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EMPLOYED AND UNEMPLOYED SEARCH: THE MARGINAL WILLINGNESS TO PAY FOR ATTRIBUTES IN LITHUANIA, THE US AND THE NETHERLANDS

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Employed and Unemployed Search:
The Marginal Willingness To Pay
For Attributes in Lithuania, the US and the Netherlands

Abstract. This paper introduces a method for estimating workers' marginal willingness to pay for job attributes employing data on job search activity. Worker's willingness to pay to avoid a temporary contract, which increases the risk of becoming unemployed, is derived for Lithuania. The empirical relevance of this method is further shown re-interpreting studies that examine search behaviour in the U.S. and the Netherlands. We provide estimates of workers' willingness to pay for a wide range of job attributes including the risk of becoming unemployed and promotion prospects. Further, we discuss and apply a method for estimating unemployed individuals' willingness to pay for recall opportunities and the residual entitlement period. JEL: J3, J6.

1. INTRODUCTION

Since the eighteenth century, when Adam Smith wrote "The Wealth of Nations", economists have been interested in how the theory of compensating wage differentials might explain the existence of wage differences in the labour market. One of the attractions of this theory is that it allows for the estimation of workers' marginal willingness to pay (MWP) for job attributes such as collective bargaining and the risk of becoming unemployed. This may explain the impressive number of empirical hedonic wage studies that have focused on the workers' willingness to pay for attributes. Although many studies have shown that non-wage differences between jobs can be significant to workers, the general conclusion is that non-wage differences between jobs are not very important to workers (see Brown, 1980; Rosen, 1986).

The theory of compensating wage differentials assumes that workers have complete information in a static environment. This suggests that if job outcomes are a result of a dynamic process and workers having to search for jobs, estimates for the willingness to pay for job attributes may be biased (Epple, 1987). These considerations have encouraged theoretical research that looks at the willingness to pay for attributes. In particular, it has been demonstrated

that the estimates of the conventional marginal willingness to pay for a job attribute are likely to be biased downwards if it is not acknowledged that a job is a search good and a result of a match between an employer and a worker (Hwang et al., 1992). For example, if firms differ with respect to the cost of providing nonwage job attributes, then low cost firms offer both higher wages and greater values of desirable job attributes, because they face greater opportunity costs in having job vacancies go unfilled. This example may be particularly relevant in the context of the risk of unemployment, because more profitable firms are less likely to make employees redundant.

These considerations have generated a number of studies aimed at estimating the MWP for job attributes using data on *job moving* behaviour and comparing the MWP estimates with conventional estimates (Herzog and Schlottmann, 1990; Gronberg and Reed, 1994; Van Ommeren et al., 2000).[1] These studies point to considerably higher estimates than those based on conventional hedonic wage methods. Herzog and Schlottmann (1990) and Gronberg and Reed (1994) reported higher estimates for the willingness to pay to avoid job-induced risk. Van Ommeren et al. (2000) found higher estimates for the willingness to pay to avoid commuting. Similarly, Bartik et al. (1992) compared the MWP for residential characteristics based on *residential* moving behaviour and hedonic price methods and showed that the MWP estimates for crime reduction and school quality are higher than those based on conventional estimates. In addition, McCue and Reed (1996) examined self-reported data on the workers' willingness to pay for job attributes, and concluded that "workers' valuations of nonpecuniary dimensions of work are substantially larger than previous research has indicated".

Given the frequent use of hedonic-based models to assess the benefits of environmental, health and safety regulations in the labour and housing market, these results are relevant for theoretical and applied research and policy makers. "Hedonic-based benefit estimates shou1d be used with caution, and other benefit estimation approaches should receive greater emphasis." (Bartik et al., 1992).

In this paper, we develop a method to estimate the MWP for job attributes that explicitly acknowledges that jobs are search goods. Assuming initially an elementary stationary environment in which workers search for jobs, we demonstrate that the workers' MWP for job attributes can be derived from data on *job search activity*. This estimation method is conceptually related to studies in which MWP estimates are derived from data on *job moving behaviour* by

application of search theory (Gronberg and Reed, 1994; Van Ommeren et al., 2000).

We relax the assumptions regarding the search environment. We allow workers to search in a nonstationary environment and follow the literature by presuming that workers are involuntarily separated due to firm firings (see also Gronberg and Reed, 1994). Importantly, we extend the literature by assuming that the separation rate may depend on job attributes. This extension is essential, because the current literature explicitly assumes that the *separation rate is exogenous of job attributes*. Gronberg and Reed (1994, p. 913) state "The assumption that the separation rate is exogenous of firm wage and nonwage characteristics is crucial for empirically identifying workers' marginal willingness to pay for job attributes." Khanker (1988) employs the identical exogeneity assumption in his work on compensating wage differentials. The exogeneity assumption is clearly not innocuous (Hamermesh et al., 1994) and limits the applicability of the MWP method. Moreover, by allowing the separation rate to depend on job attributes, we are able to generate estimates of the MWP for job attributes that are related to the unemployment risk (e.g. duration of employment contract).[2]

The method developed here is used primarily to estimate the workers' willingness to pay for job attributes. However we will also discuss the possibilities of applying the same method to the estimation of the *unemployed* individuals' willingness to pay for unemployment attributes. In the current paper, we will provide estimates of the unemployed individuals' willingness to pay for the expectation of being recalled from layoff and for the residual entitlement period of receiving unemployment benefit.

Our estimation method, which is based on observations of job search, will be applied to estimate the workers' MWP to avoid temporary contracts and other job attributes in Lithuania. Further, we will re-interpret four previously published studies to calculate the MWP for job attributes and unemployment attributes in the Netherlands and USA.

The outline of the paper is as follows. In section two we introduce an elementary search model. We then derive the optimal search strategy in section three and derive the workers' marginal willingness to pay for non-wage job attributes. In section four, we generalise the search

[1] In addition, these considerations have generated a number of studies to improve conventional estimates by correcting for mobility bias (see, for example, Kim, 1992).

[2] In the literature on the theory of compensating wage differentials there is a large interest in MWP estimates of the risk of becoming unemployed (Rosen, 1986). Compensating wage differentials estimation methods are plagued by the endogeneity of job riskiness (Garen, 1988; Moretti, 2000). As is well known, search theory is particularly well suited to the analysis of the effect of risk on labour market behaviour (Mortensen, 1986).

model to increase the empirical relevance of the estimation method. Section five pays special attention to the unemployed individuals' MWP for unemployment attributes. In section six, the estimation method for the MWP is discussed. The empirical relevance of the method to estimate the individuals MWP for attributes is then demonstrated in section seven. Section eight concludes the paper.

2. THE ELEMENTARY SEARCH MODEL

The point of departure in this paper is an employed individual. This individual derives utility from job attributes X. v(X) is the quasi-concave instantaneous utility function associated with a job having attributes X. The once-only loss in utility due to moving job equals c. The person searches in the labour market with effort s at a cost of k(s), $s \geq 0$. Search costs k(s) are increasing and convex in search effort s, hence $k'(s) > 0$ and $k''(s) > 0$. Jobs arrive with arrival rate p(s). The job arrival rate p is increasing and concave in s, hence $p'(s) > 0$ and $p''(s) < 0$. We suppose that the effects of the search costs on the instantaneous utility function are additive, hence $v(X,s) = v(X) - k(s)$. Job attributes offers are drawn randomly from a given distribution, which is independent of X.[3] X_0 denotes the attributes of the job offered to the job searcher. Pooling of offers is not allowed: job offers are either refused or accepted before other offers arrive.

The expected lifetime utility received from the current job is denoted as V(X). Future utility is discounted at rate ρ. V includes the possibility of offers in the future. The individual is assumed to maximise lifetime utility V. The decision whether to accept a job offer accounts for expected future offers. Discounted lifetime utility can then be written as the sum of the instantaneous utility and the expected benefit of accepting a job offer during the next time unit (we assume that workers live forever). This leads to the following equation:

$$rV(X) = v(X) - k(s) + p(s)E\max[V(X_0) - c - V(X), 0]. \qquad \textbf{(1)}$$

In this expression the expectation is taken with respect to the distribution of the job attributes X_0. The interpretation of the above formula is well known. Note that at rate p(s) a job offer will be received, and that offer will be accepted if the value of the new job exceeds that of the current

position plus the moving costs. Hence, the optimal acceptance strategy is to accept a job offer if $V(X_0)$ -c -V(X) > 0. The offer should otherwise be rejected. In the case that job moving costs c are zero, the optimal acceptance strategy can be simplified: accept a job offer if $v(X_0)$ -v(X) > 0, otherwise reject the offer.

3. THE MARGINAL WILLINGNESS TO PAY FOR JOB ATTRIBUTES

In this section, the choice of search effort s is derived using the first-order condition for the worker's optimal search effort. The optimal choice of s is obtained by differentiating equation (1) with respect to s, and setting the resultant to zero:

$$-\frac{\partial k}{\partial s}+\frac{\partial p}{\partial s}E\max[V(X_0)-c-V(X),0]=0 \qquad if\ s>0. \tag{2}$$

The interpretation of equation (2) is well known (Mortensen, 1986). The marginal search costs equals the marginal benefit of an increase in the job arrival rate. The second-order condition is that the left-hand side of equation (2) is decreasing in s. The concavity of p and the convexity of k in their arguments ensure that this condition will be satisfied. For nearly all workers, optimal search effort s will be positive, because the marginal search costs will be close to zero for the first unit of search and the marginal benefit from the first unit will be large. So, following Hey and McKenna (1979) and Van den Berg (1992), we will assume that workers are always involved in on-the-job search.

We will use equation (2) to express the marginal change of a change in a job attribute X_i on the workers' search effort. Dividing both sides by ∂p/∂s, differentiating with respect to X_i and using the envelope theorem (∂V/∂s = 0), gives:

$$\frac{\partial s}{\partial X_i}\frac{\partial}{\partial s}[-\frac{\partial k}{\partial s}/\frac{\partial p}{\partial s}]-\Pr[V(X_0)-c-V(X)>0]\frac{\partial V(X)}{\partial X_i}=0. \tag{3}$$

[3] In case the employed individual considers taking a second job, the distribution of job attributes of both jobs depends on X, so this assumption is violated. The proportion of workers that consider a second job is small. In the data we analyse later on, only 0.4% of all workers (4.3% of all searchers) search for a second job.

where $\Pr[V(X_0) - c - V(X) > 0]$ denotes the probability of accepting a job offer and where i = 1,..., n+1. Suppose that the n+1's job attribute is the wage. The workers' marginal willingness to pay for the ith nonwage job attribute (MWP_i) is then defined as the ratio of the marginal *lifetime* utility of the ith job attribute over the marginal *lifetime* utility of the wage. Hence, by using equation (3), we obtain:

$$MWP_i = \frac{\partial s}{\partial X_i} / \frac{\partial s}{\partial w}, \tag{4}$$

where i= 1,...,n. Our first result is that the workers' marginal willingness to pay for the ith non-wage job attribute, MWP_i, equals the ratio of the marginal effects of the ith non-wage attribute and the wage on search effort, conditional on search.

Our second result is also straightforward to obtain. Since V= V(v(X),s(X)), and using the envelope theorem, one can readily see that:

$$\frac{\partial V}{\partial X_i} / \frac{\partial V}{\partial w} = \frac{\partial v}{\partial X_i} / \frac{\partial v}{\partial w}. \tag{5}$$

Thus, we have shown that the ratio of the marginal *lifetime* utility of the job attributes equals the ratio of the marginal *instantaneous* utility of the job attributes. As a consequence, the MWP_i is equal to the ratio of the marginal *instantaneous* utility of the ith job attribute over the marginal *instantaneous* utility of the wage. And, therefore:

$$\frac{\partial s}{\partial X_i} / \frac{\partial s}{\partial w} = \frac{\partial v}{\partial X_i} / \frac{\partial v}{\partial w}. \tag{6}$$

Consequently, the ratio of the marginal effects of the ith non-wage attribute and the wage on the search effort equals the marginal instantaneous utility of the ith job attribute over the marginal instantaneous utility of the wage, conditional on search. In summary, given information

on $\partial s/\partial X_i/\partial s/\partial w$, one obtains (i) the MWP_i for job attribute X_i; (ii) the ratio of the marginal effects of job attribute X_i over the wage on the instantaneous utility function.

4. THE SEARCH ENVIRONMENT REVISITED

The on-the-job search model introduced in section 2 is the standard theoretical framework to understand on-the-job search. For empirical applications however, it may be too simplistic. It is therefore useful to investigate whether the results derived above still hold under weaker, and therefore more realistic, conditions.

Nonstationarity. Empirical applications of on-the-job search and job moving behaviour indicate that workers are active in a nonstationary environment. In particular, on-the-job search activities decrease with the time being in the current job (Kahn and Low, 1984; Parsons, 1991; Van Ophem, 1991). We will therefore introduce time 't' into the model, which denotes the job duration (tenure). We suppose that the structural parameters of the search environment (v, p, c, k and δ) are nonstationary and depend on t (see Van den Berg, 1990). This implies that lifetime utility is nonstationary, so V= V(t), and, therefore, search effort is nonstationary, so s = s(t).

Unemployment. It is natural to assume that the employed individuals take into account that they may become unemployed in the future. We assume that unemployed individuals will receive a benefit b. Let δ denote the involuntary separation rate of workers from jobs. We emphasise that δ may depend on job nonwage attributes X, so $\delta = \delta(X)$. For example, temporary employment contracts generally increase the risk of becoming unemployed.[4] Given the assumptions stated above, lifetime utility V(X) satisfies the following equation derived in the same way as equation (1):

$$\rho V(X) = \partial V(X)/\partial t + v(X) - k(s) + p(s)Emax[V(X_0) - c - V(X),0] + \delta(X)[U(b) - V(X)]. \qquad \textbf{(7)}$$

Equation (7) can be interpreted as follows (see, similarly, Van den Berg, 1990). The lifetime utility forgone per unit of time is equal to the sum of the appreciation of lifetime utility V at t, the

[4] In essence, MWP estimates can be derived from the relationship between X and voluntarily behaviour. As emphasised in the introduction, MWP estimates based on observations of job *moves* need to rely on the assumption that δ does not depend on X (Gronberg and Reed, 1994). This assumption is needed, because job moves consist of voluntary and involuntary job moves, which are not distinguishable. In contrast, job search is always voluntarily, even if triggered by a threat of involuntary separation.

instantaneous utility, the expected benefit of accepting a job offer and the expected loss of becoming unemployed, where U denotes the lifetime utility of an unemployed individual.

The optimal choice of s can be obtained by differentiating equation (7) with respect to s, and setting the resultant equal to zero. Going through the same mathematical steps as in the previous sections - and making use of the optimality condition that $\partial(\partial V/\partial t)\partial s = \partial(\partial V/\partial s)/\partial t = 0$ - we find again that $MWP_i = \partial s/\partial X_i/\partial s/\partial w$.

We see now that lifetime utility cannot be written as V(v(X),s(X)), but only as V(v(X),s(X),X), since lifetime utility depends directly on X via δ(X). As a consequence, $\partial V/\partial X_i/\partial V/\partial w$ cannot be written as $\partial v/\partial X_i/\partial v/\partial w$. This implies that the ratio of the marginal effects of the ith nonwage attribute and the wage on the search effort does not equal the marginal instantaneous utility of the ith job attribute over the marginal instantaneous utility of the wage. In summary, under the weaker conditions as stated in this section, one may interpret $\partial s/\partial X_i/\partial s/\partial w$ as MWP_i, however identification of the ratio of the marginal effects on the instantaneous utility function is not possible.

5. UNEMPLOYMENT

In previous sections, we have discussed the relationship between on-the-job search effort and *workers*' marginal willingness to pay for job attributes. In principle, unemployed' search can be modelled in the same way as employed' search and hence the MWP method is applicable to unemployed' search. Consequently, the *unemployed* individuals' marginal willingness to pay for non-pecuniary attributes that explicitly depend on the current state of unemployment can be derived by supposing that the n+1's attribute is the unemployment benefit b. An example of such an unemployment non-pecuniary attribute is the expectation of being recalled to the previous job from layoff. Another example is the residual entitlement period.

6. ESTIMATION METHOD

We will discuss here a method for estimating workers' MWP for job attributes given information on search behaviour. Suppose that exact information on search efforts s is not available and it is only known whether workers report that they search (s* = 1) or do not search (s* =0). Workers report that they search when search effort exceeds a (unknown) threshold value. One may then specify search activity s* by means of a latent-variable framework: $s = \beta'Y + u$, $E(u) = 0$; β is a

vector of unknown coefficients. Y represents a vector of explanatory variables and Y includes job attributes X; u is a random variable with expectation 0; s and s* are related as follows: s* = 1, if s > c (where c is an arbitrary threshold value); s* = 0 otherwise. Estimation of this discrete choice model is standard.

Let β_i the parameter associated with job attribute X_i, i= 1,..., n+1. It is then obvious that $\partial E(s)/\partial X_i = \beta_i$, i= 1,...n+1. Suppose now that the n+1's job attribute is the wage and let β_w be the parameter associated with the wage. The ratio of the marginal effects of the ith nonwage attribute and the wage on expected search effort is then equal to β_i/β_w and thus:

$$MWP_i = \beta_i/\beta_w, \quad 1, \ldots., h. \qquad \textbf{(8)}$$

In consequence, estimates of β_i/β_w can be interpreted as the workers' marginal willingness to pay for the ith non-wage job attribute (MWP_i). This result also holds when one observes search effort in a different way. For example, one may observe the number of search hours per week (leading to a truncated variable model, which, like the discrete choice model, relies on the latent variable linearly related to attributes, see Greene (2000, p. 908)) or the number of search contacts per week (leading to a Poisson model with expected number of contacts linearly related to attributes, Greene (2000, p. 880)).

The assumption that search effort depends linearly on the job attributes implies that the workers' MWP does not depend on any current wage or nonwage job attribute (see equation (8)). In most empirical applications of on-the-job search behaviour however, it is assumed that the wage determines search effort non-linearly. The most common specification is that the logarithm of the wage determines search effort. Such a specification implies that the MWP_i equals $w.\beta_i/\beta_w$. Hence, the MWP_i is proportional to the current wage. The unemployed individuals' MWP for unemployment attributes can be estimated in a similar way. On the other hand, it is common to include both linear and quadratic terms in age, tenure and working hours as explanatory variables. In such cases one has to replace β_i in (8) with $\partial E(s)/\partial X_i, = \beta_{i1} + \beta_{i2} X_i$, hence MWP_i is linear in X_i.

7. EMPIRICAL APPLICATION

7.1 Background

In this section, we estimate workers' marginal willingness to pay for a range of job attributes in Lithuania in the year 2000. Our emphasis is on the MWP for the type of contract and in particular temporary contracts. Temporary contracts are clearly associated with higher than average unemployment risk.[5]

Lithuania is one of the former Soviet republics, which by the end of 2000 was at the end of its first decade of transition to the market economy. Labour force participation in Lithuania was then close to the EU average, but the unemployment rate (16.1 percent according to ILO definition) was high by EU standards. By this time, unemployment had been continuously rising for more than two years (and kept on rising in the beginning of 2001), suggesting that workers were concerned about the risk of losing job.[6] This is a suitable case for studying MWP to avoid such a risk.

7.2 Data, descriptives and methods

Our data come from the national labour force survey (LFS) conducted by the Statistical Department of Lithuania in November 2000. 3 thousand households were randomly drawn from the household register and all (available) members of these households aged 15 and older were interviewed. From this sample of over 7.5 thousand individuals, employed individuals were selected. After excluding those working abroad and cases with missing information we are left with 2641 observations of workers. In the LFS, employed individuals are asked whether they are involved in a search activity for another job.

[5] Data for Lithuania indicate that workers with temporary contracts in 1999 have a 22% probability of being unemployed one year later. By comparison, those with permanent contracts have merely a 6% probability of being unemployed one year later. Moreover, workers with temporary contracts are also several times more likely to leave the labour force.

[6] Subjective questions on the value of work in the Baltic states (UNDP, 1998, p. 60) as well as regular public opinion polls show that job security is the most important characteristic of work for workers. The main explanation is plausibly the combination of high unemployment rates and low levels of unemployment and welfare benefits. Lithuania's unemployment benefits do not follow an insurance principle and are *maximally* two-third of the minimum wage and less than half of the average gross wage. Furhermore, most unemployed do not receive unemployment benefits. The bottom line is that the welfare loss from becoming unemployed is larger for workers in Lithuania than for EU workers.

Incidence of on-the-job search is 8.8 percent.[7] Table 1 reveals that temporary workers,[8] as well as part-time workers and those with short tenures are much more likely to search. Workers with short temporary contracts (less than 6 months) are more likely to be engaged in search than those with longer temporary contracts. The descriptives indicate that temporary and part-time males search more often than their female counterparts. Nevertheless, compared to data from West-European countries, observed search differences between males and females are relatively small. For example in the UK and in the Netherlands, part-time female workers search on average less than their full-time counterparts (e.g. Pissarides and Wadsworth, 1994), whereas in Lithuania they search more.[9] This phenomenon is thought to be related to the strong employment position of females in the Baltic labour markets (OECD, 2002). This justifies our procedure to pool initially observations for both gender.

Subjective information indicates that 53% of the searchers state that they search because they wish to improve their working conditions, 20% search because the current job is anticipated to be terminated or seen as transitional and 20% search to increase the number of the hours worked (see Table 2).[10] Table 2 also suggests that the reasons for search are similar for males and females. The data suggest that part-time work is not seen as attractive in Lithuania. Part-time workers tend to search in particular because they wish to increase the number of working hours. As one might expect, most temporary workers search because the current job will be terminated or is seen as transitional.

MWP estimates have been derived using a logit model.[11] We will also report estimates correcting for sample selection of workers using a standard Heckman correction and estimates for males and females separately.

[7] Incidence of on-the-job search including search for a second job is only slightly higher (9.2%). This percentage is comparable to EU countries. For example, the same percentage is reported in the Netherlands (Statistics Netherlands, 1992). In the UK, somewhat lower percentages are reported (Pissarides and Wadsworth, 1994).
[8] Temporary workers are defined as workers with non-permanent employment contracts. The average contract duration is just under 6 months. About 70% of the contracts are for no more than 6 months.
[9] In this sense, the Lithuanian labour market is more similar to the US labour market.
[10] Respondents were allowed to choose only one reason.
[11] Post-stratifying weights provided by the Statistical Department of Lithuania were used in the estimation process. Reported standard errors are the robust ones and allow for clustering within households. MWP estimates appear to be insensitive to the use of weights. Results without weights can be received from the authors upon request.

7.3 Results

In the baseline model we include one job attribute associated with the risk of unemployment: whether workers have, or have not, a temporary contract. Other job attributes included are night work and weekend work, which are generally thought of as attributes that reduce welfare. The number of working hours is also included. A priori, it is not clear whether the number of hours is less or more than the workers' optimal number of hours.[12]

We have estimated a range of models, including a range of control variables for worker and firm characteristics.[13] Further, we control for the local unemployment rate (county and gender specific) and the percentual annual change in employment in the industry of employment. In Table 3, the full results of the baseline model can be found. In Table 4, we provide the MWP estimates given four different specifications.[14] Specification (1) distinguishes between workers with temporary contract durations up to 6 months (short contracts) and contract durations of more than 6 months (long contracts).[15] Specification (2) includes a dummy for temporary jobs and a contract duration variable (if the job is temporary, otherwise 0). Specification (3) uses a spline contract duration variable to test whether the MWP is higher for short contracts. Specification (4) uses subjective information on the reason why the contract worker is temporarily employed, distinguishing between 'involuntarily temporary contracts' (e.g., "could not find permanent job") and 'voluntarily temporary contracts' (e.g., "Did not want a full-time job", "A contract covering a period of training" etc). In Table 5, we provide MWP estimates for males and females separately. In addition, we correct for sample selection.

According to the baseline model (Table 3), workers with temporary contracts are willing to pay about 190 percent of their monthly wage to get a permanent job. The estimate for men is higher than for women (240 and 140 percent respectively, see Table 5). MWP to avoid short temporary contracts (up to 6 months) is somewhat higher than for long contracts (see Table 4, specification (1)). Specification (2) suggests that the average MWP for an extra month contract is

[12] The descriptive data indicate that 11.5 percent of workers would prefer longer hours, while 7.4 percent want to work less; workers search more when they work part time (see Table 1).

[13] We have also experimented with the specifications by including additionally 8 occupation dummies and 23 industry dummies. The main difference is that the MWP for working hours drops from 3.4 to 2.3 percent. The other MWP estimates remain unchanged. These results can be received from the authors upon request.

[14] The full results can be received upon request from the authors. In addition, overall effect of workers' characteristics on search effort (encompassing both wage effect and effect on search conditional on wage) can also be estimated from a logit model like the one in Table 3 where individual (log) wages are replaced with residuals from the wage equation. MWP estimates do not change .

around 6 percent. Although specification (3) shows a higher value for durations up to 6 months (14.9 percent), and virtually zero afterwards. This latter result makes sense, because the value of a contract extension must increase sharply when the contract duration approaches zero. Specifications using workers' subjective information on the reason why workers hold a temporary job suggest that workers whose job is temporary job because they failed to find a permanent one attach more value to a permanent contract (Table 4; specification (4)). So, the MWP estimate for temporary jobs is higher for workers who 'involuntarily' hold a temporary contract.

Using a simple search model, we will show now that the above estimates which imply that the willingness to pay to avoid temporary contracts exceeds the wage rate are plausible. Suppose an individual has a temporary employment contract, earns wage w and anticipates becoming unemployed at rate δ and finding a permanent job at rate q. The individual discounts the future at rate ρ. So, lifetime utility V can be written as follows: $\rho V = w+\delta(U-V)+qV^p$, where V^p denotes the lifetime utility of a permanent contract. When unemployed, this individual will receive a benefit B ($B < w$) with probability $\pi < 1$ and will find again a temporary job at rate λ. In the permanent job, the individual will earn wage w^p forever, so $\rho V^p = w^p$. So, the unemployed lifetime utility U can be written as: $\rho U = b+\lambda(V-U)$, where $b = \pi B$. Lifetime utility V can then be written as:

$$V = \frac{1}{r(r+l+d)}\left(\frac{w+qV^p}{r+l}+db\right). \qquad (9)$$

The willingness to pay (WP) for δ is defined as $[V(\delta)-V(0)]/\partial V/\partial w$, so:

[15] The first group of workers has a mean temporary contract of 3.2 months, whereas the second group has a mean contract duration of 12.9 months.

$$WP = -\frac{d}{r+l}\left(w + qV^{p}\right) + db(r+l). \tag{10}$$

Now suppose that unemployment benefits levels are low, so b = 0 (which will be a reasonable assumption for Lithuania). It follows that:

$$\frac{WP}{w} = -\frac{d}{r+l}\left(1 + \frac{q}{r}\frac{w^{p}}{w}\right). \tag{11}$$

For most reasonable parameters WP/w < -1. For example, data for Lithuania (OECD, 2002) indicate that λ, the annual rate of becoming re-employed, is 0.35 and δ, the annual rate of becoming unemployed for temporary workers, is *at least* 0.22. Further, q must be larger than ρ and w^{p} must exceed w. Presuming that the discount rate is 0.10 shows now that WP/w is larger than one in absolute value, so the willingness to pay to avoid a temporary job would be more than the current wage rate.[16]

Further, the empirical results show that the MWP estimates for temporary contracts are higher for males than for females (see Table 5). When sectoral fall in employment is interacted with temporary jobs, it turns out that workers working in a declining industry attach considerably higher values to occupying a permanent position (7.5 percent of the wage per each percentage point of gender specific decline in industry employment). Presumably, in a declining industry the chances that a temporary job contract will not be renewed are higher than in other industries. Temporary workers are a lot more concerned about an employment cut in the sector they work in.

Workers are willing to pay 3.4 percent of their monthly wage rate for an extra working hour, on average.[17] This might also encompass a risk premium, as longer hours tend to imply more stable jobs. This estimate is a decreasing function of the number of working hours. For example, a worker with a 20 hours a week contract values a marginal working hour at 6.0 percent

[16] Likely, the willingness to pay is higher, because we presume in this simple model that workers are risk neutral, and we ignore that temporary workers are more likely to leave the labour force.
[17] If we include search for a second job in the definition of on-the-job search, it appears that this estimate is slightly higher (4.3%).

of the monthly wage,[18] so almost twice as much as the average worker. For workers who work more than 62 hours a week the MWP for working hours becomes negative.

About half of all workers work either at nights or on weekends (or both), and these job attributes proved to have an impact on search activities (especially for men). Workers on average are willing to pay up to 56 percent of their wage to avoid night work and 34 percent to avoid weekends (not night) work. These estimates suggest that the discomfort of working at night (or in the weekends) is substantial. In the literature, estimates based on hedonic models are normally lower (usually up to 20%), see Kostiuk (1990) and Lanfranchi et al. (2002). When we re-estimated the model defining search activity including search for a second job the MWP for night work drops to -49, but the MWP for weekend work becomes statistically insignificant. Interpretation of the MWP for night weekend work is therefore hazardous. A larger data set is clearly needed to investigate this further.

Finally, we will discuss the other determinants of the search decision. Results reported in Table 3 show that likelihood of search decreases with age and tenure (for tenures up to 22 years).[19] Females and ethnic minorities are less likely to search, although both effects are not statistically significant. Like it was found for the UK by Pissarides and Wadsworth, 1994, tertiary education promotes on-the-job search. Workers search more when their industry employment falls. To the extent that a decline in industry employment can be interpreted as a job attribute the estimates suggest that workers value each percentage point of (gender-and-industry-specific) fall in annual average employment at 2 percent of the monthly pay, but the precision of this estimate is too low to interpret this result.

We find that the local unemployment rate has a positive impact on workers' search effort[20]. If the probability of finding a job is low, workers are expected to search less, however the period of search is increased as the probability of being accepted is small. As a consequence, the effect of the local unemployment rate on the probability of being observed searching is theoretically ambiguous a priori.[21]

[18] Other specifications suggest that men value extra hours higher than women, although the difference is not statistically significant (Table 5).
[19] Impact of tenure becomes statistically significant when the sample is restricted to fulltime workers.
[20] Positive relationship between local unemployment and on-the-job search was recently documented for the UK by Fuentes (2002), but see Pissarides and Wadsworth (1994) and Mekkelholt (1993) for the Netherlands.
[21] Our measure of unemployment is gender-specific county level rate based on the LFS conducted one year ago. Using such predetermined variable avoids possible endogeneity (those who search on-the-job compete for vacancies with unemployed job-seekers, thus prolonging their unemployment spells).

7.4 Hedonic wage estimates

As a exploratory exercise, we have also estimated a standard hedonic wage model using the same regressors (see Table 6). In line with a range of other empirical studies, it appears that a hedonic wage model based on cross section data gives a theoretically incorrect MWP estimate for a job attribute associated with the risk of unemployment (see Moretti, 2000, for a review). Temporary workers are paid (other things equal) less than employees on permanent jobs. The inability to control for individual characteristics associated with higher probability of unemployment in a cross section leads clearly to a downward bias in the estimate of compensating differentials. In a cross section, workers of higher unmeasured ability may earn higher wages and suffer less unemployment, so that the observed differentials may be wrong-signed (Card, 1987; Moretti, 2000). Workers in declining industries are also underpaid (differential is significant at 1% level). Monthly pay increases with hours worked (at a rate of 1.5 percent per hour at 38 hours per week), reaching a maximum at 57 hours per week, so there is no wage compensation for working shorter hours. Night work does not have a significant impact on earnings.[22] Weekend workers are paid less than other workers, in contrast with the MWP estimates of the job search model.

These results seem to indicate, in line with a range of other studies (Herzog and Schlottman, 1990; Gronberg and Need, 1994; Van Ommeren et al., 2000) that workers attach substantial value to non-wage differences for which they are not compensated. While we do not know which set of estimates is better, as also argued by Gronberg and Reed (1994), we emphasise that hedonic wage methods presume that workers are fully compensated and are perfectly mobile. In particular, these methods exclude the possibility that it takes time for non-compensated workers to leave the current employer (for example due to lack of information about other job alternatives). This presumption does not hold in our sample where a substantial proportion of workers are involved in on-the-job search. In particular, 53% of male workers with temporary contracts in our sample search for other jobs. So, the hedonic wage method presumes that workers are in jobs that maximise their utility, but in our sample, the *majority* of male workers with temporary contracts are involved in on-the-job search and are therefore not in jobs

[22] We have also re-estimated the model including additional controls for industry and occupation. When industry and occupation are controlled for, the results are identical, except for a positive compensating differential for night work, but it is much smaller in size (just 9 percent) than the MWP from the search model. If we correct for selectivity, the night work premium is higher, close to 20 percent, as also shown by Lanfranchi et al. (2002) and Kostiuk (1990), but still less than the MWP based on the search model.

that maximise utility. This indicates that the assumption of perfect mobility is not innocuous. So, this raises the question why individuals *accept* temporary contracts when they know beforehand that they are not compensated? Presumably, one answer is that temporary contracts are mainly accepted by *unemployed* searchers (and job searchers with temporary contracts of shorter duration). By accepting a temporary job, unemployed searchers do not forego the opportunity to search for a permanent job. In equilibrium, these temporary jobs are more likely offered by less profitable firms, which have lower opportunity costs in having vacancies go unfilled, and which offer temporary jobs and lower wages (Huang et al., 1992).[23]

8. ESTIMATES BASED ON PREVIOUS STUDIES

In this study, we also make use of previous studies of on-the-job search behaviour in order to derive MWP estimates. We make use of one study that has examined workers' job search behaviour in the U.S.: Parsons (1991) and one in the Netherlands: Van Ophem (1991). Given the estimates of the determinants of job search activity as reported in these studies, we derive workers' MWP for job attributes (section 8.1). Results published by Barron and Mellow (1979) for the U.S. and Lindeboom and Theeuwes (1993) for the Netherlands on the unemployed individuals' search behaviour are used to derive the unemployed individuals' MWP for unemployment attributes (section 8.2).

8.1. On-the-job search

We provide the estimates of the MWP for job attributes using the estimation method discussed in section 6.[24] To facilitate comparison of the results, we provide estimates of the MWP for a job attribute divided by the wage (multiplied by 100) denoted as %MWP. One advantage of this

[23] One of the consequences of non-compensating wage offers, which increase on-the-job search, is that the firm's job turnover increases. Although this may be an additional cost for many firms due to additional recruitment and loss of productivity costs, it is plausible that for some firms (e.g. those with a temporary increase in the demand for their products or those which expect a decrease in future demand) it is advantageous that workers leave after a specified period.

[24] The variance of the estimated MWP_i is derived using the delta method, so $Var(\beta_i/\beta_w)$ is calculated as $[Var(\beta_i)+(\beta_i/\beta_w)^2.Var(\beta_w)-2.(\beta_i/\beta_w)Cov(\beta_i/\beta_w)]/\beta_w^2$. As it is common practice not to report the covariance matrix of the coefficients, we suppose that $Cov(\beta_i,\beta_w)$ is zero. Hence, the reported precision of the MWP estimates is somewhat inaccurate. For the current application, this is not problematic. In the case that MWP_i equals zero, variances of the MWP_i estimates are exact, so one may test the hypothesis that MWP_i equals zero using a standard t-test. In addition, when MWP_i is positive, the bias in the variance is small, even for a relatively high correlation between β_i and β_w. For example, if the correlation between β_i and β_w is 0.2, which is high in this type of application, then the relative bias in the standard error can be shown to be less than 10%.

measure is that it is the closest to the empirical specifications employed by the studies discussed here.

Parsons (1991). Parsons (1991) used the 1980-1981 National Longitudinal Survey of Youth to study the employed workers' choice among employed search, unemployed search, and not searching for a new job. Ordered probit models are employed for men and women. The wage rates are specified in logarithms. The results show that current wages, promotion prospects, job tenure (for men, not for women) and full-time work are each negatively associated with search intensity (see Table 7).

Parsons (1991) found that workers searched more if they worked part-time. Thus the average part-time worker prefers to work more hours than available on their current job. Hence, the marginal rate of substitution of wage for leisure is less than the current wage rate (the marginal rate of substitution of wage for leisure is equal to the wage rate provided that search activity is (i) affected by the hourly wage rate and (ii) not affected by the number of work hours). The results also indicate that the MWP for a full-time position is higher for men (65% of the current wage) than for women (35% of the current wage). Such a finding is consistent with the notion that men in the US (generally prefer to) work more hours than women. The results clearly show that workers search significantly less if they expect to be promoted. Promotion opportunities are highly valued by workers. 'Very good promotion prospects' are valued at 176% for men and 225% for women. 'Good promotion prospects' are valued at 112% for men and 144% for women. 'Not so good promotion prospects' are valued at 84% for men and at 42% for women (the latter is not significant at conventional levels of significance). We find that the MWP for promotion prospects is not so much gender dependent.

We will explain by application of search theory that these estimates of the MWP for promotion prospects are plausible. To simplify matters, we suppose again a simplified model that allows us to obtain an explicit solution for the MWP for promotion prospects. Suppose a worker earns wage w and expects to be promoted at rate β. Promoted workers receive a wage w^p forever ($w^p > w$). The worker discounts the future at rate ρ. Job-to-job mobility is ignored. Lifetime utility V can then be written as $(\rho w+\beta w^p)/\rho(\rho+\beta)$ and the MWP for β equals $(w^p-w)/(\rho+\beta)$ (since $\partial V/\partial w = 1/(\rho+\beta)$ and $\partial V/\partial \beta = (w^p-w)/(\rho+\beta)^2$). So, the MWP for β is positive, and decreasing and concave in β.

In Parson's (1991) empirical specification of job search behaviour, dummies for various

levels of promotion prospects are included. Each dummy indicates a different level of β. So, the MWP for a dummy can be interpreted as the willingness to pay (WP) for a certain level of β. The willingness to pay for β, defined as $[V(\beta)-V(0)]/\partial V/\partial w$, can be written as $\beta(w^p-w)/\rho$. Now suppose that the determinant 'very good promotion prospects' implies that β/ρ is 5. This seems quite reasonable, for example, the yearly promotion rate β might be 0.50 and the yearly discount rate ρ is 0.10. Since Parsons' (1991) empirical results indicate that the percent MWP for 'very good promotion prospects' is about 200%, promoted workers receive a wage increase of 40%. Such an estimate seems plausible (see Murphy, 1985; van Gameren, 1999).

Van Ophem (1991). Van Ophem (1991) used the 1985 OSA Labour Market Survey to study the importance of nonwage attributes on the search decision of Dutch employees. The results show that on-the-job search activity increases with unemployment expectations and unpaid overtime, but decreases with wage and good promotion prospects. Different measures for the predicted wage are used using a 'structural form' and 'reduced form' model. In the current paper, we report the MWP estimates for job attributes based on the 'reduced form' model (see Table 8) (the MWP estimates based on the 'structural form' model are larger in absolute value). Van Ophem (1991) reports that workers search more if they work unpaid overtime and expect to become unemployed within a year. Workers search less if they have good promotion prospects.[25] The workers' MWP for the absence of unpaid overtime (measured in hours per week) is 4.75% of the weekly wage. In the Netherlands, the average employed individual works about 35 hours per week. Thus, the percent MWP for the absence of one hour unpaid overtime is not significantly different from 100 (at the 5% level), which implies that the marginal rate of substitution of wage for leisure equals the wage rate.

[25] Van Ophem (1991) also includes commuting time. Unfortunately, the standard error of the MWP estimate is so high that interpretation is hazardous.

Van Ophem (1991) includes a determinant of search activity defined as 'the expectation of becoming unemp1oyed within 12 months'. The MWP for the absence of this expectation is about 147% of the wage rate. Consequently, Dutch workers who expect to become unemployed within 12 months anticipate a substantial loss. This is likely to be due to low re-employment probabilities in the Dutch labour market in the mid eighties, since the direct loss in income is relatively small in the Netherlands. Using again the simple search model (see section 7.3), we will show that the Dutch estimates are plausible.

We assume now that the annual value of λ is 0.66 and of ρ is 0.10. The expected duration of being unemployed after losing the job is then 1.5 years, which corresponds to the average Dutch unemployment duration during the period 1983-1987 (Gorter et al., 1990). The determinant 'the expectation of becoming unemployed within a year' seems to indicate a large yearly separation rate. We assume that the annual value of δ is three (the probability of becoming unemployed within a year is then 0.90). In 1985, the Dutch unemployment insurance payment initially amounted to 80% of the most recently earned wage, however, this was reduced after a maximum of six months. Thus, the loss in earnings is much more than 20%. For simplicity, we assume that the loss is 30%. Given these assumptions, the WP for the absence of δ is 118%. Such a number is not too far from the 147% implied by the results reported by Van Ophem (1991). Clearly, higher re-employment rates λ imply a lower WP for the absence of δ, since the expected duration of being unemployed is shorter. In the case that λ is one, the WP for the absence of δ is 59%. Furthermore, the *marginal* willingness to pay for the absence of δ is decreasing in λ and in δ. In the case that λ is 0.66 and ρ is 0.10, the percent MWP for the absence of the yearly separation rate δ decreases from 39% ($\delta = 0$) to 2.5% ($\delta = 10$). Finally, and in line with the estimates implied by the results of Parsons (1991), Van Ophem reports that 'good promotion prospects' are valued at about the current wage rate (MWP is 97% of the wage). As argued before, this result is plausible. Given the noted differences between the studies, this suggests that the estimation method generates robust results among different studies.

8.2. Unemployed job search

Barron and Mellow (1979). Barron and Mellow (1979) used a special survey among a sample of the unemployed respondents in the May 1976 U.S. Current Population Survey, to study the unemployed individual's choice of how much time to devote to searching for a job. Particular

attention was paid to the role of unemployment insurance benefits and to individuals who have recently been (temporarily) laid off. Regression models are employed for the full sample and for a sample restricted to individuals entering unemployment from prior jobs. The weekly insurance benefits are assumed to affect search time linearly. Dummies are used for 'expected recall, within 30 days' and 'expected recall, no period specified'. The results show that the unemployment insurance benefits and recall expectations reduce unemployment search time. In Table 3, the results are given for the sample restricted to individuals entering unemployment from prior jobs (the results for the full sample imply somewhat higher – but less significant – MWP estimates).

The MWP for expected recall within 30 days is 313 dollars, which is about four times the average weekly benefit (for the 31% receiving benefits, the mean is 77 dollars). Therefore unemployed individuals are willing to forgo benefit for a month to receive a recall within a month. The MWP for expected recall when the period is not known is 144 dollars, almost twice the average weekly benefit.

Lindeboom and Theeuwes (1993). Lindeboom and Theeuwes (1993) used a random sample drawn from the 1982-1984 administrative records of the Dutch unemployment benefit administration for the Leiden district to study the determinants of search effort. One of the determinants is the residual entitlement period of receiving unemployment benefit. Under the Dutch Unemployment Act, the benefit level considered is approximately 80% of gross earnings before unemployment. The length in days of unemployment benefit entitlement depends on the number of days worked. The maximum benefit duration is 26 weeks. At the end of the unemployment entitlement period, the benefit drops to 94% of the benefit level (75% of previous earnings). Depending on the length of the prior job, the unemployed will receive this benefit for a certain period. Ultimately, the unemployed receive welfare, which is generally substantially less than the benefit and which does not depend on previous income. Search effort is measured by the number of search contacts. The analysis is based on a Poisson model. The results show that search effort declines significantly with increasing benefit levels, and rises over the residual entitlement period (see Table 4).

These results imply that the MWP as a percentage of the benefit for residual entitlement (in weeks) is equal to 13.33% -0.26% times the residual entitlement period. Hence, the MWP for residual entitlement is positive over the entitlement period (maximally 26 weeks) and increases at a weekly rate of 0.26% of the benefit as the end of the entitlement period comes near. The

willingness to pay for one week extra residual entitlement, at the beginning of entitlement, is 6% of the benefit level. The willingness to pay for one week extra residual entitlement, at the end of the entitlement period, is 13% of the benefit level. The empirical outcomes seem quite plausible, since, as explained above, at the end of the entitlement period considered, unemployed individuals lose at least six percent of the benefit for a certain period and, after this period the benefit will be reduced to welfare level.

9. CONCLUSION

In this paper, we have demonstrated that the marginal willingness to pay for job attributes can be derived from data on on-the-job search activity. The main advantage of this estimation method, compared to estimation methods based on job moves, is that one needs less restrictive assumptions on the search environment. The empirical relevance of the search approach to estimate the workers' marginal willingness to pay is applied to observations from Lithuania and further demonstrated based on a number of studies in the U.S. and the Netherlands. We have provided evidence, that workers attach substantial value to non-wage differences between jobs like unpaid overtime, risk of becoming unemployed and promotion prospects. Furthermore, we demonstrate that data on unemployed individuals' search behaviour may be useful in obtaining information on the value of unemployment attributes such as recall opportunities and the residual entitlement period of receiving unemployment benefit.

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Table 1. Incidence of on-the-job search in Lithuania, 2000.

Employees		All	Men	Women
All		8.8	9.6	8.1
Full-time		7.2	8.3	5.9
of which:	Tenure ≤ 6 months	13.9	14.4	13.1
	Tenure 2 – 5 years	7.3	8.5	5.9
	Tenure ≥ 15 years	2.9	2.8	3.0
Part-time (less than 32 hours)		23.6	34.7	19.0
Temporary contracts		45.5	53.1	32.6
Of which:	≤ 6 months	47.9	55.0	35.9
	> 6 months	39.2	48.1	23.6
Night work		9.3	10.9	6.5
Weekends work (excluding nights)		10.6	11.4	10.8

Source: LFS data and own calculation.

Table 2. On-the-job search by reason in Lithuania, 2000.

percent

		Employees				
	Reasons	All	Men	Women	Part-time	Temporary
1	Risk or certainty of loss of present job / transitional job	19.3	21.5	17.0	17.0	64.7
2	Seeking more hours	20.3	20.0	20.6	39.3	7.2
3	Wish to have better working conditions, pay etc.	52.7	51.8	53.6	42.1	25.2
4	Other	7.8	6.8	8.8	1.6	2.9

Source: LFS data and own calculation.

Table 3. Determinants of on-the-job search and MWP estimates for job attributes in Lithuania, 2000.

Variables	Mean	Coeff.	s.e	MWP	
				% wage	s.e.
Job attributes					
Log wage[a]	6.354	-1.042	(0.193)***		
Temporary job	0.041	1.946	(0.299)***	-187***	(46.1)
Night work[b]	0.165	0.584	(0.269)**	-56**	(28.0)
Weekends[c]	0.320	0.358	(0.204)*	-34*	(20.5)
Working hours	38.2	-0.092	(0.030)***	3.4***	(1.4)
Hours sq. (coef. × 100)	1512	0.075	(0.035)**		
Firm characteristics					
Plant size 1-10 empl.	0.173	0.102	(0.240)		
Plant size 11-19 empl.	0.089	0.125	(0.292)		
Plant size 20-49 empl.	0.154	0.084	(0.278)		
Search environment					
Decline of sector employment[d], %	2.2	0.019	(0.011)*		
Local unemployment rate[e], %	15.1	0.085	(0.028)***		
Worker characteristics					
Tenure	8.0	-0.033	(0.032)		
Tenure sq. (coef. × 100)	145.8	0.072	(0.001)		
Age	39.7	-0.031	(0.009)***		
Resident of capital city	0.197	0.960	(0.233)***		
Rural resident	0.199	-0.619	(0.284)**		
Female	0.517	-0.226	(0.198)		
Ethnic minority	0.156	-0.383	(0.257)		
Single	0.191	-0.342	(0.244)		
Divorced	0.102	0.511	(0.299)*		
Tertiary education	0.515	0.999	(0.432)**		
Secondary education	0.368	0.524	(0.433)		
Vocational education	0.035	0.652	(0.527)		
Constant		5.219	(1.499)***		
Number of observations			2641; F(23,1580) = 9.21		

Notes: Standard errors in parentheses. [a]Net monthly earnings in Litas (1 Litas =0.25USD). [b]Night work = 1 if respondents works at night (sometimes or usually). [c]Weekends = 1 if respondent works on Saturdays or Sundays usually or occasionally, and does not work at nights. The reason for excluding nights is that most respondents who work at night also work on weekends. [d]Gender specific percentage fall (+) or growth (-) in employment by 15 major NACE sectors, according to Labour Exchange data. Varies from –28 to 22. [e]Gender specific unemployment rate (percent) according to November 1999 LFS in the county where respondent's main job is located. Varies from 10 to 24. Other specifications (unemployment rate in the county of residence; registered rather than LFS unemployment rate) give similar but less significant results.

***, **, * - estimates significant respectively at 1%, 5%, 10% level.

Table 4. MWP for job attributes, percent of net monthly wage in Lithuania, 2000. Alternative specifications.

Job attributes	**Mean**	**(1)**	**(2)**	**(3)**	**(4)**
Temporary job	0.041		-187[b]	-163	
			(47)	(53)	
Temporary job ('involuntarily')	0.026				-188[b]
					(55)
Temporary job ('voluntarily')	0.015				-82[b]
					(50)
Temp. job: short contract (≤ 6 months)	0.030	-191	-202[b]	-202	-256[b]
		(51)	(53)	(53)	(64)
Temp. job: long contract (> 6 months)	0.011	-177	-146[b]	-151	-193[b]
		(52)	(51)	(47)	(53)
Temp. job × contract duration	5.9[a]		6.0		
			(5.4)		
Short contract × contract duration	6.8			14.9	26.6
				(14.6)	(16.9)
Long contract × contract duration				1.5	-0.7
				(17.4)	(18.7)
Number of observations: 2641, of which 208 search.					

Notes: Specification (1) distinguishes between workers with temporary contract durations up to 6 months and contract durations of more than 6 months (long contracts). Specification (2) includes a dummy for temporary jobs and a contract duration variable. Specification (3) uses a spline contract duration variable to test whether the MWP is higher for short contracts. Specification (4) uses subjective information on the reason why the contract worker is temporarily employed, distinguishing between 'involuntarily temporary contracts' and 'voluntarily temporary contracts'. Standard errors in parentheses. [a]Mean is given for temporary contracts. [b]The MWP estimates for temporary jobs (less than 6 months and more than 6 months) in specifications (2) and (3) are based on estimated coefficients for temporary job dummy and contract duration variables, and mean contract durations for each category. These specifications do not include dummies for temporary job with contract duration 6 months and > 6; respective MWP.

Table 5. MWP for job attributes by gender, percent of net monthly wage in Lithuania, 2000.

	Men and women		Men		Women	
Job attributes	(1)[a]	(2)[b]	(3)	(4)[b]	(5)	(6)[b]
Temporary job	-187***	-183***	-238***	-255***	-142***	-128***
Night work	-56**	-52**	-70**	-82**	-25	-26
Weekends (excl. Nights)	-34*	-28	-44	-41	-29	-22
Hours worked	3.4**	3.0**	2.9**	3.7**	2.5*	2.0
Correlation		-0.30		0.96		-0.57
(std. Err.)		(0.37)		0.07		(0.23)
No. of observations	2641	3174	1266	1568	1375	1606
No. of employed	2641	2641	1266	1266	1375	1375
of which searchers	208	208	112	112	96	96

Notes: [a]Estimates are identical to the estimates reported in Table 3. [b]Estimates from bivariate probit models with sample selection from labour force into employment. Instruments used in the selection equation in models (2), (4), (6) include dummies for non-manual workers, ethnic minority (removed from search equation) and additional education categories. Additional instruments used for robustness check: dummy for being born abroad in model (2); dummy for ongoing education or training in models (2), (4) (6); dummy for living in hostel in model (6) did not change the results. Correlations reported show that unobserved characteristics, which promote employability have a negative effect on probability of search for female employees and very strong positive effect for male employees (in both cases significant at 5%). In the pooled sample correlation is negative but not significant.

***, **, * - estimates significant respectively at 1%, 5%, 10% level.

Table 6. Hedonic wage model for Lithuania, 2000.

	Coeff.	Std. Error	
Job attributes:			
Temporary job	-0.1404	0.0555	***
Night work	0.0092	0.0283	
Weekends	-0.0446	0.0230	**
Hours	0.0450	0.0059	***
Hours sq. (coef. × 100)	-0.0393	0.0001	***
Firm characteristics			
Plant size 1-10 empl.	-0.1873	0.0260	***
Plant size 11-19 empl.	-0.1234	0.0329	***
Plant size 20-49 empl.	-0.0940	0.0284	***
Search environment			
Decline of sector employment 1999-2000, %	-0.0046	0.0012	***
Local unemployment rate, %	-0.0017	0.0029	
Worker characteristics			
Tenure	0.0183	0.0035	***
Tenure sq. (coef. × 100)	-0.0003	0.0001	***
Age	-0.0024	0.0010	
Resident of capital city	0.1956	0.0308	***
Rural resident	-0.0917	0.0243	***
Female	-0.2206	0.0220	***
Ethnic minority	-0.1124	0.0300	***
Single	-0.0320	0.0269	
Divorced	0.0018	0.0302	
Tertiary education	0.4792	0.0341	***
Secondary education	0.1796	0.0316	***
Vocational education	-0.0460	0.0605	
Constant	5.1433	0.1403	***
Number of observations	2641		
R^2	0.3576		

Notes: Wages specified in logarithm. See, further, the notes from Table 3.
***, **, * - estimates significant respectively at 1%, 5%, 10% level.

Table 7. Coefficients of search activity (encompasses search plans) with respect to attributes, out-of-school, aged between 17 and 23, U.S., 1980-1981 (based on Parsons, 1991) and the marginal willingness to pay for job attributes.

Men			**Women**	
Variables	Coefficient	% MWP	Coefficient	% MWP
log wage rate	-0.573		-0.484	
	(0.129)[a]		(0.165)[a]	
part-time	0.373	-65.2	0.202	-35.3
	(0.173)[b]	(33.6)[c]	(0.108)[c]	(20.5)[c]
promotion prospects:				
very good	-1.008	175.9	-1.289	225.0
	(0.174)[a]	(49.9)[a]	(0.163)[a]	(58.1)[a]
good	-0.640	111.6	-0.827	144.3
	(0.180)[a]	(40.2)[a]	(0.141)[a]	(40.8)[a]
not good	-0.479	83.6	-0.240	41.9
	(0.183)[a]	(37.1)[b]	(0.149)	(27.7)
absent				

Notes: standard errors in parentheses. a: significantly different from zero at the 0.01 level; b: significantly different from zero at the 0.05 level; c: significantly different from zero at the 0.10 level.

Table 8. Coefficients of search with respect to job attributes, The Netherlands, 1985 (based on Van Ophem, 1991) and the present marginal willingness to pay for job attributes.

Variables	coefficient	%MWP
log wage rate	-0.442	
	(0.165)[a]	
unpaid overtime (hours per week)	0.021	-4.75
	(0.013)[b]	(2.83)[b]
unemployment expectation	0.650	-147.06
	(0.106)[a]	(46.24)[a]
good promotion prospects	-0.428	96.83
	(0.087)[a]	(16.93)[a]

Notes: standard errors in parentheses. a: significantly different from zero at the 0.01 level; b: significantly different from zero at the 0.10 level.

Table 9. Coefficients of search time (search hours per week) with respect to attributes of unemployed individuals entering unemployment from prior jobs in 1976, U.S. (based on Barron and Mellow, 1979) and the marginal willingness to pay for unemployment attributes.

variables	Coefficient	MWP/AWB	MWP
weekly insurance benefit	-0.019		
	(0.006)[a]		
expected recall,	-5.950	4.065	313.158
within 30 days	(1.390)[a]	(1.674)[b]	(129.117)[b]
expected recall,	-2.740	1.873	144.211
no period specified	(0.739)[a]	(0.811)[b]	(62.442)[b]

Notes: Standard errors in parentheses. a: significantly different from zero at the 0.01 level; b: significantly different from zero at the 0.05 level. AWB: average weekly benefit.

Table 10. Coefficients of search contacts of unemployed individuals receiving unemployment benefit in 1982-1984, Leiden, The Netherlands (based on Lindeboom and Theeuwes, 1993) and the marginal willingness to pay for unemployment attributes as a percentage of the benefit.

variables	Coefficient	%MWP
log benefit	-0.27	
	(0.031)[a]	
residual entitlement period	-0.036	13.33
	(0.005)[a]	(2.40)[a]
(residual entitlement period)^2	0.0007	-0.26*residual entitlement period
	(0.0002)[a]	(0.08)[a]

Notes: Standard errors in parentheses. a: insignificantly different from zero at the 0.01 level. Residual entitlement period measured in weeks.

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