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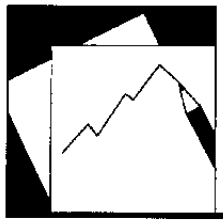

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The Determinants of On-the-Job Search: An Empirical Exploration

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The Determinants of On-the-Job Search: An Empirical Exploration

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Abstract

There was a surge in the number of employed people looking for another job in the United Kingdom in the 1980s. In this paper, we present a panel analysis of aggregate data and a cross-section analysis of individual data on on-the-job search in the United Kingdom. We find evidence that the availability of jobs and wage dispersion increase on-the-job search. The importance of these results is twofold. First, to the extent that on-the-job search responds to the tightness of the labor market, it can contribute to explaining the observed cyclical behavior of the unemployment outflow rate. Second, as shown in Fuentes (2002), to the extent that changes in on-the-job search can be explained by factors other than labor market tightness, such as wage dispersion, these shift the unemployment-vacancies relationship (the Beveridge curve) and therefore have a role to play in the determination of unemployment.

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[1] Work on this paper started when Andrés Fuentes was at the Institute of Economics and Statistics, Wolfson College and Oriel College in the University of Oxford (United Kingdom). The paper benefited from comments by Stephen Bond and Stephen Nickell, and from comments received at an IMF Institute seminar. All errors remain the author's.

Contents Page

I. INTRODUCTION

In this paper, we present an empirical study of the determinants of on-the-job search. This is to address two questions: First, to what extent does on-the-job search respond to changes in labor market tightness? In industrialized countries the unemployment inflow rate shows more cyclical variation than the unemployment outflow rate, whereas persistent unemployment is accompanied by low unemployment outflow rates. If on-the-job search is itself subject to cyclical influences, it can explain the cyclical behavior of the unemployment outflow rate (Burgess, 1993).

The idea behind this is as follows: As an economy recovers from a recession, the increase in available jobs encourages on-the-job search. A larger number of on-the-job searchers provides competition for the unemployed and therefore reduces ceteris paribus the unemployment outflow rate, neutralizing the effect of a larger number of hires on the unemployment outflow rate. Conversely, at the onset of a recession job availability declines, discouraging on-the-job search. The decline in on-the-job search mitigates the resulting fall in the unemployment outflow rate.

Burgess (1993) has also noted that the unemployment outflow rate is poorly correlated with hires in the United Kingdom and that this may result from the endogeneity of on-the-job search to labor market tightness. Even though much empirical research has been based on search models of the labor market, and considerable effort has been put into explaining the unemployment outflow rate in Britain (e.g. Burgess, 1993; Layard, Nickell, and Jackman, 1991, Chapter 5), there is as yet no empirical investigation of aggregate on-the-job search data in a time-series context.

Second, we investigate which other factors, apart from labor market tightness, play a role in the determination of on-the-job search. We have shown elsewhere (Fuentes, 2002) how these can produce shifts of the unemployment-vacancies relationship (the Beveridge curve) and that on-the-job search appears to have shifted the U.K. Beveridge curve outwards, implying an increase in unemployment at any given level of vacancies.

We present both a cross-section analysis of individual data and a panel data analysis of aggregate data on the determinants of on-the-job search. The cross-section investigation will allow us to determine the role of demographic and job characteristics, including wages, in the search decision. The analysis of aggregate data will allow us to address the cyclical behavior of on-the-job search and to estimate the effects of changes in the dispersion of the wage distribution over time. The aggregate data cover 10 British regions over a period of 12 years.

Section II gives an overview of the insights the job search literature offers into the determinants of on-the-job search. In Section III we present the cross-section, and in Section IV, the panel data study. Section V concludes.

II. MODELS OF ON-THE-JOB SEARCH

In this section we first review the insights that search theory has provided into the determinants of an employed individual's decision to search following Burgess (1992) and Mortensen (1986). Second, we review the empirical literature on on-the-job search which is so far confined to cross-section investigations.

In search models individuals maximize the expected present value of income over an infinite time horizon, and engage in costly search in order to receive wage offers. The arrival of wage offers over time is determined by a known random process. The distribution of the random wage level in each wage offer is also known and is identical to the distribution of wages across existing jobs.

Following Burgess (1992), employed workers face an exogenous probability of job loss and face a cost when changing jobs. When unemployed, they receive an unemployment benefit.

On the basis of these assumptions a unique reservation wage is derived at which employed workers are indifferent between searching and not searching. The proportion of all employed workers who decide to search on the job is given by the proportion of workers earning a wage below the reservation wage level. The reservation wage depends on the unemployment benefit, the costliness of search, mobility costs, the wage received on the current job, the offer arrival rate, the probability of job loss and the discount rate.

We next review the comparative statistics of these models. The higher the offer arrival rate, the more likely employed people are to decide to search on the job. Search costs incurred on the job reduce search, whereas higher search costs for the unemployed tend to increase the incidence of on-the-job search by inducing unemployed workers to take up relatively unattractive jobs and continue to search on-the-job. Similarly, lower unemployment benefit increases the probability of on-the-job search.

A mean-preserving increase in the spread of the wage offer distribution makes job search more likely. This is because a wider spread of the wage distribution increases the chance to obtain a job at the top tail of the wage distribution and therefore increases the expected gains from search. An increase in the mean of the wage distribution, holding all its other moments as well as the wage currently earned constant, also increases the probability that an employed worker will engage in on-the-job search.

An increase in the probability of job loss has an ambiguous effect on on-the-job search. On the one hand, an increase in the probability of loss of a worker's current job, given the aggregate layoff probability, adds to the benefits of searching on the job and therefore increases its incidence. The intuition of this is that on-the-job search provides insurance against the eventuality of becoming unemployed. However, a higher probability of job loss in a newly found job reduces the value of searching (Burgess and Nickell, 1990). Finally, higher costs of changing jobs deter on-the-job search.

We summarize these findings in the following equation relating the probability that an employed person will search on the job to the forces driving the search decision: $\phi = \phi(x, W, Z)$ where x denotes the offer arrival rate, W the vector of parameters describing the wage distribution and the individual's wage, and Z the remainder of the variables influencing the reservation wage at which the employed are indifferent between searching and not searching.

Given the increase in wage dispersion in the United Kingdom in the 1980s, the effect of wages on the job search decision of employed workers is of primary interest. Few studies have however investigated the relationship between wages and on-the-job search. No consensus has yet emerged on wage effects from the literature. Moreover, existing work on the issue is based on cross-section data which does not allow the determination of the effects of changes in the dispersion of the wage distribution on the incidence of on-the-job search.

Black (1981), using data from the United States, finds that an individual's potential wage gain is one of the main determinants of the decision of whether or not to search for another job. He estimates the potential wage gain as the residual from a regression of the wage on human capital variables, local labor market conditions and demographic variables. Hartog and van Ophem (1994) also model the probability of an employed worker's decision to search on-the-job as a function of a wage residual, where the wage residual is derived from a wage equation with labor market experience and years of education as explanatory variables. They conduct several cross-section regressions for various years and find a significant negative effect of the wage residual on the search decision in some but not all of them. They include a variable in their estimating equation indicating the interviewee's assessment of promotion possibilities, with a significant negative effect of good promotion possibilities.

Allen and van der Velden (2001) investigate job search decisions of a sample of a cohort of graduates from tertiary education, and find insignificant negative effects of the workers' own wages on the job search decision, controlling for a range of self-reported job amenities which turn out to be statistically significant. Pissarides and Wadsworth (1994) have conducted a cross-section analysis of on-the-job search based, as this study, on the U.K. *Labour Force Survey*, at a time, however, when it did not include data on wages.[2]

Further cross section studies of on-the-job search in the United States show the positive effect of advance notification of redundancies on the on-the-job search decision (for example, Schwer and Wadoups, 1996).

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[2] Pissarides and Wadsworth's approach is to estimate, in a first stage, the probability that an employed person decides to search, no matter whether search is undertaken by quitting into unemployment or by searching on-the-job, and, in a second stage, the probability of an individual searching on-the-job, conditional on the decision to search using a sample selection framework. We find however that the number of voluntary quits into unemployment is very small so we do not follow this procedure.

The empirical studies reviewed above show a negative effect of age and a positive effect of temporary employment on on-the-job search. Workers with high qualifications and short job tenures appear to be more likely to search. Studies which include variables capturing labor market conditions (Black, 1981; Pissarides and Wadsworth, 1994) find that vacancies have a positive and unemployment a negative effect, although the latter is not significant in Black's equation. Some studies control for the probability of job loss, through survey respondents' self assessment (Hartog and van Ophem, 1994) or, if less comprehensively, with a dummy for employment in the construction industry (Black, 1981), and find a positive effect of higher job insecurity. However, in all studies the data do not permit separately to control for job availability and the probability of job loss.

III. Cross-Section Analysis

A. The Empirical Model

We model on-the-job search as a dichotomous variable which assumes the value 1 if an employed person is searching and 0 if not. This is the dependent variable of a *probit* equation. The explanatory variables, outlined below, capture the determinants of an employed individual's probability to search, following the insights provided by the literature reviewed in the previous section.

- *The potential wage gain*: To approximate an employed worker's position in the wage distribution of potential wage offers, given the worker's skills and experience, we estimate a Mincerian earnings equation in which the logarithm of earnings is regressed on human capital and job amenity variables. Our explanatory variable in the on-the-job search equation is the difference between the logarithm of actual earnings and the earnings predicted by the earnings equation.

- *Mobility costs*: The *Labour Force Survey* allows the inclusion of dummies on the presence of dependent children in the household, marital status, and home ownership to capture mobility costs.

- *Offer arrival rates, job loss probabilities, search costs*: We have no data separately to identify occupation-specific and region-specific offer arrival rates, job loss probabilities and differences in search costs. To control for all three variables, we include occupational, qualificational, and regional dummies (the latter however turned out to be insignificant). In addition, since job turnover is particularly strong in construction, we include a dummy for workers employed in this industry.

- *Job tenure variables*: In the presence of "last in, first out" layoff rules, job tenure reduces the probability of layoff in the current job relative to other jobs, resulting in a negative effect of long job tenure on on-the-job search incidence given the potential wage gain.

- *Small firm*: To allow for the possibility that internal labor markets in big corporations reduce the need for job mobility we also control for the proportion of workers in firms with 10 or fewer employees.

- *Age*: Older workers have a shorter remaining working life ahead and therefore benefit less from a job change.

B. The Data

The source is the Spring 1996 U.K. *Labour Force Survey*. The sample includes about 8000 employees but excludes the self-employed. Interviewees are asked about their earnings only in one of the five interviews conducted with every household throughout the year. Therefore, only 20 percent of the full *Labour Force Survey* sample can be exploited, and panel data analysis is not possible with data at the individual level.

We estimate separate equations for full-time and part-time employees and present results for earnings rather than wages in order to determine the effect of pay on on-the-job search. The analysis was however conducted for both wages and earnings with no substantial differences in the results. Observations containing missing values for any of the variables were dropped from the sample. This reduced the sample by 0.5 percent.

C. The Earnings Equation

The variables that capture human capital in the earnings equation are age, qualification, and job tenure variables. The job tenure variables reflect the returns to the accumulation of firm-specific human capital. Since firm-specific human capital does not contribute to potential earnings in an alternative job the residuals from the earnings equation are computed setting the coefficients on the tenure variables to zero. In addition we control for racial differences in pay.

Topel (1991) and Ruhm (1990) point out that job tenure is likely to be correlated with the error term in the earnings equation because the length of job tenure reveals the quality of the match between employer and worker. Match quality thus results in an upward bias on the coefficient of the job tenure variable as an indicator of firm specific human capital. The instruments proposed by Ruhm and Topel require a panel data set which is not available for this study. However we are not interested in the returns to firm-specific human capital. For the purposes of this study, we note that unobservable effects on wages picked up by job tenure, such as match quality, are not likely to contribute to a worker's potential wage outside their current job.

We also control for job amenities to the extent that the *Labour Force Survey* provides information on these. Since on-the-job searchers are likely to confine their search to jobs in occupations similar to their current one, we need to control for broad occupational groups in order to calculate workers' potential earnings. We have also added dummies for 20 regions to our earnings equation to control for differences in regional price levels. The estimation results for full-time employees are reproduced in Table 1.

Table 1. Earnings Equations for Full-Time Employees

(Sample: men: 3,926 observations; women: 2,344 observations
dependent variable: log weekly earnings)

	Men		Women	
	Coefficient	*t-Statistic*	Coefficient	*t-Statistic*
Age	0.030	15.1	0.026	10.0
Age squared/100	-0.036	-14.8	-0.031	-9.3
Married	0.053	7.0	-0.013	-1.5
Permanent job	0.021	1.3	0.038	2.2
Evening work	0.059	7.4	0.049	4.8
Night work	0.049	5.8	0.041	3.3
Saturday work	-0.010	-1.1	-0.018	-1.8
Sunday work	0.020	2.5	0.037	3.0
Shift work	-0.009	-1.0	-0.025	-1.9
Degree	0.222	15.9	0.216	12.1
Below degree	0.155	10.7	0.152	8.5
Teacher	0.060	1.1	0.151	5.4
A levels	0.125	9.7	0.107	6.5
Apprenticeship	0.079	6.8	0.091	4.0
GCSE high	0.082	6.6	0.092	6.6
GCSE low	0.046	3.5	0.033	2.2
White	0.074	3.3	0.052	1.9
Black	0.003	0.1	0.015	0.4
Tenure <1 year	-0.110	-8.8	-0.151	-8.0
1 yr <tenure <2 yrs	-0.099	-7.5	-0.162	-8.3
2 yr <tenure <5 yrs	-0.071	-6.1	-0.099	-5.5
5 yr <tenure <10 yrs	-0.029	-2.7	-0.070	-4.1
Tenure<20 yrs	-0.026	-2.5	-0.054	-3.1
Dummies on				
Occupations	yes		yes	
Regions	yes		yes	

Notes: A levels and GCSE are high school diplomas received after 13 and 10 years of schooling, respectively. GCSE "high" and "low" refers to the number of subjects covered in the diploma; and "tenure" refers to job tenure.

D. The On-the-Job Search Equation

Before turning to the results of the on-the-job search equation, we discuss two identification issues. First, the residual in the earnings equation is likely to be correlated with the error term in the on-the-job search equation. We expect the earnings residual to be *positively* correlated with the error in the on-the-job search equation because those unobserved individual characteristics that make workers keener searches are also likely to make them earn a higher wage. The effect is to bias the coefficient on the wage residual towards zero. If we find a statistically significant negative effect of wages on on-the-job search, we can therefore conclude that wages play a role in the probability to search on-the-job.

Second, since heteroskedasticity renders the *probit* coefficients inconsistent, we have included an LM test against heteroskedasticity, where the alternative hypothesis for the variance of the error ε_i of observation i is of the form:

$$Var(\varepsilon_i) = \left(e^{\alpha' w_i}\right)^2 \qquad (1)$$

where w_i is a vector of variables including earnings, qualifications and occupations. This test also has power against misspecification of the functional form of the estimating equation (see Godfrey, 1988).

The results of the on-the-job search *probit* equation are presented in Table 2. The results of the on-the-job search equation for female part-time workers are reproduced in Table 3. In the regressions for male and female full time employees the LM test failed to reject the null hypothesis of homoskedastic disturbances.

The coefficient on the residual of log earnings is significant for men. For women the effect of earnings on the search decision is on average smaller than for men, making it insignificantly different from zero. This result was robust to using weekly or hourly earnings. Inclusion and exclusion of occupational and regional dummies also made little difference. A 1 percent increase in the earnings residual increases the probability of search by 6 percent in the case of men and by 4 percent in the case of women.

Since the earnings residual approximates the expected present value of future gains in earnings to be made from changing jobs, a smaller effect for women is to be expected. Since women are more likely to interrupt their working careers than men, their expected gain from changing jobs is smaller, for any given wage differential, than it is for men.

We also estimated the on-the-job search equation interacting the earnings residual variable with dummies for groups of qualifications attained (results not presented). Dividing workers into three skill groups, we found that unskilled workers displayed the strongest effect of the wage residual on search, but the differences were not statistically significant.

Table 2. On-the-Job Search: Binomial Probit Model, Full-Time Employees

(Sample: men: 3,926 observations; women: 2,344 observations
dichotomous dependent variable: on-the-job search)

	Men		Women	
	Coefficient	*t-Statistic*	Coefficient	*t-Statistic*
Earnings residual	-0.53	-3.2	-0.36	-1.5
Married	-0.23	-2.8	-0.22	-2.4
Permanent job	-0.68	-5.6	-0.72	-5.0
Age	0.03	1.5	0.07	2.3
Age squared/100	-0.05	-1.8	-0.11	-2.5
Degree	0.02	2.3	0.04	0.3
No qualifications	-0.20	-1.7	-0.01	-0.07
Owner occupier	-0.06	-0.9	0.03	0.3
No children	-0.10	-1.3	-0.04	-0.4
Construction	-0.01	-0.1	0.39	1.4
Tenure < 1year	0.48	4.3	0.45	2.9
1 yr <tenure <2 yrs	0.62	5.5	0.54	3.4
2 yr <tenure <5 yrs	0.65	6.6	0.42	2.9
5 yr <tenure <10 yrs	0.33	3.4	0.30	2.1
Small firm	-0.03	-0.4	-0.06	-0.5
Occupations	yes		yes	
Regions	no		no	
LM(17)	23.7		21.9	

Notes: "owner occupier" is a dummy for individuals living in households whose dwelling is owned by one of its occupiers; "no children" refers to the presence of dependent children; and "tenure" refers to job tenure.

Naturally, a temporary job makes on-the-job search more likely. So does short job tenure. As far as costs of changing jobs are concerned, neither ownership of the dwelling nor the presence of dependent children in the household deter search much. Married men and women are less likely to search for a job than single full-time employees. Employment in the construction industry, despite its large amount of job turnover, does not make on-the-job search more likely, nor does employment in a small firm.

Turning to the coefficients on the age variables, the two coefficients, although not significantly different from zero singly in the equation on men, are jointly significant at the 10 percent level. The magnitudes of the coefficients for both men and women imply that the marginal effect of age is positive below the age of 34 years and negative above it.

Our results on the wage residual are in line with Black's (1981). Two reasons may account for the fact that Hartog and van Ophem's (1994) and Allen and van der Velden's (2001) results differ from ours on this point: First, Hartog and van Ophem include a highly significant dummy variable on the promotion possibilities individuals think they have on their current job. If workers feel that promotion possibilities are not as good as they should be this likely indicates that they are in fact paid less than they are able to earn elsewhere. Second, Allen and van der Velden limit their sample to highly skilled workers. Our own results suggest that highly skilled workers have the smallest wage effects on search, a finding that our panel data results (Section IV) confirm. Our results thus provide an explanation for the lack of conclusive wage effects in the literature.

In the on-the-job search equation for part-time working women (see Table 3), the LM test gave evidence of heteroskedasticity (LM=62, 17 degrees of freedom). Therefore, we allowed the variance of the error term to depend on age, pay, occupations and qualifications. Most variables are not significant at the 5 percent level. The earnings residual has the expected negative sign. The point estimate is larger in absolute magnitude than in the full-time equations for men and women. However, the estimated marginal effect on the probability of search is considerably smaller at 1.1 percent. Differences in the coefficients across skill groups were similar to the ones found for full-time employees and again were not statistically significant.

Table 3. On-the-Job Search: Binomial Probit Model, Part-Time Employees—Women

(Sample: 1,919 observations dichotomous dependent variable: on-the-job search heteroskedastic error term)

	Coefficient	*t-Statistic*
Earnings residual	-0.60	-1.3
Married	-0.15	-1.3
Permanent job	-0.27	-1.4
Age	0.06	1.7
Age squared/100	-0.06	-1.5
Degree	0.01	0.1
No qualifications	-0.20	-1.1
Owner occupier	-0.27	-1.6
No children	0.25	1.5
Construction industry	0.09	0.3
Tenure < 1year	0.51	1.5
1 yr <tenure <2 yrs	0.56	1.6
2 yr <tenure <5 yrs	0.49	1.5
5 yr <tenure <10 yrs	0.36	1.5
Small firm	0.03	0.4
Dummies on:		
Occupations	yes	
Regions	no	
LM (17)	63.0	

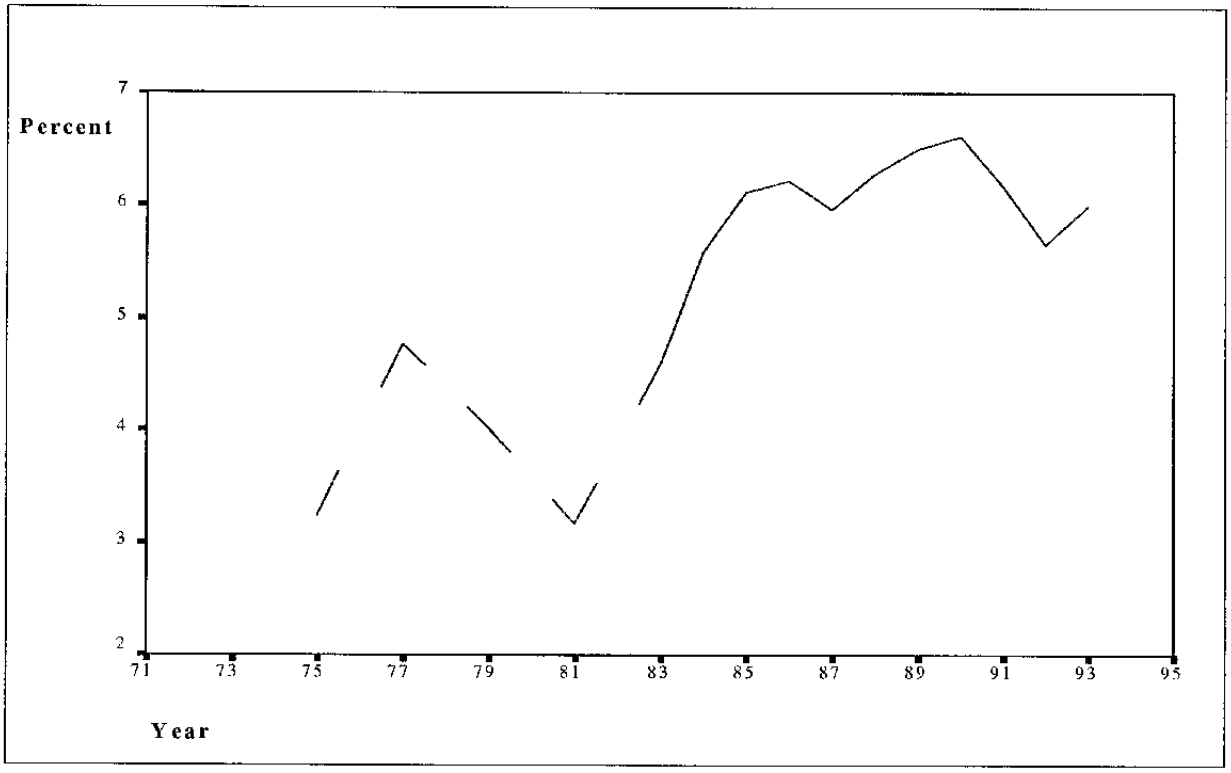
Notes: "owner occupier" is a dummy for individuals living in households whose dwelling is owned by one of its occupiers: "no children" to the presence of dependent children; and "tenure" refers to job tenure.

IV. Panel Data Analysis

A. The Empirical Model

Figure 1 shows how on-the-job search has evolved in the United Kingdom between 1975 and 1993. Some cyclicity is evident, and there is a marked increase in on-the-job search over the sample period. The purpose of this section is to explore both the apparent cyclical behavior of on-the-job search and its dramatic rise in the 1980s. In order to do so we will use a panel of data of 10 British regions over the period 1982 to 1993.

Figure 1. Employed Workers Looking for a Job



Source: United Kingdom, *Labour Force Survey*, own computations.

As the theory reviewed in Section I indicated, both the availability of jobs, reflected in the job offer arrival rate, as well as the probability of job loss influence on-the-job search.

While the aggregate probability of job loss can be approximated with the number of redundancies as a proportion of employment, the offer arrival rate is not observable. This variable will be modeled as follows (see Burgess, 1993). The proportion of job offers searchers accept is assumed to be exogenous and constant, and, without further loss of

generality, is set to one. The number of job offers in any period of time is therefore equal to the number of hires. The number of hires in turn is determined through the matching process as a function of the number of job searchers and vacancies:

$$H = h(JS,V) \ , \tag{2}$$

where V is the number of vacancies. Assuming that on-the-job searchers and unemployed searchers are perfectly substitutable from the employers' point of view, the offer arrival rate, x, is the ratio of hires to the number of searchers (both employed and unemployed).

$$x = \frac{H}{JS}, \tag{3}$$

where JS is the total number of job searchers: $JS = U + \phi(L-U)$. U and L denote unemployment and the labor force, respectively.

Since regional hiring data are not available, we need to substitute the matching function of equation (2) in equation (3). The proportion of on-the-job searchers can, then, be rewritten as a function of vacancies (V) and unemployment (U):

$$\phi = \Phi(U,V,Z,W). \tag{4}$$

We contrast the assumption of perfect substitutability between searchers with the assumption that employers rank employed applicants above unemployed workers. Ranking of employed and unemployed searchers is first proposed by Anderson and Burgess (1995) and is described in more detail in Fuentes (2002). Under the ranking assumption the presence of unemployed searchers is, by definition, irrelevant for an employed searcher's odds of receiving a job offer. The number of job offers going to employed job searchers can therefore be calculated by evaluating the matching function at the number of *employed* job searchers rather than the total number of job searchers. The offer arrival rate is then given by

$$x = \frac{h(OJS,V)}{OJS}. \tag{5}$$

In both the assumptions of substitutable searchers and ranking we make the underlying assumption that employed and unemployed job searchers apply for the same vacancies. Table 4 gives an indication to what extent this is the case, comparing employed and unemployed job searchers by broad socioeconomic groups according to their current job, or, in case of the unemployed searchers, according to their last job. The skill composition of both groups of searchers appears to be sufficiently similar for the unemployed to face competition from employed job searchers for jobs. This conclusion is also born out in van Ours' (1995) study which shows that 65 percent of vacancies to which the unemployed apply receive applications from employed job searchers as well.

Table 4. Skill Composition of Employed and Unemployed Job Searchers
(percent)

Socioeconomic group	Employed Job Searchers	Unemployed
Professional, managerial	22.2	10.1
Nonmanual	44.1	19.8
Skilled manual	12.3	19.1
Low-skilled manual	21.4	27.2
None (no previous job)	--	23.8

Sources: United Kingdom, *Labour Force Survey*, own computations.

Z in equation 4 includes variables found to be significant in the cross-section study. In addition we include the following variables:

- *The redundancy rate*: This captures the probability of job loss.
- *The replacement ratio*: This is unemployment benefits[3] divided by the 10 percentile of male full-time earnings.

We do not include the proportion of employees in temporary jobs in the list of explanatory variables because it is not available for the whole sample period. The available observations however suggest that there has been very little variation over time in it, so that this omission is unlikely to produce a significant bias.

Turning to our explanatory variable capturing wage dispersion, it would be desirable to use a measure of wage dispersion within skill groups. Wage differentials between skill groups are not relevant for our analysis because job search will likely be confined to jobs with similar skill requirements. Since the *Labour Force Survey* does not include data on wages over the sample period we use the dispersion of earnings across all male workers in full-time employment which includes within and between skill group variation. Machin (1996) gives evidence that the rise in British wage differentials is as pronounced within skill groups as between skill groups, and that both have evolved over time in a similar pattern. Table 5 compares the overall wage dispersion with the dispersion of the wage residual from a regression of hourly earnings on educational attainment and age variables, as estimated by Machin.

[3] Earnings related supplement is added when applicable.

Table 5. 90 to 10 Percentile Ratios of Log Hourly Earnings in the United Kingdom
(index, 1982=100)

	1982	1986	1990	1993
Gross wage	100	110.6234	117.5593	117.9104
Wage residual	100	109.9253	113.7673	120.1708

Source: Own calculations based on Machin (1996), Table 6.

To capture changes in the wage distribution we include two measures of wage dispersion: The first, capturing dispersion at the low end of the distribution, is the difference between average male full-time earnings and the 10 percentile of male full-time earnings, normalized on average earnings. The second is defined as the difference of the 90 percentile of male full-time earnings from average earnings, again normalized by average earnings. It captures wage dispersion at the high end of the wage distribution.

We next turn to identification issues in the estimation. Since changes in on-the-job search can shift the Beveridge curve, both unemployment and vacancies have to be treated as endogenous to the proportion of on-the-job searchers. We constructed the vacancy rate in a way to ensure it is predetermined.[4] Lagged unemployment variables and lagged redundancies were used as instruments for the unemployment variables. Simultaneity problems also arise for the job tenure variable: the more workers decide to search on-the-job the more frequent are short job tenures. Job tenure was therefore instrumented with lagged job tenure.

The instrument set includes variables which are predetermined but not strictly exogenous (that is, they may be correlated with the error term in previous time periods). The presence of predetermined variables makes the fixed effects estimate inconsistent if the number of time periods is small. We have therefore estimated the on-the-job search equation in first differences, instrumenting all predetermined variables, combining the Arellano and Bond estimator (see Arellano and Bond, 1991) with the Anderson and Hsiao estimator (see Hsiao, 1986). We use Arellano and Bond's instrumentation technique for the lagged dependent variable and the vacancies variable, whereas we use the standard lagged levels, following Anderson and Hsiao, for the change in vacancies, the job tenure, the redundancies and the unemployment variables. This choice reflects the lack of degrees of freedom which prevents the use of Arellano and Bond's estimator for all variables. Moreover we had to limit the

[4] This is the number of vacancies divided by employment in the two months preceding the survey reference period.

number of lags to two for the lagged dependent variable and to one for vacancies to limit the number of overidentifying restrictions.

B. The Data

The data consist of a panel of the 10 British standard regions covering the years from 1982 until 1993. Most data were taken from the *Labour Force Survey*. *Labour Force Survey* data are available biannually from 1975 till 1983 and annually from 1983 onwards. Observations for the missing year 1982 were interpolated. Values of the job tenure variables in 1992 and 1993 values were set equal to the 1991 value because of changes in the construction of this variable in 1992. To measure the skill composition of employment we divided workers into four groups: Professional and managerial, intermediate and junior nonmanual, skilled manual, and semiskilled and unskilled manual workers. A description of the data is contained in Fuentes (2002).

C. Results

Table 6 shows the results of three regressions. All three models fail to reject the presence of first order serial correlation and reject second order serial correlation in the error term, in line with the maintained hypothesis of an MA(1) error term, so the model appears to be correctly specified. The change in vacancies variable was kept despite being insignificant to ensure the correct dynamic specification of the model. The lagged dependent variable is significant throughout.

In Model 2 we eliminate the replacement ratio even though it is significant at the 10 percent level for the following reason: Theory suggests that the replacement ratio should have a negative impact on on-the-job search. The likely cause of the positive sign of the coefficient in Model 1 is that this variable is correlated with the variable measuring wage dispersion at the lower end of the wage distribution. As expected, once the replacement ratio is dropped, the coefficient on the wage dispersion variable increases.

In Model 3 we eliminate the short-term and long-term unemployment rates in the list of regressors. The results of Models 1 and 2 indicate that a large number of unemployed searchers does not discourage on-the-job search, suggesting that the unemployed do not effectively compete with employed searchers, consistent with the ranking hypothesis. However, the fact the coefficients are positive (if insignificant), and larger for long-term unemployment than for short-term unemployment, suggests that our instruments do not entirely remove the effects of reverse causality[5]. As noted above on-the-job search may increase unemployment for a given level of vacancies. In addition, on-the-job search increases the duration of unemployment spells. This may explain that the coefficient on long-term unemployment is larger than on short-term unemployment.

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[5] A Sargan test of overidentifying restrictions fails to reject the hypothesis of instrument validity, but owing to sample size limitations this test is not based on estimators robust to heteroskedasticity.

All other variables have signs consistent with theory. The results show that on-the-job search responds to the availability of jobs, as evidenced in the significant coefficient on the vacancies variable. The fear of job loss also makes on-the-job search more likely, as the positive coefficient on the redundancy variable indicates.[6]

Table 6. Generalized Method of Moments Estimation

(Sample: 120 observations over 12 years (1982-93).
Dependent variable: proportion of employed workers searching on the job)

	Model 1		Model 2		Model 3	
Variable	Coefficient	*t-Statistic*	Coefficient	*t-Statistic*	Coefficient	*t-Statistic*
Constant	0.72	2.69	0.13	0.40	0.26	1.14
On-the-job search lagged	0.24	2.50	0.27	2.72	0.31	4.17
Vacancies	0.72	2.46	0.82	2.86	0.79	2.03
Vacancies change	-0.12	-0.66	-0.11	-0.59	-0.11	-0.46
Age 16-50	0.16	2.75	0.17	2.51	0.17	2.47
Unemployed < 2years	0.08	0.95	0.08	0.93	-	-
Unemployed > 2years	0.20	1.53	0.20	1.53	-	-
Redundancies	0.25	2.94	0.17	2.15	0.17	1.92
Wage dispersion (low end)	0.13	3.39	0.17	5.22	0.19	6.24
Wage dispersion (high end)	0.03	1.10	0.04	1.32	0.04	1.68
Professional employment	0.14	3.92	0.14	3.25	0.13	3.17
Low skill manual employment	0.10	3.14	0.12	4.05	0.13	4.50
Job tenure less than 20 years	0.23	2.47	0.23	2.26	0.26	2.42
Replacement rate	0.25	1.84	-	-	-	-
LM tests on:						
First order serial correlation		-2.09		-2.14		-2.1
Second order serial correlation		-1.51		-1.18		-1.58
Overidentifying restrictions		38		37		37

Notes: Standard errors are asymptotically heteroskedasticity consistent as the number of cross section units approaches infinity. Strictly exogenous variables: Age 16–50, wage dispersion and skill composition of employment variables, replacement rate. Additional instruments, in levels: on-the-job search lagged twice and three times, vacancy level lagged, change in vacancies lagged once and twice, short-term and long-term unemployment lagged twice (Models 1 and 2 only), redundancies lagged once, twice and three times.

[6] The procyclicality of on-the-job search is mitigated, but not offset, by the positive effect of redundancies on on-the-job search, as vacancies pick-up more of the variation in the dependent variable than redundancies.

Of the wage-dispersion variables, wage dispersion at the lower end of the distribution has the stronger effect, being highly significant throughout. Taking the estimated long-term coefficients of Model 3 at face value, the increase in wage dispersion can explain most of the increase in on-the-job search over the sample period.[7] The stronger effect of wage dispersion at the lower end of the wage distribution is also consistent with our cross-section results.

A control variable for the proportion of employed workers who are less than 50 years old captured age effects on on-the-job search best and was significant at the 5 percent level. A relatively late cut-off point of 50 years is consistent with our cross-section results which indicated that early on in the work life increases in age make on-the-job search more likely. The estimated positive effect of relatively short job tenures is also in line with both the theoretical and our cross-section results.

While the positive coefficient on the proportion of professional and managerial workers is also consistent with our cross-section results, the positive coefficient on low-skill employment is not (note that the omitted category are intermediate skills), as the cross-section results indicated that low-skill workers were ceteris paribus less likely to search on-the-job (if insignificantly so). A possible explanation is that the skill composition of employment is a procyclical variable, given that low-skill workers are much more likely to lose their jobs than high-skill workers. The vacancy variable may not fully capture the effect of labor market tightness on on-the-job search, giving rise to a positive coefficient of low-skill employment in the equation.

V. Conclusions

Three conclusions emanate from this study. First, the econometric results indicate that on-the-job search is sensitive to changes in labor market tightness. Improved job opportunities, associated with a high number of vacancies, encourage on-the-job search. The results suggest that on-the-job search needs to be taken into account when explaining the behavior of the unemployment outflow rate over time. They lend support to the hypothesis that the poor correlation between hires and the unemployment outflow is caused by endogenous on-the-job search, and that the lack of cyclical variation in the unemployment outflow rate is, at least in part, due to the endogeneity of on-the-job search.

Second, there is no evidence of high unemployment deterring on-the-job search. Our results thus lend support to the view that the unemployed do not provide effective competition to on-the-job searchers. One explanation is the ranking hypothesis outlined above.

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[7] This calculation disregards the endogeneity of unemployment, vacancies, short job tenure and redundancies. Since these variables have positive estimated effects on on-the-job search, this omission results in underestimation of the contribution of wage dispersion in explaining the variation in on-the-job search.

Third, in addition to variables reflecting labor market tightness, wage dispersion as well as the age and skill composition of employment affect the labor market through on-the-job search. The importance of this result lies in the effect on-the-job search has on the locus of unemployment-vacancies relationship (the Beveridge curve). The results in Fuentes (2002) suggest that on-the-job search has shifted the Beveridge curve in the United Kingdom outwards, implying that, for any given level of vacancies, unemployment is higher. Wage dispersion is the most noteworthy in the list of the aforementioned variables, given the large increase in wage dispersion in the United Kingdom over the 1980s and the substantial estimated positive effect of wage dispersion, especially at the lower end of the wage distribution, on on-the-job search.

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