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TEMPORARY JOBS: STEPPING STONES OR DEAD ENDS?*

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In Britain, about 7% of male employees and 10% of female employees are in temporary jobs. This proportion has been relatively stable over the 1990s. Using data from the British Household Panel Survey, we confirm the popular perception that temporary jobs are generally not desirable when compared to permanent employment. Temporary workers have lower levels of job satisfaction, receive less training and are less well-paid. There is some evidence that fixed-term contracts are a stepping stone to permanent work. Women who start in fixed-term employment and move to permanent jobs fully catch up to those who start in permanent jobs.

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 as shown in the introduction to this Symposium. 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 relatively low and fairly stable.

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, 1999; Arulampalam and Booth, 1998). Purcell *et al.* (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.

The purpose of this paper is to examine whether temporary jobs in the UK are 'dead end' jobs with poor pay and prospects or 'stepping stones' to permanent employment in good jobs. Remarkably little is currently known about temporary workers in Britain (Dex and McCulloch, 1995). Discussion of flexibility in Britain has therefore tended to rely on popular perceptions rather than systematic

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analysis. In this paper, we investigate three main issues. First, we describe who holds temporary jobs in 1990s' Britain. Second, we compare temporary jobs to permanent ones in terms of wages, and then estimate how satisfied temporary workers are with their jobs, and how much training they receive. 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 data from the first seven waves of the British Household Panel Survey (BHPS), conducted over the period 1991–7. The analysis is carried out separately for men and women in employment, and distinguishes between 'casual and seasonal workers' (where the nature of the job is temporary) and workers on 'fixed-term contracts' (where the job could, in principle, be held on a permanent basis). Since the second type is what advocates of greater labour market flexibility appear to have in mind, we explore how our results vary across the two forms of temporary work and if indeed fixed-term jobs are better than seasonal-casual employment.

Our results confirm the common – but hitherto undocumented – perception that temporary jobs are generally not desirable when compared to permanent employment. Temporary jobs typically pay less, are associated with lower satisfaction in some job components and provide less work-related training. However, we do find evidence that fixed-term contracts (but not casual/seasonal employment) 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.

Our results look at temporary jobs from a microeconomic perspective. We believe an exploration of individual-level data is an appropriate way to look at the potential costs of temporary jobs. Further, we examine a largely unregulated labour market. As the introduction to this Symposium makes clear, in economies with greater employment protection for permanent workers, there are typically more temporary jobs with potentially higher marginal costs. Our results make two important points for policy. Even in a largely unregulated labour market, the use of temporary contracts has costs in terms of less training, lower job satisfaction, and lower wages probably reflecting lower specific human capital investment. However, these costs are typically transitory, in the sense that workers on fixed-term contracts (and, to a lesser extent, in seasonal-casual work) move readily into permanent jobs and catch up either partially (for men) or fully (for women) to their counterparts who started in permanent jobs.

Section 1 presents the main hypotheses underlying our analysis. Section 2 describes the data. Section 3 provides a picture of temporary work in 1990s' Britain. In particular, we estimate who takes a temporary job, and the level of wages, satisfaction and work-related training of temporary workers compared to permanent workers. Section 4 examines the impact of an experience of a temporary job on subsequent employment and wages. The final section summarises and draws conclusions.

1. Hypotheses

Dolado *et al.* (2001) discuss the use of fixed-term contracts from a predominately macroeconomic perspective. In our discussion, as in our empirical results, we emphasise the microeconomic aspects of such contracts. Although Britain may be viewed as a benchmark case, because of her relatively mild restrictions on dismissal for redundancy or cause, it is still costly to discharge long-serving employees.[1] There is currently no limit on the number of times a fixed-term contract can be renewed, although the Fixed Term Work Directive, to be implemented in July 2002, seeks to remedy this.

Other things being equal, in a competitive labour market, workers on temporary contracts should receive a higher wage that just offsets the value of this absent employment protection. In practice, however, temporary workers may receive lower wages due to lower investment in specific human capital or because they are of lower average ability.[2] The extent of these factors may differ across gender due to standard reasons in the economics of the family. Further, for some professional jobs, general human capital is more important than specific human capital, and temporary jobs in these fields may be high wage jobs. In this section, we clarify these points to provide a basis for our empirical analysis.

First, consider temporary jobs where it is unlikely that the worker can eventually obtain a permanent contract for the job. This may be the case if the firm has a stable permanent workforce and maintains a buffer stock of temporary workers who can be readily dismissed to adjust to economic downturns. Alternatively, it may arise if the temporary job is a leave replacement for a permanent worker. In this situation, it is inefficient for workers in temporary jobs to invest in specific human capital, or for the employer to provide this training. These jobs will therefore be relatively attractive to workers who have a lower probability of wishing to remain at the firm. This includes young, single individuals who might be disinclined to make a large investment in a particular job until they are sure of their career and regional preferences. For women, the higher probability of a move to non-market employment increases this effect. Older workers might – given the shorter period of return – also be less inclined to invest in specific human capital. However, since a permanent contract is at least as desirable as a temporary one (except for any costs of investing in specific human capital), wages in this sort of temporary job should be at least as high (net of investment costs) as wages in permanent jobs. Workers should voluntarily sort into these jobs, and levels of job satisfaction should be as high as in permanent jobs. Since early investments in specific human capital should have limited impact on wages later

[1] Workers with sufficient length of service are entitled to statutory redundancy pay and can claim unfair dismissal. 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.

[2] Guell (2001) develops an efficiency wage model in which the wages of temporary workers may be low because they play no incentive role in reducing shirking. Booth (1997) raises the possibility that firing costs may be endogenous and traces through their implications for wages and temporary employment.

in their career, wages should converge for those workers who start in temporary or in permanent jobs.

However, firms may also use temporary contracts as a probation device, whether or not they wish to have a buffer stock of temporary workers to adjust to economic fluctuations or to fill in for permanent workers on leave.[3] Workers on initial temporary contracts who display high ability are later offered permanent employment at the firm.[4] If a firm knows a potential worker is of high ability, the firm might offer a permanent contract, encouraging the worker to immediately begin acquiring specific human capital.[5] If the firm is unsure, however, it may offer a temporary contract. There is an efficiency gain (in the ability to freely discharge low-ability workers at the end of their temporary contract) to offset the efficiency loss of deferred specific human capital investment for those workers who progress to permanent employment. Under these circumstances, workers who start in temporary contracts are of lower expected ability than those who start in permanent contracts, and the expected wage differentials should persist through the individuals' careers. Further, these workers – who would have preferred the permanent job – will have low levels of job satisfaction.

These competing features of temporary jobs are likely to differ in significance between men and women. The desire to defer investment in specific human capital – even for high ability individuals – is more likely to apply to women who may be deciding between market and home production (Weiss and Gronau, 1981).[6] Therefore, the data should show that initial placement in a temporary job has less of a permanent effect on observed wages for women than for men. For men, an initial temporary job is a better signal of low ability than for women. There is another important possible gender difference. Some women may wish to retain career flexibility through a significant portion of their working lives. In this case, it can be optimal to invest in a high level of general, rather than specific, human capital, and to hold a succession of temporary posts. An important example is teaching. We would therefore expect the data to show differences in the effects of temporary work across occupation and its interaction with gender.

Our discussion to this point has concerned fixed-term contracts used either as a buffer stock of employment or for leave replacements, for the purposes of probation, or for allowing a high relative investment in general rather than specific human capital. These explanations essentially apply for jobs that could, in principle, be held under permanent contracts. Seasonal/casual temporary jobs are very different. These jobs do not lend themselves to the high acquisition of

[3] A recent literature (Autor, 1999; Polivka, 1996; Abraham and Taylor, 1996; 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, firms might find it optimal to hire temporary workers only when there is an element of probation involved. Unfortunately, our data do not distinguish workers at temporary help supply firms.

[4] Models of probation include Loh (1994) and Wang and Weiss (1998).

[5] In Britain, whether or not a firm adopts a period of probation depends on individual company policy. While some companies accord permanent workers with full rights from the first day of their employment, others offer periods of probation from 3–12 months.

[6] Using BHPS data, Paull (1997) finds significantly positive returns to tenure (firm-specific human capital) that are greater for women than men. She suggests this may reflect differences in the types of contracts offered to men and women.

specific human capital, and observation suggests that they are not concentrated in fields requiring high general human capital. Under these circumstances, these jobs would be held largely by individuals with low ability to acquire human capital, and would therefore (particularly insofar as ability to acquire human capital is correlated with general ability) be low paid as well as low in job satisfaction and training, with low probability of moving into good permanent jobs. However, a possible gender difference is that some high-ability women who seek high employment flexibility and – due to possible extended absences from the labour market – may not find it optimal to invest in general human capital, may participate in seasonal/casual jobs. We would therefore expect that the wage gap between seasonal-casual and permanent work would be greater for men than women.

In summary, individuals holding temporary jobs are likely to do so either as voluntary sorting (while determining preferences over careers, location and market or home production) or involuntary (with the firm offering permanent jobs to individuals of higher perceived ability). In either case, it is inefficient to invest heavily in specific human capital, so these are likely to be lower wage jobs and – insofar as the sorting is involuntary – display low worker satisfaction. Depending on the extent to which the sorting is by ability, workers starting their careers in temporary jobs will suffer a permanent wage penalty. However, we suggest that voluntary sorting is more likely to occur for women than for men, so starting in a fixed-term job is likely to have a smaller long-term effect for women. Similarly, women should show a smaller wage differential in either fixed-term or seasonal-casual jobs than men.

2. The Data

We use the first seven waves of the BHPS, 1991–7, a nationally representative random sample survey of private households. Wave 1 interviews were conducted during the autumn of 1991, and annually thereafter. 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), reported positive hours of work, provided complete information at the interview dates, had left school and were employed at the time of the survey, and were not in the armed forces or self-employed. We have a longitudinal sample of 1,740 male and 1,981 female workers.

The data allow us to distinguish two types of temporary work: the first refers to seasonal or casual jobs; the second refers to jobs done under contract or for a fixed period of time.[7] The percentages of men and women in these two types of temporary work are given in Table 1. Over 1991–7, the average percentage of male workers in all temporary jobs is 6.8%, with 3.9% 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

[7] The precise form of the question is as follows: '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 BHPS questionnaire as well as on the sampling scheme, weighting, imputation and other survey methods used can be obtained at http://www.iser.essex.ac.uk/bhps/doc/index.htm.

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 and 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 and 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 1991–7.
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.

casual jobs and 3.3% in fixed-term contracts.[8] The distinction in types of temporary jobs is important since labour market flexibility arguments are typically based on the use of short-term contracts in what might otherwise be permanent jobs, rather than the expansion of seasonal-casual work.

Table 1 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.[9] In these raw data, men gain the highest wages in permanent work. 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 actually earned by workers on fixed-term contracts, who receive a significant £0.90 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). The two

[8] The proportions of male and female workers in seasonal and in fixed-term contracts has remained fairly stable over the sample period. Data from the Labour Force Surveys (LFS) 1991–97 show proportions of temporary male and female workers that are about 2–3 percentage points lower than those found with the BHPS data. However, the LFS figures confirm the relatively stable time trend. We are grateful to Tim Butcher for providing us with the LFS figures.

[9] The hourly wage rate is given as $\omega = PAYGU/[(30/7)(HS + \kappa HOT)]$, 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.

types of temporary work also differ in other ways. Seasonal-casual workers are concentrated in personal and protective services, sales, plant and machine operative, and other low-skill occupations as well as primary, distribution and catering industries. A sizeable group of seasonal-casual female workers are also in transport, banking and other service industries. The large share of male and female workers on fixed-term contracts, in contrast, is in professional and technical occupations across most industries. Effort – as measured by hours of work – also differs significantly between seasonal-casual workers and fixed-term contract workers. For men, mean normal weekly hours (and standard deviation) are: permanent, 45 (11); seasonal-casual, 28 (17); and fixed-term, 41 (15). For women, mean normal weekly hours are: permanent, 32 (13); seasonal-casual, 21 (13); and fixed-term, 31 (14). For both men and women, hours of work are similar between permanent and fixed-term contracts, although the latter displays greater variance. Seasonal-casual jobs entail lower hours of work.

In summary, the raw data show the need to distinguish between seasonal-casual temporary jobs (which are clearly low wage and effort jobs concentrated in low human capital occupations) and fixed-term contracts. The latter appear much like permanent jobs and, for that reason, may be buffer stock or probationary posts that can be stepping stones into good permanent jobs. The data also indicate that there are significant differences between men and women in the nature of temporary jobs, notably in the pay premium women receive under fixed-term contracts.

3. A Picture of Temporary Work

We now look more closely at a number of characteristics of temporary workers, controlling for individual and workplace attributes. We examine who takes a temporary job, the levels of job satisfaction and training of temporary workers, and how wages compare to permanent jobs. In later sections, we examine the longer-run effects of holding a temporary job.

3.1. *Who Takes a Temporary Job?*

To address this question, we perform multinomial logit regressions for men and women separately, in which the dependent variable distinguishes our three categories of employment.[10] These results are not shown but can be found in Booth *et al.* (2001). Men aged 45 and older are 2–3 times more likely to be in either form of temporary work, relative to the base of men aged 35–44 years. Men with more 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

[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. 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.

risk of being on a fixed-term contract by 30%.[11] Although not significant, there is a positive coefficient on young men aged 16–24 holding fixed-term contracts. These results are consistent with the hypotheses in Section 1. Older individuals, those who are laid off and the young may not find it efficient to invest heavily in specific human capital in a new job or may be on probationary contracts.

How does the pattern differ for women? The primary difference in the results is that women in high-skilled occupations (professional, technicians and teachers) are more likely to work under fixed-term contracts. The female probability of holding a fixed-term contract is greater in local government and other public and non-profit employment. Women with older children are also more likely to hold fixed-term contracts, while younger women are more likely to hold seasonal-casual jobs.

3.2. *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). Table 2 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.[12] 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 men and women 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.

3.3. *Training Opportunities*

In 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.*[13] 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

[11] This finding is consistent with Stewart (2000) and Arulampalam (2001). Stewart (2000) argues that unemployment experience followed by low-paid unstable jobs contributes to observed low pay persistence.

[12] In the BHPS interviews, individuals 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.

[13] 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.

Table 2

Job Satisfaction[†] *of Temporary Workers*

	Men ($N = 11{,}186$)		Women ($N = 12{,}821$)	
	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)

[†] 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 cohort of entry into the labour market (5 dummies), disabled, region of residence (6), industry (6), firm size (7), number of full-time and part-time jobs ever held at the start of the panel, marital status (2 dummies), age-marital status interactions (2), number of (marital or cohabiting) partnerships, and cohort of partnership (3).

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

days of training, conditional on receiving training. Pooled tobit regressions indicate that seasonal-casual workers receive, on average, 9–12 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.

3.4. *Wages in Temporary Jobs*

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. 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.[14] These are reported in Table 4. The OLS estimates show that the gap (compared to permanent jobs) for a seasonal-casual male is 16% and, for a fixed-term male, 17%. For women, the seasonal-casual gap is 13% and the fixed-term gap 14%. Controlling for observable characteristics has brought the temporary–permanent gaps close together across gender and across the nature of the temporary work. The FE estimates 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.[15]

In summary, we find that temporary jobs are held disproportionately by the young and the old, for whom investment in specific human capital may be inefficient. There is a significant wage penalty in both types of temporary work. Seasonal-casual jobs have low job satisfaction, while fixed-term contract jobs have low satisfaction in particular components of the measure – job security and promotion prospects. There is lower training in both types of temporary jobs. The main difference across gender is that some women seem to hold fixed-term professional jobs (e.g., in teaching) on a career basis.

4. The Effects of Temporary Employment on Career Prospects

If temporary jobs are voluntarily chosen by individuals who are unsure of their career or location preferences, then taking a temporary job should have no long-run career implications. Given that mobility rates across permanent jobs are fairly high for the young, low investment in specific human capital early in the career should have little effect. In contrast, if permanent jobs are rationed to higher ability individuals, then an initial temporary job signals that an individual is of relatively low ability, and this effect should be permanent. The relative impact of these effects may differ by gender. If new female entrants are initially less committed to the workforce, placement in a temporary job may convey less of a negative signal about ability than for men. In this section, we examine what happens to

[14] The variables included in this estimation are linear and quadratic terms in years of job tenure, local unemployment rate, and dummy variables for region of residence, education, industry, occupation, sector, firm size, disability status, part-time employment, marital status, 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 to account for non-participation using a participation probit equation. This equation is performed on 2,844 women and 17,947 person-wave observations. It is identified by age, time trend dummy variables, cohort of entry in the labour market, age–marital status interactions, cohort of first partnership, number of partnerships, number of children by age group, housing tenure and individual attitudes about working women. 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.

[15] In a previous version of the paper, we also estimated random-effects wage equations. Those estimates always lie between the OLS and FE estimates reported here for both men and women. They can be found in Booth *et al.* (2001).

Table 3

Training Receipt and Training Intensity† *of Temporary Workers*

	Men (N = 11,186)		Women (N = 12,821)	
	Pooled probit	RE probit	Pooled probit	RE probit
Training receipt				
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	
	Pooled tobit	RE tobit	Pooled tobit	RE tobit
Training intensity				
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§	

† The reported numbers are marginal effects for the two types of temporary work obtained from pooled and random-effects probit regressions (top) and from pooled and random-effects tobit regressions (bottom). 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 Table 2 plus union coverage. 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. Model is χ^2 the Wald statistic for the goodness-of-fit test and is equal $-2(L_R - L_U)$ where L_R is the constant-only log-likelihood value and L_U is the log-likelihood reported in the table. 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.

temporary workers in terms of the duration of temporary jobs, whether such jobs lead to permanent work, and the long-term wage effects of holding temporary jobs.

4.1 *Job Duration and Type of Exit*

How long do temporary jobs last compared to permanent jobs? Kaplan–Meier estimates of job duration, including both completed and uncompleted spells, reveal that the median duration of seasonal-casual jobs over the 1990s is very short: it is about 3 months for men and 6 months for women.[16] The median duration of

[16] These estimates are reported in Booth *et al.* (2001).

Table 4

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^2		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^2		−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^2		−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^2		−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^2	−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^2	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)

Table 4
(continued)

	Men				Women			
	OLS		FE		OLS		FE	
Variable	[1]	[2]	[1]	[2]	[1]	[2]	[1]	[2]
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	
No. 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. Base is married or cohabiting workers with no educational qualification, without disabilities, who live in Greater London, who have received on-the-job training and performance-related pay, work in unskilled occupations, in primary industries, in the private sector, in union-covered jobs and in firms with 1,000 or more employees. 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.

fixed-term contracts is around 12 months for both men and women. Permanent jobs have a median duration of almost 3½ years for men and 2½ years for women. By five 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?[17] The destination patterns by gender are quite similar. About 71% of men and 73% of women in temporary jobs take another job with the same employer; another 26% and 24%, respectively, move to a job at a different employer; and another 3% leave the labour force. We observe virtually no transitions from either of the two types of temporary work to unemployment (Böheim and Taylor, 2000). Of those employed in a seasonal-casual job, 28% of men and 34% of women have become permanent 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 (compared to seasonal-casual jobs) for men (38%) and almost the same for women (36%). The median duration of fixed-term contracts before exit 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.[18]

Which temporary workers are most likely to exit into permanent jobs? To investigate the transition of workers from temporary to permanent employment in a multivariate setting, we specify a discrete-time proportional hazard model relating the exit process to a number of individual- and job-specific characteristics. We 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) 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 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. 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

[17] The transition from temporary to permanent jobs is also analysed *inter alia* by Blanchard and Landier (2001) for France, Guell and Petrongolo (2000) for Spain, and Holmlund and Storrie (2001) for Sweden.

[18] 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 this transition occurs within the same firm.

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 or O level	0.167	−0.168	−0.935	−0.912	0.106	0.394	1.513***	−0.017
	(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 degree	−0.099	−0.634	0.545	0.539	−0.126	−0.141	1.739***	0.454
	(0.222)	(1.037)	(1.095)	(1.094)	(0.377)	(1.197)	(3.133)	(1.434)
University degree or more	0.259	−0.781	0.842	0.893	0.290	1.376	1.521**	0.739***
	(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)

Table 5
(continued)

	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]
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 pers. 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 govt.	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)

Table 5
(continued)

Variable	Men: Exit from seasonal and casual work to permanent work [1]	Men: Exit from seasonal and casual work to permanent work [2]	Men: Exit from a fixed-term contract to permanent work [1]	Men: Exit from a fixed-term contract to permanent work [2]	Women: Exit from seasonal and casual work to permanent work [1]	Women: Exit from seasonal and casual work to permanent work [2]	Women: Exit from a fixed-term contract to permanent work [1]	Women: Exit from a fixed-term contract to permanent work [2]
Local unempl. 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†	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. Base is workers aged 35–44, with no educational qualifications, and who work in unskilled occupations, in primary industries, in the private sector, in full-time union-covered jobs, and in firms with 1000 or more employees. For the definition of Model χ^2 see footnote of Table 3. The χ^2 statistic has 35 degrees of freedom and its p-value is in square brackets.
† 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, 11,130) = 9.36 (p-value = 0.000) and F(11, 12,763) = 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.

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, but not for women, being younger than 35 years of age has a positive impact on exit to permanent work, suggesting that men on fixed-term contracts follow a natural career progression into permanency. Women in the public and non-profit sectors have a lower likelihood of exiting into permanent work. Interestingly, part-time men have a low exit rate, while part-time women do not have a significantly different exit rate than full-time women.

The transition from seasonal-casual to permanent work also displays a differential public sector effect, with public sector women less likely to exit into permanent jobs. Part-time seasonal-casual workers, regardless of gender, are less likely to move to permanency. Local labour market conditions (as measured by the unemployment–vacancy ratio) also have a negative impact on exit.

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 effort to be a crucial determinant of exit from a temporary into a permanent position at a firm. To proxy effort, we use the number of weekly unpaid overtime hours usually worked. Because of potential endogeneity, 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.[19] 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.[20]

4.2. *Wage Profiles*

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

$$\ln \omega_{ijt} = \beta_0 + \beta_1 \mathbf{X}_{ijt} + \beta_2 \mathbf{Z}_{ijt} + \mu_i + \phi_{ij} + \varepsilon_{ijt} \tag{1}$$

[19] The number of children by four age groups, dummy variables for cohort of labour market entry (5), region of residence (6) and receipt of 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.

[20] We explored the relationship between effort and exit rates by looking at two additional specifications, one that distinguishes 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 is insignificantly affected by any of the effort measures. For all the other temporary workers, though, effort matters. 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.

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).[21] 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 on-the-job training in the last 12 months, job mobility variables (indicating if they have changed job because of promotion, quit or layoff), and the average local unemployment rate. The vector $\boldsymbol{Z}$ includes the contract-related variables that are the focus of our study. Specifically, $\boldsymbol{Z}$ 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.[22] 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 (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}.[23] We treat as endogenous all the regressors in $\boldsymbol{Z}$, along with part-time employment status, part-time and full-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.[24] 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

[21] 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.

[22] 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).

[23] 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.

[24] We have performed several sensitivity tests in which other variables in $\boldsymbol{X}$ 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.

uncorrelated with the composite error term by construction, and thus they are valid instruments.[25]

Table 6 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 $\mathrm{Var}(\phi_{ij})$ is about 44% of the total variance). This may help to reconcile the differences between the raw data presented in Table 1 and the estimates found with the OLS and FE regressions. 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. For 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. For workers with ten years' 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% for men and women respectively, with ten years of full-time experience. The returns to experience capital differ strongly by contract type and 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.[26]

To describe the effect of contract type on wages further, we compute predicted log-wages paths from the column [2] estimates of Table 6 for workers with four

[25] In other regressions not reported here, we also used as instruments the number of children that each worker has during the 7-year period and the local unemployment rate (Light and McGarry, 1998). The over-identifying-restrictions tests fail to reject the hypothesis that these two additional sets of variables are valid instruments at any conventional level of significance, and they slightly improve the R^2 in the first-stage regressions. However, the estimated parameters for the variables of primary interest (*NSCJ7*, *NFTC7* and their interactions with full-time experience) are not altered when these additional instrumental variables are used. Moreover, with these 'extra' instruments, the structure underlying model (1) relies on exclusion restrictions which are hard to justify. We therefore decided to exclude such instruments from the specifications discussed below.

[26] 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 nonlinear 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.

Table 6

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)
$NSCJ7^2$	0.027**	0.033***	0.013**	0.012**
	(2.212)	(2.737)	(2.409)	(3.509)
NSCJ7 × full-time experience		−0.001		−0.007***
		(0.298)		(2.786)
NSCJ7 × full-time experience^2		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)
$NFTC7^2$	0.013**	0.014**	0.006	0.004
	(2.183)	(2.098)	(1.418)	(0.910)
NFTC7 × full-time experience		0.005		0.007
		(1.324)		(1.193)
NFTC7 × full-time experience^2		−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^2	−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	

Notes: *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 those used in the OLS and FE regressions (see text). All wage equations for women are selectivity corrected. *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.

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 assuming that 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.[27]

[27] 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 profiles in Fig. 1.

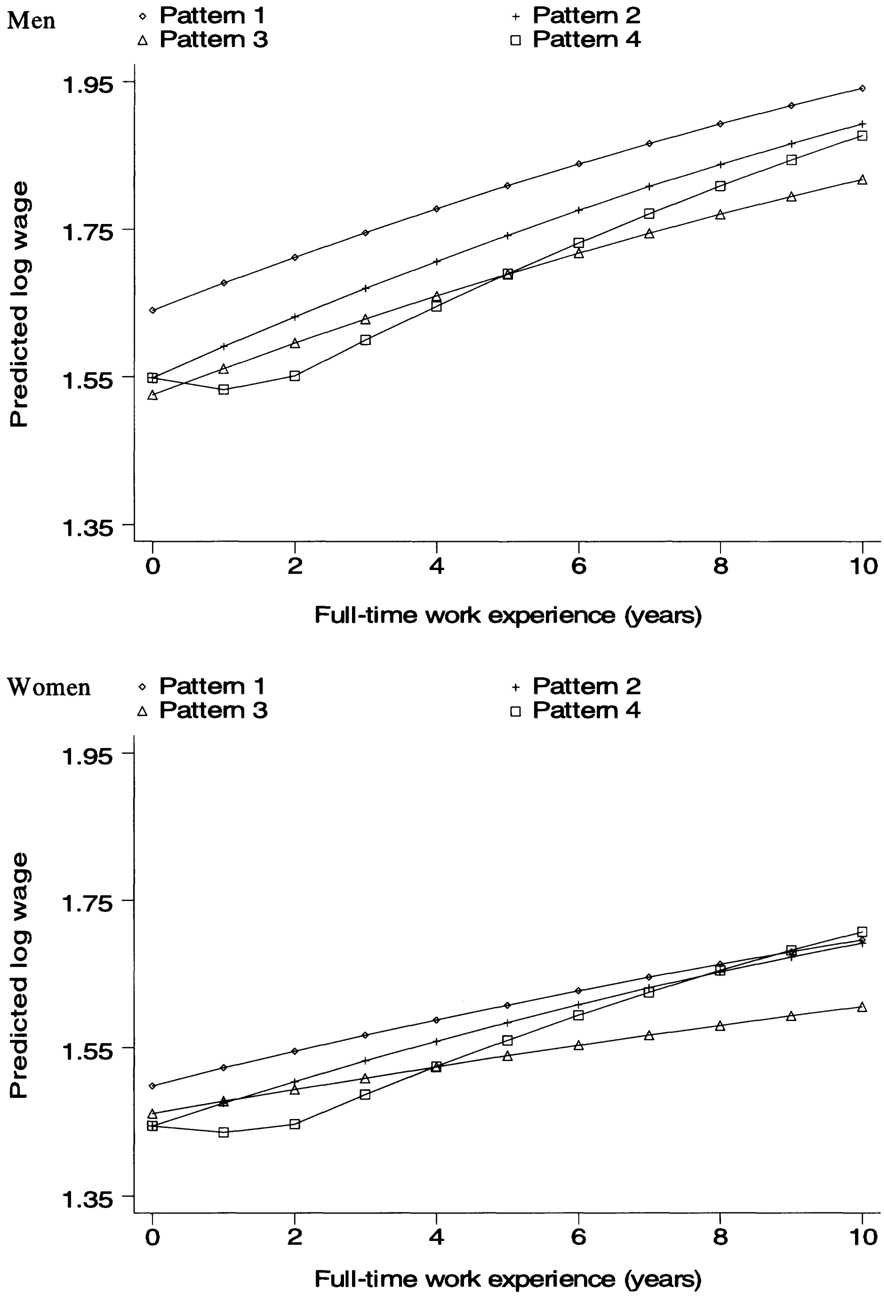

Fig. 1. *Predicted Log Wages by Experience Level and Early Employment Patterns*

Note: Based on predictions from the estimates presented in Table 5. 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

The results of this simulation are graphed in Fig. 1. Having always had a permanent job is clearly the pattern delivering the highest real wage profile over the first ten years of a man's career, with an average growth of 3% per year. Men 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. Men who started off with a seasonal-casual job, though, 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 with patterns 2 or 4. This finding holds for women too. However, for women, 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 10-year period). The wage gap for these 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'.

5. 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 (at least in some components), receive less work-related training than their counterparts in permanent employment and receive lower wages. This holds for both seasonal-casual workers and workers on fixed-term contracts, and for both men and women. Therefore, the creation of temporary jobs as a substitute for permanent jobs – in the desire to increase labour market flexibility – comes at a cost.

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 are consistent with theories of temporary work. Particularly, in seasonal-casual jobs (where there is little possibility of moving to a permanent job), there will be little training. Wages and job satisfaction will be low. For fixed-term temporary jobs, there is greater potential for moving into permanent jobs. There is evidence for this in our study. 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 compared to men who begin their careers in permanent jobs. This is consistent with the idea that these men are less able than those who immediately acquire a permanent job on entering the workforce. 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, on 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 specific human capital acquisition during the period of temporary work.

The important policy conclusions from our work is that expansion of temporary work, as a way of increasing labour market flexibility, comes at a cost. However, the cost may be transitory and workers can effectively use fixed-term jobs (and, to a lesser extent, seasonal-casual jobs) as stepping stones to permanent work.

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References

Abraham, Kathrine G. and Taylor, Susan K. (1996). 'Firms' use of outside contractors: theory and evidence', *Journal of Labor Economics*, vol. 14(3), pp. 394–424.

Altonji, Joseph G. and Shakotko, Robert A. (1987). 'Do wages rise with job seniority?', *Review of Economic Studies*, vol. 54, pp. 437–59.

Arulampalam, Wiji (2001). 'Is unemployment really scarring? Effects of unemployment experiences on wages', Economic Journal, vol. 111, November, pp. F585–606.

Arulampalam, Wiji and Booth, Alison L. (1998). 'Training and labour market flexibility: is there a trade-off?' *The British Journal of Industrial Relations*, December, vol. 36(4) pp. 521–36.

Autor, David H. (1999). 'Why do temporary help firms provide free general skills training?', mimeo, MIT, June.

Bentolila, Samuel and Bertola, Giuseppe (1990). 'Firing costs and labour demand: how bad is Eurosclerosis,' *Review of Economic Studies*, vol. 57, pp. 381–402.

Bentolila, Samuel and Saint-Paul, Gilles (1994). 'A model of labour demand with linear adjustment costs', *Labour Economics*, vol. 1, pp. 303–26.

Blanchard, Olivier and Landier, Augustin (2001). 'The perverse effects of partial labour market reform: fixed term contracts in France', Economic Journal, this issue.

Böheim, René and Taylor, Mark P. (2000). 'The search for success: do the unemployed find stable employment?', mimeo, University of Essex, February.

Booth, Alison L. (1997). 'An analysis of firing costs and their implications for unemployment policy', in (D. Snower and G. de la Dehesa, eds) *Unemployment Policy*, Cambridge: Cambridge University Press, pp. 359–88.

Booth, Alison L., Francesconi, Marco and Frank, Jeff (2001). 'Temporary jobs: who gets them, what are they worth, and do they lead anywhere?', ISER Working Paper No. 00-13. University of Essex, April.

Clark, Andrew E. (1996). 'Job satisfaction in Britain', *British Journal of Industrial Relations*, vol. 34(2), pp. 189–217.

Cramer, J. S. and Ridder, Geert (1991). 'Pooling states in the multinomial logit model', *Journal of Econometrics*, vol. 47, pp. 267–72.

Dex, Shirley and McCulloch, Andrew (1995). 'Flexible employment in Britain: a statistical analysis', Equal Opportunities Commission, Research Discussion Series No. 15.

Dolado Juan J., Garcia-Serrano, Carlos and Jimeno, Juan F. (2001). 'Drawing lessons from the boom of temporary jobs in Spain', ECONOMIC JOURNAL, this issue.

Farber, Henry S. (1999). 'Alternative and part-time employment arrangements as a response to job loss', *Journal of Labor Economics*, vol. 17(4, pt. 2), S142–69.

Guell, Maia (2001). 'Fixed-term contracts and unemployment: an efficiency wage analysis', Princeton University IRS Working Paper No. 433.

Guell, Maia and Petrongolo, Barbara (2000). 'The transition of workers from temporary to permanent employment: the Spanish case', London School of Economics CEP DP 438.

Hausman, Jerry A. and Taylor, William E. (1981). 'Panel data and unobservable individual effects', *Econometrica*, vol. 49(6), pp. 1377–98.

Holmlund, Bertil and Storrie, Donald (2001). 'Temporary work in turbulent times: the Swedish experience', ECONOMIC JOURNAL, this issue.

Houseman, Susan N. and Polivka, Anne E. (1999). 'The implications of flexible staffing arrangements for job stability', Upjohn Institute Staff Working Paper No. 99-056, May.

Light, Audrey and McGarry, Kathleen (1998). 'Job change patterns and the wages of young men', *Review of Economics and Statistics*, vol. 80, pp. 276–86.

Loh, Eng Seng (1994). 'Employment probation as a sorting mechanism', *Industrial and Labor Relations Review*, vol. 47(3), pp. 471–86.

Paull, Gillian (1997). 'Low pay and wage growth in Britain: the returns to experience and tenure', mimeo, Institute for Fiscal Studies, March.

Polivka, Anne E. (1996). 'Are temporary help agency workers substitutes for direct hire temps? Searching for an alternative explanation of growth in the temporary help industry', mimeo, Bureau of Labor Statistics, Washington, DC.

Purcell, Kate, Hogart, Terence and Simm, Claire (1999). *Whose Flexibility? The Costs and Benefits of Non-standard Working Arrangements and Contractual Relations*, Joseph Rowntree Foundation, York: York Publishing Services, September.

Segal, Lewis M. and Sullivan, Daniel G. (1997). 'The growth of temporary services work', *Journal of Economic Perspectives*, vol. 11(2), Spring, pp. 117–36.

Stewart, Mark B (2000). 'The inter-related dynamics of unemployment and low pay', mimeo, University of Warwick.

Wang, Ruqu and Weiss, Andrew (1998). 'Probation, layoffs, and wage-tenure profiles: a sorting explanation', *Labour Economics*, vol. 5(3), pp. 359–83.

Weiss, Yoram and Gronau, Reuben (1981). 'Expected interruptions in labour force participation and sex-related differences in earnings growth', *Review of Economic Studies*, vol. 48, pp. 607–19.