Unemployment duration and exit states in Britain

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Abstract: This paper presents new evidence on the determinants of unemployment duration for men and women in Britain in the 1990s, using a nationally representative data set. It examines the impact of individual and local labour market characteristics on the probability of unemployment spells ending with moves into full and part-time employment, self-employment and economic inactivity. The data show that the median duration of unemployment spells among men, at 5 months, is almost double that for women, although much of this differential is explained by exits to part-time work and economic inactivity among women. Multivariate analysis suggests that policies to reduce unemployment duration and encourage full-time work, especially among men, should be targeted towards those aged 25 and over on entering unemployment and on increasing education levels. Mothers are found to have significantly lower exit rates into full-time work than both men and childless women.

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Non-technical Summary

This paper presents new evidence on the determinants of unemployment duration for men and women in Britain in the 1990s, using a nationally representative data set. It investigates the probability of unemployment spells ending with moves into full-time or part-time work, self-employment, and economic inactivity. The impact of individual heterogeneity and local labour market characteristics on each exit destination is studied. This paper builds on previous work by Narendranathan and Stewart (1993a) by estimating the probability of leaving unemployment in a discrete time independent competing risks framework with flexible baseline hazard rates.

Individual level data from Waves 1 to 7 of the British Household Panel Survey (BHPS) are used, which provide accurate information on unemployment spell duration over the period 1991-1997. These allow the probability of an unemployment spell terminating in full-time or part-time employment, self-employment or economic inactivity to be examined separately for men and women. Most previous studies have examined data for the 1970s and 1980s, and focused on men only.

Our results suggest that, although unemployment spells experienced by women tend to be shorter than for men, much of this can be explained by exits to part-time work and labour market withdrawal. Female exits rates from unemployment to full-time employment are similar to those for men. The value of competing risk analysis clearly emerges. A number of covariates have different impacts depending on the competing risk under consideration. Age, education, spouses' employment status, dependent children, health, previous employment history and local labour market conditions have different effects on the state specific exit rates for men and women. The findings imply that policies to reduce unemployment and encourage full-time employment, especially among men, should be targeted towards those aged over 25 on entering unemployment, and on increasing education levels. We find a scarring impact of previous unemployment experience, which suggests that policies aimed at reducing unemployment incidence will have longer term effects. Further investigation of gender differences reveals that mothers have considerably lower exit rates into full-time employment relative to both men and childless women. Education reduces the hazard rate into part-time employment for women. However, highly educated women are still significantly more likely to enter part-time work from unemployment than men.

Additionally, the estimated hazard rates show that the probability of leaving unemployment for full-time employment declines after an elapsed duration of five months for men and nine months for women. This suggests that policies aimed at helping the long term unemployed back into full-time work should focus on men and women with unemployment duration exceeding these.

I. Introduction

> "....waiting for work often involves so much anxiety and worry that it causes more strain than the work itself would do." (Alfred Marshall, 1920, p.462).

Policy makers, economists and social scientists in general have long been interested in the probability of an unemployed individual finding a job after a certain length of time out of work. This interest originates from a number of different concerns, including the loss of output and the cost to the social security bill resulting from unemployment.[1] Long-term unemployment may also have undesirable impacts on society and on the financial and psychological well-being of the unemployed individual.[2] This paper presents new evidence on the determinants of unemployment duration for men and women in Britain in the 1990s, using a nationally representative data set. It investigates the probability of unemployment spells ending with moves into full-time or part-time work, self-employment, and economic inactivity, and the impact of individual heterogeneity and local labour market characteristics on each.

Numerous studies using British data have examined the influences of personal characteristics and the length of the spell on the probability of leaving unemployment (early studies include Mackay and Reid, 1972; Nickell, 1979a,b; Lancaster and Nickell, 1980).[3] Many of these studies impose restrictive parametric specifications on the hazard functions, which potentially bias the results of interest or do not distinguish between exit states. The estimated coefficients are therefore some weighted average of the impact of covariates on the probability of exiting unemployment to each labour market state, rather than on the probability of finding a job.

More recent research has explicitly focused on the probability of an unemployed individual finding full-time work. For example Narendranathan and Stewart (1993a) and Arulampalam and

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[1] In 1995 Unemployment Benefit payments accounted for more than £1 in every £100 spent by the government on social security in Britain (Office for National Statistics, 1999, Table 10.21).

[2] The psychological well-being of the unemployed is something that economists do not traditionally study. Oswald (1998) however reports that the attempted suicide rate among young men in Edinburgh in 1982 was twice as high for those unemployed for over a year than for those unemployed for less than 4 weeks, and nearly twenty times higher than for those in work. Interestingly, Clark and Oswald (1994) report that people unemployed for a long time show less mental distress than those losing their jobs more recently. A number of recent studies have also documented a positive correlation between unemployment and crime, these include Freeman (1996) and Burdett, Lagos and Wright (1999).

[3] Devine and Kiefer (1991) provide a summary of studies from Britain, the United States, and other developed economies.

Stewart (1995) suggest that the probability of entering full-time work falls with age and voluntary separation from the previous job and increases with predicted earnings in employment. Other studies have examined the impact of unemployment insurance on unemployment duration (Nickell, 1979a; Atkinson et al, 1984; Narendranathan et al, 1985; Narendranathan and Stewart, 1993b). The impact of targeted assistance schemes on unemployment duration has also been examined. Dolton and O'Neill (1996) analyse the success of the Restart programme in helping the long-term unemployed return to the labour market.[4] Their results show that a Restart interview increased the probability of individuals signing off the unemployment benefit register (which does not necessarily equate to finding employment).

The relationship between the duration of an unemployment spell and the probability of finding a job (duration dependence) is itself of interest. The basic search model predicts that the distribution of unemployment spell duration is exponential, and the probability of finding a job is independent of elapsed duration. Time varying covariates may change this result. For example an exhaustive benefit entitlement system could cause the probability of finding work to increase with elapsed duration (Mortensen, 1977). An offer arrival rate that falls with elapsed duration (due to systematic search or human capital decay) has the opposite effect. In empirical studies for Britain, the probability of finding work is typically found to decline with duration in unemployment (negative duration dependence), even when controlling for different individual characteristics (Lancaster, 1979; Nickell 1979a; Atkinson et al, 1984; Jackman and Layard, 1991; Van den Berg and Van Ours, 1994; Hildreth et al, 1998). This may be due to reduced search intensity and discouragement on behalf of the worker, the erosion of human capital, or because unemployment duration is used as a signal of low productivity by employers.

Studies of unemployment duration among women are rare, Lynch (1989) and Hildreth et al (1998) are two exceptions. The reason for this undoubtedly lies in the difficulty observing unemployment among women. However, given the large increase in female labour force participation in recent years, it is clearly important to identify the determinants of unemployment duration among women.

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[4] The Restart programme was introduced in April 1987 to review the position of the long-term unemployed. Individuals were interviewed at six monthly intervals to be offered advice on benefits, job search methods and potential training. One aim of the process was to identify individuals who were either not available to work or not seeking work in order to reduce the number claiming unemployment benefit.

This paper builds on previous work by Narendranathan and Stewart (1993a) by estimating the probability of leaving unemployment in a discrete time independent competing risks framework with flexible baseline hazard rates. This procedure yields more robust results than those obtained from parametric approaches, and follows the Prentice-Gloeckler (1978) model also used by Arulampalam and Stewart (1995). Although our estimation does not allow for unobserved individual heterogeneity, we argue that the possible misspecifications and distortions from doing so are as serious as ignoring it completely. Recent research suggests that failing to take unobserved heterogeneity into account does not seriously bias results given a fully flexible baseline hazard specification (see, for example, Ridder, 1987; Han and Hausman, 1990; Meyer, 1990; Narendranathan and Stewart, 1993a; Böheim, 1999).

Individual level data from waves 1 to 7 of the British Household Panel Survey (BHPS) are used, which provide accurate information on unemployment spell duration over the period 1991-1997. These allow separate examination of male and female probabilities of an unemployment spell terminating in full-time or part-time employment, self-employment or economic inactivity. Most previous studies have examined data for the 1970s and 1980s, concentrating on men only. Our study is therefore of direct policy relevance, given that the composition of the unemployment stock altered considerably between the recessions of the 1980s and the 1990s (Turnbull, 1998), and that the proportion of working age men and women in part-time employment and self-employment, and men in economic inactivity, has been increasing in recent years (Schmitt and Wadsworth, 1993; Naylor, 1994; Taylor, 1997; Hakim, 1998).[5] Many other studies have focused on specific groups of the labour market, for example on young persons finding employment (amongst others see Lynch, 1989; Bradley and Taylor; 1991). Our data are representative for the British population in the 1990s.

Our results suggest that, although unemployment spells experienced by women tend to be shorter than for men, much of this can be explained by exits to part-time work and labour market withdrawal among women. Female exits rates from unemployment to full-time work are similar to those for men. The value of competing risk analysis clearly emerges. A number of covariates have different impacts depending on the competing risk under consideration. Age, education,

[5] Schmitt and Wadsworth (1993) describe a fall in the economic activity rate among the non-student male population of working age from 96% in 1977 to 88% in 1992. Naylor (1994) shows an increase in the proportion of employment accounted for by part-time jobs from 15% in 1971 to 28% in 1994, while Taylor (1997) presents figures illustrating the increase in the self-employment rate from 7% in 1979 to 13% in 1996.

spouses' employment status, dependent children, health, previous employment history and local labour market conditions have different effects on the state specific exit rates for men and women. The findings imply that policies to reduce unemployment and encourage full-time employment, especially among men, should be targeted towards those aged over 25 on entering unemployment, and on increasing education levels. The scarring impact of previous unemployment experience would suggest that policies aimed at reducing unemployment incidence will have longer term effects. The estimated hazard rates show that the probability of leaving unemployment for full-time employment declines after an elapsed duration of five months for men and nine months for women. This suggests that policies aimed at helping the long term unemployed back into full-time work should focus on men and women with unemployment duration exceeding these.

II. Framework

The typical framework used in the empirical analysis of unemployment duration is the job search approach (see for example, Lancaster, 1990). The duration of unemployment is modelled by specifying the conditional probability of leaving unemployment, the hazard function. The hazard function can be viewed as the product of two probabilities, the probability of receiving a job offer and the probability that the offer is acceptable.[6] Economic theory, however, is not informative about the relationship between the hazard rate and elapsed duration in unemployment. Erroneous assumptions on the form of the baseline hazard can potentially bias the estimated effects. The widely used Weibull specification allows only for hazard rates that monotonically increase or decrease with duration. Han and Hausman (1990) reject this specification using US data from the Panel Study of Income Dynamics as too restrictive. Likewise, using UK data, Narendranathan and Stewart (1993a) conclude that the Weibull model does not provide a satisfactory representation of the empirical hazard rate from unemployment.

A flexible specification of the baseline hazard rate allows for non-monotonic variation with duration, and therefore a wider range of possible effects of duration on the hazard rate are captured. This is important if non-stationarities exist in some element of the job search

[6] It should be noted that we essentially estimate reduced form specifications as we do not model the separate probabilities of receiving a job offer, and the offer being acceptable.

environment. For example, if benefits could be drawn over a limited time period only, a person optimising the expected returns from search would change his/her behaviour as the duration in unemployment lengthens. According to some models, hazard rates out of unemployment increase when benefit exhaustion approaches (Mortensen, 1977). The empirical evidence on this is mixed, see Devine and Kiefer (1991) for an overview.

The discrete time hazard rate for person i in the time interval j to leave a certain state can be written as:

$$h_j(X_{ij})=1-\exp\{\exp[X_{ij}'\beta + \theta(t)]\} \qquad [1]$$

where X_{ij} is a set of covariates, β are the coefficients to be estimated, and $\theta(t)$ is some functional form of how the duration of the spell affects the hazard rate. We assume that for each time interval there is a specific parameter that is constant over that period (Prentice and Gloeckler, 1978). In other words, we employ a fully flexible specification of the baseline hazard with an interval specific parameter $\gamma(t)$. This parameter can be interpreted as the logarithm of the integral of the baseline hazard over the relevant time interval.[7]

The extension of the standard single risk model to two or more independent exit destinations is referred to in the statistics literature as an independent competing risks model, where the log-likelihood can be split into the sum of its risk-specific hazards (Lancaster, 1990). In such a model observations which exit to a different destination are treated as censored. In our data, four competing risks are identified: unemployment can end with finding a full-time job, a part-time job, entering self-employment, or with a spell of economic inactivity.[8]

Meyer (1990) extended this model using a gamma-distributed random variable to allow for unobserved heterogeneity. However the inclusion of an error term which is independent of both observed heterogeneity and time has been criticised by Narendranathan and Stewart (1993a). Ridder (1987), Han and Hausman (1990) and Meyer (1990) conclude that the bias in the

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[7] Note that the hazard rate is restricted to the [0,1] interval by the functional form in [1].
[8] Although different states of economic inactivity can be identified in the data, for example retirement or looking after the family, these are grouped into one category.

parameters caused by omitting unobserved heterogeneity is negligible with a sufficiently flexible specification of the baseline hazard.[9]

Even stronger assumptions regarding the unobserved heterogeneity term are required if considering more than one exit destination. Two approaches have typically been adopted in the empirical literature. The first involves introducing a random disturbance term in each of the cause-specific hazards (Katz and Meyer, 1990), requiring the assumption of independence across terms. The second approach assumes a disturbance term common to all cause-specific hazards, or terms proportional to each other (Flinn and Heckman, 1982; Pickles and Davis, 1985). Narendranathan and Stewart (1993a) argue that introducing possible misspecifications through the unobserved heterogeneity term could bias the results of interest. In particular, they argue that there is no reason for any resulting distortions to be less serious than those caused by ignoring unobserved heterogeneity. Given these findings we, like Arulampalam and Stewart (1995), estimate models without accounting for unobserved heterogeneity.[10]

III. The Data Source

The data used in the analysis are from waves 1 to 7 of the British Household Panel Survey (BHPS). The first wave was designed as a nationally representative random sample of the population of Great Britain living in private (non-institutional) households in the Autumn of 1991, consisting of 5,500 households covering approaching 10,000 individuals. These original respondents have been followed and they, and any adult co-residents, are interviewed at annual intervals. Children in original sample households are also interviewed when they reach the age of sixteen. The sample therefore remains broadly representative of the British population as it changes through the decade.[11] Our sub-sample consists of individuals who provided complete responses to interviews in at least two consecutive waves and are below the state retirement age.[12]

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[9] In addition, van den Berg and van Ours (1994) find no unobserved heterogeneity effects but strong negative duration dependence for British men using aggregate data, similar to Jackman and Layard (1991).
[10] Experimenting with models that include unobserved heterogeneity, following Meyer (1990), suggest that its exclusion does not change the main results. These results are available on request from the authors.
[11] In addition, weights are provided in the data to keep the sample representative of the British population.
[12] This includes men under the age of 65 and women under the age of 60. Individuals are excluded from the analysis when they reach these ages.

At each interview, respondents are asked detailed questions relating to their current employment status, and their household composition, individual demographics and income. Respondents are also asked about any other labour market spells experienced since September one year previously. In particular, respondents are asked to recall the start dates of each new job, self-employment spell, spell of unemployment and looking for work or other labour market spell.[13] Various related characteristics are collected for each job spell experienced, including the type of employment (full-time, part-time or self-employed), occupation, industry, and the reason for leaving the job. Therefore for each unemployment spell experienced we have information on spell length, details of the previous job including reason for job separation, and information on the subsequent labour market spell.

To each unemployment spell experienced by a sample member we have attached a vector of demographic characteristics, the values for which are determined at the previous date of interview. All unemployment spells that start prior to the Wave one date of interview are discarded. This is to ensure that the vector of demographic characteristics is exogenous to the unemployment spell. In addition, we have matched in the unemployment rate in each individual's travel-to-work area at the date of interview prior to the *end* of the spell to provide information on local labour market conditions.[14] It should be noted that although the definition of unemployment is left to the respondent, the definition used by the interviewer is unemployed *and looking for work*. It is also clear that the annual recall of labour market transitions is subject to reliability problems, and Paull (1996) provides a discussion of the issues relevant to these data.[15] However, we argue that the data provide the best available information for Britain on unemployment spell duration and relevant individual demographics for men and women in the 1990s.

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[13] The respondents are given a calender to help remember precise dates of spells of paid work, unemployment, retirement, maternity leave, looking after the family or home, in full-time education, long-term sick or disabled, government training scheme or something else.

[14] Note that the local unemployment rate is measured at the date of interview prior to the end of the spell rather than that prior to the start of the spell. This is because we consider the local labour market conditions to be completely exogenous. The local labour market information is taken from the National Online Manpower Information Service (NOMIS), and is matched into the BHPS by date of interview and travel-to-work area.

[15] Paull (1996) in particular suggests that spells of unemployment are less likely to be recalled accurately than other types of spells, and that there is a tendency for them to be redefined as time out of the labour force. Higher attrition rates among the unemployed are countered by not requiring individuals to have been interviewed at all waves.

Note that our focus on spells, rather than individuals, is problematic if there is correlation between the spells experienced by one person. We condition the estimation on previous unemployment experience which should overcome this problem. Furthermore, we have re-estimated the regressions using individual's first unemployment spell observed in the sample period. The results do not change the main conclusions.

IV. Descriptive statistics

The BHPS annual employment histories provide information on 1,744 unemployment spells, experienced by 1,158 individuals since the Wave 1 date of interview. The unemployment spells are not therefore concentrated among a small number of individuals. Table 1 shows that the majority of the spells are experienced by men (1,074, or 62%). The mean duration of these spells is 8 months (9 months for men and 6 months for women). The median durations are about one half the mean, with 50% of spells among men lasting less than 5 months, and among women less than 3 months.[16]

Table 2 decomposes spells by destination state and gender. Among men, 57% of unemployment spells end with a transition into full-time work, and these have the lowest mean and median duration (at 7 months and 4 months). Economic inactivity accounts for the next largest proportion of unemployment spells among men (11%). Spells terminating with economic inactivity have the longest mean and median duration among completed spells (at 10 and 6 months).[17] Exits to part-time work and self-employment account for 6% and 8% of male exits from unemployment. Narendranathan and Stewart (1993a), using data from the late 1970s, find that 3.5% of unemployment spells terminate with self-employment, and 2% with part-time work. This is further evidence of the increased importance of part-time and self-employment in Britain. Unemployment spells ending in part-time employment have a higher mean and median

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[16] These descriptive statistics are biased downwards because the data restrict the maximum length of an observed unemployment spell to be 80 months (the maximum number of months between the Wave 1 date of interview and the Wave 7 date of interview). Very long unemployment spells are therefore missed or censored. Narendranathan and Stewart (1993a) report a mean unemployment duration of 23.3 weeks, although their data relate to the late 1970s. Similarly, Narendranathan and Elias (1993) use data covering the mid to late 1970s for a sample of men aged 23 in 1981, and report a mean duration for completed spells of 5 months. Hildreth et al (1998) find an average unemployment duration of approaching 10 months from retrospective data covering the period 1980-1993.

[17] The relatively long duration of unemployment spells terminating in economic inactivity among men can perhaps be explained by viewing labour market withdrawal as a form of continuing unemployment. This withdrawal from the labour market may be the result of discouragement caused by the worker's unemployment experience.

duration than those ending in full-time employment, while those ending in self-employment have a higher mean and a similar median.

Among women, 43% of unemployment spells end with a transition into full-time work, substantially fewer than for men. Again, however, such unemployment spells have the shortest mean and median duration. Almost one quarter of unemployment spells experienced by women end in part-time employment, with a mean duration of 6 months (median of 4 months). Only 2% terminate in self-employment, and these tend to have a relatively long duration with a mean (median) of 8 (5.5) months. A larger proportion of unemployment spells end in economic inactivity for women than for men (16.5%), and these have a mean (median) duration of 6 (3) months, considerably shorter than for men.

Table 3 presents life table estimates for unemployment duration by gender and destination state. It shows the proportion of unemployment spells surviving for a given period of time. The first row considers all destination states for men, and shows that 8% of unemployment spells experienced in the 1990s last one month or less, and 20% last under two months. One half of spells last less than 5 months, 56% less than 6 months and 76% less than one year. According to these data, 24% of unemployment spells experienced by working aged men in the 1990s fall into the typical definition of long-term unemployment (lasting 12 months or more). A rather different picture emerges if we concentrate only on exits into full-time work. The second row shows that 25% of unemployment spells experienced by men terminate in full-time work within 3 months, 43% within six months, and 60% within a year. Almost 30% of male unemployment spells have not terminated in full-time employment after two years. The third row considers exits into any employment; full-time or part-time employment, or self-employment. This shows that 30% of unemployment spells have ended in employment within 3 months, 50% within 6 months, 68% within one year and 80% within two years.

Among women, 14% of unemployment spells experienced in the 1990s last one month or less, and 30% last under two months. Two thirds of spells last less than 6 months and 86% less than one year. According to these data, 14% of unemployment spells experienced by working aged women in the 1990s are long-term. The second row shows that 26% of unemployment spells experienced by women terminate into full-time work within 3 months, 43% within six months, and 59% within a year. One third of female unemployment spells have not terminated into full-time employment after two years. These proportions are very similar to those for men. The third row shows that 36% of unemployment spells have ended in employment within 3 months, 59% within 6 months, 77% within one year and 86% within two years.

These tables have shown that, although unemployment spells experienced by women typically tend to be shorter than those experienced by men, much of this can be explained by exits into part-time employment and economic inactivity. Female exit rates from unemployment into full-time employment are very similar to those for men.[18]

Table 4 provides some descriptive statistics for the characteristics of men and women that are measured at the date of interview prior to the start of the unemployment spell.[19] This shows that approaching one half of unemployment spells are experienced by men and women under the age of 31, and a large proportion by those aged 25 or under (31% of spells among both men and women).[20] Three fifths of spells are experienced by married individuals, and 65% by those with no children.[21] Unemployment is just as likely to be experienced by the highly educated as those with lower education levels. Over 50% of unemployment spells among men are experienced by those educated to A-Level standard or higher. Those with no formal qualifications experience between 15% and 20% of unemployment spells. There is evidence of unemployment persistence, especially among men. Less than one third of unemployment spells are experienced by men with no prior experience of unemployment (48% for women), while 17% (14% for women) are experienced by those with three or more previous unemployment spells.[22] The gender bias in self-employment and part-time employment is reflected, with 21% of unemployment spells experienced by men with previous self-employment experience compared with 7% among women. Over one half of unemployment spells are experienced by women with part-time employment experience, compared with 13% among men.

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[18] While recognising the problems differentiating between unemployment and economic inactivity for women, particularly when relying on recall data, we believe that it is important to do so for policy purposes.

[19] We use the information collected at the date of interview prior to the start of the unemployment spell to ensure that that all variables are predetermined and therefore exogenous. Excluding those with missing data on any relevant variables, there are 490 individuals who experience more than one unemployment spell in the sample period. This table therefore refers to spells rather than to individuals.

[20] Turnbull (1998), using data from the Labour Force Survey, reports that 30% of the unemployed in 1993 were under 25, and more than two-thirds are under the age of 40. Our figures are consistent with these.

[21] Throughout the paper, married refers to legally married or cohabitating (living as a couple). Thus any reference to a husband or wife in employment also applies to a live-in partner.

V. Estimation Results

The results of the discrete time hazard models are presented in Table 5 (men) and 6 (women).[23] We also estimated a single risk model, but the log-likelihood tests reject pooling the exit states and these results are not presented here. The figures reported are the estimated coefficients. Note that the proportionate impact of each variable on the state-specific hazard rate can be calculated by taking the exponent of the coefficient. Figures 1, 2 and 3 illustrate the resulting estimated baseline hazard rates out of unemployment for men and women. There are insufficient exits from unemployment to self-employment among women to determine robust estimates.

Men

The first column of Table 5 shows the effects of the covariates measured at the date of interview prior to the start of the unemployment spell on the hazard from unemployment to full-time employment. Age clearly has a large and well determined impact, with men over the age of 30 less likely to leave unemployment for full-time employment than those aged 25 and under. Indeed, the hazard rate to full-time employment is approximately halved for men over the age of 40.[24] This inverse relationship between age and the probability of leaving unemployment is a common finding in the British literature (see, for example, Pissarides and Wadsworth, 1992; Narendranathan and Stewart, 1993a; Arulampalam and Stewart, 1995; Dolton and O'Neill, 1996).[25] Narendranathan and Stewart (1993a), for example, find that men aged 45 and over have the lowest probability of leaving unemployment for full-time work.

Men of Afro-Caribbean origin have significantly lower hazard rates into full-time work than whites (at the 10% level of significance), while those of Pakistani or Bangladeshi origin have higher rates. Men with a working spouse have a higher hazard rate into full-time employment, suggesting some interdependence in husband and wife labour supply choices (Booth, Garcia-

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[22] Arulampalam, Booth and Taylor (1999) investigate unemployment persistence among working age men, and find strong evidence of state dependence in unemployment.

[23] Estimation used Jenkins (1997) Stata procedure. Means of variables included in the models are available in Appendix Table 1.

[24] The quantitative impact of a variable is calculated by taking the exponent of the coefficient.

[25] Garcia-Serrano (1996) and Bover et al (1999) find a similar relationship for Spanish men, although Lambert (1992) concludes that long term unemployment is more prevalent among young men in Australia. In the U.S., Han and Hausman (1990) find that increasing age reduces the duration of unemployment spells terminating in recalls, but increases that for spells terminating in new jobs.

Serrano and Jenkins, 1996, and Dolton and O'Neill, 1996, come to similar conclusions). The quantitative impact is large, increasing the hazard rate by some 87%. There is also some evidence that the exit rate into full-time work increases with education, although only the coefficient on the A-Level or equivalent variable is well determined.

A health condition that limits the type or amount of work possible, a source of non-labour income, and being a local authority tenant all significantly reduce the probability of leaving unemployment.[26] The effects of health and housing tenure are consistent with previous studies (see for example Narendranathan and Stewart, 1993a). Recent research has found that local authority tenants in general are more likely to be unemployed and persistently unemployed, are less likely to move, particularly for job reasons, and if they do move are more likely to move shorter distances (Hughes and McCormick, 1990; Coleman and Salt, 1992; Wadsworth, 1998; Böheim and Taylor, 1999; Arulampalam et al, 1999). Oswald (1996, 1998) conjectures that home-ownership in today's climate reduces workers' residential mobility, contributing to higher levels of unemployment. Microeconometric evidence from the 1990s, however, suggests that home owners have higher exit rates from unemployment than renters. The effect of non-labour income is consistent with search theory, with a longer search duration implied by the higher reservation wage caused by non-labour income. Individuals will be more selective in accepting job offers if they have out-of-work income.

There is some evidence that previous unemployment experience reduces the exit rate from unemployment to full-time work.[27] In particular, men who have experienced three previous unemployment spells have a 26% lower probability of finding full-time work (this is significant at the 12% level). This indicates the presence of what Heckman and Borjas (1980) refer to as 'occurrence dependence', suggesting that unemployment has a scarring effect and a long term impact on future labour market behaviour.[28] Previous experience of self-employment reduces

[26] Non-labour income refers to income from state provided benefits, private transfers and investments. This could be considered endogenous, in that the unemployed will receive social security benefits. However, the results shown are robust to the exclusion of this variable.

[27] These experience variables are calculated from the lifetime employment history collected at Wave 2 of the survey. A dummy variable is in each model to control for individuals not interviewed at Wave 2. Although the coefficients on these variables are suggestive of state dependence effects, we do not take initial conditions or unobserved heterogeneity into account and therefore we can not be certain of the exact mechanism at work here.

[28] Arulampalam et al (1999) study unemployment persistence among men in Britain in some detail, and conclude that previous unemployment has a scarring effect. Phelps (1972), Lockwood (1991), Pissarides (1992, 1994) and Blanchard and Diamond (1994) suggest reasons why previous unemployment may impact upon future labour market outcomes.

the exit rate to full-time work for men by 40%. This may indicate that potential employers view previous self-employment as a negative signal, that individuals with self-employment experience have higher reservation wages or are otherwise more selective in the employment that they are willing to accept.[29] The unemployment rate in the individual's travel-to-work area at the previous date of interview has little effect on the probability of leaving unemployment for full-time work, all things equal.[30]

Most previous studies of unemployment duration focus either on the single risk of exit or on finding full-time employment. However, given the increasing importance of part-time employment and self-employment, it is valuable to examine these competing risks. The second column of Table 5 focuses on exits to part-time employment for men. Married men whose partner is in work or has dependent children have significantly higher rates into part-time work than men with no family. This again suggests some interdependence in husband and wife labour supply choices. There is also some evidence suggesting that men in rented accommodation (either private or social) are less likely to enter part-time work from unemployment than owner-occupiers (this is significant at the 10% level). Men with previous part-time experience have a higher probability of leaving unemployment for part-time work. The size of this effect is large, increasing the hazard rate by a factor of three. Local labour market conditions appear to have little impact on the hazard rate into part-time employment for men.

The third column of Table 5 focuses on transitions from unemployment to self-employment for men. Men aged 25 and under have the lowest probability of making this transition, in contrast to exits to full and part-time work. Also, the employment status of a partner, health and non-labour income have no significant impact. Men with formal qualifications are more likely to enter self-employment than those with no qualifications, although again the coefficients are in general not statistically significant. Men who are laid off from their previous job are less likely to enter self-employment relative to those who leave for health reasons or to look after the family or home (significant at the 10% level). Previous unemployment spells reduce the hazard rate into self-employment, with the impact of two previous unemployment spells being particularly large and

[29] Williams (1998) assesses the impact of self-employment experience on future wages of men and women in the United States. He identifies a potential wage cost to self-employed women who return to the wage and salary sector that does not exist for men.
[30] Coles (1997) develops a model where, in equilibrium, the average duration in unemployment increases as the unemployment stock increases.

well-determined, reducing the rate by almost 60%. Previous self-employment experience increases the hazard by a factor of three. Again, local labour market conditions appear to have little significant impact on the exit rate to self-employment.

The final columns consider transitions from unemployment into economic inactivity. Few statistically significant relationships emerge here. Men with a health condition that limits the type or amount of work possible have a higher exit rate from unemployment into economic inactivity (by a factor of 2). Private tenants have lower transition rates into inactivity. This probably reflects the typical age and mobility (both labour market and geographical) associated with this particular housing tenure. The positive and well-determined impact of the local unemployment rate suggests that there is some discouraged worker effect.

Figure 1 plots the estimated baseline hazard rate from unemployment into full-time work for men. The hazard rate is evaluated at the means of the covariates (it is calculated for the average man in the sample). This shows evidence of a non-monotonic hazard rate, which increases up to the sixth month of unemployment and peaks at about 0.17. It then declines to the end of the first year, and fluctuates around 0.05. The probability of leaving unemployment for full-time employment therefore declines after six months, suggesting policies should be targeted specifically towards men with elapsed unemployment duration of more than six months. This evidence of negative duration dependence in the hazard rate from unemployment to full-time work is consistent with previous work (for example, Jackman and Layard, 1991; van den Berg and van Ours, 1994; Hildreth et al, 1998).

Figure 2 shows the estimated baseline hazard rates from unemployment to part-time employment, self-employment and economic inactivity for men. All show evidence of a non-monotonic relationship with elapsed duration. The rate into part-time work initially increases to 0.012 after two months, and then to 0.018 after seven months. It then fluctuates around 0.005. The baseline hazard from unemployment into self-employment for men suggests an initially increasing rate to the third month of unemployment (at 0.006), followed by a decline. A further peak occurs at 9 months (again at 0.006), after which the rate fluctuates around 0.002. The low transition rates into self-employment from unemployment may be because the unemployed are unable to raise the capital necessary to establish a business.[31] The baseline hazard from unemployment to economic inactivity illustrates a relatively constant hazard. It fluctuates from 0.002 to 0.007.

Women

Table 6 presents the competing risks estimates for women, and the first column tabulates the hazard rate from unemployment to full-time work. The results suggest that, as for men, a working spouse increases the probability of leaving unemployment for full-time work, although the size of the impact is smaller and not significant for women. Therefore the presence of a working spouse increases the probability of men and women entering full-time work from unemployment. The presence of dependent children reduces the probability of a woman leaving unemployment for full-time work, although the impacts are not statistically significant from zero at conventional levels. Lynch (1989) finds a similar result for young women in the U.S.. This may reflect the disincentive effect of entering work caused by a loss of state benefits, or the costs of childcare. Education in general increases the hazard rate into full-time work, with the 'A'-Level and 'O'-Level terms being particularly large and well-determined. As for men, a source of non-labour income reduces the probability of entering full-time employment (by 50%) consistent with the job search model. The local unemployment rate increases the rate into full-time employment for women.[32]

The second column of Table 6 focuses on transitions from unemployment to part-time employment for women. Women aged between 41 and 50 are less likely to enter part-time work than those aged 25 or under, while those educated to O-Level standard or who have a limiting health condition also have lower hazard rates into part-time work. Indeed, education in general has a negative effect. There is some evidence that married women, and those with working husbands, have higher hazard rates from unemployment into part-time work. The number of children have a positive impact, but again not statistically significant at conventional levels. The association between these family characteristics and the exit rate to part-time employment are different to that to full-time employment, suggesting that family responsibilities are an important factor in the choice of hours to supply to the labour market. This is consistent with other

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[31] Blanchflower and Oswald (1998) and Taylor (1999) provide evidence on the relationship between access to capital and self-employment.
[32] This may be caused by women changing their job search behaviour by, for example, increasing their search intensity or reducing their reservation wage, in response to an increase in the supply of locally available labour.

research showing that part-time work is used by many women for brief periods of time, and is highly correlated with changes in household demographics (see, for example, Blank, 1994).

The final column of Table 6 shows the estimates for transitions from unemployment to economic inactivity, which are rather poorly determined.[33] A few notable relationships emerge however. Dependent children appear to reduce the probability of women exiting unemployment to economic inactivity, although only the effect of two children is statistically significant at conventional levels. Local authority tenants also have lower hazard rates into economic inactivity from unemployment relative to owner-occupiers. Previous part-time employment and a higher local unemployment rate increase the hazard rate, the latter indicative of a discouraged worker effect.

Figure 3 plots the estimated baseline hazard rates from unemployment to full- and part-time work and economic inactivity for women. This shows an estimated hazard into full-time work increases to an initial peak at 0.05 after 3 months duration, with further peaks at 7 and 10 months. The rate then clearly declines. Again, this suggests the presence of a non-monotonic relationship and negative duration dependence in the hazard rate out of unemployment into full-time work. The nature of the relationship between elapsed duration and the exit rate results in less clear policy implications for women than for men, although it suggests that policies should be particularly directed to women unemployed for more than nine months. The baseline hazard from unemployment to part-time work for women initially declines to 0.02 in the fourth month. It then increases to approaching 0.06 by the eighth month before declining again. A further peak occurs in the eleventh month of an unemployment spell, after which the baseline hazard clearly falls. The estimated hazard rate into economic inactivity is relatively flat, fluctuating around 0.01. Despite this, some evidence of non-monotonicity emerges, with peaks in the hazard rate after three months and nine months of unemployment.

Investigating gender differences

To investigate further the differences in the probability of leaving unemployment between men and women, we also estimate an independent competing risks model where we pool both sexes and allow for various interactions of covariates by gender. The results from doing so are

[33] Note that there are too few observations to estimate the impact of variables on the transition from unemployment to self-employment with any confidence for women.

summarised in Table 7. Specification (i) is the same model reported in Tables 5 and 6, but includes a gender dummy, while specification (ii) includes additional gender interactions.

Estimation of specification (i) shows that women are significantly more likely to enter part-time work and economic inactivity from unemployment than men, and are less likely to enter self-employment. This corresponds with our descriptions of the raw data in Table 2. Specification (ii) shows that women aged under 25 with no qualifications and no children are significantly less likely than men to enter full-time employment. The coefficients on the age and education interaction terms are quantitively small and statistically insignificant. However, children further reduce the hazard rate from unemployment into full-time work for women. For example, a woman aged in her late twenties with one child has a coefficient of (-1.116-0.286-0.538)= -1.94 relative to men, and with two children a coefficient of -2.05, compared with (-1.116-0.286)= -1.402 for a similar woman with no children. Children are therefore a big hindrance in escaping unemployment to full-time work for mothers. The interaction term with the local unemployment rate is positive and well determined, suggesting that low local labour demand increases the exit rate into full-time work for women relative to men.

Young women with no qualifications and no children have a higher probability of leaving unemployment for part-time work relative to men, with a point estimate of 1.886, implying an exit rate some seven times greater than for men. Again, age has little influence on this result. However, education has a significant negative effect on the hazard rate. For example, a woman in her thirties with a degree has a coefficient of (1.886+0.784-1.540)=1.13 relative to men, compared to (1.886+0.784)=2.67 for a similarly aged woman with no qualifications. Although qualifications reduce the probability of women entering part-time employment from unemployment, they do not off-set the large gender difference. Highly educated women have a higher hazard rate into part-time work from unemployment than men. It is interesting to note that the presence of children has no differential impact on the hazard rate into part-time work for men and women.

Specification (ii) reveals little about gender differences in the hazard rates into self-employment and economic inactivity. The estimated coefficients in general are statistically insignificant. Women do however have a higher chance of escaping unemployment in a deteriorating labour market. The coefficients on the unemployment rate and gender interaction terms are positive for all exits (but significant only for exits to full-time work). But remember that women start from a worse situation relative to men. The coefficient on the gender dummy indicates a huge penalty for being female in entering full-time work.

Summary of results

These results have emphasised the importance of analysing hazard rates from unemployment by destination state, as a number of covariates have different impacts depending on the independent competing risk studied. Among men, for example, the probability of leaving unemployment for full-time work is highest for those aged 25 and under, while that for entering self-employment is lowest for that age group. Some interdependence among husbands and wives labour supply choices emerges. The exit rate from unemployment to full- and part-time work among men is higher for those with an employed wife at the start of the spell. Women have a higher exit rate to full- and part-time employment if they have an employed husband at the spell start. Dependent children reduce the hazard rate into full-time work and economic inactivity for women, but in general increase that into part-time work. This may suggest that working mothers enjoy the flexibility offered by such jobs, or that they find it difficult to find full-time work. Men with dependent children have a higher exit rate into part-time work.

Education generally has a positive impact on the probability of leaving unemployment for men, while living in local authority accommodation has a negative effect. The latter probably reflects the relative (im)mobility of council tenants. These associations are less apparent among women, Women with a limiting health condition are less likely to leave unemployment in general. Previous labour market experience is less important in escaping unemployment for women than for men and, in particular, previous unemployment appears to have less of a scarring effect. The local unemployment rate is inversely related to the probability of entering any kind of employment for men but increases that of entering full-time work for women. Further investigation of gender differences reveal that children reduce further the hazard rate into full-time work for women, while education reduces the hazard from unemployment into part-time work.

VI. Conclusions

This paper has analysed unemployment duration among men and women in Britain in the 1990s using a nationally representative longitudinal data set, the British Household Panel Survey. The data show that the median duration of unemployment spells among men, at 5 months, is almost double that for women. However, much of this duration differential can be explained by exits into part-time employment and economic inactivity. Female exit rates from unemployment into full-time employment are very similar to those for men, but account for a lower proportion of transitions.

Multivariate analysis in a discrete time independent competing risk framework with flexible baseline hazard rates provides important results. Firstly, the non-monotonicity of the estimated baseline hazard rates suggest that using parametric specifications may bias the coefficients of interest. Secondly, covariates are found to have very different impacts on the state-specific exit rates from unemployment for both men and women. The results suggest that policies to reduce unemployment and encourage full-time employment, especially among men, should be targeted towards those aged over 25 on entering unemployment, and on increasing education levels. The scarring impact of previous unemployment implies that such policies aimed at reducing short-term unemployment incidence will have longer run effects. The estimated hazard rates suggest that policies to increase the flow from unemployment to full-time employment should be targeted at men unemployed for more than six months, and at women unemployed for more than nine months. Mothers are found to have considerably lower exit rates into full-time employment relative to both childless women and men. Education reduces the hazard rate into part-time employment for women, although highly educated women are still significantly more likely to enter part-time employment from unemployment than men.

Of equal importance to policy makers is the duration of the subsequent employment spell, and in particular, the impact of previous unemployment duration. This is an obvious and important area for future research.

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Table 1: Number and duration of unemployment spells by gender: BHPS Waves 1-7

	Men	Women	Total
Number of spells	1074	670	1744
Mean duration (months)	8.83	6.35	7.87
Standard. Deviation	11.15	9.23	10.51
Median duration (months)	4.70	3.19	4.04

Notes: Only spells starting since Wave 1 date of interview. Men aged under 65 and women aged under 60. Incomplete (right hand censored) spells included.

Table 2: Duration of unemployment spell by destination state and gender: BHPS Waves 1-7 (months).

Destination state	Men			Women			Total		
	%[a]	Mean	Median	%[a]	Mean	Median	%[a]	Mean	Median
Spell incomplete[b]	7.5	16.9	10.3	8.4	12.2	4.7	7.8	15.0	7.4
Destination state missing	10.9	14.3	8.0	6.8	13.6	10.4	9.4	14.1	8.3
Self-employment	8.0	8.6	3.9	2.1	7.9	5.5	5.7	8.5	4.2
Part-time employment	6.0	7.5	5.7	23.7	5.7	3.8	12.8	6.2	4.1
Full-time employment	56.8	6.7	3.9	42.5	4.7	2.8	51.3	6.1	3.5
Inactive	10.8	9.5	6.3	16.5	5.8	3.0	13.0	7.8	4.2
Total	100	8.8	4.7	100	6.4	3.2	100	7.9	4.0

Notes: Only spells starting since the Wave 1 date of interview. Men aged under 65 and women aged under 60. [a]Column percentages [b]This includes spells censored due to attrition and current unemployment spells at wave 7.

Table 3: Unemployment duration by gender and destination state (surviving percent).

Destination state	Duration (months)									
	1	2	3	4	5	6	9	12	18	24
Men										
All states	92	80	66	58	50	44	30	24	16	12
Full-time employment	94	84	75	69	62	57	45	40	33	28
Any employment	93	82	70	63	56	50	38	32	24	20
Women										
All states	86	70	56	46	39	32	21	14	9	7
Full-time employment	91	83	74	68	62	57	48	41	34	33
Any employment	88	76	64	56	49	41	29	23	17	14

Notes: Only spells starting since the Wave 1 date of interview. Men aged under 65 and women aged under 60. Any employment includes full-time, part-time and self-employment.

Table 4: Descriptive Statistics

Variable	Men Mean	Men Std Dev.	Women Mean	Women Std. Dev
Age				
Under 25	0.309		0.306	
26-30	0.154		0.149	
31-40	0.250		0.267	
41-50	0.171		0.200	
>50	0.116		0.078	
Ethnicity				
Afro-caribbean	0.007		0.014	
Indian	0.013		0.016	
Pakistani/Bangladeshi	0.012			
Family				
Married	0.597		0.602	
Spouse in work	0.398		0.517	
No children	0.635		0.666	
One child	0.139		0.168	
Two children	0.156		0.122	
Three or more children	0.070		0.044	
Education				
Degree or above	0.102		0.145	
‘A’-Levels or equivalent	0.435		0.320	
‘O’-Levels or equivalent	0.205		0.237	
Qualification below ‘O’ Level	0.094		0.113	
No qualifications	0.164		0.185	
Other demographics				
Health limits work possible	0.109		0.097	
Has non-labour income	0.690		0.789	
Housing tenure				
Owner-occupier	0.665		0.632	
Local Authority tenant	0.196		0.216	
Private tenant	0.139		0.152	
Previous employment				
Laid off from previous job	0.268		0.234	
Years previously spent unemployed	0.294	0.881	0.166	0.537
No previous unemployment	0.321		0.481	
One prior unemployment spell	0.332		0.290	
Two prior unemployment spells	0.174		0.138	
Three + prior unemp. spells	0.173		0.108	
Previous self-employment?	0.213		0.067	
Previous part-time employment?	0.129		0.517	
Local labour market demand				
Unemployment rate	8.605	2.769	8.441	2.381

Notes: Descriptive statistics of unemployment spells starting since the Wave 1 date of interview, men aged under 65 and women aged under 60. Sample limited to non-missing observations. Variables measured at date of interview prior to the start of the spell, with the exception of the local unemployment rate which is measured at the date of interview prior to the end of the spell. N=435 women and 594 men.

Table 5: Discrete time proportional hazard regression results for unemployment duration in a competing risks framework: Men

Variable	Unemployment spell ends with move to: *Full-time work*		*Part-time work*		*Self-employed*		*Economic inactive*	
Age[a]								
26-30	-0.062	*0.36*	-0.476	*0.81*	1.300	*2.97*	0.466	*1.15*
31-40	-0.232	*1.51*	-0.647	*1.18*	0.435	*0.96*	0.124	*0.31*
41-50	-0.569	*3.02*	-0.422	*0.75*	0.597	*1.31*	-0.076	*0.17*
>50	-0.596	*2.57*	0.521	*0.88*	0.357	*0.61*	0.740	*1.57*
Ethnicity								
Afro-caribbean	-1.813	*1.80*	0.675	*0.64*	—	—	0.852	*1.07*
Indian	0.521	*1.20*	—	—	—	—	0.863	*1.03*
Pakistani/Bangladeshi	0.918	*1.97*	—	—	1.531	*1.30*	0.310	*0.28*
Family								
Married	-0.151	*0.78*	-0.780	*1.29*	0.372	*0.86*	-0.232	*0.57*
Spouse in work	0.628	*3.85*	0.994	*2.11*	0.023	*0.06*	-0.085	*0.24*
One child[b]	0.017	*0.09*	1.033	*2.39*	0.441	*1.16*	-0.820	*1.52*
Two children[b]	0.106	*0.59*	1.040	*1.79*	0.138	*0.33*	-0.183	*0.40*
Three or more children[b]	0.234	*0.96*	0.597	*0.70*	-0.517	*0.89*	0.380	*0.75*
Education[c]								
Degree or above	0.124	*0.58*	0.585	*0.98*	0.441	*0.78*	0.063	*0.12*
‘A’-Levels or equivalent	0.298	*1.90*	-0.088	*0.19*	0.747	*1.78*	0.467	*1.37*
‘O’-Levels or equivalent	0.212	*1.21*	0.163	*0.33*	0.700	*1.55*	-0.199	*0.47*
Qualification below ‘O’ Level	-0.092	*0.41*	0.806	*1.42*	0.837	*1.53*	-0.193	*0.40*
Other demographics								
Health limits work possible	-0.558	*2.84*	0.195	*0.47*	-0.360	*0.85*	0.779	*2.58*
Has non-labour income	-0.300	*2.67*	0.356	*0.90*	-0.156	*0.56*	-0.173	*0.61*
Housing tenure[d]								
Local Authority tenant	-0.387	*2.60*	-0.758	*1.67*	-0.354	*1.01*	-0.404	*1.29*
Private tenant	0.223	*1.46*	-1.753	*1.68*	0.036	*0.09*	-0.946	*1.74*
Previous employment								
Laid off from previous job	0.004	*0.04*	0.150	*0.42*	-0.577	*1.77*	-0.212	*0.70*
Years previously unemployed	-0.103	*1.14*	-0.200	*0.62*	0.036	*0.27*	0.170	*1.50*
One prior unemployment spell[e]	-0.104	*0.77*	-0.718	*1.54*	0.012	*0.04*	-0.093	*0.28*
Two prior unemployment spells[e]	-0.155	*0.95*	-0.213	*0.43*	-0.824	*1.91*	-0.349	*0.81*
Three + prior unemp. spells[e]	-0.296	*1.55*	0.137	*0.27*	-0.634	*1.40*	-0.137	*0.32*
Previous self-employment	-0.514	*3.63*	0.021	*0.05*	1.166	*4.47*	-0.446	*1.42*
Previous part-time employment	0.090	*0.55*	1.031	*2.62*	0.488	*1.21*	0.050	*0.13*
Local labour market demand								
Unemployment rate	0.034	*0.17*	-0.009	*0.01*	0.012	*0.02*	1.324	*2.63*
Number of observations (person months)	6983							
Log-likelihood	-1409		-214		-348		-398	
χ^2	263.3		96.0		87.1		61.1	

Notes: Figures reported are the coefficients. t-statistics in italics. Only spells since the Wave 1 date of interview. Month dummies also included. [a] Aged 25 or under is the base category. [b] No children is the base category, [c] No qualifications is the base category, [d] Owner-occupation is the base category, [e] No previous unemployment is the base category. A dash, —, indicates that the variable has been dropped due to collinearity. Variables measured at date of interview prior to the start of the spell except for the local unemployment rate which is measured at the date of interview prior to the end of the spell.

Table 6: Discrete time proportional hazard regression results for unemployment duration: Women

	Unemployment spell ends with move to:					
Variable	*Full-time work*		*Part-time work*		*Economic inactive*	
Age[a]						
26-30	-0.250	*0.97*	-0.402	*1.20*	0.136	*0.36*
31-40	0.271	*1.20*	0.132	*0.45*	-0.100	*0.25*
41-50	0.138	*0.56*	-0.688	*1.91*	-0.374	*0.82*
>50	-0.336	*0.89*	-0.310	*0.71*	-0.132	*0.24*
Ethnicity						
Afro-caribbean	-0.085	*0.16*	-0.734	*0.71*	—	—
Indian	0.157	*0.28*	-1.015	*1.26*	-0.505	*0.48*
Family						
Married	-0.229	*0.79*	0.575	*1.61*	-0.023	*0.05*
Spouse in work	0.398	*1.33*	0.462	*1.41*	-0.185	*0.39*
One child[b]	-0.315	*1.29*	0.332	*1.12*	-0.047	*0.12*
Two children[b]	-0.309	*1.08*	0.104	*0.32*	-1.467	*2.21*
Three or more children[b]	-0.517	*1.09*	0.453	*1.04*	-0.408	*0.59*
Education[c]						
Degree or above	0.214	*0.72*	-0.512	*1.24*	0.606	*1.29*
'A'-Levels or equivalent	0.728	*2.81*	-0.053	*0.18*	0.417	*0.97*
'O'-Levels or equivalent	0.609	*2.38*	-0.776	*2.54*	-0.078	*0.19*
Qualification below 'O' Level	0.441	*1.42*	0.009	*0.03*	-0.300	*0.55*
Other demographics						
Health limits work possible	-0.027	*0.12*	-1.117	*2.74*	0.068	*0.18*
Has non-labour income	-0.668	*3.63*	-0.281	*1.00*	0.337	*0.99*
Housing tenure[d]						
Local Authority tenant	0.123	*0.62*	0.025	*0.10*	-0.662	*1.80*
Private tenant	0.143	*0.66*	0.019	*0.06*	0.056	*0.15*
Previous employment						
Laid off from previous job	-0.032	*0.19*	0.093	*0.43*	-0.452	*1.45*
Years previously unemployed	-0.002	*0.01*	-0.032	*0.11*	0.043	*0.18*
One prior unemployment spell[e]	-0.079	*0.43*	-0.328	*1.48*	-0.337	*1.02*
Two prior unemployment spells[e]	0.173	*0.69*	-0.504	*1.45*	0.570	*1.48*
Three + prior unemp. spells[e]	0.181	*0.63*	-0.311	*0.70*	-0.133	*0.27*
Previous self-employment?	-0.190	*0.58*	0.189	*0.50*	-0.520	*0.70*
Previous part-time employment?	-0.203	*1.21*	0.271	*1.22*	0.462	*1.60*
Local labour market demand						
Unemployment rate	0.974	*3.19*	0.412	*1.36*	1.433	*2.81*
Number of observations (person months)			3370			
Log-likelihood	-705		-503		-328	
χ^2	156.6		108.6		93.8	

Notes: Insufficient observations to estimate self-employment equations. Figures reported are the coefficients. t-statistics in italics. Only spells since the Wave 1 date of interview. Month dummies also included. [a] Aged 25 or under is the base category. [b] No children is the base category, [c] No qualifications is the base category, [d] Owner-occupation is the base category, [e] No previous unemployment is the base category. A dash, —, indicates that the variable has been dropped due to collinearity. Variables measured at date of interview prior to the start of the spell except for the local unemployment rate which is measured at the date of interview prior to the end of the spell.

Table 7: Gender differences in Exits?

	Full-time		Part-time		Self-employment		Economic Inactivity	
i) Dummy for Female								
Female	-0.065	*0.64*	1.529	*6.83*	-1.004	*2.81*	0.800	*4.03*
log likelihood	-2151		-753		-433		-757	
ii) Dummy, Interactions								
Female	-1.116	*2.49*	1.886	*2.22*	-1.188	*0.68*	0.548	*0.65*
Female*Age:								
Age 26-30	-0.286	*0.99*	-0.080	*0.12*	0.813	*0.67*	-0.243	*0.48*
Age 31-40	0.305	*1.21*	0.784	*1.36*	0.697	*0.55*	-0.287	*0.55*
Age 41-50	0.377	*1.41*	-0.224	*0.39*	0.529	*0.42*	-0.444	*0.80*
Age 50+	-0.179	*0.45*	-1.025	*1.68*	–	–	-0.976	*1.63*
Female*Education:								
Degree	0.136	*0.38*	-1.540	*2.21*	-0.336	*0.29*	0.624	*0.97*
A-levels	0.247	*0.85*	-0.153	*0.27*	-0.701	*0.78*	0.210	*0.41*
O-levels	0.191	*0.63*	-1.251	*2.21*	-1.601	*1.55*	0.019	*0.03*
Other qualification	0.381	*1.02*	-0.950	*1.50*	–	–	-0.168	*0.24*
Female*Number of children:								
One	-0.538	*1.89*	-0.534	*1.10*	-0.128	*0.16*	0.636	*1.01*
Two	-0.648	*2.12*	-0.376	*0.67*	0.038	*0.04*	-1.114	*1.49*
Three	-0.869	*1.72*	0.548	*0.62*	–	–	-0.673	*0.84*
Female*Labour market:								
Unemployment rate	1.076	*3.07*	0.537	*0.78*	0.921	*0.78*	0.442	*0.66*
log-likelihood	-2140		-743		-426		-751	

Note: See notes for tables 5 and 6. Pooled male and female samples. N=10,353 person months. Both specifications include the covariates listed in tables 5 and 6.

Appendix

Table 1: Means of covariates by gender

Variable	*Men*		*Women*	
	Mean	SD	Mean	SD
Duration dummies				
Month 1	0.103		0.149	
Month 2	0.090		0.120	
Month 3	0.075		0.097	
Month 4	0.064		0.075	
Month 5	0.057		0.064	
Month 6	0.049		0.054	
Month 7	0.043		0.044	
Month 8	0.037		0.037	
Month 9	0.034		0.032	
Month 10	0.030		0.025	
Month 11	0.028		0.022	
Month 12	0.026		0.019	
Month 13-15	0.068		0.044	
Month 16-18	0.056		0.036	
Month 19-21	0.045		0.031	
Month 21-24	0.038		0.027	
Month 25-30	0.058		0.041	
Month 31-36	0.034		0.031	
Month 36+	0.065		0.051	
Age				
26-30	0.128		0.139	
31-40	0.270		0.249	
41-50	0.206		0.224	
>50	0.122		0.117	
Ethnicity				
Afro-caribbean	0.014		0.012	
Indian	0.006		0.015	
Pakistani/Bangladeshi	0.005			
Other ethnic minority	0.006		0.017	
Family				
Married	0.603		0.581	
Spouse in work	0.306		0.447	
One child	0.123		0.183	
Two children	0.180		0.110	
Three or more children	0.087		0.048	
Education				
Degree or above	0.103		0.141	
‘A’-Levels or equivalent	0.337		0.224	
‘O’-Levels or equivalent	0.210		0.296	
Qualification below ‘O’ Level	0.113		0.133	
Other demographics				
Health limits work possible	0.141		0.178	
Has non-labour income	0.739		0.869	
Housing tenure				
Local Authority tenant	0.314		0.280	

Variable	*Men*		*Women*	
	Mean	SD	Mean	SD
Private tenant	0.116		0.147	
Previous employment				
Quit previous job	0.081		0.069	
Laid off from previous job	0.268		0.207	
Years previously unemployed	0.387	0.993	0.175	0.513
One prior unemployment spell	0.328		0.332	
Two prior unemployment spells	0.203		0.119	
Three + prior unemp. spells	0.218		0.113	
Previous self-employment?	0.263		0.070	
Previous part-time employment?	0.099		0.545	
Local labour market demand				
Unemployment rate (x10)	0.827	0.280	0.759	0.272
Number of observations (person months)	6983		3370	

Notes: Variables measured at date of interview prior to the start of the spell except for the local unemployment rate which is measured at the date of interview prior to the end of the spell.

Figure 1: Estimated baseline hazard into full-time work for men

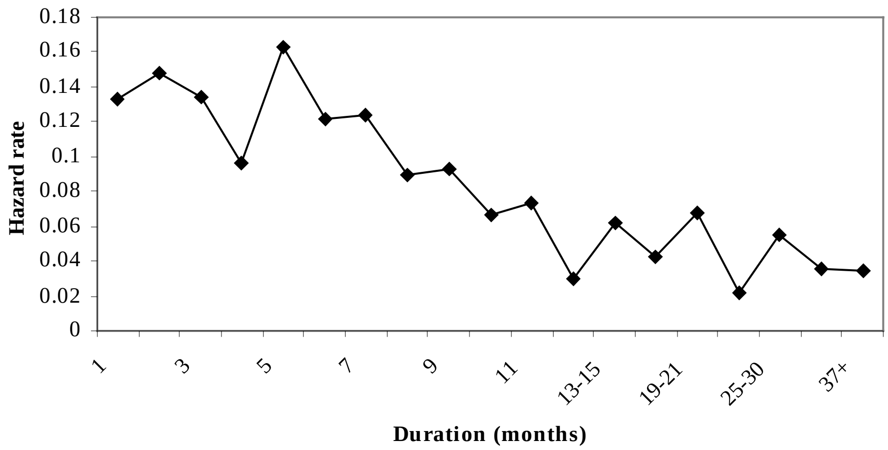

Figure 2: Estimated baseline hazard rates into part-time employment, self-employment and inactivity for men

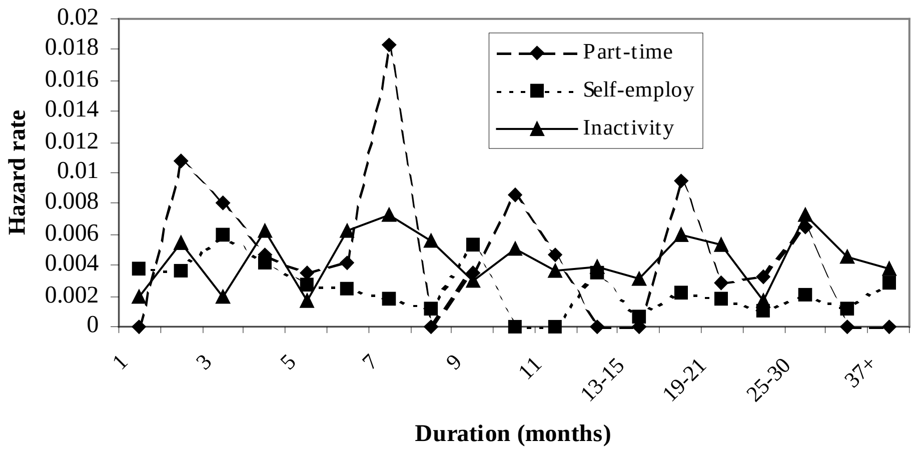

Figure 3: Estimated baseline hazard rates into full- and part-time work and inactivity for women