

Scale Effects in Markets with Search*

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Abstract

This paper tests for the presence of scale effects in markets with search frictions. It is theoretically plausible that scale effects exist in one or more aspects of the matching process, but may be offset by the optimal search strategies of firms and workers, and therefore are not picked up by reduced-form estimates of the matching function. We estimate a structural job search model on British data, letting both the arrival rate of job offers and the mean wage offer vary with personal worker characteristics and local market size, proxied by the number of existing job openings. We find evidence of scale effects in the mean wage offer. Such scale effects turn out to be largely offset by the endogenous response of worker reservation wages, so that larger markets tend to deliver higher realized wages rather than a faster exit rate from unemployment.

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1 Introduction

Scale effects in matching models have featured prominently in the economics literature since Peter Diamond's (1982) claim that the complementarities in search models are strong enough to generate multiple equilibria. Theoretically, scale effects appeared "plausible," both to Diamond and others. Yet, despite a small number of exceptions, empirical work has largely supported constant returns (Petrongolo and Pissarides, 2001). The exceptions are not ones that consistently apply to some cases or some periods. Rather, a few diverse estimates support increasing returns, with the vast majority supporting constant returns.¹

Empirical testing, however, usually proceeds by estimating aggregate matching functions or hazard rates for unemployed individuals. The test for constant returns in the aggregate matching functions is whether a proportional increase in the inputs of both firms and workers into search (at the most basic level, in the total number of vacancies and unemployed workers) increases total matches by a bigger proportion. In hazard-function estimation the test is the equivalent one of whether a proportional increase in local-area unemployment and vacancies increases a typical individual's hazard rate. In both cases, however, the estimation is on reduced forms. Both aggregate matching functions and individual hazard rates conceal more than one structural dimension. For example, they are both a composite of the mechanics of the meeting technology, and the willingness of firms and workers to accept the other side's offer.

Our main claim in this paper is that it is feasible for constant returns in aggregate matching functions and hazard rates to coexist with increasing returns at one of the micro levels, because the responses of firms and workers

¹Some may claim that the statement in the text is unduly strong. For example, estimates using translog matching functions (like the ones by Warren, 1996 and Yashiv, 2000) are more supportive of increasing returns than estimates using loglinear functions. Also, estimates restricted to manufacturing are more supportive of increasing returns than estimates that use whole-economy data (see Blanchard and Diamond, 1990, and again, Warren, 1996). But it would be premature to generalize from this small number of examples and claim that increasing returns are a feature of all such cases.

to the increasing returns can cancel out their effects at the aggregate level. The most clear case where this can happen is in the quality of the job match. There may be increasing returns in the quality of job matches, with better matches occurring in larger markets. But if reservation wages increase in proportion to the improvement in the quality of job matches, the aggregate matching function should be independent of scale. The increasing returns would be associated with higher post-unemployment wages but not with shorter durations of unemployment.

We outline a standard model of search and make the theoretical case for the co-existence of increasing returns at the structural level and constant returns at the aggregate level. We show that at the structural level scale effects could be observed at two levels: at the level of the arrival of job offers and at the level of the quality of the job match. If scale effects are due to faster arrival of offers, reservation wages could offset completely their effects on the aggregate hazard rate (and consequently on the aggregate matching rate). But whether they do or not depends on the properties of the distribution of job offers, as previously shown by Burdett (1981) and others. If the scale effects are in the quality of the job match, theory predicts that reservation wages change to offset their impact on the hazard rate only if the worker's unemployment income is small.

We estimate our model and look for scale effects at the structural level by making use of a British sample of 3,000 unemployed individuals. We can do this because in addition to the usual variables (personal and local labor-market characteristics at the county level, censored and uncensored durations of unemployment and post-unemployment wages) it also contains information on reservation wages. We decompose hazard functions into the probability of receiving an offer and the probability of accepting it, and estimate the influence of personal and local labor-market characteristics on each.

We find scale effects in the quality of the job match (proxied by the mean of the wage offer distribution) but not in the arrival of job offers. But we also find that reservation wages increase to offset the impact of increasing returns on the unemployment hazard rate. Because in larger markets work-

ers search with higher reservation wages, the effect of scale shows up as a higher post-unemployment wage and not as a shorter duration of unemployment. Aggregate matching functions derived from our sample would confirm constant returns.

Theory suggests that in this case reservation wages should compensate the effects of scale only if unemployment income (net of search costs) is small. We show that our estimates imply that even conventionally “high” replacement ratios, 40 percent of the mean wage rate net of search costs, imply that the scale effects in the quality of job matches should be reflected primarily on post-unemployment wages. A small effect on hazard rates remains but we argue that it is sufficiently small that reduced-form estimation is not likely to pick it up.

There are 66 counties in Great Britain, our local market, with mean employment level 322,285 people and range 6,000 to 3,515,400. As a test of robustness of our results we re-estimate the model by excluding all observations from London, about 12 per cent of our sample. London is counted as a single unit and it is an outlier in the county size distribution. Compared with London’s 3.5m, employment in Birmingham, the second largest county, is 1.1m. We find no scale effects in the smaller sample, although there is some weak evidence of scale effects in the arrival of job offers. This opens up the possibility that scale effects in the quality of matches operate only in very large markets, which offer a large choice of diverse occupations. But it can also mean that what we have estimated is a peculiarly “London effect,” and otherwise the constant returns estimated with the reduced forms pervade all levels of search markets. Further tests are needed before either constant returns or non-monotonicity in scale effects can be established as a more general property of search markets. We speculate further on these possibilities in the concluding section of the paper.

The paper is organized as follows. In Section 2 we outline an infinite-horizon search model and show the effect of changes in both the arrival rate of job offers and the mean wage offer on the job-finding rate. Section 3 describes our data set and presents some preliminary evidence. In Section 4

we specify the likelihood function and estimate the model, letting both the arrival rate of job offers and the mean wage offer depend on individual and local labor market characteristics. The results are discussed in section 5 and general conclusions brought together in section 6.

2 Model

We begin by describing the empirical problem. An unemployed individual searches for a job in a local labor market. Offers arrive randomly according to a Poisson distribution with parameter $p(\mathbf{x})$, where \mathbf{x} is a vector of personal and local labor-market characteristics. When an offer arrives, the individual has the option of accepting a wage which is randomly drawn from the known and fixed distribution $F(w)$. We assume that $F(w)$ is lognormal with mean $\mu(\mathbf{z})$, where \mathbf{z} is another vector of personal and local labor-market characteristics, and standard deviation σ_w , which is a fixed parameter. If the worker accepts the offer she leaves unemployment and earns w for the duration of the job. If she rejects the offer she waits for a new offer to arrive, and on average one does after $1/p(\mathbf{x})$ periods. The stopping rule is governed by the reservation wage w^* , which is a choice variable. The unemployment hazard rate is h , defined by:

$$h = h(\mathbf{x}, \mathbf{z}) = p(\mathbf{x})(1 - F(w^*; \mu(\mathbf{z}), \sigma_w)). \quad (1)$$

Our data contains information on w^* and unemployment durations for each individual (from which we can make inferences about h), and on a variety of personal and local labor-market characteristics, which are candidates for the vectors \mathbf{x} and \mathbf{z} . This allows us to estimate $p(\mathbf{x})$, $\mu(\mathbf{z})$ and σ_w for each individual, conditional on the distribution F and on an optimizing search model. Our primary interest is two fold. First, to identify separately whether scale effects affect the offer arrival rate or the mean of the offer distribution; formally, whether the vectors \mathbf{x} and \mathbf{z} contain variables for the size of the local labor market. Second, to compute the reaction of reservation wages to the vectors \mathbf{x} and \mathbf{z} , and from these to obtain the reduced form

hazard function. The key question is then whether the size variables can influence either p or μ , where intuition about scale effects normally applies, but not the hazard rate, because of the dependence of the latter on reservation wages.

The search model is a conventional continuous-time model for an unemployed individual with infinite horizon looking for a permanent job. The labor market environment is stationary and the unemployment hazard independent of duration. The model is partial, in the sense that both p and $F(w)$ are exogenous, neither depends on time or search duration and successive wage offers are independently distributed. A job is an absorbing state, once one is accepted search stops.

During search, unemployed individuals enjoy some flow real return b (typically including the value of leisure and unemployment insurance benefits, net of the cost of search) and discount future incomes at the instantaneous discount rate r . Under these assumptions we can use Bellman equations to derive stationary values for unemployment and employment, respectively denoted by V_n and $V_e(w)$. The Bellman equation satisfied by the value of unemployment is

$$rV_n = b + p \left\{ \int \max [V_n, V_e(w)] dF(w) - V_n \right\}. \quad (2)$$

The value of employment is given by

$$rV_e(w) = w. \quad (3)$$

Trivially, the choice between V_n and $V_e(w)$ inside the integral of (2) can be described by a reservation rule. There exists a unique reservation wage w^* such that $V_n = V_e(w^*) = w^*/r$. The reservation wage satisfies an equation derived from (2),

$$w^* = rV_n = b + p \int_{w^*} [V_e(w) - V_n] dF(w) \quad (4)$$

$$= b + \frac{p}{r} E(w - w^* | w \geq w^*) \Pr(w \geq w^*), \quad (5)$$

which simplifies to

$$w^* = \frac{1}{r+h} (rb + hy(w^*)), \quad (6)$$

with h given by (1) and $y(\cdot)$ defined by the conditional expectation (or “truncated mean function” of F ; see Burdett, 1981),

$$\begin{aligned} y(w^*) &= E(w \mid w \geq w^*) \\ &= \int_{w^*} w dF(w) / (1 - F(w^*)). \end{aligned} \quad (7)$$

Let now x be an element of the vector \mathbf{x} , a parameter that influences p but not F , and without loss of generality let $p'(x) > 0$. Differentiation of (1) with respect to x yields

$$\frac{\partial h}{\partial x} = (1 - F(w^*))p'(x) - pF'(w^*)\frac{\partial w^*}{\partial x}, \quad (8)$$

and from (6),

$$\frac{\partial w^*}{\partial x} = \frac{p'(x)}{r+h} (1 - F(w^*))(y(w^*) - w^*) > 0. \quad (9)$$

In general, $\partial h/\partial x$ is ambiguous, because the two terms in (8) have opposite sign: the worker increases his reservation wage when the arrival of offers improves and this can offset its direct positive effect on the hazard rate.

Substitution from (9) into (8) yields

$$\frac{\partial h}{\partial x} = (1 - F(w^*))p'(x) \left(1 - \frac{h}{r+h} \frac{F'(w^*)}{1 - F(w^*)} (y(w^*) - w^*) \right). \quad (10)$$

But making use again of (6) we find that

$$\frac{F'(w^*)}{1 - F(w^*)} (y(w^*) - w^*) \quad (11)$$

is the slope of the truncated mean function, denoted by $y'(w^*)$. Thus, a sufficient condition for a positive effect of the offer arrival rate on the hazard rate is that the slope of the truncated mean function be less than 1. The

condition, however, is not necessary (although it becomes necessary as $r \rightarrow 0$). Burdett (1981) argues that this property is satisfied by a large class of distributions, and the literature usually assumes that the property holds, so an improvement in the offer arrival rate decreases the mean duration of unemployment.

Consider next a parameter z that improves the wage offer distribution, i.e., let F depend on z such that $F(w; z)$ stochastically dominates $F(w; z')$ if $z > z'$. The effect of a small displacement in z on the hazard rate is

$$\frac{\partial h}{\partial z} = -p \left(\frac{\partial F(w^*)}{\partial z} + F'(w^*) \frac{\partial w^*}{\partial z} \right). \quad (12)$$

By the stochastic dominance assumption made, $\partial F(w^*)/\partial z < 0$, and from (6) it immediately follows that $\partial w^*/\partial z > 0$, so once again there is an ambiguity in the effects of an improvement in the offer distribution. However, as in (1), in our empirical work we restrict the estimation to shifts in the mean of the lognormal wage offer distribution, holding variance constant. The effect is still ambiguous but it can now be calculated and estimated.

The density of the lognormal is

$$f(w) = \frac{1}{w\sigma_w} \phi \left(\frac{\ln w - \mu}{\sigma_w} \right), \quad (13)$$

where $\phi(\cdot)$ is the normal density

$$\phi \left(\frac{\ln w - \mu}{\sigma_w} \right) = \frac{1}{(2\pi)^{0.5}} \exp \left[-\frac{1}{2} \left(\frac{\ln w - \mu}{\sigma_w} \right)^2 \right]. \quad (14)$$

The cumulative density of the lognormal is

$$F(w^*) = \Phi \left(\frac{\ln w^* - \mu}{\sigma_w} \right), \quad (15)$$

where $\Phi(\cdot)$ is the cumulative normal density. Finally, by integration we can derive

$$\int_{w^*} w dF(w) = \exp\left(\frac{1}{2}\sigma_w^2 + \mu\right) \left[1 - \Phi\left(\frac{\ln w^* - (\sigma_w^2 + \mu)}{\sigma_w}\right)\right]. \quad (16)$$

With these expressions the reservation wage equation (6) becomes

$$w^* = \frac{rb + p \exp\left(\frac{1}{2}\sigma_w^2 + \mu\right) \left[1 - \Phi\left(\frac{\ln w^* - (\sigma_w^2 + \mu)}{\sigma_w}\right)\right]}{r + p \left[1 - \Phi\left(\frac{\ln w^* - \mu}{\sigma_w}\right)\right]} \quad (17)$$

For any parameter z that influences the mean for the lognormal, (15) yields

$$\frac{\partial F(w^*)}{\partial z} = -\frac{\mu'(z)}{\sigma_w} \phi\left(\frac{\ln w^* - \mu}{\sigma_w}\right) = \frac{F'(w^*)}{w^*} \mu'(z), \quad (18)$$

so (12) becomes:

$$\frac{\partial h}{\partial z} = pF'(w^*)w^*\mu'(z) \left(1 - \frac{1}{w^*\mu'(z)} \frac{\partial w^*}{\partial z}\right). \quad (19)$$

To derive the response of the reservation wage to the change in z we differentiate (17) to derive

$$\frac{\partial w^*}{\partial z} = \frac{p \exp\left(\frac{1}{2}\sigma_w^2 + \mu\right) \left[1 - \Phi\left(\frac{\ln w^* - (\sigma_w^2 + \mu)}{\sigma_w}\right)\right]}{r + p \left[1 - \Phi\left(\frac{\ln w^* - \mu}{\sigma_w}\right)\right]} \mu'(z), \quad (20)$$

and so

$$\frac{1}{w^*\mu'(z)} \frac{\partial w^*}{\partial z} = \frac{p \exp\left(\frac{1}{2}\sigma_w^2 + \mu\right) \left[1 - \Phi\left(\frac{\ln w^* - (\sigma_w^2 + \mu)}{\sigma_w}\right)\right]}{rb + p \exp\left(\frac{1}{2}\sigma_w^2 + \mu\right) \left[1 - \Phi\left(\frac{\ln w^* - (\sigma_w^2 + \mu)}{\sigma_w}\right)\right]}. \quad (21)$$

Comparison of (21) with (19) shows that if $rb > 0$ the effect of higher z (noting that by assumption $\mu'(z) > 0$) is to raise the hazard; if $rb = 0$ the hazard is unaffected by changes in the mean of the distribution and if $rb < 0$ the reservation wage overreacts and the hazard is lower.

The empirical strategy is to use information on reservation wages and unemployment durations to uncover the dependence of the offer arrival rate and the mean of the wage offer distribution on the size of the local market and other parameters. We explain how (1) and (6) can be used to construct a likelihood function after a description of the data.

3 Data

The data used for this study come from the UK Survey of Incomes In and Out of Work (SIOW). This was a one-off survey that collected individual information on a representative sample of men and women who started a spell of unemployment, and registered at any of the 88 Unemployment Benefit Offices (UBO) selected, in the four weeks starting March 16, 1987. Information on survey participants was collected from two separate personal interviews. The first interviews were carried out shortly after unemployment began, between April and July 1987, and a total of 3003 interviews were completed. The second interviews were held about nine months later, in January 1988, on respondents who had been interviewed in 1987 and had consented to a second interview. A total of 2146 interviews were completed at this second stage. We use available information on all respondents interviewed once or twice, by assuming that attrition between the first and second interview is random.

The first interview focused on individuals' personal characteristics and their employment history during the 12 months prior to the interview, including employment and unemployment income, type of jobs held and job search activities while unemployed. The follow-up interview covered individuals' employment history since their first interview.

The data contain three types of unemployment spells. *Completed spells*, by respondents who had found jobs by the time of the first or second interview. Completed spells are measured by the number of weeks between the date the worker signed at the UBO and the date he or she re-entered employment. The longest completed spells in the sample are between nine and

ten months. *Censored spells*, by respondents still unemployed at the time of the second interview (or the first interview for those who only had one interview), measured by the number of weeks between the date the respondent registered at the UBO and the date of the interview. *Censored spells*, by respondents who left the register without finding a job and who were out of the labor force at the time of interview. This type of censored spell is measured by the number of weeks that the respondent was on the unemployment register.² We call the third type of spell as a censored spell following the logic of a competing risk duration model. Exits into jobs compete with exits into other states but given that our focus is on the determinants of the exit into jobs, all unemployment durations finishing with destinations other than jobs are treated as censored (see Narendranathan and Stewart, 1993)

In addition to data on unemployment spells, we use information on worker reservation wages and on their post-unemployment wages. The information on reservation wages comes from the question “what is the lowest weekly take-home pay you might consider accepting”, which is asked of all unemployed workers, or the question “what is the lowest weekly take-home pay you might have considered accepting”, which is asked of those already employed at the time of the first interview. We then obtain hourly reservation wages by using information on the expected number of hours to be worked each week. Post-unemployment hourly wages are constructed from a question on the usual weekly take-home pay in the first job after the unemployment spell and a question on the usual hours worked. Although for our purposes it would be more appropriate to estimate the parameters of the pre-tax wage distribution, better representing the productivity distribution across firms, we have no choice but to estimate the distribution of take-home pay, as information on the (subjective) pre-tax reservation wage is not available (and constructing a tax schedule for each individual is also not feasible).

The empirical p.d.f. of the post-unemployment (log) wage distribution is

²We assume that once a respondent leaves the unemployment register active search ceases, since once on the register, no active searcher has an incentive to leave it, even if entitlement to benefit ceases. The register for non-recipients of benefit provides free job information which the job seeker is free to use or ignore.

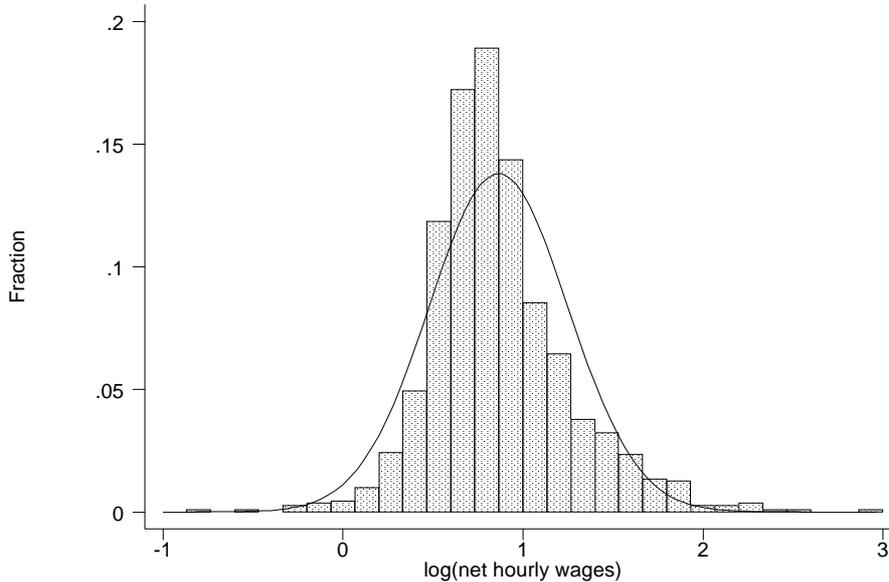


Figure 1: The empirical p.d.f. of the post-unemployment (log) net hourly wages. Source: SIIOW.

shown in Figure 1. Although the observed distribution appears more skewed than the normal, typical alternatives such as the exponential or the Pareto distributions would not fit the data better. In particular, the heavy right tail implied by the Pareto distribution is not present in the empirical wage distribution.

Using self-reported information on reservation wages involves a problem, namely that it is not guaranteed that the reservation wage falls always between net unemployment income and the post-unemployment wage, as required by our model.³ We find that self-reported reservation wages are higher than post-unemployment wages for 16% of observations in our sample. We explain in the next section how we deal with this apparent discrepancy be-

³In some cases, e.g. when having a job increases the entitlement to unemployment compensation, it may be optimal to set the reservation wage below actual unemployment income. See Mortensen (1977).

tween theory and observation. It is a lot more difficult to compare reported wages with income during unemployment. In the absence of information on the cost of search, we cannot directly compare reservation wages with *net* unemployment income. A comparison of reservation wages with reported unemployment income shows that unemployment benefits exceed reported reservation wages in only 5% of our sample.⁴

The information on hourly reservation wages is missing for 773 workers. 1445 workers in the sample found a job within the survey period, while 1558 were still jobless at the time of the second interview or had left the unemployment register. Among those who found jobs, the information on the hourly take-home pay is missing in 330 cases. The final sample consists of 2229 respondents, the missing ones being the 773 with no reservation wage information and 1 observation with no age information. The 330 cases with missing post-unemployment wage are included in the estimation by making use of the information that they still convey; that they have had an offer exceeding their reservation wage. Some summary statistics of the samples used are presented in Table 1.

As far as the characterization of local labor markets is concerned, we merge individual records from the SIIOW with county-level data extracted from the NOMIS database. For confidentiality reasons the SIIOW does not attach explicit geographic identifiers to interviewees. The only geographical information that is provided is the code of the UBO at which the worker is registered. Using NOMIS information, we achieved a mapping between UBOs and counties. Out of the 66 British counties, respondents in the SIIOW reside

⁴Further tests on the reliability of the reservation wage information in the SIIOW were carried out by Manning and Thomas (1997), who estimated both wage regressions and unemployment duration models on these data. They showed that, consistent with our search model, both post unemployment wages and unemployment duration depend positively affected on self-reported reservation wages. For more general discussion of the problems and benefits involved in the use of self-reported reservation wage data see Lancaster and Chesher (1983), who make use of two British surveys of unemployed workers (the P.E.P survey of 1973 and the Oxford survey of 1971). More recently a number of authors have used Dutch data on self-reported reservation wages, where econometric procedures are also discussed, e.g., Van den Berg and Gorter (1997), Van den Berg (1990) and Bloemen and Stancanelli (2001).

Table 1: Sample characteristics of the unemployment inflow

Variables	Whole sample			Excluding London		
	Mean	St.dev.	No.obs.	Mean	St.dev.	No.obs
% uncensored	52.7		2229	53.1		1962
% censored	47.3		2229	46.9		1962
uncensored duration	12.3	11.1	1174	12.2	11.0	1042
censored duration	23.9	17.4	1055	24.1	17.5	920
females	37.7		2229	37.2		1962
age	36.9	11.5	2229	36.8	11.5	1962
% skilled	43.4		2229	44.4		1962
hourly res. wage	2.38	0.98	2229	2.33	0.94	1962
hourly take-home pay	2.57	1.30	927	2.56	1.30	839

Notes. *% Skilled*: includes all those who attended school or vocational training courses until the age of 18, plus those with higher education. *Hourly res. wage*: denotes the lowest weekly take-home pay that the worker considers accepting, divided by the expected number of hours worked. Source: SIIOW.

in 43 of them. The county level information that we use in our estimates is reported in Table 2.

A preliminary picture of the relationship between market size and the quality of job matches can be gathered by regressing county mean wages on market size. We compute mean wages by 2 educational groups and 43 counties, and regress them on an education dummy, the local labor market tightness (denoted by θ), and the number of vacancies (also disaggregated into 2 occupational groups, denoted by $V_{by\ skill}$) in each county. The results are reported in Table 3. Local wages are positively correlated with the number of job openings, proxying market size, although the size effect is only significant for the whole sample.

We repeat the same exercise using mean completed unemployment duration (by county) as the dependent variable. The resulting estimates are reported in Table 4. In neither specification can any size effect be detected.

Table 2: Local labor markets in Britain

Variables	Whole sample		Excluding London	
	Mean	St.dev.	Mean	St.dev.
Unemployed	59986	68698	52380	47817
Vacancies	4140	5639	3349	2229
Tightness	0.08	0.03	0.08	0.03
Skilled vacancies	1072	1614	846	648
Unskilled vacancies	2710	3397	2242	1484
Firms	26118	30463	21867	12432
Average firm size	15.8	4.3	15.7	4.3
Area size (acres)	347258	431079	351767	435277
No. of observations	43		42	

Notes. *Unemployed*: number of claimant unemployed, April 1987. *Vacancies*: vacancies notified at Job Centres, April 1987. *Tightness*: vacancies/unemployed. *Skilled vacancies*: vacancies notified at Job Centres, March 1987, in the following KOS occupations: managerial; professional: supporting; professional (education, welfare); literary, artistic, sports; professional (science, engineering); managerial (excluding general); clerical and related. *Unskilled vacancies*: vacancies notified at Job Centres, March 1987, in the following KOS occupations: selling; security and protective; catering, cleaning, etc.; farming, fishing and related; processing (excl. metal); making/repairing; processing (metal./elect.); repetitive assembling, etc.; construction, mining; transport operating; miscellaneous. *Firms*: stock of VAT registered businesses at the end of 1986. *Average firm size*: employment/firms. *Area size*: area of county, in acres. Source: NOMIS.

Table 3: Mean wages and labor market size

	Whole sample	Excluding London
constant	0.237 (0.166)	0.377 (0.194)
skilled	0.247 (0.033)	0.233 (0.038)
$\log(\theta)$	-0.099 (0.035)	-0.110 (0.036)
$\log(V_{by\ skill})$	0.042 (0.015)	0.018 (0.023)
R^2	0.45	0.45
No. Obs.	85	83

Notes. The dependent variable is the (log) mean wage across 2 educational groups and 43 counties. Source: SIIOW and NOMIS.

Table 4: Mean durations and labor market size

	Whole sample	Excluding London
constant	1.873 (0.310)	2.143 (0.404)
skilled	-0.169 (0.062)	-0.192 (0.072)
$\log(\theta)$	-0.168 (0.066)	-0.186 (0.067)
$\log(V_{by\ skill})$	0.021 (0.028)	-0.023 (0.043)
R^2	0.17	0.16
No. Obs.	85	83

Notes. The dependent variable is the (log) mean completed unemployment duration across 2 educational groups and 43 counties. Source: SIIOW and NOMIS.

4 Estimation specification

The likelihood contribution of an individual with an unemployment spell length of d_i , and, in the case the spell is completed, a wage w_i is

$$\begin{aligned} L_i &= \exp[-p \Pr(w \geq w^*) d_i] [p \Pr(w \geq w^* | w_i) \Pr(w_i)]^{c_i} \\ &= \exp[-p \Pr(w \geq w^*) d_i] [pf(w_i)]^{c_i}, \end{aligned} \quad (22)$$

where c_i is a censoring indicator that takes value 1 if the unemployment spell is completed and 0 otherwise (we ignore for the moment workers with completed spells but missing post-unemployment wage). Under the log-normality assumptions, (22) becomes

$$L_i = \exp \left\{ -p \left[1 - \Phi \left(\frac{\ln w^* - \mu}{\sigma_w} \right) \right] d_i \right\} \left[p \frac{1}{w_i \sigma_w} \phi \left(\frac{\ln w_i - \mu}{\sigma_w} \right) \right]^{c_i}. \quad (23)$$

The parameters of the model can be estimated by maximizing the log likelihood of a sample of n observations, $\log L = \sum_{i=1}^n \log L_i$, with L_i given by (23), with respect to p, w^*, μ and σ_w , under the restriction imposed by (17) and $w^* > 0$. The availability of data on reservation wages in our data set avoids a problem often encountered by studies that have to estimate the reservation wage. Flinn and Heckman (1982) show that if observed wages are measured without error, the maximum likelihood estimator for w^* is the minimum accepted wage \underline{w} . But this method implies that the reservation wage can not be greater than any observed wage in the sample, so the presence of outliers in the observed wage distribution disproportionately affect the results, by attributing the distance between the observed wage and the reservation wage to unobservable or chance events.

When we use reported reservation wage data for w^* , it is no longer guaranteed that realized wages always exceed reservation wages. In the context of the empirical model an observation with $w < w^*$ has a zero likelihood, as the distribution of realized wages should be truncated from below at the reservation wage. But the inconsistency between theory and observation arises

only if both reservation wages and post-unemployment wages are measured without error. We generalize the empirical model by assuming that post-unemployment wages are measured with error, i.e., we let $\ln w^0 = \ln w + u$, where w denotes the wage offer received by the worker and w^0 our observation of the wage. The measurement error u is assumed to be normally distributed with 0 mean and variance σ_u^2 , and independent of w . Therefore observed wages w^0 are log-normally distributed, with mean μ and variance $\sigma^2 = \sigma_w^2 + \sigma_u^2$.⁵ Under these assumptions, the probability of receiving an acceptable offer remains $1 - \Phi[(\ln w^* - \mu)/\sigma_w]$, and the joint probability that the wage exceeds the reservation wage and that w^0 is observed is

$$\Pr(w \geq w^* | w^0) \Pr(w^0) = \left[1 - \Phi \left(\frac{\ln w^* - \rho^2 \ln w^0 - (1 - \rho^2)\mu}{\sigma_u} \right) \right] p \frac{1}{w^0 \sigma} \phi \left(\frac{\ln w^0 - \mu}{\sigma} \right), \quad (24)$$

where $\rho^2 = \sigma_w^2/\sigma^2$ represents the share of observed wage variation which is not explained by the measurement error. The resulting likelihood is

$$L_i = \exp \left\{ -p \left[1 - \Phi \left(\frac{\ln w^* - \mu}{\sigma_w} \right) \right] d_i \right\} \left[1 - \Phi \left(\frac{\ln w^* - \rho^2 \ln w^0 - (1 - \rho^2)\mu}{\sigma_u} \right) \right]^{c_i} \left[p \frac{1}{w^0 \sigma} \phi \left(\frac{\ln w^0 - \mu}{\sigma} \right) \right]^{c_i}. \quad (25)$$

Finally we need to allow for the existence of respondents who complete an unemployment spell but do not provide information on their post-unemployment wage. The information that is conveyed by these observations is that they have had an offer exceeding their reservation wage, so, taking this into ac-

⁵The assumption that wages are measured with error is used in the estimation of structural search models by Wolpin (1987), Christensen and Kiefer (1994) and Eckstein and Wolpin (1995). An alternative to introducing measurement error in wages is to assume that the utility derived from jobs is determined by the wage and some non-monetary attributes, i.e. utility = $\log w + u$, where u is the non-monetary component to the utility from the job, normally distributed with mean 0 and variance σ_u^2 . This latter approach, adopted by Manning and Thomas (1997) for estimating the predictions of a shirking model à la Shapiro and Stiglitz (1984), delivers a very similar likelihood function to the one estimated here.

count, our likelihood function generalizes to

$$L_i = \exp \left\{ -p \left[1 - \Phi \left(\frac{\ln w^* - \mu}{\sigma_w} \right) \right] d_i \right\} \left\{ p \left[1 - \Phi \left(\frac{\ln w^* - \mu}{\sigma_w} \right) \right] \right\}^{c_i} \left[1 - \Phi \left(\frac{\ln w^* - \rho^2 \ln w^0 - (1 - \rho^2)\mu}{\sigma_u} \right) \right]^{\tilde{c}_i} \left[p \frac{1}{w^0 \sigma} \phi \left(\frac{\ln w^0 - \mu}{\sigma} \right) \right]^{\tilde{c}_i}, \quad (26)$$

where c_i is equal to one for all completed spells and zero otherwise, and \tilde{c}_i is equal to one for all completed spells with a non-missing wage and zero otherwise.

Equation (26) is maximized with respect to p, μ, σ_w and σ_u . Note that in order to deliver both reservation wage and realized wage heterogeneity the model needs to allow for individual heterogeneity in at least one of the parameters $p, \mu, \sigma_w, \sigma_u$. We introduce heterogeneity in both p and μ , as explained in the next Section.

Data on both unemployment duration and post-unemployment wages allow us to separately identify the effect of variables included in p, μ or both (see Flinn and Heckman, 1982, and Wolpin, 1987, for detailed discussions of identification issues in stationary search models). In practice, however, identification may turn out to be a delicate issue when the same covariates are included in the specification of both p and μ , because of missing information on post-unemployment wages due to censoring or non-reporting. With this caveat in mind, we present alternative specifications for p and μ as a check of the robustness of our estimates.

5 Results

The estimates presented here are based on the likelihood function (26), in which σ_w and σ_u are estimated as constant parameters, and p and μ are functions of both individual and labor market characteristics.⁶ Either theory

⁶We attempted to include scale effects in the variance of the wage offer distribution, σ_w , but our estimation programme did not achieve convergence. Note, however, that under the log-normal assumptions, the variance of wages depends positively on the mean log wage, i.e. $Var(w) = \exp(2\mu + \sigma_w^2)[\exp(\sigma_w^2) - 1]$. If there are scale effects in the log of the mean wage offer, these also show up in the dispersion of the level of wages.

or well-established empirical regularities help determine which labor market variables should affect p and which μ . Search theory predicts that the arrival rate of job offers should depend on labor market tightness $\theta = V/U$, which is therefore included in the determination of p . A well known stylized fact is the employer size-wage effect, according to which large firms pay higher wages than smaller firms.⁷ As we cannot track down individual information on employer size, we capture the size-wage effect by including the local average firm size in the determination of μ . We estimate the effect of market size with four alternative measures, the number of vacancies by broad skill category, the total number of vacancies, employment, and the number of firms. Our specification of p and μ is

$$\begin{aligned} p &= \exp(\alpha_0 + \alpha_1 \text{female} + \alpha_2 \text{skilled} + \alpha_3 \log \text{age} + \alpha_4 \log \theta + \alpha_5 \log \text{size}); \\ \mu &= \beta_0 + \beta_1 \text{female} + \beta_2 \text{skilled} + \beta_3 \log \text{age} + \beta_4 \log \text{firm size} + \beta_5 \log \text{size}. \end{aligned}$$

Note that our estimated model is only semi-structural in the sense that no structural model is imposed to specify p and μ . In our estimation strategy, we tried in fact to rely as much as possible on observables, using a structural approach only when crucial for addressing the main point of this work, namely the presence of scale effects in different stages of the job search process. In our specification, we restrict the arrival rate of job offers to be non negative, and its log-linear relationship with market tightness bears close resemblance with most existing matching function estimates. Wage offers are specified as log-linear functions of human capital variables, as it is typically the case in Mincerian wage equations, to which we add size controls.

The results are presented in Table 5. In column 1 we do not include size indicators in either p or μ , and we find a fairly familiar picture of the determinants of the arrival rate of job offers and wage distributions. Men, the highly educated and older workers sample wage offers from a distribution that stochastically dominates that for women, the less skilled and the young, respectively. Markets in which the average firm size is larger are associated

⁷The literature is surveyed by Brown and Medoff (1989).

Table 5: Estimation results - Whole sample

Variables	1	2	3	4	5	6
μ						
constant	-0.378 (0.198)	-0.379 (0.199)	-0.613 (0.186)	-0.582 (0.190)	-0.374 (0.193)	-0.307 (0.469)
female	-0.256 (0.032)	-0.256 (0.032)	-0.259 (0.031)	-0.259 (0.031)	-0.256 (0.032)	-0.259 (0.031)
skilled	0.201 (0.037)	0.200 (0.037)	0.248 (0.037)	0.245 (0.035)	0.200 (0.036)	0.248 (0.036)
log(age)	0.386 (0.042)	0.387 (0.042)	0.391 (0.041)	0.391 (0.041)	0.387 (0.042)	0.390 (0.042)
log(firmsize)	0.057 (0.047)	0.056 (0.047)	0.010 (0.048)	0.008 (0.049)	0.054 (0.047)	0.021 (0.065)
log($V_{by\ skill}$)			0.043 (0.016)	0.040 (0.015)		0.040 (0.014)
log(acres)						-0.015 (0.021)
p						
constant	-1.732 (0.446)	-1.531 (0.464)	-1.381 (0.567)	-1.638 (0.473)	-1.618 (0.560)	-1.656 (0.470)
female	0.009 (0.077)	0.012 (0.079)	0.012 (0.078)	0.013 (0.077)	0.012 (0.077)	0.014 (0.076)
skilled	0.484 (0.085)	0.899 (0.181)	0.872 (0.181)	0.868 (0.185)	0.896 (0.180)	0.865 (0.185)
log(age)	-0.273 (0.102)	-0.268 (0.103)	-0.269 (0.102)	-0.267 (0.103)	-0.267 (0.103)	-0.265 (0.102)
log(θ)	0.293 (0.073)	0.305 (0.074)	0.284 (0.075)	0.273 (0.076)	0.301 (0.076)	0.268 (0.075)
log($V_{by\ skill}/V$)		0.431 (0.177)	0.432 (0.178)	0.400 (0.181)	0.420 (0.179)	0.399 (0.180)
log($V_{by\ skill}$)			-0.026 (0.042)		0.009 (0.039)	
σ_w	0.368 (0.016)	0.368 (0.016)	0.367 (0.016)	0.367 (0.016)	0.368 (0.016)	0.367 (0.016)
σ_u	0.301 (0.016)	0.301 (0.016)	0.297 (0.014)	0.297 (0.014)	0.301 (0.016)	0.297 (0.014)
log(lik)	-7120.5	-7115.9	-7109.9	-7110.2	-7115.8	-7109.8

Notes. Robust standard errors (for clustered data) reported in brackets. No of observations: 2229. Source: SIIOW and NOMIS.

with higher wage offers on average, although this effect is not significantly different from zero. Arrival rates of job offers are higher for the highly educated and for younger workers. In line with much of the matching-function literature, job offers positively depend on labor market tightness, and the elasticity of p with respect to θ , close to 0.3, is comparable with the results obtained by several estimates based on aggregate British data.

It may be argued that the relevant tightness measure is not the aggregate one, simply computed as the number of total vacancies to total unemployment in the local market, but one which is skill-specific. In the presence of market segmentation, with skilled and unskilled workers applying to different sets of jobs, the relevant tightness measure for a given worker should be given by the vacancy/unemployment ratio in the relevant skill segment. Although we have data on vacancies disaggregated by occupation, data on unemployed workers disaggregated by skills are not available at the county level. We therefore tried to pick the effect of tightness by skill by including a measure of relative tightness in p , given by $V_{by\ skill}/V$. This variable is included in column 2 and is highly significant. As our measure of market size is based on the number of vacancies in each worker's skill segment, controlling for relative market tightness is crucial if we are to attain a consistent estimate of the size effect.

The effect of market size on arrival rates and mean wage offers is obtained from the estimates in columns 3-5. Column 3 includes the number of vacancies among the determinants of both p and μ . Vacancies here are disaggregated into two broad occupational groups, skilled and unskilled (see notes to Table 2). We find that local labor market size has a positive effect on the mean wage offer distribution, but not on the arrival rate of job offers. In columns 4 and 5 we test for the effect of vacancies on p and μ separately. The effect of size on μ stays positive and practically unchanged from column 3, while the one on p is virtually zero.

We further investigate whether the effects that we estimated are not due to the absolute size of the local market but to its density. In column 6 we drop the size effect from p , and include both the number of vacancies and

the geographical size of the local market in μ . If density matters, we expect a negative and significant coefficient on $\log(\text{acres})$, once size is accounted for by $\log(V_{by\ skill})$. If *only* density matters, as opposed to size, the coefficients on $\log(V_{by\ skill})$ and $\log(\text{acres})$ should not differ from each other in absolute value. We find that the effect of $\log(V_{by\ skill})$ on μ remains largely unchanged from the one in column 4, and that the one on $\log(\text{acres})$ is negative, but not significantly different from zero. It should also be noted that the coefficients on $\log(V_{by\ skill})$ and $\log(\text{acres})$ do not differ significantly from each other in absolute value (with a p -value of 0.19). But we do not consider this to be evidence that density matters more than size because of the high standard error on the coefficient on $\log(\text{acres})$, which admits a large range of parameters not significantly different from it. As a final check, we included size and density separately, proxied by $\log(V_{by\ skill})$ and $\log(V_{by\ skill}/\text{acres})$ respectively (results not reported). Although neither of them was significant at conventional levels, the size effect was more important than the density effect, with t -statistics of 1.3 and 0.9 respectively.⁸

Given that we have data on a number of county level indicators, we also measured size effects by making use of the alternative variables of total number of vacancies (not disaggregated by occupation), the number of registered businesses and the employment level. The total number of vacancies was significant but the results were very similar to those reported in Table 5. The other variables gave imprecise estimates.

We noted that London is an outlier in our cross-section of counties. In order to check the robustness of the estimated size effect we perform the same set of estimates in Table 5 on a sub-sample which excludes London. The results obtained are reported in Table 6. When we do not include any

⁸Size and density effects in economic activity have been previously studied by Ciccone and Hall (1996), who estimate the effect of both county size (proxied by output) and county density (proxied by output per acre) on output per worker in the United States. Our study differs in the measurement of the variables of interest, but also in the results, as Ciccone and Hall find that density effects are (slightly) more important than size effects. Density effects were found to be significant by Coles and Smith (1996) in the estimation of a matching function for travel-to-work areas in England and Wales.

Table 6: Estimation results - Excluding London

Variables	1	2	3	4	5
μ					
constant	-0.386 (0.200)	-0.388 (0.199)	-0.329 (0.227)	-0.348 (0.232)	-0.384 (0.197)
female	-0.278 (0.032)	-0.277 (0.032)	-0.277 (0.032)	-0.277 (0.032)	-0.277 (0.032)
skilled	0.202 (0.028)	0.202 (0.028)	0.193 (0.031)	0.195 (0.031)	0.202 (0.032)
log(age)	0.402 (0.042)	0.404 (0.042)	0.403 (0.042)	0.403 (0.043)	0.404 (0.042)
log(firmsize)	0.032 (0.043)	0.031 (0.043)	0.036 (0.045)	0.036 (0.043)	0.029 (0.042)
log($V_{by\ skill}$)			-0.009 (0.017)	-0.007 (0.017)	
log(firms)					
p					
constant	-1.733 (0.486)	-1.488 (0.495)	-1.660 (0.668)	-1.482 (0.494)	-1.612 (0.653)
female	0.057 (0.078)	0.061 (0.079)	0.061 (0.079)	0.061 (0.079)	0.060 (0.080)
skilled	0.509 (0.090)	0.989 (0.185)	0.994 (0.186)	0.994 (0.186)	0.988 (0.185)
log(age)	-0.291 (0.113)	-0.287 (0.114)	-0.286 (0.114)	-0.287 (0.114)	-0.286 (0.114)
log(θ)	0.276 (0.079)	0.294 (0.079)	0.298 (0.079)	0.297 (0.079)	0.296 (0.079)
log($V_{by\ skill}/V$)		0.487 (0.178)	0.470 (0.190)	0.493 (0.179)	0.470 (0.190)
log($V_{by\ skill}$)			0.022 (0.057)		0.015 (0.054)
σ_w	0.354 (0.015)	0.353 (0.015)	0.353 (0.015)	0.353 (0.015)	0.353 (0.015)
σ_u	0.283 (0.014)	0.283 (0.014)	0.283 (0.014)	0.283 (0.014)	0.283 (0.014)
log(lik)	-6287.4	-6281.7	-6281.5	-6281.6	-6281.6

Notes. Robust standard errors (for clustered data) reported in brackets. No of observations: 1962. Source: SIIOW and NOMIS.

size indicator in either p or μ (columns 1 and 2), the results are similar to those obtained on the whole sample. But when we include vacancies (disaggregated by occupation) as a proxy for market size, we do not find any size effect in matching rates, coming either through the mean wage offer or the arrival rate of job offers.

Our estimates would seem to imply that there are scale effects in wage offers, which are largely absorbed by reservation wages, but they show up only when there are large differences in size across markets.

Making use of estimates from the regression 4 of Table 5, which is our preferred specification, we compute the predicted arrival rates and mean wage offers for markets of different sizes. Also, with these estimates we compute the reservation wage that is implied by the optimal search strategy that we used in our estimation, as given by equation (17). Our predictions are computed setting $r = 0.005$ for the weekly discount rate,⁹ for two alternative values of b ($b = 0$ and $b = 40\%$ of the average wage). With data on predicted reservation wages, we compute acceptance rates, hazard rates and realized wages for the average market and for the largest market in our sample. The results are reported in Table 7.

The table shows that, when $b = 0$, moving from the average to the largest market size raises the mean wage offer by 9.2%. As predicted by the model of Section 2 for $rb = 0$, the consequent increase in reservation wages completely offsets any effect of better job offers on the re-employment hazard. Higher job offers are simply translated into an equiproportional increase in realized wages. When b is equal to 40% of the average wage, and therefore $rb > 0$, higher job offers translate into a 8.8% increase in realized wages and a 1.9% increase in the re-employment hazard. Noting that b is measur-

⁹This implies an annual rate of about 30 per cent. In our simple model, the discount factor is the interest rate, however in models with limited job durations it is the sum of the interest rate and the job separation rate. New jobs last about five years in the UK, but because this group of workers is less skilled durations may even be shorter. So an annual job separation rate for these workers of 20 to 25 per cent is reasonable. In case 0.005 is regarded as too high, we note that the smaller the weekly discount rate that we use, the more support there is for the points made in the text that follows.

ing unemployment income that has to be given up when moving to a job, net of search costs, a number such as 40% is high and above the average replacement ratio for the UK in the late 1980s. Yet, the split between a post-unemployment wage effect and a duration effect of scale is firmly in favor of the post-unemployment wage effect.

The split of the effects of scale in favor of post-unemployment wages may explain why scale effects do not show up in matching function estimation, or indeed in hazard rate estimation. At reasonable benefit replacement ratios, net of search costs, the effect of scale on the hazard is too small, relative to the observed cross-sectional variations in hazard rates, to be picked up in reduced-form estimates. The effect on wages, however, is larger and should be picked up in reduced-form estimates of regional wages (see Glaeser and Maré 2001 for the US and Combes et al. 2002 for France).

6 Conclusions

In this paper we argued that the fact that the vast majority of empirical estimates find that there are no scale effects in aggregate matching functions does not necessarily mean that they are not present at the micro level. We have shown that scale effects in the quality of matches or in the arrival rate of offers can coexist with constant returns at the aggregate level. Specifically, workers raise their reservation wages in markets characterized by scale effects so as to offset their impact on the aggregate matching function. Because of this, the impact of scale effects is primarily on the mean level of accepted wages, and this should be picked up in wage regressions. We have not done these tests ourselves but our findings are consistent with the empirical literature that finds local size effects on wages and labor productivity.

Our results should be qualified by noting that scale effects disappear in a subsample which excludes London (representing 12% of observations). One possibility is that London is an outlier, namely that the effects that we estimated are a peculiarly London phenomenon. But our estimates admit also the interpretation that scale effects in the quality of job matches emerge only

Table 7: Comparative statics for the effect of market size

Variables	$r = 0.005$ $b = 0$		$r = 0.005$ $b = 0.4E(w w > w^*)$	
	Local market:		Local market:	
	Average size	London	Average size	London
Mean wage offer	3.773 (0.770)	4.121 (0.849)	3.773 (0.770)	4.121 (0.849)
Arrival rate	0.044 (0.020)	0.044 (0.020)	0.044 (0.020)	0.044 (0.020)
Acceptance rate	0.485 (0.087)	0.485 (0.087)	0.395 (0.073)	0.403 (0.074)
Hazard rate	0.020 (0.006)	0.020 (0.006)	0.016 (0.007)	0.017 (0.008)
Realized wage	4.706 (1.024)	5.141 (1.131)	4.905 (1.018)	5.338 (1.125)

Notes. All predictions are based on the estimates of column 4 in Table 5. The average market size is calibrated using the average number vacancies across counties (1072 skilled and 2710 unskilled vacancies). The size of London is calibrated using the local number of vacancies (10559 skilled and 22335 unskilled vacancies). Arrival, acceptance and hazard rates are values per week. Mean wage offers and realized wages are values per hour. Standard errors are reported in brackets.

in very large markets, where choice is really superior to the choice available in smaller markets. More specifically, comparing say a county of 1 million employed people with one of 0.2 million (say Birmingham and Southampton) we may not be able to find that the bigger choice of jobs available in Birmingham really makes much of a difference to the quality of job matches between it and Southampton. But in a city that supports employment of 3 million people (in reality even more within travelling distance) the available choice is more likely to be rich enough to accommodate specialist talents and push up average wages. More research on different data sets and countries is needed here to uncover the true causes of scale effects, if indeed they exist.

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