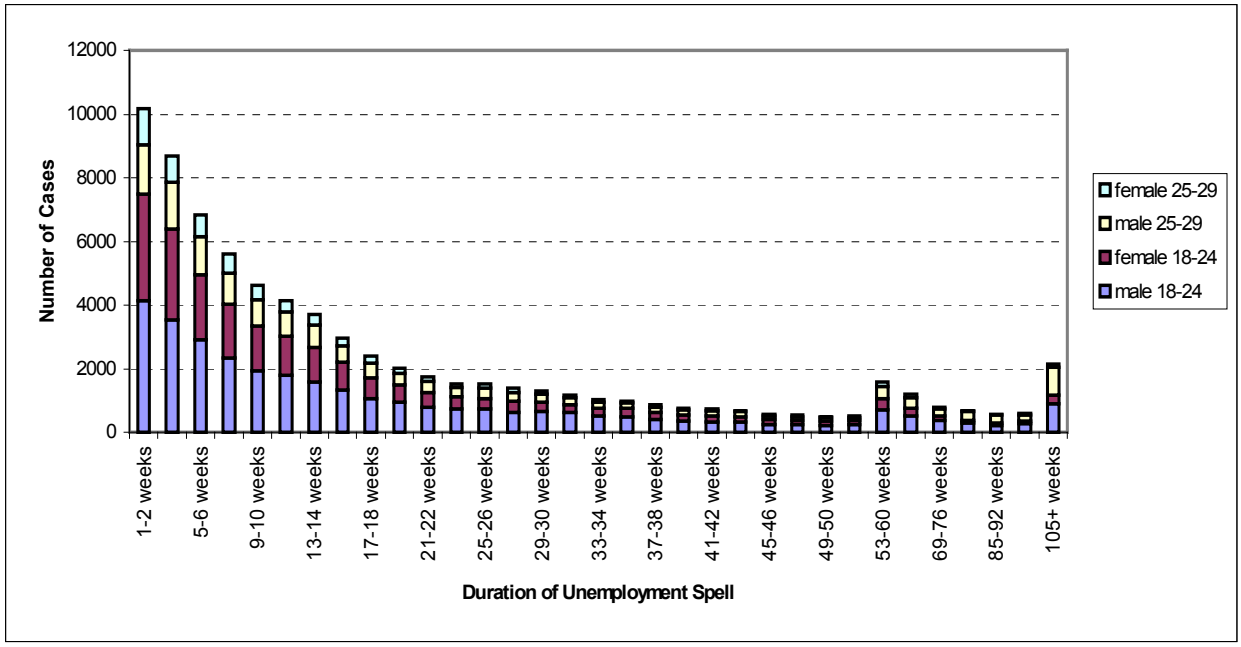
Figure 8: Frequency Distribution of Spell Duration, 20% Sample, PG Model Duration Groupings, January 1995-July 2001, Ages 16-30



Source: NIERC/DETI.

Figure 9 breaks down the distribution of duration of spells by age group (18-24s, 25-29s) and gender. Males aged 18-24 years are the largest group in all duration categories, but males aged 25-29 years are over-represented at the longer durations compared to their weight in the sample overall. In other words, Figure 9 suggests older males on entry to unemployment are more likely to experience unemployment durations of one year or more than younger males. Females in both age groups are skewed towards the shorter durations, relative to males.

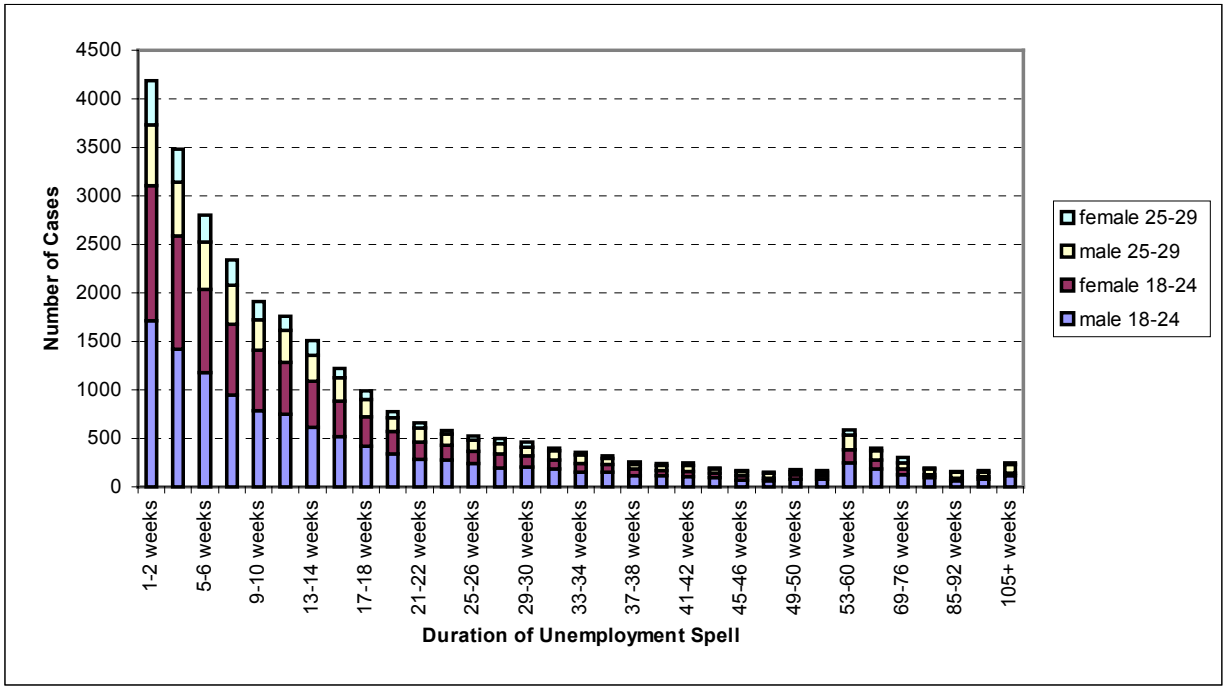
Figure 9: Frequency Distribution of Spell Duration, 20% Sample, PG Model Duration Groupings, January 1995-July 2001, by Age Group and Gender



Source: NIERC/DETI. Notes: Age on entry to unemployment. Includes right-censored spells.

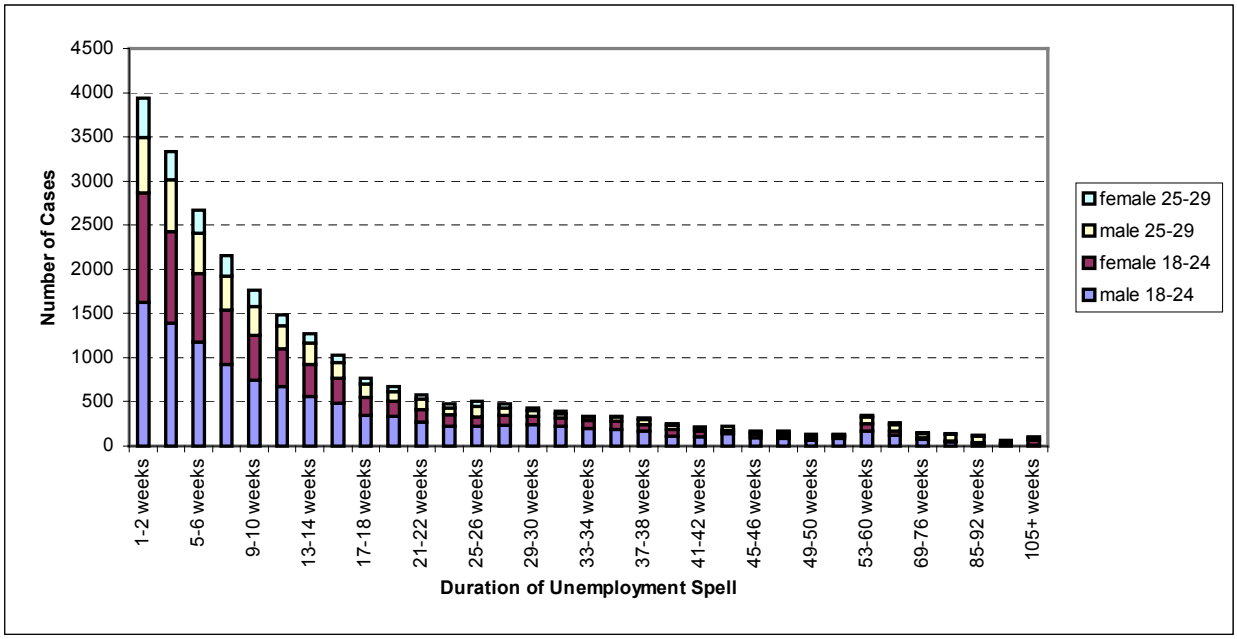
Figures 10 and 11 repeat the exercise of Figure 9 for the sample split into pre and post ND18-24. There is little immediately apparent contrast between the two figures – both are similar to Figure 9. There is, however, slightly more weight in the tail of the distribution for the pre ND18-24 sample. In other words, durations of one year or longer were more common pre ND18-24 than since the introduction of ND18-24. This appears to be not only true of 18-24 year olds but of both age groups (and both genders), which suggests it may not be directly related to ND18-24.

Figure 10: Frequency Distribution of Spell Duration, 20% Sample, PG Model Duration Groupings, by Age Group and Gender, Pre ND18-24



Source: NIERC/DETI. Note: Right censored spells are omitted from the sample**.**

Figure 11: Frequency Distribution of Spell Duration, 20% Sample, PG Model Duration Groupings, by Age Group and Gender, Post ND18-24



Source: NIERC/DETI. Note: Right censored spells omitted from sample.

As in Table 2, we can identify the number of unemployment spells that end in particular types of exits for the 20% sample, not only by age group, but also by gender and separately for pre and post ND18-24. This is shown in Table 4.

Table 4: Number of Unemployment Spells Ending in Exits to Particular Destinations, by Age Group, Gender, Pre and Post ND18-24, 20% Sample

	Full Sample	Male 18-24 Pre ND18-24	Male 18-24 Post ND18-24	Male 25-29 Pre ND18-24	Male 25-29 post ND18-24	Female 18-24 Pre ND18-24	Female 18-24 Post ND18-24	Female 25-29 Pre ND18-24	Female 25-29 Post ND18-24
Employment	40,756 (47%)	6,589 (50%)	5,363 (44%)	3,267 (57%)	2,533 (52%)	4,509 (49%)	3,753 (49%)	1,604 (54%)	1,447 (58%)
Education/ training	7,893 (9%)	1,846 (14%)	1,339 (11%)	339 (6%)	182 (4%)	1,516 (16%)	947 (12%)	146 (5%)	89 (4%)
Other benefits	9,504 (11%)	542 (4%)	664 (5%)	379 (7%)	445 (9%)	545 (6%)	551 (7%)	265 (9%)	166 (7%)
Other exits	28,812 (33%)	4,145 (32%)	4,850 (40%)	1,716 (30%)	1,728 (35%)	2,668 (29%)	2,385 (31%)	936 (32%)	808 (32%)

Source: NIERC/DETI. Note: Figures in parentheses give proportion of total exits for each group.

Perhaps the most striking pattern for males is the increase in the proportion of exits to the 'other exits' category (e.g. failure to sign on, destination unknown, moved abroad, prison etc.) following the introduction of ND18-24. This is true of both age groups but is more dramatic for the target age group. 18-24 year old females show only a small relative increase in the proportion of exits to other destinations. For males, the proportions of exits that are to employment and to education and training fall for both age groups and the proportion of exits to other benefits rise for both age groups. For 18-24 year old females the proportion of exits to employment remains unchanged but rises for 25-29 year olds, the proportions of exits to education and training fall for both age groups and the proportion of exits to other benefits rises for the target age group but falls for the older age group.

The raw data therefore suggest the possibility of a positive ND18-24 effect on the hazard rate for exits to other destinations for both genders and to other benefits for females, but little in the way of positive ND18-24 effects on exits to employment or to

education and training. Of course, if hazard rates overall have increased for the target age group relative to the comparison group then even a reduced proportion of exits might correspond to a higher hazard rate. Also, by not distinguishing ND18-24 effects at different unemployment durations, an unchanged proportion of exits might reflect increased hazard rates over some durations and decreased hazard rates over others.

Given the longitudinal nature of the data it is also possible to identify individuals who return to the unemployment register following particular exits. Exits to employment have therefore been disaggregated into exits to employment with a return to the register within three months, six months, 12 months and no return within 12 months. Around one fifth of exits to employment return to the register within three months, around a quarter return within six months and around a third return within 12 months. These proportions remain largely unchanged following the introduction of ND18-24. In other words, there is no apparent pattern in the raw data to suggest employers are using ND18-24 for temporary employment only while the subsidy lasts.

Existing literature – both theoretical and empirical – suggests a set of generally observable factors that may influence unemployment duration and should therefore enter into any model of the hazard rate for exit from unemployment (see, e.g. McVicar and Podivinsky, 2001). Many of these factors are contained within the DETI unemployment register database in a form suitable for use and are therefore included in our model as covariates (explanatory variables). To this we add a binary dummy variable indicating whether the unemployment spell began before or after the introduction of JSA (after JSA = 1) and the Multiple Deprivation Index (MDI) score for the electoral ward in which sample members live. These are listed in Table 5, together with sample means and standard deviations, for the whole 20% sample and for age groups and genders separately, both pre and post ND18-24.

The list of factors observable in the DETI data is not exhaustive, however, and information on other potentially significant factors, such as qualifications held, is not available.[11] The nature and level of benefits received during an unemployment spell is also argued to be a potentially significant predictor of unemployment duration, but

[11] For example, McVicar and Podivinsky (2001) find qualifications held to be a significant predictor of the hazard rate for exits to education and training for young people in NI, using survey data.

information on this, although available to an extent, is not in a readily usable form in the DETI unemployment register database. Macroeconomic trends are also omitted from the models although the use of the 25-29 year old comparison group is intended to control for these. We also include a linear time trend as an additional control.

The covariates listed in Table 5 are defined in detail in Appendix A.

The average age of the sample on entry to the unemployment register is 22.3 years (the sample – and population – is skewed towards the younger end of the age range). Pre and post ND18-24 sample means and standard deviations for age on entry are almost identical for all groups.

Around a sixth of the sample describe themselves on entry to the unemployment register as seeking managerial, professional or related employment (skewed towards females and the older age group) and a slightly smaller proportion describe themselves as seeking skilled manual employment (skewed towards males and the older age group). It is possible that these indicators act as proxies for qualification levels (this is discussed further in Section 5). These proportions are also stable around the introduction of ND18-24.

There is a small drop in the sample means for the *living together with partner* variable for all groups following the introduction of ND18-24. In other words, post ND18-24, individuals on the unemployment register are less likely to live with a partner. Given this occurs in both age groups, it is unlikely to be an effect of ND18-24 itself, but may be picking up an effect of the introduction of JSA or some other policy or macroeconomic factor.

JSA1 is a binary dummy to indicate whether the unemployment spell began before or after the introduction of JSA – unsurprisingly the sample mean for all post ND18-24 groups is therefore equal to one by definition.

There is a small but interesting contrast in the pre and post ND18-24 sample means for the MDI covariate for the young age group – MDI is higher post ND18-24. The MDI figures are almost identical pre and post ND18-24 for the older age group which

suggests this may be picking up a direct ND18-24 effect – 18-24 year olds entering the unemployment register post ND18-24 tend to come from slightly more deprived areas than those entering pre ND18-24.

Finally, there is an unsurprising contrast in pre and post ND18-24 sample statistics for the number of unemployment spells (and number of spells squared), given that spells starting before January 1995 are not recorded (left-censoring). This is true of both age groups, however, so is unlikely to reflect an effect of ND18-24 itself.

Table 5: Sample Means and Standard Deviations of Observed Covariates, by Gender, Age Group, Pre and Post ND18-24, 20% Sample

	Full Sample	Males 18-24 pre ND18-24	Males 18-24 Post ND18-24	Males 25-29 Pre ND18-24	Males 25-29 Post ND18-24	Females 18-24 Pre ND18-24	Females 18-24 Post ND18-24	Females 25-29 Pre ND18-24	Females 25-29 Post ND18-24
Age	22.3 (3.26)	20.6 (1.91)	20.5 (1.89)	26.8 (1.42)	26.8 (1.41)	20.5 (1.92)	20.5 (1.91)	26.7 (1.41)	26.7 (1.41)
Man sought	.175 (.380)	.133 (.340)	.123 (.328)	.161 (.368)	.167 (.373)	.228 (.420)	.198 (.398)	.342 (.475)	.362 (.481)
Skill sought	.136 (.343)	.203 (.402)	.171 (.377)	.236 (.425)	.203 (.403)	.018 (.133)	.010 (.100)	.027 (.163)	.021 (.144)
No. of spells	2.21 (1.81)	2.00 (1.36)	2.83 (2.17)	2.24 (1.57)	3.93 (2.72)	1.67 (.996)	2.07 (1.43)	2.10 (1.49)	3.37 (2.61)
Live tog.	.077 (.266)	.020 (.141)	.015 (.120)	.172 (.378)	.120 (.325)	.042 (.200)	.027 (.163)	.331 (.471)	.257 (.437)
JSA1	.574 (.495)	.372 (.483)	1.00 (0)	.380 (.485)	1.00 (0)	.380 (.485)	1.00 (0)	.396 (.490)	1.00 (0)
MDI	25.7 (17.6)	25.8 (17.8)	27.6 (18.7)	26.2 (17.5)	26.6 (17.7)	23.5 (16.6)	24.5 (16.5)	21.7 (15.6)	21.7 (15.3)
Spells^2	8.17 (19.8)	5.84 (11.9)	12.72 (26.6)	7.47 (16.1)	22.9 (39.4)	3.77 (5.99)	6.33 (11.0)	6.62 (11.2)	18.2 (32.4)
No. of obs.	86,965	13,122	12,216	5,701	4,888	9,238	7,636	2,951	2,510

Source: NIERC/DETI.

Table 5 suggests that young people entering the unemployment register post ND18-24 have similar observed characteristics as young people entering the unemployment

register post ND18-24. In other words, there is very little evidence to suggest the employability of those young people entering unemployment is significantly lower post ND18-24 than pre ND18-24, i.e. our third assumption (see Section 3) appears to hold.

To summarise Section 4:

- The data we use in this study are a 20% random sample taken from the computerised unemployment register for 16-30 year olds in NI. The data span January 1995-July 2001. There are 86,965 unemployment spells in total in the sample.
- We identify exits to employment, exits to education and training, exits to other benefits and other exits separately. Exits to employment are the most common. Exits to education and training are more common than exits to other benefits for the 18-24 age group, but the opposite is true for the 25-29 age group.
- There are approximately twice as many unemployment spells observed for males as for females, reflecting the gender balance in unemployment rates.
- Unemployment duration is grouped fortnightly for the first year, and further aggregated at longer durations for the purposes of estimation.
- The data contain a number of observable factors likely to influence unemployment duration, included in our study as covariates. To this we add a binary JSA dummy to indicate the introduction of the JSA in October 1996, and the MDI score for the electoral ward in which sample members live.
- Young people on the unemployment register before and after ND18-24 have broadly similar observed characteristics.

5. Unemployment Duration Pre and Post ND18-24

Before estimating competing risks PG models on the 20% sample the data is first expanded by the duration groups defined in Table 3. In other words, each individual spell of unemployment now corresponds to one or more rows in the data file depending on the spell length (see Jenkins 1997). A spell of 3 days or less (0 weeks) corresponds to a single row of data in the expanded data set, a spell of 1-2 weeks corresponds to two rows of data, a spell of 3-4 weeks corresponds to three rows of data, and so on. This is necessary for PG estimation and also allows us to specify a more precise JSA dummy variable (JSA2), which takes the value 'one' for all spells or parts of spells from October 1996 onwards and 'zero' otherwise.

Models are estimated separately by gender and age group for exits to employment. Results for these models are presented and discussed in Sections 5.1 and 5.2. Models for exits to education and training, exits to other benefits and other exits are estimated separately by age group but on both genders together (because of the relatively small number of such exits compared to exits to employment). Results from these models are presented and discussed in Section 5.3. All hazard functions shown are for the *average sample member* in each case, i.e. with (continuous) covariates set to their sample means and binary dummy covariates set to zero.

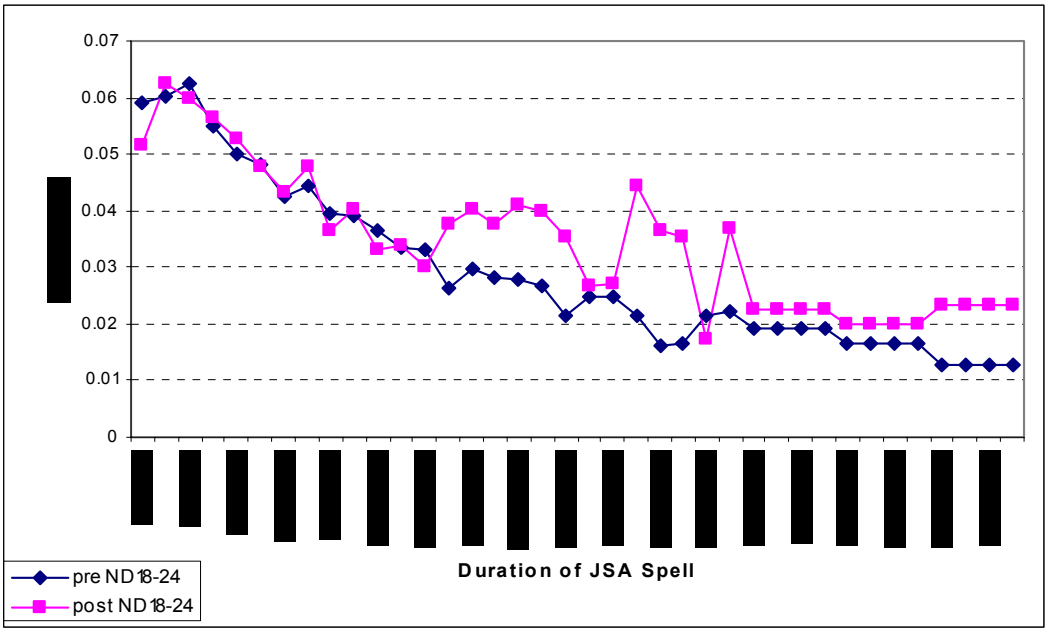
Estimating PG models explicitly allowing for (parametric) unobserved heterogeneity is a straightforward extension to the standard estimation procedure (see Appendix B for technical details). However, given the size of the data set and the nature of the estimation process for such models, it is extremely time consuming: it can take anything between four and eight days for a single estimation to run on a fast PC. For this reason, PG models with unobserved heterogeneity have not been estimated in all cases, but for a representative subset of the various models (five models in all). In each case where unobserved heterogeneity has been specified, it is found to be an insignificant addition to the standard model. In other words, there is no evidence of significant unobserved heterogeneity in the PG models of unemployment duration set out in the remainder of this section. This is a common finding when estimating PG models (see, e.g. McVicar and Podivinsky, 2001; Boheim and Taylor, 2000; Carling

et al., 1996). For this reason, we report only those models that do not incorporate unobserved heterogeneity in their specification.

5.1. Exits to Employment, Males

First consider exits from unemployment to employment for males. Figure 12 presents hazard functions pre and post ND18-24 for exits to employment for the 18-24 age group.[12] Figure 13 presents similar hazard functions for the 25-29 age group.[13]

Figure 12: Hazard Functions for Exits to Employment, Males, 18-24 Year Olds, Pre and Post ND18-24



The hazard functions indicate the probability of exit to employment, for an 'average' unemployed male in the 18-24 age group, after a particular duration of unemployment. For example, on entering the third week of unemployment, an average male has a probability of around .06 (i.e. a 6% chance) of getting a job during the following two weeks, both pre and post ND18-24. If the same individual is still unemployed after 25 weeks, his probability of getting a job during the 25th and 26th

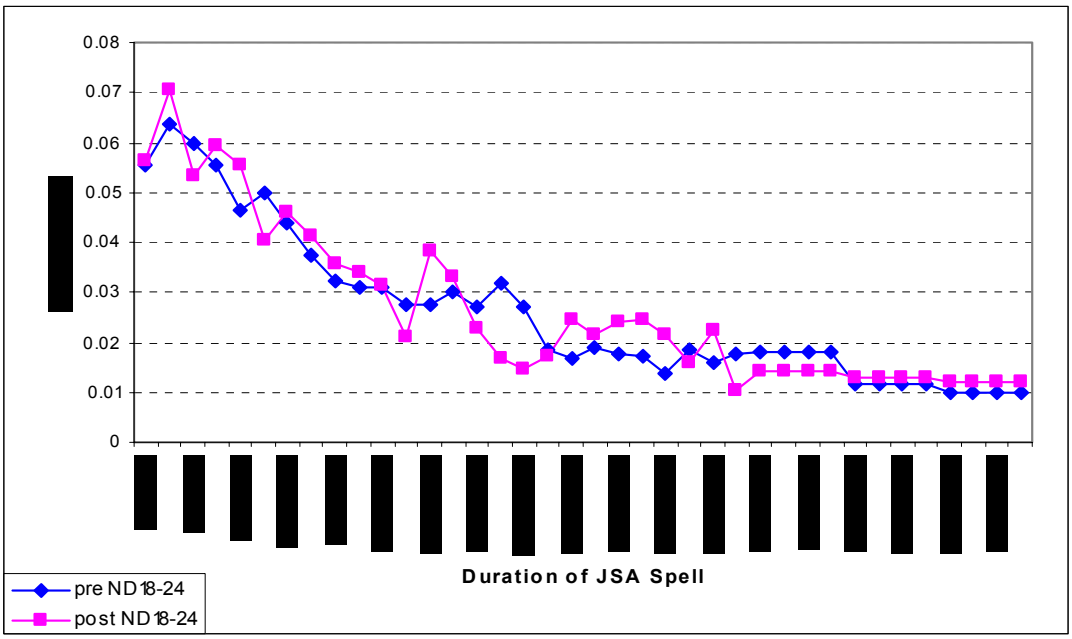
[12] For estimation purposes, the 18-24 age group is defined by the ELIGIBLE indicator (see Appendix A for further details). This includes young people aged 17 on entry to unemployment but who are aged 18 following six months of unemployment. It excludes young people aged 24 on entry to unemployment but who are aged 25 following six months of unemployment.
[13] The 25-29 age group is defined simply by age on entry to unemployment.

week of unemployment is considerably lower, at around .03. This pattern of falling probability of getting a job in a particular space of time the longer the unemployment spell has gone on is known as *negative duration dependence* and is shown by a downward sloping hazard function (see Section 3 for a discussion). The longer a young person in NI is unemployed, the less likely they are to get a job.

After the introduction of ND18-24, however, the hazard function for 18-24 year old males slopes upwards between six and seven months duration before levelling off and falling back to the level of the pre ND18-24 hazard function at around twelve months duration, i.e. the pattern of negative duration dependence is broken for ND18-24 participants. During the 31st and 32nd week of an unemployment spell, for example, the pre ND18-24 probability of getting a job is around .03 compared to a post ND18-24 probability of around .04.

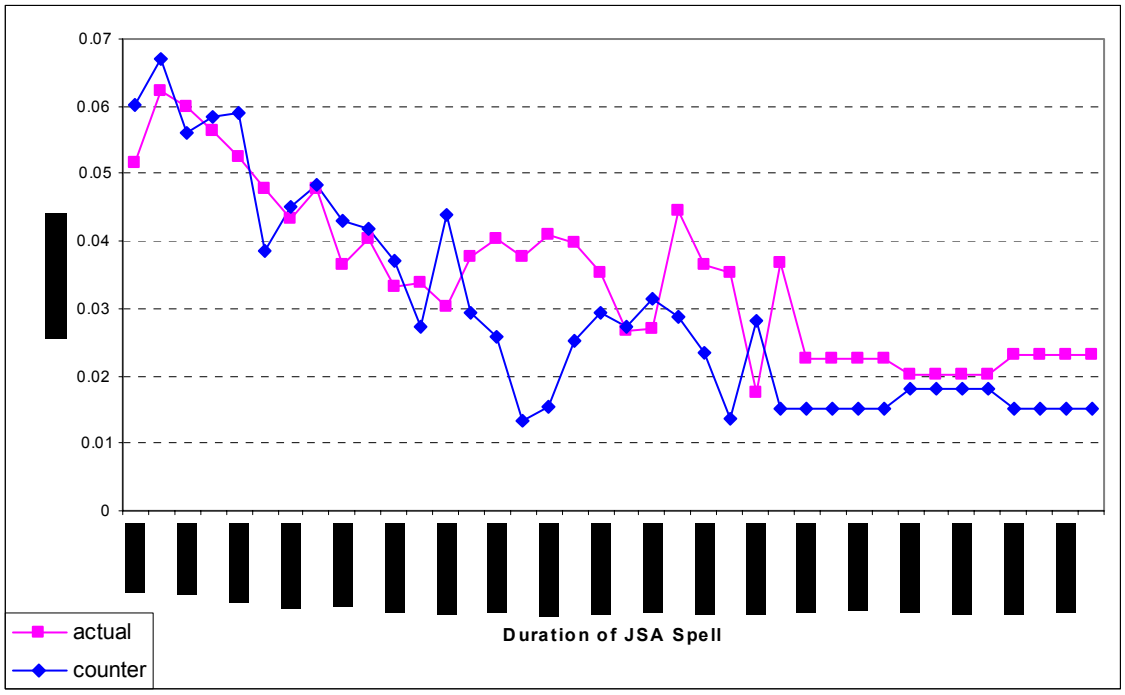
Although there is this apparent difference in the hazard functions pre and post ND18-24 for the target age group, we also have to examine the hazard functions pre and post ND18-24 for the comparison group of 25-29 year olds males before we can conclude whether this effect is caused by ND18-24 or by some other background trend. For this older age group, Figure 13 shows little contrast between the pre and post ND18-24 hazard functions. In other words, an average member of our comparison group is just as likely to get a job after various lengths of unemployment after the introduction of ND18-24 as before the introduction of ND18-24. Therefore we can conclude that the contrast in the hazard functions for the target age group shown in Figure 12 is a result of ND18-24.

Figure 13: Hazard Functions for Exits to Employment, Males, 25-29 Year Olds, Pre and Post ND18-24



It is straightforward to quantify the ND18-24 effect on the chance of 18-24 year olds getting a job within different durations of unemployment by using the 25-29 age group hazard rates to calculate a *counterfactual* hazard rate for the 18-24 age group had ND18-24 *not* been introduced. This is shown along with the actual post ND18-24 hazard function for the target age group in Figure 14. There is no ND18-24 effect at unemployment durations below six months or unemployment durations above twelve months. Between six and twelve months, however, the introduction of ND18-24 has increased the hazard rate for exit to employment by a factor of almost 40% on average, i.e. from around .025 to around .035. The timing of this effect suggests it is capturing both increased job search during Gateway and, for those not finding an unsubsidised job, entry to the ND18-24 employment options (subsidised employment, voluntary sector and environmental taskforce placements). This is broadly consistent with existing findings for GB and in terms of magnitude, is towards the top end of the range of existing estimates for GB (e.g. Van Reenen, 2001; Blundell et al., 2001).

Fig 14: Actual and Counterfactual Hazard Functions, Exits to Employment, Males, 18-24 Year Olds



If Assumption (ii) does not hold, i.e. if there are substitution effects across age groups as a result of ND18-24, then these estimates are likely to represent an upper bound on the true effects of ND18-24.

Table 6 presents the estimated coefficients for the covariates in the PG models corresponding to the hazard functions in Figure 12, i.e. pre and post ND18-24 hazard functions for 18-24 year old males. Significant negative coefficients indicate a negative effect on the hazard rate, i.e. a negative effect on the probability of exiting unemployment for employment. Significant positive coefficients indicate the opposite. Although it is straightforward to see the direction of the marginal effect of each covariate on the hazard rate, it is not necessarily straightforward to calculate their *magnitudes*, especially for binary dummy variables (see Appendix B). We derive an approximation of magnitudes by calculating the hazard rate for different values of the covariates around their sample means.

Table 6: Estimated Effects of Covariates on Hazard Rate, Pre and Post ND18-24, Males, 18-24 Year Olds, Exits to Employment

	Coefficient Pre ND18-24	Coefficient Post ND18-24
Age on entry	.041* (.006)	.054* (.008)
Managerial job sought	.175* (.036)	.197* (.040)
Skilled manual job sought	.037 (.029)	.001 (.035)
No. of unemployed spells	.005 (.018)	-.066* (.013)
Live with partner	-.053 (.072)	-.001 (.099)
MDI	-.012* (.001)	-.012* (.001)
Number of spells^2	.005* (.002)	.005* (.001)
JSA	.176* (.048)	n/a
Time trend	-.014* (.002)	-.005* (.001)
Constant	-4.38* (.141)	-5.05* (.185)
Log likelihood	-29289	-23254
χ^2 statistic	1724 #	1542#

Notes: The χ^2 statistic is a test of the explanatory power of the model compared to an intercept-only model (i.e. not including any covariates). Rejection of the intercept-only model is denoted by #. Covariates that have statistically significant effects on the hazard rate (at the standard 5% level) are denoted *. Standard errors are given in parentheses. JSA is dropped in the post ND18-24 sample since it takes the value one at all times.

All the covariates have the same signs and (where significant) similar magnitudes for the pre and post ND18-24 models.[14] Age is positively related to the hazard rate for 18-24 year olds, so that the hazard rate for a 23 year old is higher than that for a 19 year old by a factor of 17%. Those describing themselves as seeking skilled manual employment have higher hazard rates than others by a factor of around 18%. Neither seeking a skilled manual job, number of previous unemployment spells, nor living with a partner has any significant effect on the hazard rate. MDI is a government composite index of deprivation derived from weighted indicators of the local labour

[14] This is also the case for the 25-29 year old age group.

market, housing, qualifications and other factors (see Appendix A for more details). Young people living in electoral wards with a high MDI (more deprived) score have lower hazard rates for exit to employment. The hazard rate for a young unemployed male from Wallace Park (the least deprived ward) is over twice as high as that for a young unemployed male from Crumlin (the most deprived ward), other things being equal. The squared number of previous unemployment spells is positively related to the hazard rate. There are likely to be two contrasting effects here: a higher number of previous spells might indicate poor job chances and therefore be negatively related to the hazard rate, whilst conversely it might indicate fluidity, i.e. dipping in and out of unemployment frequently, which we might expect to show a positive relationship with the hazard rate. These effects are either insignificant or in balance in levels, but the squared term suggests the fluidity interpretation dominates for high numbers of previous spells.

The introduction of JSA in October 1996 coincides with significantly increased hazard rates for exits to employment for 18-24 year old unemployed males. We estimate the introduction of JSA to have increased the hazard rate for exits to employment for young males by a factor of around 17%.

5.2. Exits to Employment, Females

Figure 15 presents hazard functions pre and post ND18-24 for female exits to employment for the 18-24 age group. Figure 16 presents similar hazard functions for the 25-29 age group. As for males pre ND18-24 hazard functions for both age groups are generally downward sloping, indicating negative duration dependence. For the older age group the post ND18-24 hazard function is also downward sloping. Once again, however, negative duration dependence is broken between six and twelve months for the post ND18-24 hazard function for the target age group.

Figure 15: Hazard Functions for Exits to Employment, Females, 18-24 Year Olds, Pre and Post ND18-24

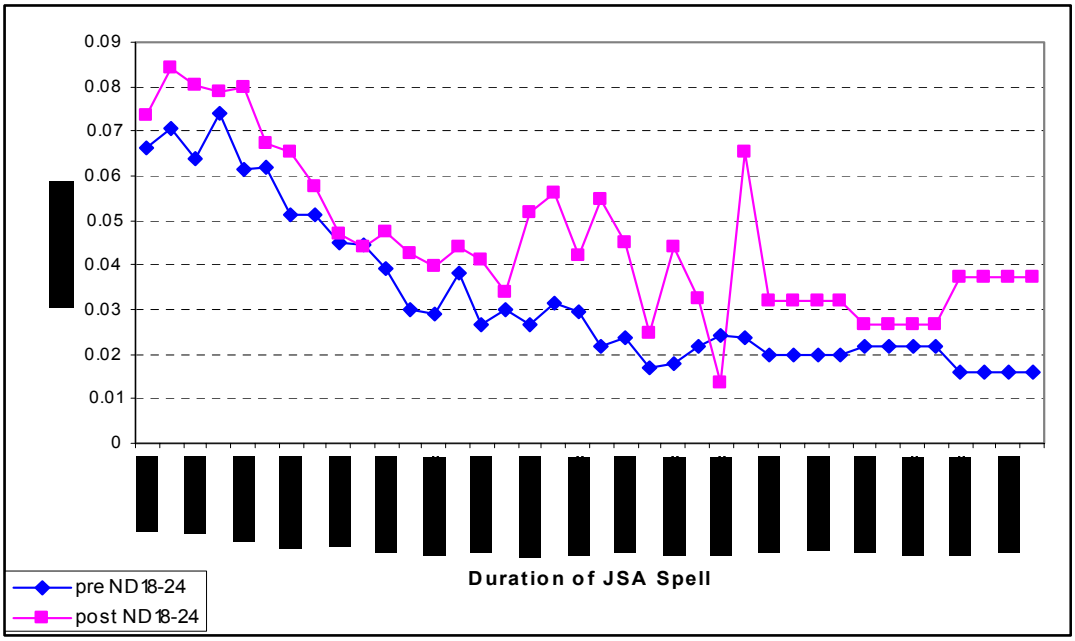


Figure 16: Hazard Functions for Exits to Employment, Females, 25-29 Year Olds, Pre and Post ND18-24

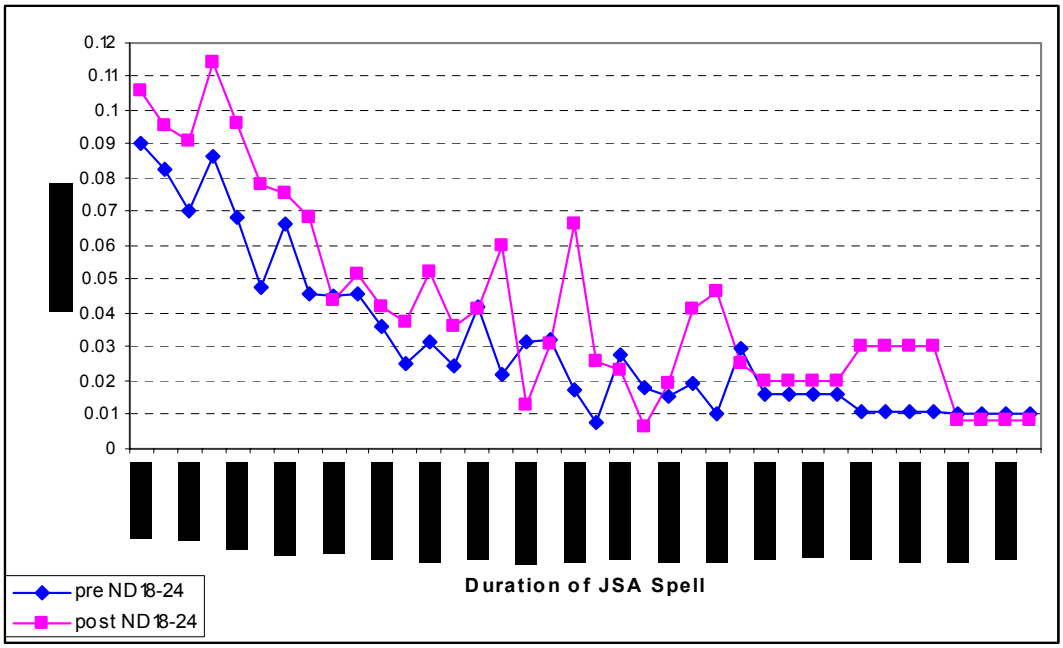
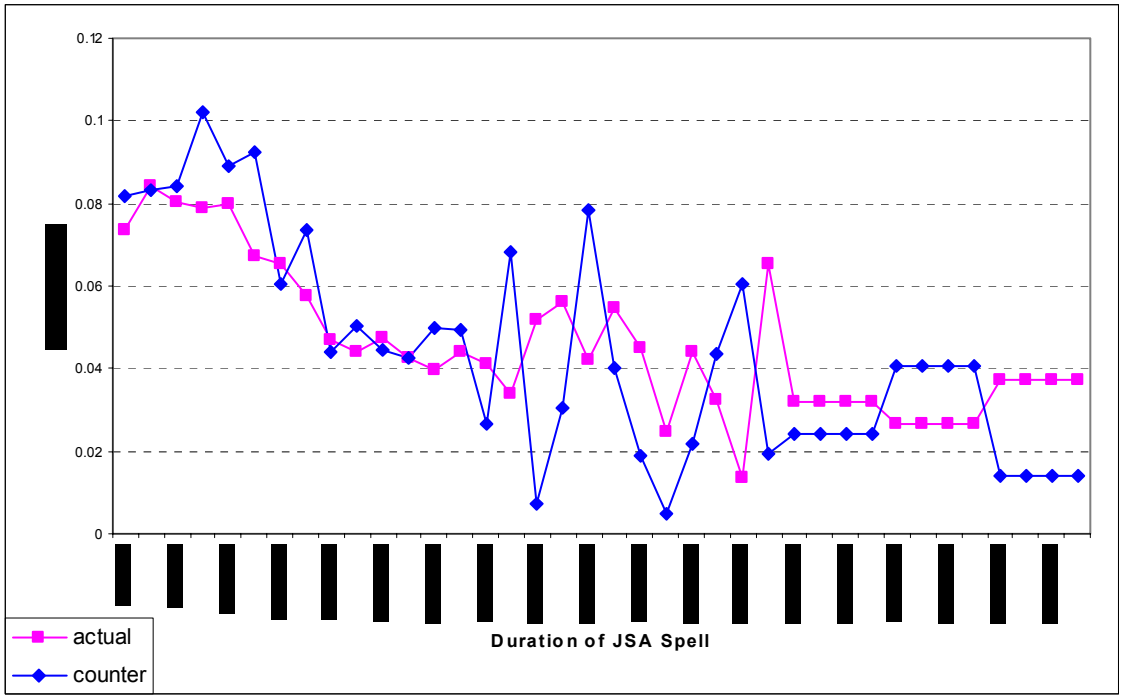


Figure 17 shows the actual and counterfactual female hazard functions for exits to employment. The pattern is similar to that for males – ND18-24 has increased the

hazard rate at unemployment durations between six and twelve months, i.e. for programme participants. The magnitude of the effect is smaller than for males, however, with the hazard rate increasing by a factor of around 20% on average. Again this is consistent with earlier findings for GB that suggest the female ND18-24 effect on exits to employment is weaker than the male effect.

Fig 17: Actual and Counterfactual Hazard Functions, Exits to Employment, Females, 18-24 Year Olds



There is also a small apparent *negative* ND18-24 effect on the hazard rate at short unemployment durations, with the hazard rate lower on average by a factor of 10% during the first six months of an unemployment spell. There are two possible explanations for this weak negative effect. First, unemployed females may feel the programme offers better quality employment options or more help with overcoming particular barriers to employment than are available outside the programme, and may therefore *reduce* job search prior to entry to the programme. Second, there may be substitution effects whereby programme participants take jobs that would otherwise have been taken by those unemployed for shorter durations. This effect has not generally been found by previous GB studies and may reflect the somewhat weaker labour market in NI compared to GB.

Table 7 presents the estimated coefficients for the covariates in the PG models corresponding to the hazard functions in Figure 15. The picture is similar to that for males and there is little difference between the pre and post ND18-24 estimates. For females, the number of previous spells is negatively and significantly related to the hazard rate, and squared number of previous spells is again positively related to the hazard rate. This is consistent with fluidity beginning to dominate occurrence dependence for higher numbers of previous spells, as suggested above. Living with a partner is negatively and significantly related to the hazard rate for females, with hazard rates lower by a factor of 16% on average.

Table 7: Estimated Effects of Covariates on Hazard Rate, Pre and Post ND18-24, Females, 18-24 Year Olds, Exits to Employment

	Coefficient Pre ND18-24	**Coefficient Post ND18-24**
Age on entry	.087* (.008)	.132* (.009)
Managerial job sought	.207* (.034)	.138* (.040)
Skilled manual job sought	.078 (.107)	-.067 (.146)
No. of unemployed spells	-.027 (.031)	-.122* (.025)
Live with partner	-.181* (.073)	-.276* (.099)
MDI	-.009* (.001)	-.015* (.001)
Number of spells^2	.011* (.005)	.014* (.003)
JSA	.217* (.057)	n/a
Time trend	-.007* (.003)	-.006* (.001)
Constant	-5.25* (.175)	-6.16* (.222)
Log likelihood	-18881	-14931
χ^2 statistic	1387#	1614#

Notes: The χ^2 statistic is a test of the explanatory power of the model compared to an intercept-only model (i.e. not including any covariates). Rejection of the intercept-only model is denoted by #. Covariates that have statistically significant effects on the hazard rate (at the standard 5% level) are denoted *. Standard errors are given in parentheses. JSA is dropped in the post ND18-24 sample since it takes the value one at all times.

5.3. Exits to Education and Training and Exits to Other Benefits

Because exits to education and training, to other benefits and other exits are less common than exits to employment, we estimate hazard functions for both genders together.

Figures 18 and 19 give pre and post ND18-24 hazard functions for exits to education and training for the two age groups. Notice first that hazard rates are generally lower for exits to education and training than for exits to employment – such exits are less common. For the 18-24 year olds age group, the hazard rate is significantly higher following the introduction of ND18-24 at unemployment durations beyond six months. At shorter durations, the hazard rate is slightly lower, i.e. the hazard function appears to have pivoted around the six month point. There is no such pattern for the comparison group, with hazard rates generally lower following the introduction of ND18-24.

Figure 18: Hazard Functions for Exits to Education and Training, Males and Females, 18-24 Year Olds, Pre and Post ND18-24

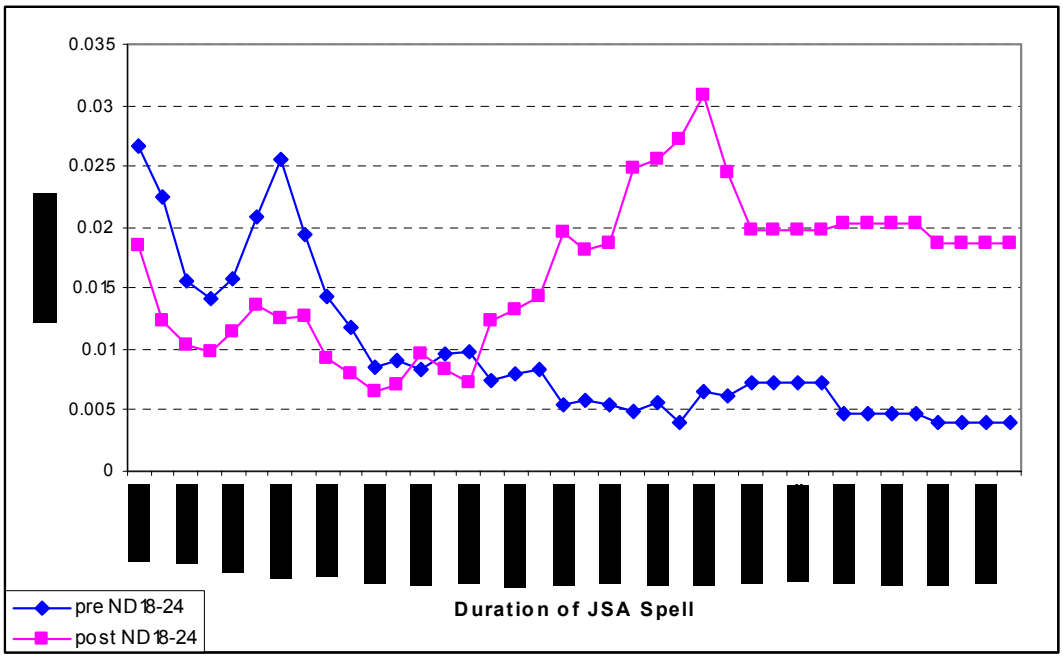


Figure 19: Hazard Functions for Exits to Education and Training, Males and Females, 25-29 Year Olds, Pre and Post ND18-24

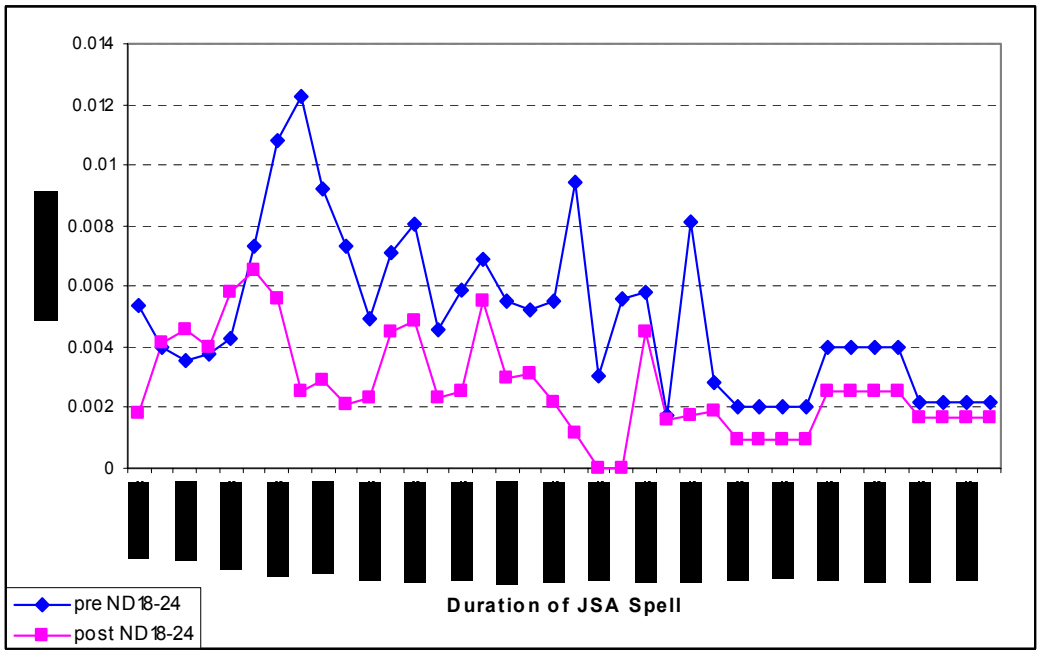


Figure 20 puts Figures 18 and 19 together to show the actual post ND18-24 and counterfactual hazard functions for the target age group. Beyond six months unemployment duration there is a very strong ND18-24 effect on the hazard function, with hazard rates higher on average by a factor of around 300%. As for exits to employment the timing of this effect indicates that it is programme participation that increases the probability of exits to education and training, likely to be through both ND18-24 education and training option placements and take up of other education and training places outside the programme during Gateway. Despite the very large *relative* increase in the hazard rate, however, given the relative numbers of exits to employment and exits to education, in *absolute* terms the ND18-24 effect on exits to education and training is of similar magnitude to the effect on exits to employment.

At shorter unemployment durations, there is some evidence of a small negative ND18-24 effect on the hazard rate for exits to education and training, with the hazard rate lower on average by a factor of around 20%. Again, this may be suggestive of a weak anticipation effect, where unemployed young people that would previously have taken up education or training places shortly after becoming unemployed now wait to enter ND18-24, perhaps because of employment opportunities or education and

training opportunities offered by the programme. Neither the large positive ND18-24 effect on exits to education and training at longer unemployment durations nor the small negative effect at shorter unemployment durations is found by existing GB studies because these studies do not separately identify such exits. These results, therefore, may have significant implications for policy makers across the whole UK. If NI is at all representative of other UK regions, then ND18-24 may well have shifted a large number – not previously quantified – of young people out of unemployment into education and training nationally.

Figure 20: Actual and Counterfactual Hazard Functions, Exits to Education and Training, Males and Females, 18-24 Year Olds

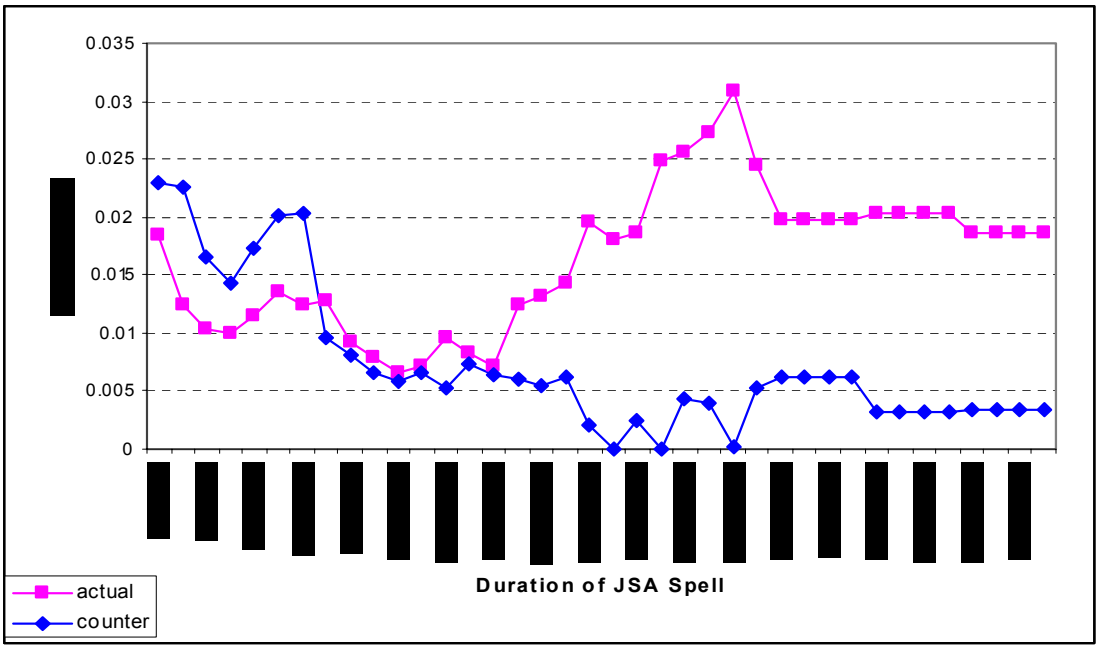


Table 8 presents the estimated coefficients for the covariates in the PG models corresponding to the hazard functions in Figure 18. First consider the pre ND18-24 estimates. There are some covariates that have similar effects on both exits to education and training and exits to employment, e.g. those seeking managerial employment have higher hazard functions for both exits, those living in more deprived wards have lower hazard functions for both exits, those living with a partner have lower hazard rates for both exits (remember genders are combined here) and number of previous spells of unemployment (previous spells squared) is negatively (positively) related to both hazard rates. Some other covariates have opposite effects

on exits to education and training and exits to employment. Age is negatively related to the hazard rate for exits to education and training as we might expect. The introduction of JSA reduced hazard rates for exits to education and training. Finally, two relationships are shown that are not reflected by exits to employment. First, those seeking skilled manual employment have lower hazard rates for exits to education and training. Second, given that we estimate on both genders here, we include a male dummy whose coefficient is negatively related to the hazard rate for exits to education and training. Other things being equal, males are less likely than females to exit to education and training by a factor of around 14%.

Table 8: Estimated Effects of Covariates on Hazard Rate, Pre and Post ND18-24, Males and Females, 18-24 Year Olds, Exits to Education and Training

	Coefficient Pre ND18-24	Coefficient Post ND18-24
Age on entry	-.149* (.009)	-.073* (.013)
Managerial job sought	.309* (.046)	.127* (.063)
Skilled manual job sought	-.468* (.063)	-.320 (.077)
No. of unemployed spells	-.326* (.039)	-.053* (.024)
Male	-.152* (.035)	-.254* (.043)
Live with partner	-.805* (.160)	-.717* (.221)
MDI	-.006* (.001)	-.003* (.001)
Number of spells^2	.020* (.006)	.003 (.002)
JSA	-.651* (.067)	n/a
Time trend	.014* (.003)	-.006* (.002)
Constant	-.544* (.198)	-2.21* (.277)
Log likelihood	-18099	-12628
χ^2 statistic	1754#	744#

Notes: The χ^2 statistic is a test of the explanatory power of the model compared to an intercept-only model (i.e. not including any covariates). Rejection of the intercept-only model is denoted by #. Covariates that have statistically significant effects on the hazard rate (at the standard 5% level) are

denoted *. Standard errors are given in parentheses. JSA is dropped in the post ND18-24 sample since it takes the value one at all times.

Post ND18-24 coefficients are similar in terms of sign but not always in terms of magnitude to those for the pre ND18-24 model.[15] Three covariates stand out in this respect. First, the age effect appears to have weakened following the introduction of ND18-24, suggesting the programme may have increased exits to education and training for the older end of the 18-24 age group to a greater extent than for the younger end of the age group. Second, the effect of seeking managerial, professional or related employment has weakened, perhaps suggesting young people at the less academic end of the spectrum have been encouraged to enter education and training more than those at the more academic end of the spectrum. Third, the negative effect of being male has strengthened, perhaps suggesting ND18-24 has increased the hazard rate for exits to education and training to a greater extent for females than for males.

Figures 21 and 22 give pre and post ND18-24 hazard functions for exits to other benefits for both age groups. Given the small number of such exits hazard rates are considerably lower in general than those for exits to employment. For the 18-24 year old age group, the post ND18-24 hazard function is everywhere above the pre ND18-24 hazard function, but again more so between six and twelve months unemployment durations. For the 25-29 year old age group, however, this is not the case. At first glance, therefore, ND18-24 appears to have increased the hazard rate for exits to other benefits.

[15] These differences between pre and post ND18-24 coefficients for the target age group are not reflected (at least not to the same extent) by the coefficients for the older age group.

Figure 21: Hazard Functions for Exits to Other Benefits, Males and Females, 18-24 Year Olds, Pre and Post ND18-24

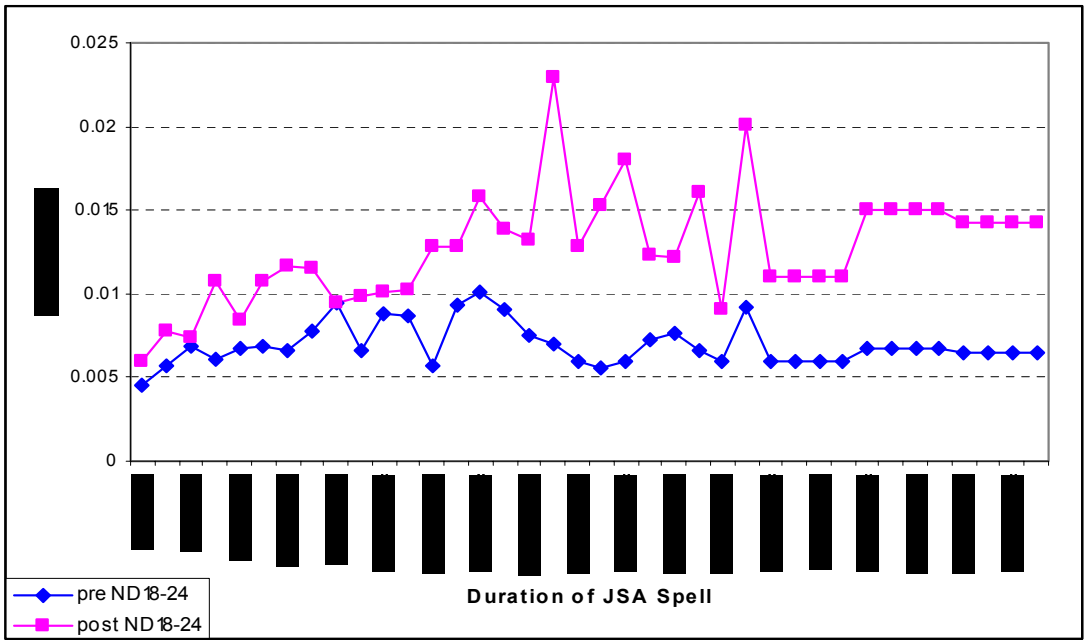


Figure 22: Hazard Functions for Exits to Other Benefits, Males and Females, 25-29 Year Olds, Pre and Post ND18-24

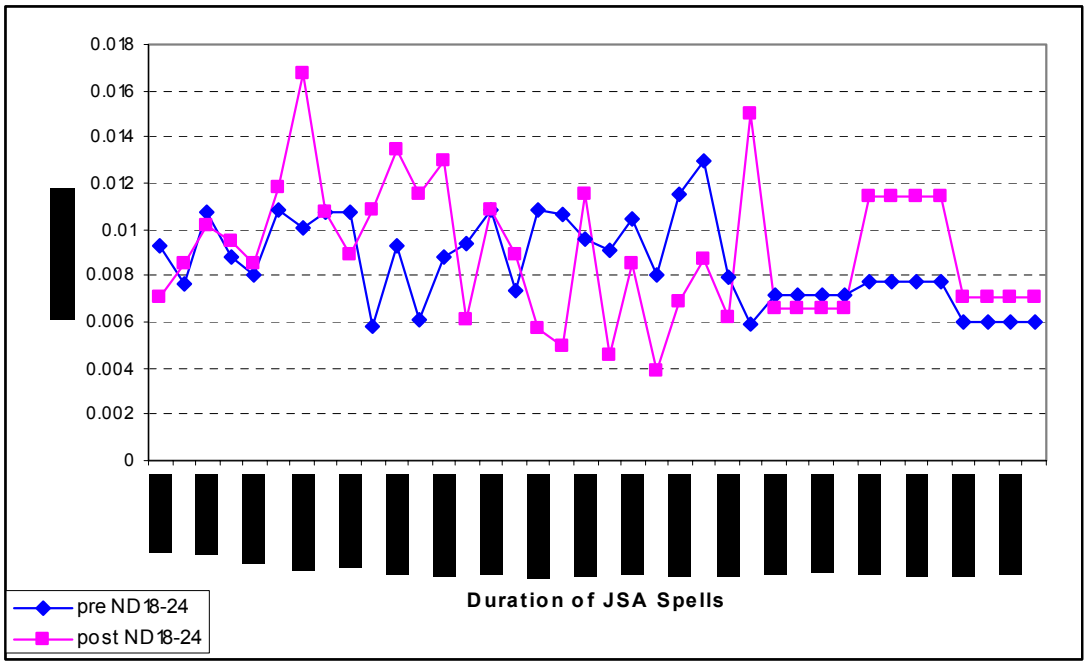


Figure 23 shows the actual and counterfactual hazard functions for exits to other benefits. The ND18-24 effect suggested by Figures 21 and 22 is clearly shown. At unemployment durations between six and twelve months, the actual post ND18-24

hazard rate is on average around twice the counterfactual. There is little apparent effect at durations less than six months. Again this suggests a programme participation effect, rather than anticipation of avoidance effects. In other words, it appears to be on entry to and during Gateway that the hazard rate for exits to other benefits increases rather than in the months leading up to entry to ND18-24. Despite the strength of this ND18-24 effect on exits to other benefits, because of the relatively small number of such exits the ND18-24 effects on exits to employment and to education and training far outweigh this other benefits effect. As for exits to education and training, previous GB studies have not separately identified such exits, therefore our results have implications for policy makers across the UK. These young people would appear to be being driven further from the labour market rather than helped into the labour market.

Figure 23: Actual and Counterfactual Hazard Functions, Exits to Other Benefits, Males and Females, 18-24 Year Olds

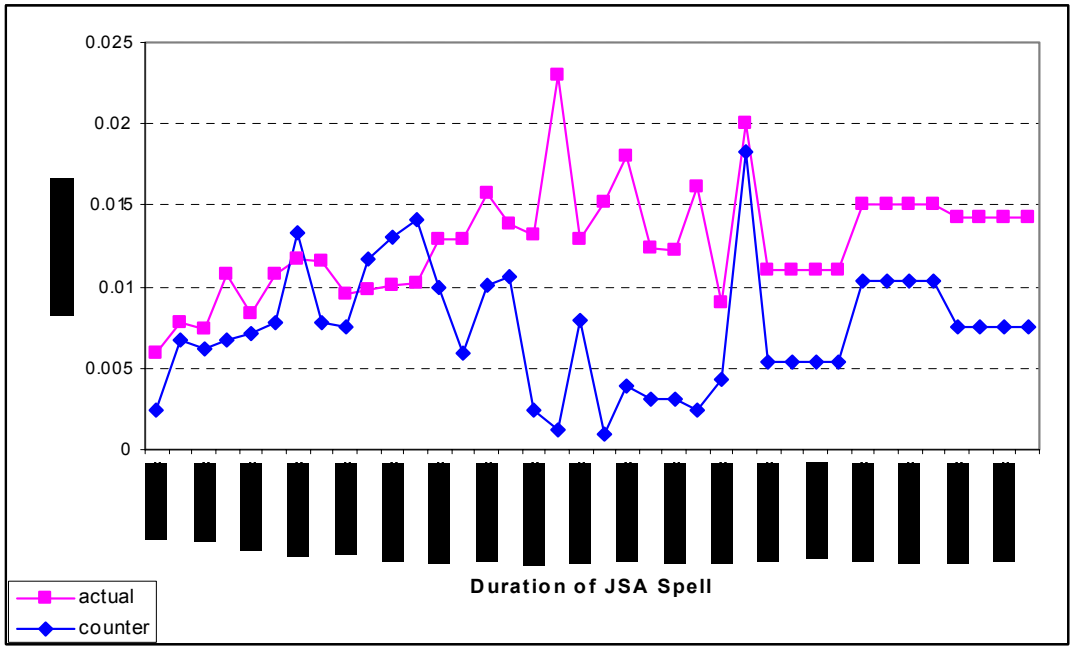


Table 9 presents the estimated coefficients for the covariates in the PG models corresponding to the hazard functions in Figure 21. Exits to other benefits are (unsurprisingly) more likely for those living in deprived wards, those not seeking managerial or related employment and those living with a partner. Females are around twice as likely to exit to other benefits as males. For this age group the introduction of

JSA did not significantly increase exits to other benefits. The estimated coefficients for the post ND18-24 sample are broadly similar to those for the pre ND18-24 sample. Where there are differences between pre and post ND18-24 coefficients, they are generally also reflected in the 25-29 age group pre and post ND18-24 models.

Table 9: Estimated Effects of Covariates on Hazard Rate, Pre and Post ND18-24, Males and Females, 18-24 Year Olds, Exits to Other Benefits

	Coefficient Pre ND18-24	Coefficient Post ND18-24
Age on entry	-.007 (.016)	-.035* (.017)
Managerial job sought	-.817* (.123)	-.617* (.121)
Skilled manual job sought	.078 (.088)	.044 (.091)
No. of unemployed spells	.162* (.055)	.173* (.039)
Male	-.856* (.063)	-.677* (.060)
Live with partner	.836* (.114)	.685* (.155)
MDI	.009* (.002)	.010* (.001)
Number of spells^2	-.005 (.007)	-.007* (.004)
JSA	.071 (.119)	n/a
Time trend	-.016* (.005)	.002 (.003)
Constant	-5.87* (.348)	-6.23* (.410)
Log likelihood	-7271	-7417
χ^2 statistic	683#	685#

Notes: The χ^2 statistic is a test of the explanatory power of the model compared to an intercept-only model (i.e. not including any covariates). Rejection of the intercept-only model is denoted by #. Covariates that have statistically significant effects on the hazard rate (at the standard 5% level) are denoted *. Standard errors are given in parentheses. JSA is dropped in the post ND18-24 sample since it takes the value one at all times.

Figures 24 and 25 show pre and post ND18-24 hazard functions for other exits (i.e. failure to sign on, unknown destination, move abroad etc.) for both age groups. It is generally believed (e.g. National Audit Office, 2002) that around half of such exits are

in fact exits to employment but not recorded as such. As for exits to employment, hazard rates before the introduction of ND18-24 are generally downward sloping, indicating negative duration dependence. Again as for exits to employment this is also true for the 25-29 year old post ND18-24 hazard function but not for the 18-24 year old post ND18-24 hazard function. For the target age group, the pattern of negative duration dependence is broken between six and twelve months duration following the introduction of ND18-24.

Figure 24: Hazard Functions for Other Exits, Males and Females, 18-24 Year Olds, Pre and Post ND18-24

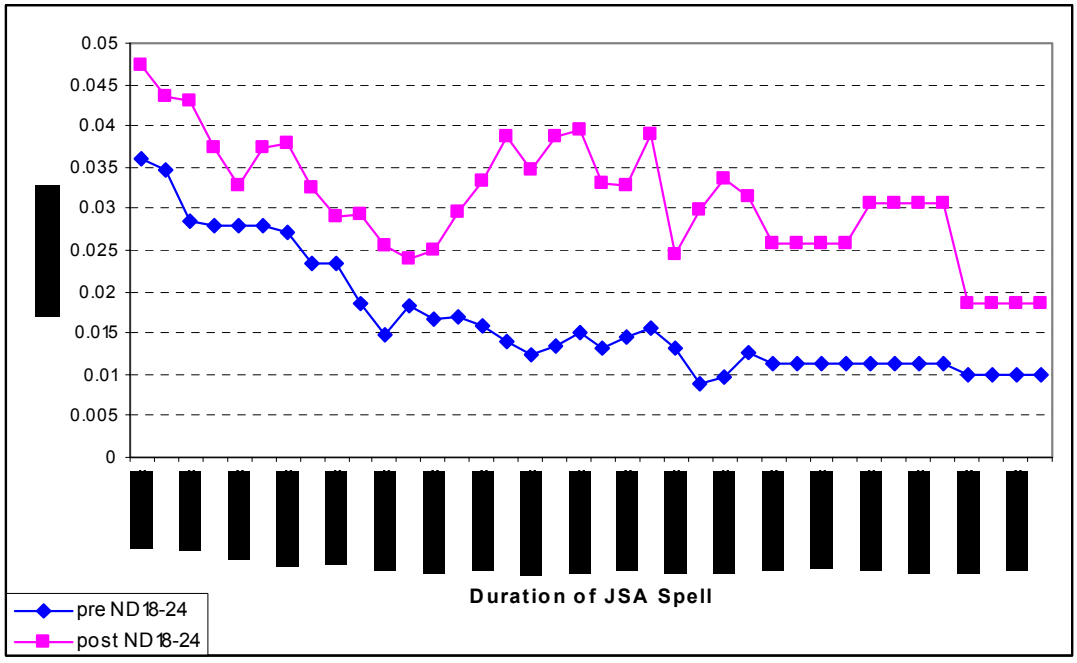


Figure 25: Hazard Functions for Other Exits, Males and Females, 25-29 Year Olds, Pre and Post ND18-24

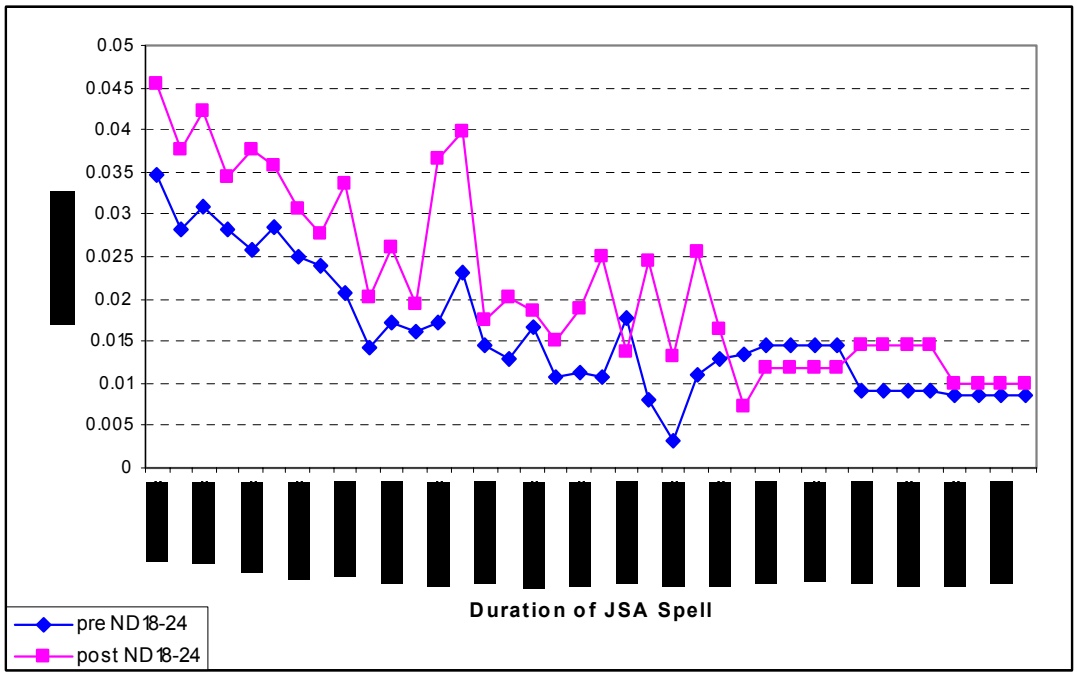
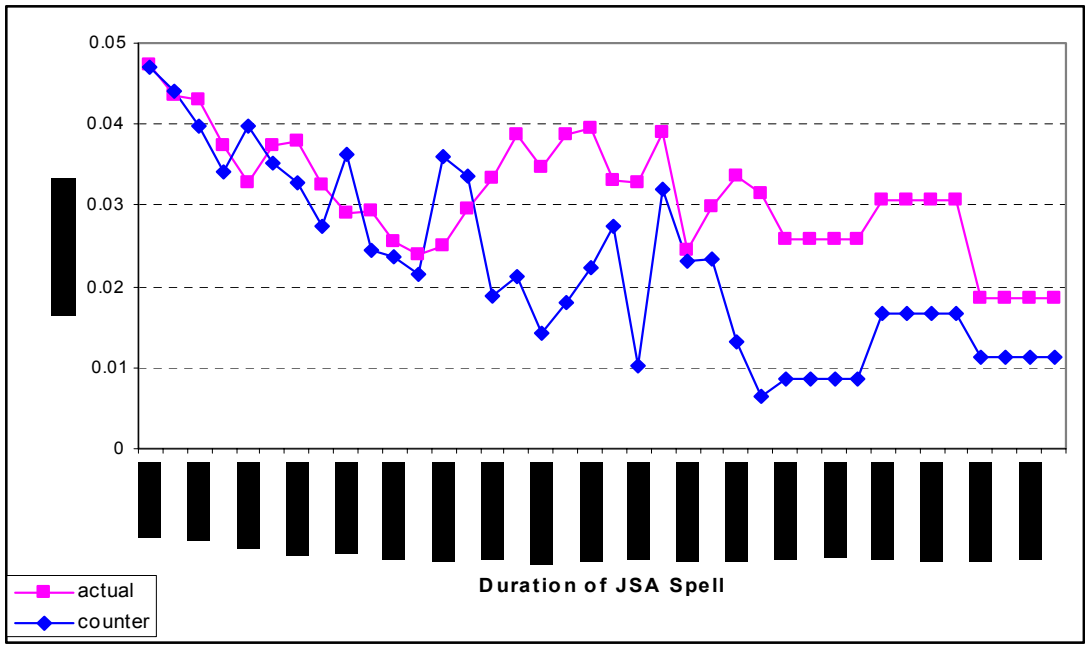


Figure 26 shows the actual and counterfactual hazard functions corresponding to Figures 24 and 25. Once again we see little ND18-24 effect at durations shorter than six months but a strong positive effect at durations between six and twelve months, i.e. a programme participation effect. Over this range of unemployment durations, the actual post ND18-24 hazard function is higher on average than the counterfactual by a factor of around 70%. This is larger in relative terms (although smaller in absolute terms) than the estimated ND18-24 effect on exits to employment presented in Sections 5.1 and 5.2, perhaps suggesting there is some other effect being captured here beyond unrecorded exits to employment. Again this has implications for policy makers across the UK. Without knowing where these young people are going on leaving unemployment, it is difficult to fully evaluate the effects of the programme. Estimated coefficients for the covariates in the models for other exits are (encouragingly) broadly similar to those for exits to employment reported in Table 6 (which supports the suggestion that many of these exits might be unrecorded exits to employment) so we do not report them separately here.

Figure 26: Actual and Counterfactual Hazard Functions, Other Exits, Males and Females, 18-24 Year Olds



To summarise Sections 5.1-5.3:

- ND18-24 has increased the male hazard rate for exits to employment by a factor of around 40% for those unemployed between six and twelve months, i.e. programme participants.
- ND18-24 has increased the female hazard rate for exits to employment by a factor of around 20% for those unemployed between six and twelve months.
- ND18-24 has had a small negative effect on the hazard rate for exits to employment for females unemployed less than six months, i.e. a small number of unemployed females appear to be waiting to enter the programme rather than finding jobs.
- ND18-24 has increased the hazard rate for both genders for exits to education and training from six months of unemployment onwards. This effect is of similar magnitude in absolute terms to the ND18-24 effect on exits to employment.
- ND18-24 has a small negative effect on the hazard rate for exits to education and training at shorter unemployment durations, again suggesting some young people may be waiting to enter the programme rather than taking up the first education or training place offered.

- ND18-24 has increased the hazard rate for both genders for exits to other benefits for those unemployed between six and twelve months
- ND18-24 has increased the hazard rate for other (unidentified) exits from unemployment for both genders for those unemployed longer than six months.
- The introduction of JSA in October 1996 increased hazard rates for exits to employment for 18-24 year olds.

5.4. Discussion: The Effects of ND18-24 on Unemployment Duration for 18-24 Year Olds in NI

Following the introduction of ND18-24 in NI in April 1998, hazard rates for all exits from unemployment have increased for 18-24 year olds relative to those for the best available comparison group, i.e. unemployed 25-29 year olds. This suggests the introduction of ND18-24 itself is the cause of these changes. This is true for exits to employment, exits to education and training, exits to other benefits and other exits. In all cases the increases in hazard rates occur at unemployment durations of six months or more, i.e. ND18-24 increases hazard rates for its participants. There is little evidence, beyond a small negative effect for exits to education and training and for female exits to employment, of ND18-24 effects at unemployment durations shorter than six months. In other words, only in a few cases is there evidence of any ND18-24 effect on young people yet to enter the programme. Where comparable analysis has been carried out, our findings for NI are generally consistent with earlier findings for GB.

The effects of the introduction of ND18-24 on the hazard rate for exits to employment for programme participants have not been uniform across *gender*. ND18-24 has increased the hazard rate for exits to employment for young male participants by a factor of around 40%. For females the corresponding figure is an increase of around 20%. In other words, the ND18-24 effect on exits to employment is stronger for males than for females. This is also consistent with existing evidence for GB. There is suggestive evidence for NI that the opposite may have been the case for exits to education and training.

Neither have the effects of ND18-24 been uniform across *different types of exit* from unemployment. Relative to the counterfactual, the largest percentage increase in hazard rates is for exits to education and training, followed by exits to other benefits, followed by other exits. Exits to employment display the smallest relative increase in hazard rates. Given the number of the various exits, however, in *absolute* terms the ND18-24 effects on exits to employment and exits to education and training are the primary effects of ND18-24 and are of roughly equal magnitude, together accounting for over three quarters of the overall ND18-24 effect.

Given ND18-24 has increased hazard rates for all types of exit from unemployment at durations of six months or more, the chances of young people in the 18-24 age group remaining unemployed for more than six months are lower following the introduction of the programme. Table 10 shows actual post ND18-24 and counterfactual (what would have happened had ND18-24 not been introduced) probabilities of remaining unemployed for three months, six months, nine months, twelve months and 18 months. For males ND18-24 does not affect unemployment duration below six months, but only 7% of spells now last one year compared to a counterfactual proportion of 13%. For females there is a slight increase in the probability of a spell lasting up to six months (20% compared to 17% under the counterfactual) but only 4% of female unemployment spells last for one year compared to 7% under the counterfactual. In other words, as a result of ND18-24, the chance of an average unemployed 18-24 year old male or female remaining in unemployment for a continuous twelve months spell is just over half what it would have been had ND18-24 not been introduced. **Without ND18-24, around *one in eight* 18-24 year old males entering unemployment would still have been unemployed a year later. Having introduced ND18-24, the actual figure is *one in fourteen.* Corresponding figures for females are *one in fourteen* and *one in twenty five.***

Clearly ND18-24 has had a significant effect to reduce the chances of young people experiencing long-term unemployment in NI. These effects are mostly identifiable – additional exits from the unemployment register to unknown destinations account for only a small proportion of the overall ND18-24 effect. But a fundamental question remains, i.e. why these effects are not stronger? If ND18-24 has been implemented according to the 'rules' there should be almost no young people in the target age

group that experience unemployment durations of a year or more. In other words, the introduction of ND18-24 should have reduced the chances of year-long unemployment spells by close to 100% rather than closer to 50%. One contributing factor is likely to be that Gateway often extends beyond four months (in 20% of cases) and sometimes considerably beyond four months (there have been some spells in Gateway lasting over a year). There may of course be good reasons for these extended stays in Gateway, e.g. particular barriers to employment that cannot be overcome within four months. Second, anecdotal evidence suggests that, even in 2001, some long-term unemployed young people in the target age group had still not been invited for ND18-24 interview, for whatever reason. If confirmed to be the case, this latter point would appear to require urgent policy attention.

Table 10: Actual and Counterfactual Survival Rates (% Still Unemployed)

	Males Actual	Males Counterfactual	Females Actual	Females Counterfactual
After 3 months	49%	48%	41%	37%
After 6 months	27%	26%	20%	17%
After 9 months	14%	17%	9%	10%
After 12 months	7%	13%	4%	7%
After 18 months	3%	8%	1%	4%

6. Conclusions

Our main conclusions relate to the five questions posed in the introduction to this report.

1. **How has ND18-24 affected the duration of unemployment spells for young people aged 18-24?**

ND18-24 has significantly reduced unemployment duration for young people in NI aged 18-24. Females are 40-45% less likely to experience an unemployment spell of one year or more. Males are 45-50% less likely to experience an unemployment spell of one year or more.

2. **How large are any changes in the hazard rate resulting from ND18-24?**

Hazard rates for exits to employment have increased by up to 20% for females and up to 40% for males. Hazard rates for exits to education and training have increased by up to 300%. Hazard rates for exits to other benefits have increased by up to 100%. Hazard rates for other (unidentifiable) exits have increased by up to 70%.

3. **After what duration of unemployment are the effects of ND18-24 on the hazard rate strongest?**

The effects of ND18-24 on hazard rates for all types of exits occur primarily over unemployment durations between six and twelve months, i.e. ND18-24 primarily affects programme participants.

4. **Are there any unemployment durations over which ND18-24 has reduced the hazard rate for exit from unemployment?**

In most cases there are no negative ND18-24 effects on hazard rates. This suggests that generally there are no strong programme anticipation or avoidance effects and no strong substitution effects where those yet to enter the programme lose out in the job

market to programme participants. Exceptions to this are a small negative effect at short unemployment durations for female exits to employment and a small negative effect at short durations for exits to education and training.

5. **Have the effects of ND18-24 on unemployment duration been uniform across gender, age within the 18-24 age group, and location, or have they been heterogeneous?**

ND18-24 has had stronger effects on exits to employment for males than for females. Other than this, our results do not suggest ND18-24 effects have been significantly stronger for certain 'types' of young people in any consistent way, although older young people and those with fewer qualifications may have benefited more from ND18-24 in terms of increased exits to education and training than their younger or more qualified counterparts. We regard this latter conclusion as tentative.

By focussing on outflows from unemployment and the duration of unemployment spells, the study has not considered any ND18-24 effects on *inflows* to unemployment. In addition, we have not examined what happens to young people *after* their initial exits from unemployment. Further, the study itself raises interesting questions relating to the *implementation* of ND18-24. We therefore make five recommendations for further evaluation of ND18-24 in NI:

1. To merge the New Deal database held by DEL with the JSA-register database held by DETI.

2. Using this merged database, to evaluate the effects of ND18-24 on inflows to unemployment in NI.

3. To evaluate what happens to young people that have passed through ND18-24 in the longer term, i.e. beyond their first destination on leaving the programme.

4. To compare the effects of ND18-24 in NI with those in other UK regions.

5. To investigate the implementation of ND18-24 in NI, in particular the causes and consequences of overstaying on Gateway.

References

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 nonparametric 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.

Blundell, R., Costas Dias, M., Meghir, C. and Van Reenen, J. (2001). 'Evaluating the employment impact of a mandatory job search assistance program.' Institute for Fiscal Studies, Working Paper 01/20, London.

Boheim, R. and Taylor, M. (2000). 'Unemployment duration and exit states in Britain.' Institute for Social and Economic Research, Working Paper 2000-1, University of Essex.

Bratberg, E. and Vaage, K. (2000). 'Spell durations with long unemployment insurance periods.' Labour Economics, 7, 153-180.

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.

Cox, D. R. (1975). 'Partial likelihood.' Biometrika, 62, 269-276.

DEL (2001). *New Deal for 18-24 Year Olds: Survey of Participants. Stage 1 Report*. Department for Employment and Learning, Belfast.

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

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

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

Florens, J.-P. (1990). 'Models and inference for duration data: A survey.' Ch. 3 in J.-P. Florens, M. Ivaldi, J.-J. Laffont and F. Laisney (eds), *Microeconometrics: Surveys and Applications*, Blackwell Publishers, Oxford.

Gourieroux, C. (2000). *Econometrics of Qualitative Dependent Variables*, Cambridge University Press, Cambridge.

Greene, W. H. (2000). *Econometric Analysis,* 4th edition, Prentice-Hall International, London.

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. (1995). 'Easy estimation methods for discrete-time duration models.' Oxford Bulletin of Economics and Statistics, 57, 129-138.

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

Jenkins, S. P. (2000). 'Introduction to the analysis of spell duration data.' mimeo, Institute for Social and Economic Research, University of Essex.

Kiefer, N. M. (1988). 'Economic duration models and hazard functions.' Journal of Economic Literature, 26, 646-679.

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

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, Macreconomic 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). 'Long-term unemployment in Europe,' in Ashenfelter, O. and Card, D. (eds) *Handbook of Labor Economics*, Volume 3C, North-Holland, Amsterdam.

McVicar, D., Loudon, R., McCready, S., Armstrong, D. and Rees, G. (2000). *Young People and Social Exclusion in Northern Ireland: Status 0 Four Years On*. Training & Employment Agency, Department for Employment and Learning, Belfast.

McVicar, D. and Podivinsky, J. M. (2001). 'Duration dependence and routes out of joblessness for young people.' NIERC Working Paper No. 66, Belfast.

Meyer, B. D. (1990). 'Unemployment insurance and unemployment spells.' Econometrica, 58, 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 O. Ashenfelter and R. Layard (eds), *Handbook of Labor Economics*, Volume 2. 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 Stewart, M. (1993). 'Modelling the probability of leaving unemployment: Competing risks models with flexible baseline hazards.' Applied Statistics, 42, 1, 63-83.

National Audit Office (2002). *The New Deal for Young People*. London, The Stationary Office.

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

Neumann, G. R. (1997). 'Search models and duration data.' Ch. 7 in M. H. Pesaran and P. Schmidt (eds), *Handbook of Applied Econometrics, Vol. II: Microeconometrics*, Blackwell Publishers, Oxford.

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.

Riley, R. and Young, G. (2001a). 'Does welfare-to-work policy increase employment? Evidence from the UK New Deal for 18-24 Year Olds.' Discussion Paper 183, National Institute of Economic and Social Research, London.

Riley, R. and Young, G. (2001b). 'The macroeconomic impact of the New Deal for 18-24 Year Olds.' Discussion Paper 184, National Institute of Economic and Social Research, London.

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.

Stata Corporation (2001). *Stata 7 Reference Manual*, Stata Press, College Station, TX (www.stata.com).

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 Reenen, J. (2001). 'No more skivvy schemes? Active labour market policies and the British New Deal for the young unemployed in context.' Working Paper 01/09, Institute for Fiscal Studies, London.

Appendix A: Data Details

This section lists and defines the variables used in this study, all taken directly or derived from the JSA-register data held by DETI. First we define variables used as covariates in the estimation and secondly we define the other variables used in estimation, e.g. censored indicators, information on destination on leaving unemployment, information of unemployment duration etc.

Age on entry to unemployment

Each unemployment spell in the DETI data has a start date expressed in day/month/year format. The individual experiencing the unemployment spell has a birth date expressed in day/month/year format. Age on entry to unemployment is simply start date less birth date, expressed in complete years.

Seeking managerial, professional or related employment

This is a binary dummy indicating whether the unemployed individual describes him or herself as seeking managerial, professional or related work when signing on to the unemployment register (SOC Groups 1-3). The dummy =1 for an individual described as seeking managerial, professional or related employment and =0 otherwise.

Seeking skilled manual employment

This is a binary dummy indicating whether the unemployed individual describes him or herself as seeking skilled manual work when signing on to the unemployment register (SOC Group 5). The dummy =1 for an individual described as seeking skilled manual employment and =0 otherwise.

Number of unemployment spells

This variable is a simple count of the number of separate spells of unemployment experienced by an individual since January 1995 and prior to the current spell. A further variable, **number of spells^2**, is the squared number of unemployment spells, included to capture any possible non-linearity in the relationship between number of unemployment spells and hazard rates.

Male

A simple binary dummy (male=1, female=0) indicating gender of claimant (at beginning of unemployment spell).

Live with partner

A simple binary dummy (live with partner=1, 0 otherwise) indicating whether an individual lives with a partner (at beginning of spell). Categories covered by this definition include cohabiting and married and cohabiting with partner but not married.

MDI

MDI is a local government electoral ward-level composite index of deprivation, expressed as a number ranging from 1.1 for the least deprived ward in NI to 78.07 for the most deprived ward in NI. The MDI score is derived from separate indices for various aspects of disadvantage including ward-level unemployment levels, health indicators, housing indicators, income deprivation measures, education and skills measures and measures of access to services. The ward-level MDI score is available to download from the NISRA website (www.nisra.gov.uk), which also gives details of the construction of the composite index and component indices. This MDI score is matched to the ward in which each unemployed individual on the JSA register lives. The MDI score, sometimes also called the 'Noble' index, is based on deprivation measures for the year 1999 and essentially updates the 1994 Robson index of relative deprivation for NI.

JSA1 and JSA2

JSA1 is a simple binary dummy variable indicating whether an unemployment spell started after the introduction of JSA in October 1996 or before its introduction (1=after introduction, 0 otherwise). This simple binary dummy is used in the descriptive statistics presented in Section 4 of this report. The estimation itself uses an alternative dummy, JSA2, which can only be derived from the data following its expansion by duration (see the introduction to Section 5). This takes the value 1 for all spells of unemployment starting following the introduction of JSA, but also takes the value 1 for parts of unemployment spells falling after October 1996 that nevertheless may start before October 1996. All spells, or parts of spells before October 1996 take the value 0.

Time trend

This is simple monthly count of time from January 1995 at the beginning of an unemployment spell, i.e. January 1995=1, February 1995=2 etc.

Eligible

This is a binary indicator (1=eligible for ND18-24, 0 otherwise) based on age. It takes the value 1 for all those individuals aged 18-23 years on entry to unemployment, those individuals aged 17 at the start of the unemployment spell but who are aged 18 following six months of unemployment and those individuals aged 24 on entry to unemployment who have not turned 25 following six months of unemployment. It takes the value 0 otherwise. For estimation purposes, this variable is used to define the target age group for ND18-24.

Duration of unemployment spell

The DETI JSA register data contains start and end dates for each unemployment spell,, expressed in day/month/year format. From this the duration of the unemployment spell in days is derived. We then aggregate this, first to the nearest week (three days or less is defined as 0 weeks) and second along the lines set out in Table 3 of this report.

Destination on leaving unemployment

This information is recorded at the end of each spell of unemployment. Around 30% of destinations are categorised as 'other' in this report (see Section 4.1 for details of the destinations covered by the 'other' category).

Indicators for censored spells

Two binary indicators are constructed for spells which are incomplete at the last point of observation – one for March 1998 (i.e. the pre ND18-24 censored indicator) and one for July 2001 (the post ND18-24 censored indicator).

Appendix B: Technical Details

The following discussion sets out some of the technical details behind estimation of the PG models as presented in Section 5. Our advice to non-specialist readers is to skip this section, although certain technical points mentioned briefly in the main text of this report are enlarged upon here.

Hazard Models of Duration

In this report we focus primarily upon estimation of so-called *hazard models of duration*. These models, applied to data on *spells* (e.g. of unemployment), have been used extensively within both the econometric and applied statistics literatures. In this section we outline briefly the key concepts in duration modelling, and indicate some of the models and techniques most relevant for our analysis. The literature in this area is extensive, and various good (and generally up-to-date) published surveys exist, including Florens (1990), Gourieroux (2000, Ch. 12), Kiefer (1988), Lancaster (1990) and Neumann (1997).

We are interested in modelling the distribution of duration spells, which we consider as a random variable T with a probability *distribution function*:

$$F(t) = \text{Prob}(T < t)$$

This specifies the probability that the random variable T is less than some value t (e.g. time). We assume a corresponding probability *density function* $f(t) = \mathrm{d}F(t)/\mathrm{d}t$. As the data are durations, we can also define the *survival function*:

$$S(t) = 1 - F(t) = \text{Prob}(T >= t)$$

together with the *hazard function*:

$$\lambda(t) = f(t) / S(t)$$

with the interpretation that the *hazard* indicates the rate at which spells will be completed at duration t, given that they last until t. Alternatively, we can readily deduce that

$$\lambda(t) = - \mathrm{d}[\log S(t)]/\mathrm{d}t$$

and we say that *positive duration dependence* exists at the point t^* if

$$\mathrm{d}\lambda(t)/\mathrm{d}t > 0 \text{ at } t = t^*$$

This implies that the probability that a spell will end shortly increases as the spell increases in length. Alternatively, *negative duration dependence* exists at t^* if $\mathrm{d}\lambda(t)/\mathrm{d}t < 0$ at $t = t^*$.

Given a sample of observed data on durations of spells, interest is focussed on estimating the hazard function. There is an extensive literature on this issue, employing a variety of approaches. In this study, we are interested in the more substantial issue of estimating the hazard function in models where the hazard itself depends on a set of covariates (regressors) characterising the individuals in the observed sample. Again, there are a wide variety of alternative models that have been proposed for this.

One such class of models that is commonly used in economics is the *proportional hazard* model, where the hazard function depends on a set of (observed) covariates X with unknown parameter vector β, and the hazard function is factorised as

$$\lambda(t, X, \beta, \lambda_0) = \lambda_0(t).\varphi(X, \beta)$$

where $\lambda_0(t)$ is a "*baseline*" *hazard* that corresponds to $\varphi(X, \beta) = 1$ (possibly corresponding to the mean value of the covariates, so that λ has an interpretation as the hazard function of the mean individual in the sample). The function $\varphi(X, \beta)$ is commonly specified as $\varphi(X, \beta) = \exp(X'\beta)$, and in this case the parameter vector β can be interpreted conveniently as the constant proportional effect of X on the conditional probability of completing a spell.

The parameter vector β can be estimated in the proportional hazard model without specifying the form of the baseline hazard by using the partial-likelihood approach suggested by Cox (1975). This approach is also readily adapted to censoring in the observed data (i.e. the sample may include incomplete spells).

An alternative approach is to specify a parametric distribution for the hazard model. A wide variety of alternative specifications has been suggested (e.g. exponential, Weibull, lognormal, log-logistic, generalised Gamma, etc.), conditional on the covariates. All of these can be estimated straightforwardly by maximum likelihood methods. These specifications differ in their implication for the estimated hazard function. For example the Weibull distribution with survival function

$$S(t) = \exp\{-[\exp(-X'\beta/\gamma)t]^{\gamma}\}$$

is particularly useful in modelling data with a hazard function that is monotonically increasing or decreasing, depending upon the shape parameter γ to be estimated from the data.

Evaluation of Marginal Effects

It is common to evaluate the marginal effects of covariates in duration models at the sample means of the covariates. When the covariates are continuous-valued variables, computing marginal effects in this way is approximately equivalent (in sufficiently large samples) to evaluating the marginal effects at each individual observation, and then using the sample average of the individual marginal effects. However, when one of the covariates is a binary-valued variable (e.g. a 0-1 dummy variable), evaluating the corresponding marginal effect at the sample mean may be misleading (see Greene (2000, p. 817)). In this case the sample mean is not particularly representative. For example, in the binary 0-1 dummy variable case, a sample mean between zero and one does not correspond to a value of the dummy variable for *any* individual. Put another way, a 0-1 dummy variable will have a maximum possible variance when the sample mean is exactly 0.5 (i.e. when there are exactly 50% zeroes and 50% ones),

and in this case the sample mean will be *least* representative of any one individual observation. Hence a note of caution should be attached to interpretations of the magnitude of marginal effects evaluated from binary covariates in this manner.

Prentice-Gloeckler (PG) Estimation of Hazard Functions

The approach adopted in this study to estimating the hazard function is to consider a particular form of discrete time duration model introduced to the empirical literature by Prentice and Gloeckler (1978). This choice of model allows considerable flexibility in modelling the hazard function. The exposition in this section follows Jenkins (1997).

We consider n individuals $i = 1,...,n$, who are assumed to enter a state (e.g. unemployment) at time $t = 0$. The (instantaneous) hazard rate function for individual i at time $t > 0$ is assumed to take the proportional hazards form

$$\lambda_{it} = \lambda_0(t).\exp(X_{it}'\beta)$$

where $\lambda_0(t)$ is the baseline hazard function, and exp denotes the exponential function. This baseline hazard may take either a parametric or a non-parametric form. The covariates (regressors) characterising individual i at time t are represented by the vector X_{it} with associated unknown coefficient or parameter vector β. The focus of the estimation is assumed to be estimation of this vector β.

We assume that we observe the underlying continuous durations only for disjoint intervals through time, where the intervals are separated by time points a_1, a_2, a_3,... . Let $S(a_{j-1};X_{it})$ denote the survivor function at the start of the jth interval $[a_{j-1},a_j)$, i.e.:

$$S(a_{j-1};X_{it}) = \text{Prob}\{T \geq a_{j-1}\}.$$

Thus the probability of exit for individual i in the j-th interval $[a_{j-1},a_j)$ is given by

$$\text{Prob}\{T \in [a_{j-1},a_j)\} = S(a_{j-1};X_{it}) - S(a_j;X_{it})$$

The hazard of exit in the *j*th interval can be written as

$$h_j(\mathrm{X}_{it}) \equiv \text{Prob}\{T \in [a_{j-1},a_j) \mid T \geq a_{j-1}\} = 1 - [S(a_j;\mathrm{X}_{it})/S(a_{j-1};\mathrm{X}_{it})].$$

Given the assumption of proportional hazards, the survivor function $S(a_{j-1};\mathrm{X}_{it})$ can be expressed as:

$$S(a_j;X_{it}) = \exp[-\exp(X_{it}'\beta + \delta_j)] \qquad \text{for } j = 1,...,k,$$

where $\delta_j = \log(H_{it})$ and $H_t = \int_0^t \lambda_0(\tau)d\tau$ denotes the integrated baseline hazard at time *t*. If, following Jenkins (1997), we simplify further by assuming:

(a) all time intervals are of unit length (e.g., one week), so the recorded duration for individual *i* can be denoted by the interval $[t_i - 1, t_i)$;
(b) individuals have either left the state (e.g. unemployment) during the interval $[t_i - 1, t_i)$ (denoted by the censoring indicator $c_i = 1$) or remain in the state (denoted by $c_i = 0$, indicating a right-censored spell).
(c) the number of intervals comprising a censored spell is defined to include the last interval within which the individual is observed.

Under these assumptions, the log-likelihood function for the whole sample can be written in terms of the survivor function S(.) as

$$\log L(\beta,\delta) = \sum_{i=1}^{n} \left\{ c_i \log[S(t_i - 1; X_{it}) - S(t_i; X_{it})] - (1 - c_i) \log S(t_i; X_{it}) \right\}$$

Alternatively, the log-likelihood function can be expressed in terms of the hazard function as:

$$\log L = \sum_{i=1}^{n} \left\{ c_i \log \left\{ h_{ti}(X_{it_i}) \prod_{s=1}^{t_i - 1} [1 - h_s(X_{is})] \right\} + (1 - c_i) \log \left\{ \prod_{s=1}^{t_i} [1 - h_s(X_{is})] \right\} \right\}$$

where the discrete time hazard $h_j(X_{ij})$ in the jth interval $[a_{j-1},a_j)$ is given by

$$h_j(X_{ij}) = 1 - \exp[-\exp(X_{ij}'\beta + \gamma_j)]$$

and $\gamma_j = \log \int_{a_{j-1}}^{a_j} \lambda_0(\tau) d\tau$ denotes the natural logarithm of the integral of the baseline hazard over the jth interval $[a_{j-1},a_j)$.

With this PG model, we can assume a fully non-parametric baseline hazard with a separate parameter γ_j for each duration interval in the sample. Thus we need not make any (possibly misspecified) assumptions about the particular form of the baseline hazard function. Of course, alternatively we could choose to specify some semi-parametric or parametric function to describe the γ_j parameters.

A computationally convenient representation (based upon Jenkins (1995)) of the above log-likelihood function is estimated by maximum likelihood in Jenkin's (1997) *pgmhaz* routine implemented in the Stata package (Stata Corporation (2001)).

Frailty (Unobserved Heterogeneity)

The PG model can be extended to allow for any unobserved (or omitted) heterogeneity between individuals. Unobserved heterogeneity (or frailty) is always a potential issue in empirical models of duration (see, e.g., Meyer (1990) for an illustration), since the chosen set of covariates may not fully represent all the heterogeneity between individuals in the sample. If there is indeed unobserved heterogeneity and this is not modelled, then the estimates of the parameters on observed heterogeneity in the data (i.e. the parameter vector β) will be biased. Among other things neglected frailty can lead to over-estimation of the degree of negative duration dependence in the baseline hazard. Thus it is important to be able to test for

possible unobserved heterogeneity and, if detected, to allow for it in estimation. Jenkins (1997) allows for Meyer's (1990) suggested extension of the PG model by modelling frailty using a Gamma distributed random variable.

With the same structure as specified in the previous section, the (instantaneous) hazard rate is now given as

$$\lambda_{it} = \lambda_0(t).\varepsilon_i.\exp(X_{it}'\beta) = \lambda_0(t).\exp[X_{it}'\beta + \log(\varepsilon_i)]$$

where as before $\lambda_0(t)$ is the baseline hazard function, X_{it} denotes the covariates with unknown parameter vector β, and now where ε_i is a Gamma distributed random variate with unit mean and variance $\sigma^2 \equiv v$. This has the corresponding discrete-time hazard function:

$$h_j(X_{ij}) = 1 - \exp\{-\exp[X_{ij}'\beta + \gamma_j + \log(\varepsilon_i)]\}.$$

where again γ_j denotes the natural logarithm of the integral of the baseline hazard over the jth interval. Note that this hazard function takes a similar form to the hazard for the PG model. Similarly, the survivor function, and hence the log-likelihood function, are both augmented versions of the PG survivor and log-likelihood functions respectively.

Allowing for Gamma distributed frailty, the log-likelihood function for the full sample can be written as:

$$\log L = \sum_{i=1}^{n} \log\{(1-c_i).A_i + c_i.B_i\}$$

where

$$A_i = \left[1 + v \sum_{j=1}^{t_i} \exp\left[X_{ij}'\beta + \theta(j)\right]\right]^{-(1/v)},$$

and

$$B_i = \left[1 + v \sum_{j=1}^{t_i - 1} \exp\left[X_{ij}' \beta + \theta(j) \right] \right]^{-(1/v)} - A_i \text{, if } t_i > 1, \text{ or}$$

$$= 1 - A_i \text{, if } t_i = 1,$$

where again the interval $[t_i - 1, t_i)$ denotes the recorded duration for individual i , and where $\theta(j)$ denotes a function describing duration dependence in the hazard rate (including the non-parametric baseline hazard specification). This functional form for $\theta(j)$ may be specified by an appropriate choice of covariates. Again, Jenkins (1997) implements maximum likelihood estimation of this model with Gamma distributed frailty.

Note that the PG model's log-likelihood function is obtained as the limiting case as the Gamma variance $v \to 0$. This suggests that the two models may be tested against each other using a likelihood ratio test. This test can then be interpreted as a test of the presence of unobserved heterogeneity, conditional upon the assumed specification (e.g. Gamma) of frailty. We strongly reject the presence of unobserved heterogeneity in all cases where we allow for it in this way in our PG models of unemployment duration.[16]

Alternative distributional models for frailty may be specified. For example, an alternative to the commonly used Gamma distribution might be an Inverse Gaussian representation of frailty. Jenkins (2000) summarises the current experience with alternative parametric models for frailty, and notes that frailty models themselves can often be somewhat fragile, as their estimation can be susceptible to numerical difficulties.

A further alternative to the Meyer (1990) approach of specifying a particular parametric family (e.g. Gamma) to represent unobserved heterogeneity, is to adopt a

[16] However, as Jenkins (1997) points out, the distribution of such a likelihood ratio test is nonstandard (even asymptotically, as the sample size n becomes infinitely large), since properly the PG model is not nested within the extended model that allows for frailty.

nonparametric approach to modelling heterogeneity. Such an approach, properly implemented, would have the advantage of not relying upon a (perhaps misspecified) parametric distribution to represent any unobserved heterogeneity. Perhaps the most commonly used such approach is the nonparametric maximum likelihood estimator suggested by Heckman and Singer (1984). Baker and Melino (2000) investigate the performance of Heckman and Singer's (1984) nonparametric maximum likelihood estimator (NPMLE) for this mass point specification, in a range of different cases that they investigate using Monte Carlo simulation methods. They find that the NPMLE is well behaved even in moderately sized samples as long as the duration dependence is estimated parametrically (through the hazard function). Unfortunately in this study, this is not the case, and when both the duration dependence and the unobserved heterogeneity are estimated nonparametrically, then the estimators are subject to systematic bias. As a consequence of this, they caution against using the Heckman-Singer NPMLE when both duration dependence and unobserved heterogeneity are modelled nonparametrically.