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TEMPORARY JOBS: STEPPING STONES OR DEAD ENDS? *

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Abstract

In Britain about 7% of male employees and 10% of female employees are in temporary jobs. In contrast to much of continental Europe - with stricter employment protection provisions - this proportion has been relatively stable over the 1990s. Using data from the British Household Panel Survey, and informed by relevant theory relating to probation, sorting and human capital investment, we find that temporary workers report lower levels of job satisfaction, receive less work-related training, and are less well-paid than their counterparts in permanent employment. However, there is some evidence that fixed-term contracts are a stepping stone to permanent work. Women (but not men) who start in fixed-term employment and move to permanent jobs fully catch up to those who start in permanent jobs.

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Keywords: temporary jobs, fixed term contracts, individual unobserved heterogeneity, job-specific effects.

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I. INTRODUCTION

Temporary contracts are often regarded as an important component of labour market flexibility. Temporary workers can be laid off without incurring statutory redundancy payments or restrictions imposed by employment rights legislation. This may explain the dramatic growth in temporary jobs in France, Italy and Spain, countries characterised by high levels of employment protection. The proportion of temporary workers in these countries doubled between 1985 and 1997. In contrast, in the United States and United Kingdom, which have relatively little employment protection regulation, the proportion of the workforce on fixed term contracts has been fairly stable.[1]

While temporary contracts can avoid some labour market inflexibilities (see for example Bentolila and Bertola, 1990; Bentolila and Saint Paul, 1994; and Booth, 1997), there are potential costs. Some commentators have expressed concern about the quality of the stock of jobs and the lack of opportunities for career advancement associated with temporary or flexible work (Farber, 1997 and 1999; Arulampalam and Booth, 1998). Purcell, Hogarth and Simm (1999) have also found case study evidence from 50 British firms of decreasing *employer* enthusiasm for temporary contracts, owing to the low levels of retention and motivation of such staff.

These issues are particularly important since governments are moving away from welfare benefits towards 'workfare' systems. Temporary jobs – in the UK, subsidised by the government (see Dickens, Gregg and Wadsworth, 2000) – provide an important potential route for welfare recipients to enter the permanent workforce. The desirability of policies such as the subsidisation of temporary jobs depends upon whether they are 'dead end' jobs with poor pay and prospects, satisfactory careers in their own right, or 'stepping stones' to permanent employment in good jobs.

Remarkably little is known about temporary workers in Britain (Dex and McCulloch, 1995), and it is therefore important to improve our understanding of their career opportunities

[1] The proportion of workers in fixed term contracts in France, Italy and Spain increased respectively from 4.7, 4.8 and 15.6 percent in 1985 to 13.1, 8.2 and 33.6 percent in 1997. In the UK, the proportion of workers in fixed term contracts was 7 percent in 1985 (thus, higher than in Italy and France at that time) but remained stable over time and reached 7.4 percent in 1997. In 1997, of the 150 million workers in the European Union, about 12 percent were employed on fixed-term contracts (European Commission, 1999).

and to assess the extent and impact of this form labour market flexibility. In this paper we investigate three main issues. First, we describe who holds temporary jobs in 1990s Britain. Second, we investigate how satisfied temporary workers are with their jobs, and how much training they receive compared with permanent workers. Third, we estimate how long it takes temporary workers to move into permanent jobs, which workers will be successful in this way, and how the wage profiles of workers who have ever held a temporary job compare with permanent workers over time. We address these issues using longitudinal data from the first seven waves of the British Household Panel Survey (BHPS), conducted over the period 1991-1997. The analysis is carried out separately for men and women in employment and distinguishes between 'casual and seasonal workers' and workers on 'fixed-term contracts'.

Our results confirm that temporary jobs are not desirable as long-term careers. They typically pay less than corresponding permanent employment, and are associated with lower job satisfaction and work-related training. However, we do find evidence that fixed-term contracts are effective stepping-stones to permanent jobs. Furthermore, women who start with a fixed-term job and then move to permanent work fully catch-up to the wage level earned by women who start in permanent work. Men suffer a long-term 5% loss in wages from starting with a fixed-term contract.

In the following section, we present the main hypotheses underlying our analysis. In Section III, we describe the data source and examine the raw data to see the extent of temporary job holding in the British labour market. In Section IV, we provide a picture of temporary work in 1990s Britain. In particular, we estimate who gets a temporary job, and the level of satisfaction and the work-related training of temporary workers compared to permanent workers. Section V examines the impact of an experience of a temporary job on subsequent employment and wages. The final section summarises and draws conclusions.

II. HYPOTHESES

Even in the UK – which has relatively mild restrictions on dismissal for redundancy or cause – it is costly to discharge long-serving employees. Workers with sufficient length of service are entitled to statutory redundancy pay and can claim unfair dismissal.[2] Insofar as these are simple transfers from the firm to the separating worker, there is no particular reason to avoid permanent appointments.[3] A worker on a temporary contract will – in a competitive labour market – receive a higher wage that just offsets the loss of the expected value of redundancy pay. However, severance costs can contain a deadweight element. There is a considerable cost in time and expense – as well as in overall industrial relations – to a firm in being brought before an industrial tribunal to defend an unfair dismissal claim. For these reasons, firms might prefer to have a cushion of workers without employment rights who can be freely discharged in the event of adverse market conditions, even if the firm must pay a wage premium to these workers.

There are a number of reasons why temporary workers may not in practice receive a compensating differential in the form of a higher wage than permanent workers. It is not efficient for workers in temporary employment to invest heavily in specific human capital. This leads to a lower wage, but also has implications for the characteristics of workers holding temporary jobs. The workers holding these jobs will be the ones for whom there is either a greater probability of wishing to separate (either to change occupation or geographical location) or a higher cost (or lower benefit) to acquiring specific human capital. Young, single individuals might be disinclined to make a large investment in a particular job until they are sure of their career and regional preferences. If it is believed – as in Lazear and Rosen (1990) – that women are more likely to move to non-market employment, then women will also be more likely to hold temporary posts. Older workers might – given the shorter period of return – also be less inclined to invest in specific human capital.

Temporary workers may also differ from permanent workers in ability. The ability level of temporary workers very much depends upon the future job prospects of a post. It is possible that firms maintain a high-turnover, low ability pool of temporary workers to adjust employment to match market conditions. Individuals with low ability to acquire specific human capital will then go through a succession of low-paid, temporary jobs. Alternatively,

[2] The length of service needed to obtain most of these employment rights has recently been lowered from two years to one year. The maximum sum awardable for unfair dismissal has also recently been increased to £50,000. For women and ethnic minorities there is no limit on the sum. This may be a deterrent for firms to appoint women and workers from ethnic minorities to permanent posts.

[3] Indeed, as argued in Booth (1997), government imposed redundancy pay can substitute for incomplete private contracting and sustain more efficient investments in specific human capital, since workers receive job protection.

firms may view the initial temporary contract as a probationary stage – subject to job performance and employment demand, workers will move into permanent employment at the firm. If the likelihood of eventual permanency is sufficiently high, the temporary job can be attractive to a worker of high ability even if it is low-paid. As argued by Loh (1994) and by Wang and Weiss (1998), firms may seek to have the right workers self-select into probationary jobs by instituting a wide differential paid to the successful workers when they achieve permanency.

There are other reasons why wages of temporary workers may be low. Booth and Frank (1996) find evidence that some unions are more concerned about longer serving members, and agree contracts with steep returns to seniority. Temporary workers may be an extreme case of outsiders, who receive a low wage compared to permanent workers. However, there are also situations where temporary jobs might have high wages. If productivity is positively correlated with the returns to general (rather than specific) human capital, then highly productive workers may prefer to be employed in a succession of temporary jobs. This may hold for high skill jobs such as computer systems experts who may in fact view high-paid temporary jobs as a form of self-employment.

It may be possible to distinguish between these alternative scenarios by considering the dynamic aspects of temporary jobs. In jobs where the temporary contract is a form of probation, there will be a high probability of obtaining a permanent post at the current employer. In this case, the low wage during the temporary contract period will be compensated for by high future wages. There should be little overall career loss to starting with a temporary post. In contrast, if temporary jobs are held by individuals with low ability to acquire specific human capital, there will be a large, permanent career loss to these individuals.[4]

[4] A recent literature (Autor, 1999; Polivka, 1996; Abraham and Taylor, 1996; and Houseman and Polivka, 1999) makes the further point that firms can hire temporary workers from temporary help supply firms who have economies of scale in screening and training temporary workers. In view of this possibility, firms might find it optimal to only hire temporary workers when there is an element of probation involved. Unfortunately, our data do not distinguish workers at temporary help supply firms.

III. THE DATA

The data used in our analysis are the first seven waves of the British Household Panel Survey (BHPS), 1991-1997. This is a nationally representative random sample survey of private households in Britain. Wave 1 interviews were conducted during the autumn of 1991, and annually thereafter (see Appendix A). Our analysis is based on the sub-sample of white men and women who were born after 1936 (thus aged at most 60 in 1997), who reported positive hours of work, who provided complete information at the interview dates, who had left school and were employed at the time of the survey, and who were not in the armed forces or self-employed. We have a longitudinal sample of 1,740 male workers and 1,981 female workers.

The data allow us to distinguish two types of temporary work. The first type refers to seasonal or casual jobs; the second type refers to jobs done under contract or for a fixed period of time. The precise form of the question asked in the BHPS interviews is given in Appendix A. The percentages of workers in these two types of temporary work are given in Table 1, where individuals are disaggregated by gender. Over the seven-year period, the average percentage of male workers in all temporary jobs is 6.8%, with 3.9% of them being in seasonal and casual jobs and 2.9% in jobs involving fixed-term contracts. The proportion of women in temporary work is higher, with 6.3% of all women employees being in seasonal and casual jobs and 3.3% in fixed-term contracts.[5] Table 1 also reports the male and female average hourly wages disaggregated by type of contract (permanent, seasonal and casual, or fixed-term contract), the wage differences by contract and their significance.[6] For men, permanent work always provides higher wages. The largest wage gap is between permanent and seasonal-casual workers, averaging £3.76 over the period, a highly significant 78% wage gap. The hourly pay differential between permanent and fixed-term contract workers is also significant over the seven-year period, but it is only £1.17 (a 16% wage gap). For women, the highest wages are earned by workers on fixed-term contracts, who receive a significant £0.90

[5] The proportions of male and female workers in seasonal and in fixed-term contracts has remained fairly stable over the sample period.
[6] The hourly wage rate is given as ω=PAYGU/[(30/7)(HS+κHOT)], where PAYGU is the usual gross pay per month in the current job (deflated by the 1997 Retail Price Index), HS is standard weekly hours, HOT is paid overtime hours per week, and κ is the overtime premium. We set κ at 1.5, the standard overtime rate, but all our results below are robust to alternative values of κ ranging between 1 and 2.

per hour (a 13% wage gap) more than permanent workers. The wage gap between seasonal-casual workers and workers in fixed-term contracts is a significant £2.27 (46% wage gap). Thus the data suggest that temporary workers are heterogeneous in terms of their remuneration, with fixed-term contract workers receiving significantly higher wages than seasonal-casual workers.

Temporary jobs differ from permanent jobs in the hours of work as well as wages. For men in permanent jobs, the mean of normal hours worked per week is 45, with a standard deviation of 11; for men in seasonal-casual jobs these figures are 28 and 17, respectively, while for those on fixed-term contracts, they are 41 and 15. The data show an overall greater dispersion for women. Their mean weekly hours of work is 32 with a standard deviation of 13 if they are in permanent jobs, while the corresponding figures are 21 and 13 if they are in seasonal-casual jobs, and 31 and 14 if they are on fixed-term contracts.[7] There is a bimodal distribution (reflecting part-time working) of normal hours for male seasonal-workers, and for all types of female workers.

The differences between workers in seasonal-casual jobs and workers on fixed-term contracts also emerge when we consider their distribution by occupation and industry. There is a large concentration of seasonal-casual male and female workers in personal and protective services, sales, plant and machine operative and other low-skill occupations and in primary, distribution and catering industries.[8] But the largest share of male and female workers on fixed-term contracts is in professional and technical occupations across almost all industries.

The raw data also allow us to have a preliminary look at the longer-term effects of having had a temporary job. From the 1997 wave of the BHPS, we select a sample of permanent workers who were in one of the following labour market states in 1991, the first survey year: permanent job, seasonal-casual job, fixed-term contract, unemployed, or out of the labour force. Men in a permanent job in both years earn an hourly wage of £9.50 in 1997. Men who were on a fixed-term contract in 1991 and are in a permanent job seven years later earn about £1 less than those who have been permanent at both interview dates, but this difference is not statistically significant. The difference is larger (around £2 per hour) and

[7] Segal and Sullivan (1997) also find that US temporary workers display higher standard deviations in hours worked than permanent workers.

significant for those who were either in a seasonal-casual job or were unemployed in 1991. Those who were out of the labour force in 1991 earn almost £4 per hour less in 1997. Interestingly, women who experienced a fixed-term contract in 1991 and are in a permanent job in 1997 have the highest female wages in that year. But the difference between their wages and those of women who have been in permanent jobs in both survey years is small (about £0.70) and not significant. The wage gap for women who were unemployed in 1991 is £2.30 per hour and highly significant.[9]

Overall, the two types of temporary work seem to identify two distinct groups of workers. On one hand, we have workers in seasonal-casual jobs, who, on average, receive lower wages and put in less effort (as measured by hours) than permanent workers, are more concentrated in low-skill occupations, and have poor future career prospects. On the other hand, we have workers on fixed-term contracts, who are better paid than seasonal-casual workers (and, in the case of women, receive higher wages than permanent workers), work longer hours and are primarily concentrated in professional and technical occupations. In the raw data, there is no significant negative career effect of having held a fixed-term contract. In the following sections, we investigate the extent to which differences between temporary and permanent workers persist after controlling for individual and workplace characteristics, and the consequences of various contract types for labour market transitions, remuneration and wage dynamics.

IV. A PICTURE OF TEMPORARY WORK

We now look more closely at a number of characteristics of temporary work, controlling for worker and workplace characteristics. We examine who gets a temporary job and the levels of job satisfaction and training of temporary workers compared to those of permanent workers.

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[8] A sizeable group of seasonal-casual female workers are also in transport, banking and other service industries.
[9] The post-displacement wages of workers who have been unemployed is the focus of an extensive literature (see *inter alia* Arulmapalam (2000) and the references therein), and therefore will not be a focus of the current study.

IV.1 Who gets a temporary job?

To address this question in a multivariate setting, we perform multinomial logit regressions for men and women separately.[10] Table 2 reports the risk ratios of being in seasonal-casual work and on a fixed-term contract relative to being in permanent work. For men, relative to the base of individuals aged 35 to 44, men aged 25 to 34 are significantly less likely to be in a seasonal-casual job, while men aged 45 and over are between two and three times more likely to be in either form of temporary work.[11] Higher educational attainment (A-level and higher degrees) is associated with seasonal/casual work.[12] This finding is striking given that our sample excludes full-time students, and thus this association cannot be accounted for by young people's employment while studying, in the holidays, or in a gap year before entering university. Managers, professionals, and skilled workers are less likely to be in casual/seasonal work. Experience in the labour market (full or part-time) also has a negative association with casual/seasonal work.[13] An additional year of experience significantly reduces the risk of being in seasonal-casual jobs by about 31% and the risk of being on fixed-term contracts by another 22%. Workers with a high number of layoffs are more likely to be in temporary work. For an average male worker, an additional layoff increases the risk of being in a season-casual job by 49% and the risk of being on a fixed-term contract by 30%.[14] For men, there is a clear pattern that both types of temporary work represent a secondary

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[10] We performed several pooled (men and women) regressions. Despite the higher raw percentages (see Table 1), the regression results show that women are less likely than men to be in any type of temporary work, after controlling for demographic and labour market characteristics. We always rejected pooling by gender. For example, when we use the same specification as in Table 2 plus "female" (the gender dummy variable), the χ^2 test on "female" being zero in the two types of temporary work is 6.37 with a *p*-value of 0.0413. Higher χ^2 values (and smaller *p*-values) were obtained after "female" was interacted with other explanatory variables included in Table 2. We also performed a test for pooling the two types of temporary work, a test for pooling permanent work and seasonal-casual work, and a test for pooling permanent work and fixed-term contracts using the procedure suggested by Cramer and Ridder (1991). The three tests strongly rejected pooling.

[11] "Age" refers to the individual's age when he/she entered the survey.

[12] For non-British readers, those who completed their education in 1988 (born in 1971-72) were the first to study for the General Certificate of Secondary Education (GCSE) qualification; earlier cohorts would have studied for "O-(Ordinary)-level" qualifications. GCSE and O-level qualifications roughly correspond to a high-school diploma; "A(Advanced)-level" qualifications correspond to education beyond high-school, but short of a university degree. "Vocational degree" includes qualifications such as teaching and nursing qualifications, City and Guilds certificates, Higher National Certificate/Diploma, and University Diploma. Some of these qualifications may not require A-level qualifications.

[13] All the experience variables (measured in years) and the number of layoffs are constructed using the retrospective work history data collected in wave 3 of the BHPS and the wave-on-wave work history information collected at every survey. See Booth, Francesconi and Garcia-Serrano (1999).

[14] This finding is consistent with Stewart (2000) and Arulampalam (2000). Stewart (2000) argues that unemployment experience followed by low paid unstable jobs contributes to observed low pay persistence.

labour market inhabited by the young and old, who do not have as much attachment to the labour force as measured by work experience and the avoidance of layoffs.

How does the pattern differ for women? The effects described above for men largely continue to hold for women in seasonal-casual jobs. The main differences concern fixed-term contracts. Women with older children (5-18), with more education, in professional/technical/teaching employment, and in the public sector are significantly more likely to hold fixed-term contracts. In these ways, fixed-term contracts represent less of a secondary labour market for women.

IV.2 Job satisfaction

Despite its measurement problems, job satisfaction may offer a useful perspective on many aspects of the labour market, through its correlation with job separations, effort and productivity (Clark, 1996). Panel A of Table 3 reports estimates of an ordered probit model of seven different components of job satisfaction, as well as an overall measure, for men and women separately.[15] Each aspect of job satisfaction is measured on a scale from 1 to 7, where a value of 1 corresponds to “not satisfied at all” and a value of 7 corresponds to “completely satisfied”. The overall measure reveals that seasonal-casual (both male and female) workers are significantly less likely to be satisfied with their jobs than permanent workers. However, no difference in overall job satisfaction emerges between workers in permanent jobs and workers on fixed-term contracts. When we consider the different aspects of job satisfaction separately, we find that workers in both types of temporary work are less satisfied than permanent workers with their promotion prospects and job security.

IV.3 Training opportunities

Temporary and permanent workers may also differ in their receipt or take-up of on-the-job training. In Panel B of Table 3, the pooled probit regression estimates show that the male probability of receiving work-related training is 12% lower for workers on fixed-term contracts and 20% lower for men on seasonal-casual contracts, relative to permanent workers, *ceteris paribus*.[16] Female workers on fixed term contracts have a 7% lower probability than permanent workers of being trained, while seasonal-casual females have a 15% lower probability. Training intensity measures the number of days of training. The pooled tobit regressions show that seasonal-casual workers receive, on average, between 9 and 12 fewer training days per year than permanent workers, but there is no differential training intensity between permanent workers and fixed-term workers. Controlling for unobserved heterogeneity reduces the effects on both training incidence and training intensity only marginally. In summary, we find that both types of temporary jobs are – relative to permanent jobs – of low quality, as measured by job satisfaction, and work-related training opportunities. These gaps seem to be larger for seasonal/casual jobs than fixed-term temporary jobs. This may be the case since fixed-term contracts may not have the same negative career effects as do seasonal/casual jobs. We investigate these career prospects in the following section.

V. THE EFFECTS OF TEMPORARY EMPLOYMENT ON CAREER PROSPECTS

If temporary jobs are probationary in nature, successful workers should eventually move into permanent employment without suffering long-term negative wage effects. In this section, we examine what happens to temporary workers in terms of the duration of temporary jobs, whether or not such jobs lead to permanent work, and the long-term wage effects of holding temporary jobs.

V.1 Job duration

How long do temporary jobs last compared to permanent jobs? Kaplan-Meier estimates of job duration, including both completed and uncompleted spells, are given in Table 4, where job tenure is defined as months in the same job with the same employer and not involving a promotion. The estimates show that the median duration of seasonal-casual jobs over the

[15] In the BHPS interviews, individuals are asked to report their satisfaction level for each of the seven aspects of their job first, and then, in a separate question, they are asked about their overall satisfaction. The pooled (men and women) regressions reveal that women are significantly more satisfied than men in all but two aspects of their job (promotion prospects and initiative). Clark (1996) reports similar results and discusses a number of plausible explanations.

[16] Our measure of training incidence takes the value of unity if the worker has received training in the past 12 months to increase or improve their skills in the current job. The measure of training intensity is the number of days spent in skill-enhancing training during the last 12 months in the current job. Using the same definition of training, Arulampalam and Booth (1998) find a similar result for the first five waves of the BHPS.

1990s is very short: it is about 3 months for men and 6 months for women. The median duration of fixed-term contracts is higher and around 12 months for both men and women. But permanent jobs last for a substantially longer time, with a median duration of almost 3½ years for men and 2½ years for women. By 5 years, almost all male and female temporary jobs have finished, as compared with 64% of male and 73% of female permanent jobs.

Where do workers go at the conclusion of a temporary job? Table 4 reveals that the destination patterns by gender are quite similar (last column). About 71% of men and 73% of women in temporary jobs go to another job at the same employer; another 26% and 24%, respectively, go to a job at a different employer; and another 3% leaves the labour force. We observe virtually no transitions from either of the two types of temporary work to unemployment, and therefore these are not reported in the table (see Boheim and Taylor, 2000). Table 4 shows that, of those employed in a seasonal-casual job, 28% of men and 34% of women have become permanent workers between 1991 and 1997. About 1 in 7 workers did so within the first three months of their job. However, the median seasonal-casual job duration before exit into permanency is 18 months for men and 26 months for women. For workers on fixed-term contracts, the transition rate to permanency is significantly higher for men (38%) and almost the same for women (36%). The median duration of fixed-term contracts before turning into permanent jobs is about 3 years for men and 3½ years for women. Finally, regardless of the type of temporary employment and gender, about 70% of workers gaining permanency continue working for the same employer.[17]

To investigate the transition of workers from temporary to permanent employment in a multivariate setting, we specify a discrete-time proportional hazard model that relates the exit process to a number of individual- and job-specific characteristics. We fully exploit the time variation of job tenure by using a monthly measure. The time-varying regressors for which we have precise information (such as occupation, industry, sector and firm size) therefore also differ by month, while other time-varying regressors (for example, union coverage and local labour market conditions) take the same value for all months between interviews. Because we condition the estimating sample on temporary workers, the number of transitions is too small to allow estimation of competing-risks models, in which the exit

[17] Segal and Sullivan (1997) note that a majority of US temporary workers are employed in permanent jobs one year later, especially in clerical and technical occupations. In their analysis, however, they do not specify whether or not this transition occurs within the same firm.

process into permanency gained in the same firm differs from that into permanency gained in another firm. We do, however, allow the determinants of exit behaviour to vary between spells starting in seasonal-casual jobs and spells starting in fixed-term contracts. The estimation is performed both with and without a Gamma mixture distribution that is meant to capture unobserved heterogeneity between individuals.[18] Table 5 presents the estimation results, with columns [1] and [2] reporting the estimates without and with unobserved heterogeneity. For three out of the four exits, we find that including a mixing distribution is relevant and has significant effects on the coefficients of some of the covariates. It does not, however, improve the model fit in the case of the male exit from fixed-term contracts, for which the estimates in the two columns do not significantly differ from each other.

Our results show that the transition from fixed-term to permanent work differs for men and women. For men, only age, part-time employment status, and a few occupational groups (craft, sales and machine operatives) appear to be good predictors for this exit. This result suggests that, conditional on being on a fixed-term contract, the timing of the entry into permanency is likely to follow a well determined temporal pattern, which has little to do with either observed personal and firm-specific characteristics or worker's unobservables. The evidence for women is rather different. The strong positive effect of any educational qualification on this exit rate is likely to be spurious, as it disappears (except for higher and university degrees) once unobserved heterogeneity is controlled for. Also the negative effect of being employed in a part-time job may not be genuine for the same reason.[19] However, women employed in any organisation of the public sector have a much lower exit rate than those employed in the private sector, even after controlling for education and occupation. A higher number of previous layoffs increases the exit rate into permanency. This may capture vintage effects, as suggested by the lower risk of exit for the youngest cohort of workers. Finally, the local U/V ratio has a significant negative effect on the hazard rate of leaving a fixed-term contract and gaining permanency. A higher unemployment rate could be associated with a lower availability of permanent jobs in the labour market, while fixed-term

[18] We estimated two distinct specifications of the baseline hazard function. The first specification constrained the baseline to be of the commonly adopted Weibull type, while the second is non-parametric to avoid the potential biases caused by mis-specification of a parametric baseline [Meyer (1990), Han and Hausman (1990) and Dolton and van der Klaauw (1995)]. The Weibull model was found to be mis-specified, with upward sloping baseline estimates suggesting positive duration dependence, clearly rejected by the data. We therefore only present the non-parametric estimates.

contracts may provide firms with an additional instrument to face adjustments in their product demand.

Regardless of a worker's gender, both part-time work and living in an area with adverse labour market conditions reduce the chance of exiting seasonal-casual work into permanency. Again, there is evidence of a positive vintage effect. Table 5 also documents some striking gender differences. For men, we find a strong occupational gradient, with workers in managerial, technical and craft occupations having higher risk of leaving seasonal and casual work than workers in semi-skilled and unskilled occupations. For women, instead, the occupational gradient is clearly less pronounced, while other observables play a major role. In particular, those employed in the local government sector and non-profit organisations are significantly less likely to gain permanency than those employed in the private sector, and so are workers in the youngest age group compared to those in the 35-44 age group. Interestingly, women (but not men) who work in union-covered organisations have a higher chance of leaving their seasonal-casual jobs.[20]

A natural hypothesis is that workers' effort will be used by employers to screen out the more able or hard-working temporary workers for retention. We would therefore expect the amount of effort to be a crucial determinant of exit from temporary contracts into a permanent position at a firm. As a proxy for effort, we use the number of unpaid overtime hours usually worked in a week. Because of endogeneity problems, we use predicted (rather than actual) unpaid overtime hours, whose identification is achieved through exclusion restrictions. These estimates are reported at the bottom of Table 5.[21] The estimates show that, after controlling for unobserved heterogeneity, a higher number of hours of unpaid overtime work increases women's chances of exiting from any type of temporary work. This is, however, not the case for men.[22]

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[19] Notice that 57% of the observed spells (measured in months) on fixed-term contracts for women are in part-time jobs, as compared to 18% for men.

[20] The hazard rates of leaving any type of temporary work do not significantly differ by industry. Instead, we do find evidence of firm-size effects. Typically, workers (both men and women) in small firms are more likely to end any type of temporary work into a permanent job than workers of larger establishments.

[21] The number of children by four age groups, dummy variables for cohort of entry in the labour market (5), region of residence (6), and whether a worker receives a performance-related pay are assumed to affect an individual's exit propensity only through their effect on unpaid overtime hours. Inclusion of *actual* unpaid overtime hours does not significantly change the results, and thus we do not report those estimates.

[22] We explored the relationship between effort and exit rates by looking at two additional specifications, one in which we distinguish the effect of total hours of overtime work from that of paid overtime hours, and another specification in which we only include the number of hours of overtime work. All the other covariates enter the regressions as in Table 5. Again, the exit into permanency for men on fixed-term contracts does not seem to be

V.2 Wage profiles

OLS and FE estimates

The raw data in Table 1 showed that the permanent-temporary wage gap was between 16% (fixed-term contract) and 78% (seasonal-casual employment) for men. For women, we detected a 46% wage penalty in the case of seasonal-casual workers and a 13% wage premium for contract workers. But perhaps part of these differences is driven by differences in endowments of human capital or by differences in work motivation and other unobserved individual components. For this reason, we estimate ordinary least squares (OLS) and fixed-effects (FE) wage regressions to measure the effects of being in a seasonal-casual job (*SCJ*) and in a fixed-term contract (*FTC*) on the natural logarithm of real (1997 prices) hourly wages for men and women separately, after controlling for a large set of individual- and job-specific characteristics.[23] Table 6 reports the results from two specifications. In specification [2], we add the interactions of *SCJ* and *FTC* with full-time and part-time experience to the variables used in specification [1]. We also tried other specifications that included interactions with educational groups and job tenure, but since such interactions were jointly insignificant, they are not reported.

The OLS estimates from specification [1] show that the permanent-temporary wage gap is now 16%-17% for men and 13%-14% for women; that is, controlling for a host of observable characteristics has reduced the gap for men and women in seasonal-casual jobs, and increased the gap for women on fixed-term contracts, relative to the raw wage data. The FE estimates in specification [1] show smaller (but always precisely determined) wage gaps of 11% for men in seasonal-casual jobs and women on fixed-term contracts and 7% for men on fixed-term contracts and women in seasonal-casual jobs.[24]

significantly affected by any of the effort measures. But for all the other temporary workers, effort does matter. An increase in the number of overtime hours always leads to a higher hazard of exit (in both specifications), while an increase in the number of paid overtime hours reduces the rate of exit into a permanent job.

[23] The variables included in estimation are listed in the note to Table 6. The wage equations for women are selectivity corrected to account for non-participation. The participation equation used for this correction is performed on 2,844 women with 17,947 person-wave observations. The identifying variables are also listed in the note to Table 6. The estimated coefficient of the selection term is always negative, marginally significant in the OLS regressions, and not significant in the FE regressions. The results are unaffected if the selection term is obtained from a random-effect probit model.

[24] In a previous version of the paper, we also estimated random-effects wage equations. Interestingly, those estimates always lie between the OLS and FE estimates reported in Table 6 for both men and women. They can be found in Booth, Francesconi and Frank (2000).

The estimates obtained from specification [2] reveal that the effects of *SCJ* and *FTC* on wages differ systematically with the level of part-time and full-time experience. According to the OLS estimates, men with only one year of full-time experience (and no part-time work experience) face *ceteris paribus* a 17% wage penalty if they are in seasonal-casual jobs and a 39% wage penalty if they are on fixed-term contracts. The corresponding figures according to the FE estimates are 21% and 22% respectively. For men with 10 years of full-time work experience, the wage penalty of being in seasonal-casual jobs is stationary at about 17%, but the penalty of being on fixed-term contracts is now reduced to 15% and, using the FE estimates, these two figures are even lower at 7% and 2% respectively. Similar evidence is found for women. Considering only the FE estimates, the wage penalty for women with one year of full-time experience is 17% if they are in seasonal-casual jobs and 34% if they are on fixed-term contracts. These figures are 2% and 12% respectively for women with 10 years of full-time experience. Thus, less-experienced workers are expected to face larger wage penalties than more-experienced ones. Given this high level of heterogeneity, our empirical analysis proceeds by explicitly accounting for the interactions between temporary employment and work experience.

IV/GLS estimates

We now examine wage dynamics to see if there are any monetary longer-term effects of having held temporary jobs. The general specification of the wage equation that we separately estimate for men and women follows the approach used by Hausman and Taylor (1981), Altonji and Shakotko (1987) and Light and McGarry (1998), and can be written as

$$(1) \qquad \ln \omega_{ijt} = \beta_0 + \beta_1 \boldsymbol{X}_{ijt} + \beta_2 \boldsymbol{Z}_{ijt} + \mu_i + \phi_{ij} + \varepsilon_{ijt},$$

where ln ω_{ijt} is the real (1997 prices) hourly wage for individual *i* on job *j* at time *t*, and $\boldsymbol{X}$ denotes a standard set of variables that are often included in reduced-form wage regressions (e.g., highest educational qualification, part-time and full-time work experience, job tenure, union coverage, industry and occupation).[25] The vector $\boldsymbol{X}$ also contains dummy variables indicating the workers' region of residence, marital status and disability status, the sector and size of their employing organisation, whether they have received performance-related pay and

[25] For individuals who have more than one job between one interview date and the next, we assign to that individual the hourly wage for the interview date.

on-the-job training in the last 12 months, job mobility variables (indicating whether they have changed job because of promotion, quit or layoff), and the average local unemployment rate. The vector ***Z*** includes the contract-related variables that are the focus of our study. Specifically, ***Z*** contains controls for the number of seasonal-casual jobs and the number of fixed-term contracts held over the seven years of the survey, *NSCJ7* and *NFTC7*, respectively.[26] We also include interactions between *NSCJ7* and *NFTC7* and the linear and quadratic full-time experience terms. This allows the returns to 'experience capital' to differ by contract type. We exclude from our reported specification the interactions between contract types and other human capital variables (part-time experience and job tenure), because they had no additional explanatory power and did not alter the estimates of the other variables. The error term in equation (1) contains a time-invariant individual-specific component, μ_i, a time-invariant job-specific component, ϕ_{ij}, and a white noise, ε_{ijt}. We assume that the three error components are distributed independently from each other, have zero means and finite variances.

The estimation of (1) is performed using the instrumental-variables generalised least-squares (IV/GLS) procedure used by Light and McGarry (1998). We use an IV procedure because a number of wage regressors – including work experience, job tenure and, most notably, those related to the contract type – are likely to be correlated with individual- and job-specific characteristics, which cannot be observed by the analyst and are captured by μ_i and ϕ_{ij}.[27] We treat as endogenous all the regressors in ***Z***, along with part-time employment status, part-time experience and job tenure (and their squared terms), marital status, the job-mobility variables, and the dummy variables indicating training and performance-related pay.[28] The instrumental variables used in estimation are given by: a) the deviations from within-job means of both exogenous and endogenous time-varying variables, and b) the within-job means of all exogenous variables. Because ε_{ijt} is a white noise, the deviations are uncorrelated with the composite error term by construction, and thus they are valid

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[26] For men, the conditional mean (SD) for *NSCJ7* is 1.597 (0.920) while for women it is 1.632 (1.023). For *NFTC7*, the conditional mean (SD) for men is 1.591 (1.014) while for women it is 1.710 (1.198).
[27] See Light and McGarry (1998) for a discussion of the advantages of using a random-effects GLS procedure over a fixed-effects (within-individual/within-job) procedure.
[28] We have performed several sensitivity tests in which other variables in ***X*** were treated as endogenous (namely, part-time experience, job tenure, education, union coverage, disability status, occupation and sector). Adding these variables to the list of endogenous variables did not improve the statistical fit and did not have a statistically significant effect on *NSCJ7*, *NFTC7* and their interactions with full-time work experience.

instruments. As instruments, we also use the number of children (in five age groups) that each worker has during the seven-year period and the local unemployment rate.[29]

Table 7 reports the IV/GLS wage estimates of the contract-related variables (columns [1] and [2]) and their interactions with full-time experience (column [2] only) for men and women separately. The column [1] estimates imply that men and women who had one seasonal-casual job between 1991 and 1997 experience, respectively, a wage reduction of 8.9% and 6% as compared to those who had always a permanent job over the same period. The wage penalty associated with the experience of one fixed-term contract is half that for a seasonal-casual job, at 4.6% but is significant for men, while it is insignificant and around 2.4% for women. The fraction of the residual variance that is attributable to job-specific unobservables is quite large (particularly for women, for whom Var(ϕ_{ij}) is about 44% of the total variance).This may help reconcile the differences between the raw data presented in Table 1 and the results presented in Table 6. In column [2] we control for the interactions of temporary work with full-time experience. For both men and women, we note that the direct experience effects are always strongly significant but smaller than in the previous specification. In the case of workers with one year of full-time experience, the implied penalty to one seasonal-casual job over the first seven years of the career is, ceteris paribus, 11.5% and 4.5% for men and women, respectively. In the case of workers with ten years of full-time experience, the penalty *increases* respectively to 12.3% and 8.8%, *ceteris paribus*. Turning to workers on fixed-term contracts, the wage penalty to one fixed-term contract is about 8.5% and 4.7% for men and women with one year of full-time experience, respectively. The penalty *decreases* to 5% and 0.4% in the case of male and female workers, respectively, with ten years of full-time experience. The returns to experience capital differ strongly by contract type and by gender. Experience magnifies the differences between seasonal-casual workers and those who always have been in permanent jobs, while it reduces the differences between fixed-term workers and permanent workers. Both these effects are larger for women.[30]

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[29] The overidentifying-restrictions tests cannot reject the hypothesis that these two additional sets of variables are valid instruments at any conventional level of significance, and they improve the R^2 in the first-stage regressions. But the estimated parameters for the variables of primary interest (*NSCJ7*, *NFTC7* and their interactions with full-time experience) are not substantially altered when the additional instrumental variables are left out of the analysis.

[30] As a robustness check, we estimated two additional specifications for men and women. In the first specification, we introduced two dummy variables indicating current employment in a seasonal-casual job or current employment in a job with fixed-term contract. The IV/GLS wage estimates are similar to those obtained from standard random-effects regressions. In the second specification, we tested for the presence of non-linear

To describe the effect of contract type on wages further, we compute predicted log-wages paths from the column [2] estimates of Table 7 for workers with four different employment patterns. The first pattern involves workers who are always in a full-time permanent job for the first ten years of their career. The second and third patterns are for workers who hold one-fixed term contract or one seasonal-casual job respectively in the first period (at the start of their career) and are in a permanent job for the remaining part of their career. The fourth pattern involves workers who hold three consecutive one-year fixed-term contracts in the first three years of their career and are employed on a permanent contract thereafter. The predicted wages are computed under the assumptions that the individuals work continuously full-time for the first ten years of their career, are not disabled, are unmarried and childless, live in Greater London, work in the private sector in a non-union job and begin their career in 1991.[31]

The results of this simulation are reported in Table 8 and graphed in Figure 1. Having always had a permanent job is clearly the pattern that delivers the highest real wage profile over the first ten years of a man's career, with an average growth of 3% per year. Male workers who have one or three fixed-term contracts at the beginning of their career display lower wage profiles (especially at the beginning of their work cycle) but a slightly higher wage growth. In fact, the wage gap among these three types of workers is larger at the start of the career and tapers off over time as they accumulate general work experience. But men who started off with a seasonal-casual job have the lowest wage profile and the smallest wage growth. This leads to an increase in the wage gap in comparison with the other three types of workers, which is particularly clear when we contrast pattern 3 to patterns 2 or 4. This finding holds for women too. In the case of women, however, having had one or three fixed-term contracts at the start of the career does not permanently damage the wage profile. Indeed, women following pattern 2 or pattern 4 end up with the highest wage levels and the largest wage growth (approximately 2.5% per year over a ten-year period). The wage gap for these

effects in *NSCJ7* and *NFTC7* on (ln) hourly wages. We introduced two dummy variables, the first taking the value of one if the worker held only one seasonal-casual job or fixed-term contract over the panel years; the second taking the value one if the worker held two or more seasonal-casual jobs or fixed-term contracts over the panel years. For both men and women, we found no evidence of a wage penalty beyond the first fixed-term contract. We detected, however, a worsening of the wage penalty as the number of seasonal-casual jobs increases, especially for women.

two types of workers and those who have always been in a permanent job is very large at the start of the career, but it declines over time. It is the interaction between full-time experience and fixed-term contracts that cause type 4 (and type 2) workers to overtake type 1 workers, for their productivity increases as they move to permanent jobs as a return to this ‘experience capital’. While for these women there is no wage penalty after 10 years, their total returns from employment are lower (compare the areas beneath the curves).

VI. CONCLUSIONS

In Britain about 7% of male employees and 10% of female employees are in temporary jobs. In contrast to much of continental Europe, this proportion has been relatively stable over the 1990s. Using data from the British Household Panel Survey – which disaggregates temporary work into seasonal or casual jobs and fixed-term contract jobs – we found that, on average, temporary workers report lower levels of job satisfaction and receive less work-related training than their counterparts in permanent employment. This holds for both seasonal-casual workers and workers on fixed-term contracts.

However, we also found evidence that temporary jobs are a stepping stone to permanent work. The median time in temporary work before such a transition is between 18 months and three and a half years, depending on contract type (seasonal or fixed term) and gender. Our wage growth models (which allow for potential endogeneity of many of the explanatory variables including contract type) show that the wage growth penalty associated with experience of seasonal/casual jobs is quite high for both men and women. Even with ten years of full-time experience, having held one seasonal/casual job has a wage penalty of 12.3% for men and 8.8% for women. In contrast, men with experience of one fixed-term contract suffer a much lower wage penalty, 5%, after ten years of experience. Interestingly, we find evidence that women who start off their career on fixed-term contracts may experience a high wage growth, and, within a period of 7-10 years, have fully caught up with their permanent counterparts.

Overall, our results show the importance of distinguishing between types of temporary work. Seasonal and casual jobs are unlikely to be probationary in nature. Because of the low

[31] We also assume that each individual’s occupation, industry, education, firm size, training, performance-related pay, job mobility patterns and local unemployment rate take the sample values for men and women respectively. Changing these assumptions would only alter the levels but not the relative rankings (and slopes) of the wage

human capital held by workers in these jobs, wages will be low, there will be little job satisfaction, and poor future prospects. We find evidence for this in our study. In contrast, fixed-term temporary jobs may well be stepping-stones to a future career. Although men who begin in jobs with fixed-term contracts suffer a permanent earnings loss, it is not as great as for men with experience of seasonal-casual work. This is consistent with two possibilities. The first is that these men may be less able than those who immediately acquire a permanent job on entering the workforce. The second is that they may also lose out from never quite catching up on the human capital investment foregone during the period of temporary work. In contrast, women who start with fixed-term contracts fully catch up with those who began on permanent contracts. This is consistent with a view that some women, upon entering the labour force, may take longer to decide on their career choices. Under this hypothesis, women who begin in temporary work are as able as those who begin in permanent jobs, and these women eventually make up for the lack of human capital acquisition during the period of temporary work.

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profiles in Table 8 and Figure 1.

APPENDIX A
The British Household Panel Survey and the question on "temporary" work

The first wave of the BHPS, collected in Autumn 1991, was designed as a nationally representative sample of the population of Great Britain living in private households in 1991. The achieved wave 1 sample covers 5,500 households and corresponds to a response rate of about 74% of the effective sample size. At wave 1, about 92% of eligible adults, that is just over 10,000 individuals, provided full interviews. The same individuals are re-interviewed each successive year, and if they split off from their original households to form new households, all adult members (that is, aged 16 or more) of these households are also interviewed. Similarly, children in original households are interviewed when they reach the age of 16. Thus, the sample remains broadly representative of the population of Britain as it changes through the 1990s. Of those interviewed in the first wave, 88% were successfully re-interviewed at wave 2 (Autumn 1992), and subsequent wave-on-wave response rates have consistently been around 95-98% (Taylor et al., 1998).

The core questionnaire elicits information about income, labour market behaviour, housing conditions, household composition, education and health at each annual interview. Information on changes (e.g., employment, household membership, receipt of each income source) which have occurred within the households in the period between interviews is also collected.

The second wave (1992) obtained retrospective information on complete fertility, marital, cohabitation and employment histories for all adult panel members in that year. The third wave (1993) collected detailed job history information. Both these retrospective data have been used to construct some of the variables used in this analysis (e.g, cohort of first partnership, number of partnerships, number of year of part-time and full-time work experience, and number of years of job tenure).

The information on temporary work is obtained from the Mainstage Individual Questionnaire included in all the waves (1-7) used in the analysis. At the beginning of the "Employment" section, individuals are asked whether they do any paid work. If they do, then they are immediately asked:

E4. Is your current job
  A permanent job
  A seasonal, temporary or casual job
  Or a job done under contract or for a fixed period of time?

Further information on the questionnaire as well as on the sampling scheme, weighting, imputation and other survey methods used in the BHPS can be obtained at http://www.iser.essex.ac.uk/bhps/doc/index.htm.

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Table 1: Distribution of temporary work and mean hourly wages by type of contract and gender

	Men		**Women**	
	Unweighted	Weighted	Unweighted	Weighted
Temporary contract (%)				
Seasonal-casual	3.9	3.8	6.3	6.1
Fixed-term	2.9	2.9	3.3	3.1
N	11,186	11,167	12,821	12,830
Hourly wages (£)				
Permanent [p]	8.55	8.59	6.29	6.32
Seasonal-casual [s]	4.79	4.64	4.92	4.88
Fixed-term contract [f]	7.38	7.47	7.19	7.22
Wage differences (£)				
[p] – [s]	3.76***	3.95***	1.37***	1.44***
	(11.755)	(12.023)	(8.850)	(8.867)
[p] – [f]	1.17***	1.12***	-0.90***	-0.90***
	(3.154)	(2.942)	(4.306)	(4.020)
[s] – [f]	-2.59***	-2.83***	-2.27***	-2.34***
	(4.823)	(5.592)	(7.761)	(7.768)

Source: British Household Panel Survey 1991-1997.
Notes: Weighted figures are obtained using the BHPS cross-sectional enumerated individual weights. *N* is number of person-wave observations. Wages are in constant (1997) pounds. Absolute value of the *t*-test of the wage difference is in parentheses.
*** indicates that the wage difference is significant at 0.01 level.

Table 2: Relative risk ratios of being in a temporary job by gender
(Absolute ratio of coefficent to standard error in parentheses)

	Men			Women		
Selected variables	Seasonal & casual	Fixed-term contract	Mean	Seasonal & casual	Fixed-term contract	Mean
Age dummy:						
16-24	0.827	2.062	0.250	2.031***	1.684	0.234
	(0.390)	(1.531)		(3.021)	(1.560)	
25-34	0.357**	1.057	0.326	1.007	1.014	0.301
	(2.394)	(0.143)		(0.039)	(0.054)	
45-60	2.338**	2.778***	0.165	0.806	1.942**	0.198
	(2.215)	(2.915)		(0.809)	(2.178)	
Nr. of children aged:						
0-4	0.689*	1.014	0.186	0.911	0.984	0.131
	(1.918)	(0.075)		(1.391)	(0.087)	
5-18	1.261***	0.998	0.524	1.201***	1.391***	0.564
	(2.570)	(0.023)		(2.956)	(3.709)	
Education:						
Less than GCSE or O level	0.607	1.310	0.084	0.841	3.065***	0.108
	(1.326)	(0.776)		(0.782)	(2.651)	
GCSE/O level	1.180	1.028	0.210	1.006	2.254**	0.282
	(0.554)	(0.088)		(0.030)	(2.140)	
A level	2.030**	1.195	0.162	1.496*	2.728**	0.118
	(2.263)	(0.540)		(1.946)	(2.356)	
Vocational degree	1.393	0.791	0.258	1.190	1.932	0.212
	(1.034)	(0.670)		(0.862)	(1.637)	
University degree or more	2.777***	1.169	0.145	2.332***	3.036***	0.112
	(2.631)	(0.368)		(3.172)	(2.563)	
Occupation:						
Managerial	0.013***	0.619	0.165	0.181***	1.707	0.084
	(4.101)	(1.275)		(4.025)	(1.106)	
Professional	0.145***	1.750	0.089	0.500*	7.305***	0.046
	(4.264)	(1.388)		(1.685)	(4.919)	
Technicians	0.236***	1.818	0.097	0.596	2.409**	0.065
	(3.682)	(1.585)		(1.607)	(1.987)	
Teachers	0.551	0.983	0.020	1.257	3.839***	0.058
	(1.091)	(0.031)		(0.626)	(3.063)	
Nurses	1.260	0.321	0.005	1.012	1.855	0.043
	(0.198)	(0.930)		(0.030)	(1.364)	
Clerks and secretaries	0.715	1.149	0.096	0.991	1.777	0.293
	(1.046)	(0.446)		(0.046)	(1.535)	
Craft	0.418***	1.087	0.184	1.962*	0.896	0.026
	(2.601)	(0.275)		(1.934)	(0.132)	
Protection and personal services	1.017	0.625	0.067	1.239	1.694	0.143
	(0.062)	(1.274)		(1.144)	(1.475)	
Sales	0.448***	0.849	0.053	0.570***	1.498	0.106
	(2.951)	(0.471)		(2.943)	(0.776)	

Plant & machine operatives	1.071 (0.224)	0.851 (0.453)	0.147	4.385*** (5.068)	0.713 (0.462)	0.039
Sector:						
Civil service	0.649 (0.925)	0.463 (1.546)	0.049	0.634 (1.091)	0.982 (0.045)	0.039
Local government	0.550 (1.585)	0.993 (0.021)	0.104	1.472** (1.998)	2.562*** (3.710)	0.189
Other public	0.957 (0.120)	1.437 (1.016)	0.059	0.736 (1.099)	2.164*** (2.802)	0.109
Non-profit	1.042 (0.104)	1.792 (1.349)	0.024	1.010 (0.040)	2.288*** (2.765)	0.048
Full-time exp. at start of panel	0.925*** (3.440)	0.951** (2.377)	14.586	0.988 (0.779)	0.947** (2.321)	8.687
Part-time exp. at start of panel	0.936 (0.500)	0.948 (0.251)	0.190	0.926*** (3.004)	0.947* (1.831)	3.829
Extra full-time experience	0.686*** (7.068)	0.772*** (5.416)	6.795	0.685*** (7.035)	0.737*** (5.759)	3.491
Extra part-time experience	0.669** (2.248)	1.250 (1.355)	0.046	0.694*** (8.480)	0.902* (1.769)	2.908
In part-time job	8.259*** (8.537)	2.022*** (2.614)	0.051	4.096*** (9.422)	1.305 (1.361)	0.410
Total number of layoffs	1.494*** (5.028)	1.303*** (3.233)	0.664	1.284*** (3.260)	1.187 (1.387)	0.372
Local unemployment to vacancies ratio	0.993 (1.182)	0.990 (1.359)	15.691	0.982*** (3.210)	0.986* (1.922)	15.538
Log likelihood	-2,442			-3,920		
Model χ^2	976.2 [0.0000]			1092.4 [0.0000]		
N	11,186			12,821		

Note: The relative risk ratios are obtained from exponentiated coefficients of multinomial logit regressions. Base category is "permanent job". The specification for all equations also includes cohort of entry into the labour market (5 dummies), disabled, region of residence (6), industry (9), firm size (7), and number of full-time and part-time jobs ever held at the start of the panel. Model χ^2 is the Wald statistic for the goodness-of-fit test and is equal to $-2[L_R-L_U]$ where L_R is the constant-only log-likelihood value and L_U is the log-likelihood value reported in the table. Its corresponding *p*-value is in square brackets. The χ^2 statistic has 106 degrees of freedom. The *t*-ratios are computed using standard errors that are robust to arbitrary forms of correlation within individuals. *N* is the number of person-wave observations.

* significant at 0.10 level, ** significant at 0.05 level, *** significant at 0.01 level.

Table 3: Job satisfaction and training of temporary workers

	Men (*N*=11,186)		**Women (*N*=12,821)**	
Panel A	**Job satisfaction measure**[a]			
	Seasonal & casual	Fixed-term contract	Seasonal & casual	Fixed-term contract
Overall	-0.146**	-0.042	-0.165***	-0.052
	(2.173)	(0.597)	(3.336)	(0.872)
Promotion prospects	-0.359***	-0.251***	-0.188***	-0.144**
	(6.437)	(3.923)	(4.516)	(2.240)
Total pay	0.107	-0.172**	0.082*	-0.065
	(1.636)	(2.029)	(1.718)	(0.971)
Relation with the boss	0.114*	0.142**	0.064	0.133**
	(1.732)	(2.010)	(1.396)	(2.065)
Security	-0.714***	-0.729***	-0.695***	-0.774***
	(9.449)	(9.012)	(13.147)	(10.969)
Initiative	-0.410***	-0.118*	-0.256***	-0.102
	(6.251)	(1.662)	(5.316)	(1.620)
Work itself	-0.189**	0.053	-0.174***	0.032
	(2.783)	(0.779)	(3.700)	(0.476)
Hours worked	0.023	0.053	-0.062	-0.011
	(0.351)	(0.687)	(1.282)	(0.159)

Panel B	**Training receipt**[b]			
	Pooled probit	RE probit	Pooled probit	RE probit
Seasonal & casual	-0.198***	-0.195***	-0.146***	-0.139***
	(6.509)	(6.671)	(7.015)	(7.276)
Fixed-term contract	-0.122***	-0.095***	-0.070***	-0.062***
	(4.010)	(3.658)	(2.588)	(2.964)
ρ		0.331***		0.298***
		[0.0000]		[0.0000]
Log likelihood	-6,351	-6,003	-6,732	-6,445
Model χ^2	995.7	907.5	1397.2	1244.0
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Mean of dependent vb.	0.360		0.314	
	Training intensity[b]			
	Pooled tobit	RE tobit	Pooled tobit	RE tobit
Seasonal & casual	-12.435***	-12.212***	-9.384***	-9.245***
	(5.485)	(5.459)	(6.787)	(6.725)
Fixed-term contract	-1.569	-1.262	-0.186	-0.116
	(0.846)	(0.662)	(0.135)	(0.084)
ρ		0.020***		0.012**
		[0.0027]		[0.0214]

Log likelihood	-20,926	-20,875	-21,504	-21,478
Model χ^2	1256.2	988.3	1977.4	1515.0
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Mean of dependent vb.	3.839		2.883	
	11.262[§]		9.188[§]	

[a] Panel A: Job satisfaction. Coefficients are obtained from ordered probit regressions. For each row the dependent variable is "job satisfaction" measured on a scale from 1 to 7, where a value of 1 corresponds to "not satisfied at all" and a value of 7 corresponds to "completely satisfied". The reported numbers are the coefficients (and absolute *t*-ratios from robust standard errors) on the two types of temporary work. Other variables included in each regression are all the variables used in Table 2 plus marital status (2 dummies), age-marital status interactions (2), number of (marital or cohabiting) partnerships, and cohort of partnership (3).

[b] Panel B: Training receipt and intensity. The reported numbers are marginal effects for the two type of temporary work obtained from pooled and random-effects probit regressions (top of Panel B) and from pooled and random-effects tobit regressions (bottom of Panel B). Absolute *t*-ratios (obtained from robust standard errors in the pooled probit regressions and pooled tobit regressions) are in parentheses. Other variables included in each regression are all the variables used in Panel A above plus union coverage and marital status (2 dummies). The term ρ is the fraction of total variance contributed by the panel-level variance component. The *p*-value of the likelihood ratio test of $\rho=0$ is reported in square brackets. For the definition of Model χ^2 see footnote of Table 2. The χ^2 statistic has 64 degrees of freedom and its *p*-value is in square brackets.

[§] Computed on positive values only (*N*=3,812 for men; *N*=4,023 for women)

* significant at 0.10 level, ** significant at 0.05 level, *** significant at 0.01 level. *N* is the number of person-wave observations.

Table 4: Kaplan-Meier estimates of job exit rates by type of contract and gender (cumulative percentage)

Gender and Type of contract	Job tenure (months) 3	6	12	18	24	30	36	42	48	60	90	120	N [%]
Men													
Permanent	4	12	20	29	35	42	46	52	56	64	78	85	10,427
Seasonal & casual	53	68	73	91	92	97							431
Continuing in same firm	42	56	61	79	81	89							[70.3]
Moving to another firm	12	21	23	49	53	60							[24.8]
Ending in permanency	16	27	30	48	55	63							[28.3]
Fixed-term contract	22	39	49	70	74	80	84	88	89	92			328
Continuing in same firm	17	29	37	57	61	69	72	77	79	85			[72.3]
Moving to another firm	5	14	18	29	31	35	41	44	45	48			[26.8]
Ending in permanency	7	16	21	38	39	46	51	57	58	60			[38.1]
Women													
Permanent	6	14	23	35	41	49	54	61	65	73	86	92	11,593
Seasonal & casual	40	53	61	81	83	89	90	93	94	95			805
Continuing in same firm	29	40	48	69	72	79	80	84	85	87			[70.9]
Moving to another firm	9	16	19	32	34	45	46	51	52	58			[24.1]
Ending in permanency	13	21	24	43	47	55	57	60	61	66			[34.2]
Fixed-term contract	24	40	48	66	71	79	81	87	88	93			423
Continuing in same firm	16	29	37	54	60	69	71	78	81	88			[75.7]
Moving to another firm	7	14	17	25	26	31	32	35	36	38			[22.5]
Ending in permanency	9	16	21	34	37	45	46	51	55	60			[36.2]

Note: Figures in cells with less than 30 observations are not reported. The percentage of censored observations is in square brackets. *N* is the number of person-wave observations.

Table 5: Exit from temporary work to permanent work: estimates from proportional hazard model – Non-parametric baseline hazard specification

	Men				Women			
	Exit from seasonal and casual work to permanent work		Exit from a fixed-term contract to permanent work		Exit from seasonal and casual work to permanent work		Exit from a fixed-term contract to permanent work	
Variable	[1]	[2]	[1]	[2]	[1]	[2]	[1]	[2]
Age dummy:								
16-24	0.144	0.788	1.205***	1.180***	-0.247	-0.832**	0.301	-1.069***
	(0.321)	(1.400)	(2.842)	(2.960)	(1.116)	(2.479)	(1.014)	(3.289)
25-34	0.317	0.673	1.094***	1.083***	0.029	-0.339	0.393	0.286
	(0.643)	(0.956)	(2.856)	(2.780)	(0.142)	(0.759)	(1.331)	(0.880)
45-60	-0.706	-0.514	0.581	0.562	-0.661**	-0.581	-0.127	0.938
	(1.252)	(0.554)	(1.326)	(1.302)	(2.259)	(1.530)	(0.342)	(1.641)
Education:								
Less than GCSE	0.167	-0.168	-0.935	-0.912	0.106	0.394	1.513***	-0.017
or O level	(0.331)	(1.370)	(1.644)	(1.569)	(0.369)	(0.369)	(3.061)	(0.436)
GCSE/O level	-0.479	-0.746**	-0.254	-0.238	-0.027	-0.531	1.342***	0.573
	(1.169)	(1.985)	(0.532)	(0.482)	(0.111)	(0.606)	(2.962)	(1.478)
A level	-0.342	-0.664	-0.038	-0.027	-0.106	-0.433	1.191**	0.447
	(0.799)	(1.521)	(0.075)	(0.051)	(0.379)	(0.456)	(2.425)	(1.095)
Vocational	-0.099	-0.634	0.545	0.539	-0.126	-0.141	1.739***	0.454
degree	(0.222)	(1.037)	(1.095)	(1.094)	(0.377)	(1.197)	(3.133)	(1.434)
University	0.259	-0.781	0.842	0.893	0.290	1.376	1.521**	0.739***
degree or more	(0.422)	(0.617)	(1.426)	(1.547)	(0.830)	(1.030)	(2.439)	(2.795)
Occupation:								

Managerial	7.371***	6.849**	0.248	0.319	1.085**	0.134	-0.379	0.840
	(7.201)	(2.037)	(0.333)	(0.616)	(2.132)	(0.016)	(0.568)	(0.240)
Professional	2.940***	0.836	-0.607	-0.568	-0.633	1.074	-1.305***	-0.715
	(2.873)	(0.703)	(0.913)	(1.065)	(1.154)	(0.869)	(2.598)	(0.569)
Technicians	2.949***	2.523***	-0.428	-0.410	-0.095	-1.196	0.125	0.312
	(5.713)	(2.566)	(0.827)	(0.854)	(0.245)	(1.244)	(0.265)	(0.667)
Clerks and secretaries	0.933**	-0.039	0.072	0.079	0.308	-0.817	0.356	0.791
	(2.204)	(0.031)	(0.141)	(0.161)	(1.162)	(1.144)	(0.883)	(0.215)
Craft	2.621***	2.308***	2.291***	2.278***	1.091**	-1.189	-0.351	-0.531
	(6.667)	(3.347)	(4.945)	(5.034)	(2.383)	(0.820)	(0.848)	(0.545)
Protection and personal services	-0.682	-1.530**	0.074	0.085	0.297	-1.373**	0.039	0.217
	(1.032)	(2.315)	(0.128)	(0.154)	(1.116)	(2.015)	(0.297)	(0.329)
Sales	1.032**	1.457	1.231**	1.243***	1.058***	-0.364	-0.041	0.325
	(2.299)	(1.646)	(2.401)	(2.576)	(3.809)	(0.498)	(0.075)	(0.747)
Plant & machine operatives	-0.006	0.306	0.990**	0.984**	-1.413***	-1.795***	-1.071	-0.906***
	(0.014)	(0.262)	(2.175)	(2.182)	(3.411)	(5.071)	(0.967)	(3.893)
Sector:								
Civil service	0.314	0.657	-0.433	-0.455	-0.641	-0.817	-1.581**	-1.678**
	(0.471)	(0.597)	(0.528)	(0.571)	(0.986)	(0.570)	(2.047)	(2.059)
Local government	0.029	-0.294	-0.002	0.001	-0.340	-2.444***	-1.247***	-1.245***
	(0.054)	(0.519)	(0.006)	(0.001)	(1.378)	(3.545)	(4.094)	(3.313)
Other public	-1.224	0.176	0.142	0.135	-0.854**	-1.321	-1.059***	-1.478***
	(1.506)	(0.089)	(0.355)	(0.331)	(2.510)	(1.607)	(3.005)	(3.301)
Non-profit	2.050***	1.819*	-1.495*	-1.481*	-0.297	-3.556***	-1.445***	-1.415***
	(3.547)	(1.771)	(1.904)	(1.887)	(0.710)	(3.356)	(3.470)	(4.539)
In part-time job	-1.681***	-1.473***	-1.054***	-1.062***	-0.917***	-1.658***	-0.244	-0.039
	(4.549)	(3.695)	(2.586)	(2.658)	(5.212)	(3.342)	(0.915)	(0.038)

Union coverage	-0.155	-1.006*	0.341	0.327	0.817***	1.267***	0.436*	-0.569*
	(0.592)	(1.677)	(1.434)	(1.435)	(4.939)	(4.215)	(1.918)	(1.907)
Total number of layoffs	0.137	0.661**	0.158*	0.160*	0.304	0.650**	0.442***	0.458***
	(1.576)	(2.191)	(1.729)	(1.719)	(3.754)	(2.044)	(3.478)	(3.279)
Local unemployment to vacancies ratio	-0.023**	-0.093**	-0.014	-0.014	-0.011	-0.093***	-0.005	-0.069***
	(2.050)	(2.481)	(1.361)	(1.438)	(1.639)	(4.286)	(0.533)	(5.678)
Hours of unpaid overtime work[a]	0.522**	0.102	0.013	0.016	0.079	0.315***	0.234**	0.277**
	(2.493)	(0.202)	(0.121)	(0.185)	(0.987)	(2.672)	(2.299)	(2.431)
σ^2		2.861***		1.14×10^{-4}		2.415***		3.778***
		(3.561)		(0.004)		(4.894)		(4.743)
Log likelihood	-382	-349	-406	-406	-989	-935	-549	-501
Model χ^2	395.6	412.1	348.6	352.7	471.2	494.0	230.3	236.5
	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]
Person-month obs.	5,602	5,602	4,591	4,591	12,016	12,016	6,716	6,716

Note: Absolute ratio of coefficient to standard error in parentheses. The term σ^2 is the variance of the Gamma-distributed random variable that summarises unobserved heterogeneity between individuals. All regressions also include industry (3 dummies), firm size (7), and a constant. For the definition of Model χ^2 see footnote of Table 2. The χ^2 statistic has 35 degrees of freedom and its *p*-value is in square brackets.

[a] Predicted from tobit regressions which include all the variables used in the hazard models plus number of children by four age groups, and dummy variables for cohort of entry in the labour market (5 dummies), region of residence (6), and whether worker receives a performance-related pay. The tobit regressions contain nine rather than three industry dummies. The *F*-statistics (and *p*-values) of the variables identifying hours of unpaid overtime work are $F(11,11130)=9.36$ (*p*-value=0.000) and $F(11,12763)=14.03$ (*p*-value= 0.000) for men and women, respectively.

* significant at 0.10 level, ** significant at 0.05 level, *** significant at 0.01 level.

Table 6: Ordinary Least Squares (OLS) and Fixed-Effects (FE) wage estimates
(Absolute *t*-ratio in parentheses)

	Men				Women			
	OLS		FE		OLS		FE	
Variable	[1]	[2]	[1]	[2]	[1]	[2]	[1]	[2]
SCJ	-0.155***	-0.171***	-0.107***	-0.224***	-0.126***	-0.169***	-0.075***	-0.189***
	(4.414)	(2.986)	(5.004)	(6.511)	(5.062)	(4.111)	(4.513)	(6.439)
SCJ × full-time experience		-0.003		0.018***		0.003		0.023***
		(0.297)		(2.652)		(0.467)		(4.657)
SCJ × full-time experience squared		0.0002		-0.0003		-0.0001		-0.0006***
		(0.628)		(1.616)		(0.573)		(3.494)
SCJ × part-time experience		0.033		0.101**		0.010		-0.004
		(0.382)		(2.140)		(0.734)		(0.508)
SCJ × part-time experience squared		-0.008		-0.016**		-0.0001		0.0006
		(0.638)		(2.058)		(0.157)		(1.569)
FTC	-0.171***	-0.426***	-0.069***	-0.247***	-0.144***	-0.444***	-0.109***	-0.373***
	(3.956)	(5.443)	(3.110)	(6.008)	(3.899)	(5.468)	(5.010)	(7.771)
FTC × full-time experience		0.035***		0.030***		0.031*		0.034***
		(3.379)		(4.647)		(1.768)		(3.533)
FTC × full-time experience squared		-0.0007**		-0.0007***		-0.0009		-0.0009**
		(2.536)		(3.970)		(1.141)		(2.223)
FTC × part-time experience		0.099**		0.071***		0.063***		0.034***
		(2.456)		(2.605)		(3.248)		(3.214)
FTC × part-time experience squared		-0.006***		-0.006***		-0.002**		-0.001**
		(2.590)		(3.365)		(2.331)		(2.236)

Full-time experience	0.042***	0.040***	0.111***	0.109***	0.022***	0.020***	0.112***	0.109***
	(16.683)	(15.429)	(30.328)	(29.654)	(8.053)	(7.281)	(25.034)	(24.259)
Full-time experience squared	-0.0008***	-0.0007***	-0.001***	-0.001***	-0.0004***	-0.0004***	-0.001***	-0.001***
	(12.461)	(11.562)	(13.241)	(12.627)	(4.613)	(4.090)	(8.504)	(7.784)
Part-time experience	-0.015	-0.023	0.032***	0.029***	0.003	0.0008	0.066***	0.065***
	(1.125)	(1.521)	(5.933)	(4.537)	(1.117)	(0.297)	(12.258)	(11.982)
Part-time experience squared	0.0003	0.0009	-0.007***	-0.005**	-0.00004	0.00002	-0.0004	-0.0003
	(0.296)	(0.746)	(3.067)	(1.988)	(0.332)	(0.187)	(1.558)	(1.445)
R^2	0.544	0.546	0.216	0.214	0.539	0.542	0.170	0.169
N	11,186		11,186		12,821		12,821	
Nr. of individuals			1,740				1,981	

Note: *SCJ* and *FTC* denote the state of being in a seasonal-casual job and a fixed-term contract, respectively. Each specification also includes linear and quadratic terms in years of job tenure, local unemployment rate, and dummy variables for region of residence (6), educational level (5), industry (9), occupation (8), sector (4), firm size (7), disability status, part-time employment, marital status (2), whether worker has changed job because of promotion, quit or layoff, the number of previous jobs, whether worker has received on-the-job training in the last 12 months, whether worker is union covered and whether worker receives a performance-related pay. All wage equations for women are selectivity-corrected using a participation probit equation. This equation is identifies by age, time trend (6 dummy variables), cohort of entry in the labour market (5), age-marital status interactions (2), cohort of first partnership, number of partnerships, number of children by age group (5 age groups), housing tenure (2), and individual attitudes about working women (6). The *t*-ratios in the OLS regressions are obtained from standard errors that are robust to arbitrary forms of correlation within individuals. *N* is number of person-wave observations.

** significant at 0.05 level, *** significant at 0.01 level.

Table 7: Temporary work and wages: selected estimates from IV/GLS regressions

	Men		Women	
	[1]	[2]	[1]	[2]
NSCJ7	-0.116***	-0.147***	-0.073***	-0.050**
	(3.167)	(3.681)	(3.276)	(2.522)
NSCJ7 squared	0.027**	0.033***	0.013**	0.012**
	(2.212)	(2.737)	(2.409)	(3.509)
NSCJ7 $\times$ full-time experience		-0.001		-0.007***
		(0.298)		(2.786)
NSCJ7 $\times$ full-time experience squared		0.000		0.0002**
		(0.192)		(2.289)
NFTC7	-0.059**	-0.104***	-0.030	-0.058*
	(2.195)	(2.873)	(1.273)	(1.706)
NFTC7 squared	0.013**	0.014**	0.006	0.004
	(2.183)	(2.098)	(1.418)	(0.910)
NFTC7 $\times$ full-time experience		0.005		0.007
		(1.324)		(1.193)
NFTC7 $\times$ full-time experience squared		-0.0001		-0.0002
		(1.109)		(0.850)
Full-time experience	0.045***	0.037***	0.031***	0.024***
	(5.079)	(12.643)	(4.021)	(8.062)
Full-time experience squared	-0.001***	-0.0007***	-0.0006**	-0.0004***
	(4.259)	(10.166)	(2.367)	(4.775)
Var(μ_i)	0.086	0.085	0.076	0.074
Var(ϕ_{ij})	0.054	0.054	0.093	0.091
Var(ε_{ijt})	0.047	0.044	0.044	0.042
R^2	0.544	0.550	0.482	0.534
No. person-wave observations	11,186		12,821	
N	14,156		17,006	

Note: *NSCJ7* and *NFTC7* denote the number of seasonal-casual jobs and the number of fixed-term contracts held over the seven years of the panel survey. The terms Var(μ_i), Var(ϕ_{ij}), and Var(ε_{ijt}) are the estimated variances of the individual, job, and transitory components of the residual, respectively. The other variables used in estimation are listed in the footnote of Table 6. All wage equations for women are selectivity corrected. See note of Table 6 for details on the identifications of the selection term. *N* is number of person-job-wave observations. Absolute *t*-ratios are in parentheses.

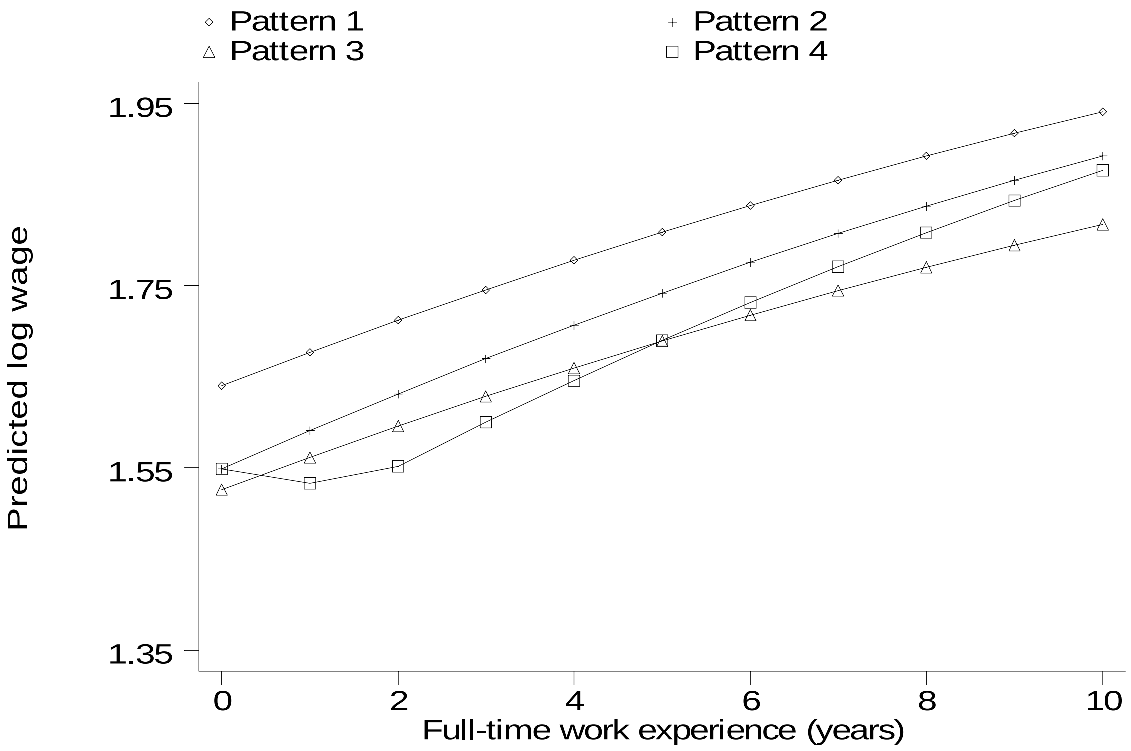
* significant at 0.10 level, ** significant at 0.05 level, *** significant at 0.01 level

Table 8: Predicted log wages by experience level and early employment patterns

	Years of full-time experience												
	0	1	2	3	4	5	6	7	8	9	10	7-0	10-0
Men													
Pattern 1	1.639	1.676	1.711	1.745	1.778	1.808	1.838	1.866	1.892	1.917	1.940	0.226	0.301
Pattern 2	1.549	1.591	1.631	1.669	1.707	1.742	1.775	1.807	1.837	1.865	1.892	0.258	0.343
Pattern 3	1.526	1.561	1.596	1.628	1.659	1.689	1.717	1.744	1.770	1.794	1.817	0.218	0.291
Pattern 4	1.549	1.533	1.552	1.600	1.646	1.690	1.731	1.771	1.808	1.843	1.877	0.222	0.327
Women													
Pattern 1	1.499	1.523	1.546	1.567	1.588	1.609	1.628	1.646	1.663	1.680	1.696	0.147	0.197
Pattern 2	1.445	1.476	1.505	1.533	1.559	1.585	1.609	1.631	1.653	1.673	1.692	0.186	0.247
Pattern 3	1.462	1.478	1.494	1.510	1.525	1.540	1.544	1.567	1.581	1.594	1.606	0.106	0.143
Pattern 4	1.445	1.436	1.447	1.487	1.525	1.561	1.594	1.626	1.655	1.682	1.707	0.181	0.262

Note: Predictions are based on the estimates in column [2] of Table 7 under the assumption that all jobs are full-time jobs. Pattern 1: worker is always employed in a permanent job. Pattern 2: worker holds one fixed-term contract in first period and is employed in permanent job thereafter. Pattern 3: worker holds one seasonal-casual job in first period and is employed in permanent job thereafter. Pattern 4: worker holds 3 fixed-term contract in first three periods and then is employed in permanent job.

Men



Women

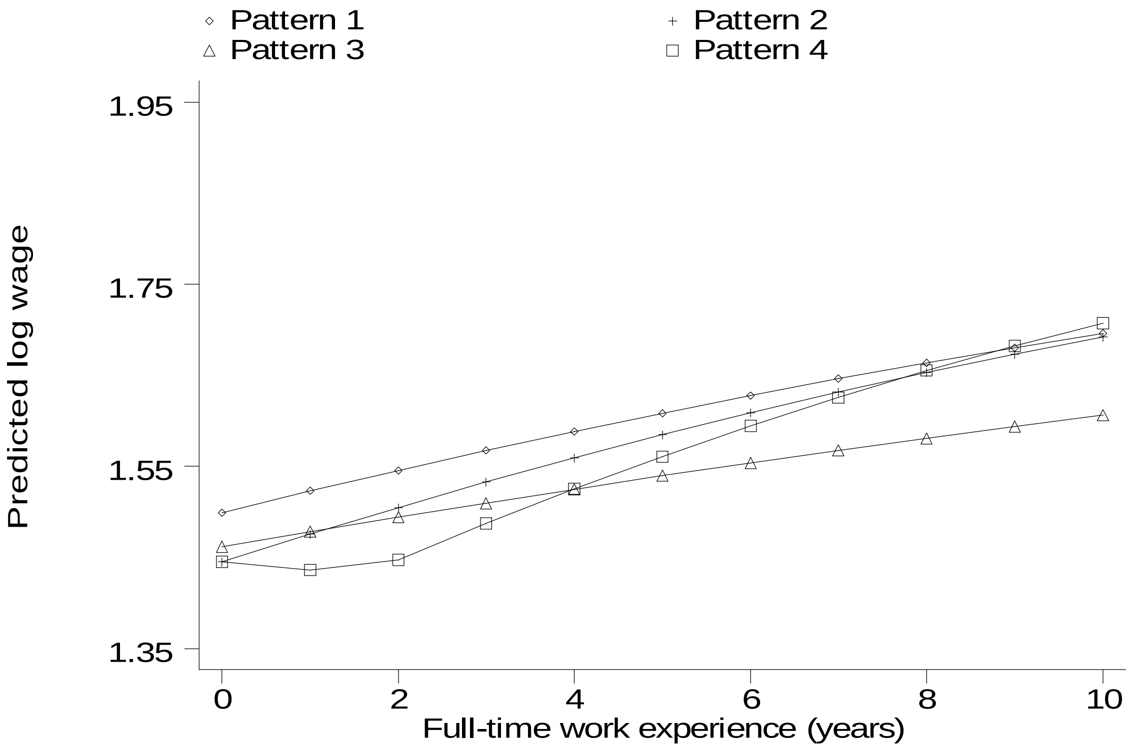


Figure 1: Predicted log wages by experience level and early employment patterns

Note: Based on predictions presented in Table 8. All jobs are full-time jobs. Pattern 1: worker is always employed in a permanent job. Pattern 2: worker holds one fixed-term contract in first period and is employed in permanent job thereafter. Pattern 3: worker holds one seasonal-casual job in first period and is employed in permanent job thereafter. Pattern 4: worker holds 3 fixed-term contract in first three periods and then is employed in permanent job.