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Temporary jobs and job search effort in Europe

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ABSTRACT

Using longitudinal data on individuals from the European Community Household Panel (ECHP) for eleven countries during 1995–2001, I investigate temporary job contract duration and job search effort. The countries are Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. I construct a search model for workers in temporary jobs which predicts that shorter duration raises search intensity. Calibration of the model to the ECHP data implies that at least 75% of the increase in search intensity over the life of a 2+ year temporary contract occurs in the last six months of the contract. I then estimate regression models for search effort that control for human capital, pay, local unemployment, and individual and time fixed effects. I find that workers on temporary jobs indeed search harder than those on permanent jobs. Moreover, search intensity increases as temporary job duration falls, and roughly 84% of this increase occurs on average in the shortest duration jobs. These results are robust to disaggregation by gender and by country. These empirical results are noteworthy, since it is not necessary to assume myopia or hyperbolic discounting in order to explain them, although the data clearly also do not rule out such explanations.

1. Introduction

A considerable volume of economic research has been devoted over the last two decades to explaining and suggesting remedies for the stubbornly high unemployment rates in a number of European countries. Among the suggested policy remedies for reducing joblessness is the relaxation of systems of employment protection by allowing firms greater freedom to create temporary jobs. These reforms presumably reflect a desire to maintain protections for workers in permanent jobs while giving firms an incentive to create new, temporary jobs, which may ultimately become permanent. And even if they don't become permanent, temporary jobs may in some cases provide employment and work experience for individuals who would otherwise have been unemployed. On the other hand, policies increasing the freedom to create temporary jobs may instead encourage firms to substitute temporary for permanent jobs (as found by Kahn 2010), and, if so, the overall exit rate from jobs may increase. The resulting higher turnover may even lead to higher equilibrium unemployment than before (Blanchard and Landier 2002; Cahuc and Postel-Vinay 2002). Moreover, temporary jobs are known to pay less, offer less training, and be less satisfying than regular jobs (Booth et al., 2002; Boeri 2011; Kahn 2007; Stancanelli 2002). Thus, reforms that encourage the creation of temporary jobs may not lower unemployment and also may not unambiguously raise employed workers' utility (Blanchard and Landier 2002; Cahuc and Postel-Vinay 2002).

Policy evaluations of reforms that encourage temporary jobs must take into account the degree to which they become stepping stones to higher paying, permanent jobs. And evidence on this question of whether temporary jobs are stepping stones to permanent jobs is mixed (Booth et al., 2002; Autor and Houseman 2010). If workers are indeed seemingly trapped in temporary jobs, this outcome could have resulted either due to the lack of availability of permanent jobs or insufficient search effort on the part of workers. Of course, a greater supply of permanent jobs is likely to encourage greater search effort. But little is known about the search effort of those currently in temporary jobs. For example, do they anticipate the end of those jobs and begin searching in advance for future work, or do they wait until the last minute to begin their job search? A similar set of questions has been asked about unemployed workers whose unemployment insurance (UI) benefits are about to expire (Katz and Meyer 1990; Mortensen 1990; Boone and van Ours 2009), and to workers who have been given advance notice of layoffs (Addison and Portugal 1987; Swaim and Podgursky 1990; Ruhm 1992; and Jones and Kuhn 1995). The answers to these questions can have important

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implications for the transition from temporary to permanent jobs and therefore for evaluations of policies that allow firms to create temporary jobs. In addition, a study of search effort is potentially important in understanding aggregate matching functions, since the probability that any vacancy receives an applicant is affected by the average intensity of job search (Cahuc and Zylberberg 2004, p. 519).

In this paper, I use European Community Household Panel (ECHP) data to study the job search behavior of workers employed in temporary jobs in several European countries over the 1995–2001 period.[1] The countries included are Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. The ECHP collects information on current job search effort among employed workers (as well as of course the unemployed). In addition, the surveys include data on the duration of one's employment contract if it is temporary, allowing one to determine the impact of contract duration on search effort. I first build a simple model of employed job search that draws from search models in Burdett (1979) and Mortensen (1990) with nonstationary reservation wages.[2] A key theoretical result is that the less time left on a temporary contract, the greater is one's search effort, a result that is not surprising. However, calibration of the model using observed transition rates to permanent work and to temporary work implies that at least 75% of the increase in search intensity over the life a 2+ year temporary contract occurs in the last six months of the contract. It is noteworthy that this result is obtained without assuming hyperbolic discounting or myopia on the workers' part, although it is also of course consistent with such behavior (see, for example, Della Vigna and Paserman, 2005 or Paserman 2008). This result is similar to Mortensen's (1990) theoretical result that almost all of the reduction in an unemployed searcher's reservation wage occurs in the period before his/her unemployment benefits expire, a conclusion based on the formation of bounds for the rate of change of reservation wages, derived using US data on the incidence of layoff unemployment (p. 77).

The spike in exits from unemployment upon benefit exhaustion previous research has found is consistent with this model (Katz and Meyer 1990; Boone and van Ours 2009). However, Boone and van Ours (2009) suggest that this finding could be due to the endogenous timing of the start of a new job which may have been chosen to correspond with benefit exhaustion rather than a sharp reduction in reservation wages. Since the ECHP has information on actual search activity, I am able to provide a more direct test of the search model than an examination of exits from unemployment would. Similar reasoning may apply to studies of the transition from temporary jobs to permanent employment or to unemployment, which also don't examine actual search activity (D'Addio and Rosholm 2005; Gagliarducci 2005).

I then estimate an individual fixed effects model of the impact of contract duration on search effort as measured in the ECHP data base. In general, those on temporary contracts search harder than those in permanent jobs, as one would expect. And search intensity increases going from the longest to the shortest duration temporary contracts, again as one would predict. Moreover, almost all of the increase in search intensity going from longest to shortest duration jobs occurs between the second shortest (6–12 months) and the shortest (less than 6 months) duration jobs, as predicted by the calibrated Burdett-Mortensen model. This finding occurs in data pooled across countries, separately by gender, and within countries analyzed individually in the aggregate and again by gender. The basic pattern thus is pervasive in European labour markets. Since I control for individual fixed effects, the findings cannot be explained by a possible correlation between an individual's fixed search propensity and the likelihood of landing a long duration job, although correlations between unmeasured changes in individual productivity and search intensity could of course still be a factor.

It thus appears that workers are indeed forward-looking in their job search behavior; however, the optimizing strategy is to not start searching intensively early in the term of one's temporary job, like that of unemployed workers with limited duration unemployment benefits. Some countries have reformed their regulations of temporary employment contracts by increasing their allowable duration (OECD 2004). Blanchard and Landier (2002) have shown that such reforms raise the threshold a worker must attain in order to be promoted from a temporary job to a permanent job. My findings suggest an additional reason why such reforms will reduce the transition rate to permanent jobs. Specifically, an implication of the results obtained here is that such policies will reduce the average search intensity of workers on temporary jobs, perhaps lessening the per period transition rate to a permanent job. Thus, while earlier research on firm behavior under these recent reforms suggests that they reduce the supply of permanent jobs (Blanchard and Landier 2002), my results imply that they also reduce the effective demand for such jobs by reducing search intensity. The dualistic labour market structure suggested by the co-existence of permanent and temporary jobs thus may be reinforced by these reforms, due both to firm and worker behavior. Nonetheless, policies allowing firms to offer longer temporary contracts can increase the total time workers in them are employed and may also give workers on temporary jobs additional opportunities to be promoted into permanent jobs.

2. A simple model of temporary jobs and search intensity

2.1. Model setup

In this section, I write down a simple model that sheds some light on the impact of a temporary contract's remaining duration on the employed worker's search intensity. Like earlier models of search intensity such as those of Della Vigna and Paserman (2005) and Paserman (2008), I use a discrete time framework and assume that a jobseeker will receive a wage offer with some probability in any given period. Moreover, one can raise this probability by searching harder and this increased search effort (e.g. putting in more time or money to the search effort) will be costly. I allow the new job offer to be either a permanent job or a temporary job, although the jobseeker doesn't know in advance what kind of job if any will be offered by a contacted firm. To simplify the analysis, I assume that permanent jobs never end and temporary jobs last T periods. By assuming that permanent jobs never end, I am in effect assuming that firing costs are high enough to deter all firing in permanent jobs. This is a stylized assumption that makes the analysis simpler but does not affect the overall conclusions about incentives to search in temporary vs. permanent jobs. As long as the probability of being able to continue in a job is lower in a temporary job than in a permanent job or in a temporary job with many periods left than with only one period left, then the qualitative conclusions of the model regarding search intensity will still hold. One can view conversions of temporary contracts into permanent jobs by a worker's incumbent firm as an outcome of the job search process. After presenting the formal search model and its calibration, I will discuss the implications of such conversions in more detail.

Let λ be the probability of receiving a permanent job offer and assume that the probability of receiving a temporary job offer is $a\lambda$,

[1] In an earlier paper (Kahn 2010), I studied the reduced form effects of reforms of European employment protection regulations on the incidence of temporary and permanent employment. In this paper, I study the transition process connecting temporary and permanent employment by examining the impact of contract duration on workers' job search intensity.

[2] See van den Berg (1990) and Card and Hyslop (2005) for additional models of nonstationary reservation wages, which like Burdett (1979) and Mortensen (1990) do not consider variable search intensity. Mortensen's (1977) earlier paper on UI does build a model of search intensity and reservation wages which predicts that search intensity will increase over the duration of coverage by UI benefits. However, he does not calibrate the rate of increase of intensity.

where a>0; following Mortensen (1990), assume that one can receive at most one job offer per period. I assume that the job seeker can affect λ at search cost c(λ), and that search costs are quadratic:[3]

$$c(\lambda) = .5\lambda^2. \tag{1}$$

Thus, by investing more resources in search costs, the job seeker raises the probability of receiving a permanent or a temporary job offer, and I make no assumptions about the relative probabilities of the two types of job offer. I only assume that greater search effort raises each probability. Wages within a job are assumed constant, for convenience. Let B be the discount factor (i.e. $B = 1/(1+r)$ where r is the discount rate). Further, let b be the value of unemployment benefits which are assumed for simplicity to be available indefinitely, and assume that search is no more efficient while unemployed than while employed in a temporary job.[4]

2.2. Main analytical results: falling reservation wages and rising search intensity with time on a temporary job

It can be shown that the jobseeker will in all cases follow a reservation wage policy. However, because the value of a wage offer x in a permanent job is different from the value of a temporary wage offer of x, the jobseeker may select a different reservation wage for permanent offers vs. temporary offers. Therefore, let R_0, and $R_n(w)$ be the reservation wages for permanent job offers facing, respectively, unemployed searchers, and temporary job holders with n periods remaining in their temporary jobs, which are assumed to pay wage w; let T_0 and $T_n(w)$ be the corresponding reservation wages for temporary job offers. The goal of this section is to study the behavior of reservation wages and search intensity with time in a temporary job. By the definition of reservation wages, we have:

$$W(R_0) = V_0 \tag{2}$$

$$V_T(T_0) = V_0, . \tag{3}$$

where V_0 is the value of being unemployed. In other words, the reservation wages for each type of offer are set so as to make the searcher indifferent between continued search and accepting the job.

To determine reservation wages for the unemployed, we need an expression for V_0. Under the assumptions of the model, this can be solved as:

$$\begin{aligned} V_0 = b + B\lambda_0 E\big[Max(W(x), V_0\big] + Ba\lambda_0 E\Big[Max\Big(V_T(x), V_0\Big] \\ + B(1-\lambda_0-a\lambda_0)V_0 - .5(\lambda_0)^2, \end{aligned} \tag{4}$$

where the expectation in each case is taken with respect to the distribution of permanent or temporary job wage offers x given an offer.

Eq. (4) states that the present value of being unemployed equals the current unemployment benefit b plus the expected discounted value of future labour market states minus current search costs. In the next period, there are three possible job offer outcomes: one can have received a permanent job offer (with endogenous probability λ_0), a temporary job offer (with probability $a\lambda_0$), or one can have not received any offers (with probability 1- λ_0 -$a\lambda_0$). If one has received a job offer of either type, one then needs to decide whether to accept it by comparing the value of accepting it with the value of continuing unemployment.

Eq. (4) can be simplified by using the reservation wage property:

$$V_0 = b + B\lambda_0 \int_{R0}^{\infty} [W(x) - V_0]dF(x) + Ba\lambda_0 \int_{T0}^{\infty} [V_T(y) - V_0]dG(y) + BV_0 - .5(\lambda_0)^2, \tag{5}$$

where F(.) and G(.) are respectively the distribution functions for wage offers in permanent and in temporary jobs. These are assumed to be given from the jobseeker's point of view, although of course in a general equilibrium framework they would be endogenous with respect to aggregate search behavior. One could also endogenize wages at the individual jobseeker-employer match level through bargaining effects, although one's incentives to search will still rise as the time remaining on one's temporary job falls. Some of these incentives may be attenuated if firms make offers to deter search. Because of this possibility, in some alternative empirical specifications discussed below, I excluded current wages on the idea that they may be endogenously determined at the firm-worker match level.

Search effort λ_0 can be calculated by maximizing (5) with respect to λ_0. Assuming an interior solution, we have:

$$\lambda_0 = B \int_{R0}^{\infty} [W(x) - V_0]dF(x) + Ba \int_{T0}^{\infty} [V_T(y) - V_0]dG(y). \tag{6}$$

Search effort is positively affected by the expected gains to an accepted job offer.[5]

Lemma 1. Reservation Wages in the Final Period of a Temporary Job are the Same as for the Unemployed Worker.

Proof. For those with one period remaining in their temporary job, which is assumed to pay a wage w, we have:

$$\begin{aligned} V_1(w) = w + B\lambda_1(w) \int_{R1(w)}^{\infty} [W(x) - V_0]dF(x) + Ba\lambda_1(w) \int_{T1(w)}^{\infty} [V_T(y) - V_0]dG(y) \\ + BV_0 - .5(\lambda_1(w))^2, \end{aligned} \tag{7}$$

where a 1 subscript on the value function, reservation wages and search intensity refers to the time remaining on the current temporary job. Since the current job will end in the next period, the value of turning down a job offer or not receiving a job offer is the same as it was for the unemployed searcher. Therefore, the reservation wages are the same in the last period of employment as they are for the unemployed searcher. By implication, so is the optimal search intensity $\lambda_1(w)$. Period 1 reservation wages and search intensity are therefore independent of the current wage. The only difference between the value of being in the last period of a temporary job and being unemployed is the wage payment w relative to the unemployment benefit b. As long as the job pays more than UI benefits, being in the last period of a temporary job is more valuable than being unemployed.

Lemma 2. The Value of Being in a Temporary Job Falls and Reservation Wages Fall as Time Left in the Job Decreases from n>1 to n=1.

Proof. See the Appendix.

This result is intuitive because a promise of n-1 periods at a wage greater than UI benefits is worth strictly more than a promise of n-2 periods at the same wage.

[3] These assumptions for search costs and the probability of an offer are made for convenience and are innocuous since we can parameterize search costs appropriately.

[4] The model could easily be modified to allow different probabilities of receiving a permanent or temporary job offer while employed than while unemployed, as in Mortensen's (1990) model of workers awaiting recall or in Kugler and Saint-Paul's (2004) model of adverse selection signaled by one's being unemployed. But the basic features would remain the same as in the simpler version presented here, and, below, I discuss the likely consequences of allowing for different offer probabilities depending on whether one is currently employed.

[5] Eq. (6) is one of the first order conditions of the multivariate maximization problem in which the jobseeker simultaneously chooses the two reservation wages and search intensity each period. A similar theoretical outcome appears in earlier models of search intensity such as Mortensen (1977) and Paserman (2008), although those papers did not consider the choice between temporary and permanent jobs.

Proposition 1. Search Intensity Rises as the End of a Temporary Job Approaches

Proof. Consider the following expressions for search intensity with n-1 and n (n>=2) periods left in a temporary job, where V_0 remains the value of being unemployed:

$$\lambda_{n-1}(w) = B\int_{Rn-1(w)}^{\infty}[W(x)-V_{n-2}(w)]dF(x) + Ba\int_{Tn-1(w)}^{\infty}[V_T(y)-V_{n-2}(w)]dG(y) \tag{8}$$

$$\lambda_n(w) = B\int_{Rn(w)}^{\infty}[W(x)-V_{n-1}(w)]dF(x) + Ba\int_{Tn(w)}^{\infty}[V_T(y)-V_{n-1}(w)]dG(y). \tag{9}$$

(these are the same first order conditions used to solve for λ_0).

By Lemma 2, for n>=2, $V_{n-1}>V_{n-2}$; therefore, the integrands are larger in Eq. (8) than in Eq. (9) and so is the range of integration. Thus, search intensity must rise as the end of the job approaches. Proposition 1 is intuitive in that the return to search rises as the time left in one's temporary job falls, since the value of not searching has fallen (e.g. a job with one period left is worth less than a job with two periods left). We are now in a position to place some bounds on the speed with which search intensity rises.

2.3. Calibration of the model

We have just shown that search intensity rises over the life of a temporary job. The goal of this subsection is to place some bounds on the increase during last period ($\lambda_1(w)-\lambda_2(w)$) relative to the total increase in search intensity over the life of a temporary job (i.e. $\lambda_1(w)-\lambda_T(w)$). I begin by proving an inequality relating to changes in the rate of increase in search intensity during the temporary job.

Lemma 3. $[\lambda_n(w)-\lambda_{n+1}(w)]/[\lambda_{n-1}(w)-\lambda_n(w)]<[V_n(w)-V_{n-1}(w)]/[V_{n-1}(w)-V_{n-2}(w)]$.

Lemma 3 establishes an upper bound for the relative rate of increase in search intensity earlier vs. later in the temporary job.

Proof: by the properties of the reservation wage, we have:

$$\lambda_{n+1}(w)>B\int_{Rn(w)}^{\infty}[W(x)-V_n(w)]dF(x) + Ba\int_{Tn(w)}^{\infty}[V_T(y)-V_n(w)]dG(y). \tag{10}$$

Inequality (10) holds because a) for all values of the permanent wage offer x less than $R_{n+1}(w)$, $W(x)<V_n(w)$ and for all values of the temporary wage offer y less than $T_{n+1}(w)$, $V_{n+1}(y)<V_n(w)$.[6] Therefore, by Eq. (9), we have:

$$0<\lambda_n(w)-\lambda_{n+1}(w)<B(V_n(w)-V_{n-1}(w))(1-F(R_n(w)) + Ba(V_n(w)-V_{n-1}(w))(1-G(T_n(w)). \tag{11}$$

By the same reasoning that led to expression (10), we have:

$$\lambda_{n-1}(w)>B\int_{Rn(w)}^{\infty}[W(x)-V_{n-2}(w)]dF(x) + Ba\int_{Tn(w)}^{\infty}[V_T(y)-V_{n-2}(w)]dG(y). \tag{12}$$

[6] These inequalities therefore hold for all wages between R_n and R_{n+1} and T_n and T_{n+1}, since $R_n<R_{n+1}$ and $T_n<T_{n+1}$.

Therefore,

$$\lambda_{n-1}(w)-\lambda_n(w)>B(V_{n-1}(w)-V_{n-2}(w))(1-F(R_n(w)) + Ba(V_{n-1}(w)-V_{n-2}(w))(1-G(T_n(w)). \tag{13}$$

Dividing inequality (11) by inequality (13) proves the desired result in Lemma 3. The first order condition for search intensity (e. g. Eq. (9)) establishes a relationship between search intensity at period n and the value of temporary employment in period n-1, and the reservation wage property allows us to form the key bounds in expressions (10) and (12) which directly lead to Lemma 3.

Proposition 2. Let E_n be the probability of not moving from a temporary job with n periods left (i.e., $E_n=\{1-\lambda_n(w)(1-F(R_n(w)+a-aG(T_n(w))\}$— recall that I assume getting a new temporary job offer and getting a new permanent job offer are mutually exclusive events). Then

$$[\lambda_n(w)-\lambda_{n+1}(w)]/[\lambda_{n-1}(w)-\lambda_n(w)]<BE_n.$$

Proof. The maximum value of being employed at wage w in a temporary job with n-1 periods left must be greater than the value of choosing at time n-1 the period n reservation wages and search intensity (assuming a unique reservation wage):

$$V_{n-1}(w)>w+B\lambda_n(w)\int_{Rn(w)}^{\infty}[W(x)-V_{n-2}(w)]dF(x) + Ba\lambda_n(w)\int_{Tn(w)}^{\infty}[V_T(y)-V_{n-2}(w)]dG(y) + BV_{n-2}-.5(\lambda_n(w))^2 \tag{14}$$

Therefore, by Lemma 3 and inequality (14), we have:

$$[\lambda_n(w)-\lambda_{n+1}(w)]/[\lambda_{n-1}(w)-\lambda_n(w)]<[V_n(w)-V_{n-1}(w)]/[V_{n-1}(w)-V_{n-2}(w)]<BE_n. \tag{15}$$

As was the case with Lemma 3, the reservation wage property directly implies the key inequality (14) which leads to the bound in expression (15).

Proposition 3. $(\lambda_1(w)-\lambda_T(w))<(\lambda_1(w)-\lambda_2(w))(1+BE_2+B^2E_2E_3+\ldots+B^{T-1}(E_2E_3\ldots E_{T-1}))$.

Proposition 3 provides the bound for the relative increase in search intensity in the last period vs. the entire duration of the temporary job.

Proof. We would like an estimate of the increase in search intensity in the last period (i.e. $\lambda_1(w)-\lambda_2(w)$) relative to the total increase in search intensity over the life of a temporary job (i.e. $\lambda_1(w)-\lambda_T(w)$). To estimate this relative increase, write ($\lambda_1(w)-\lambda_T(w)$) as:

$$(\lambda_1(w)-\lambda_T(w)) = (\lambda_{T-1}(w)-\lambda_T(w))) + (\lambda_{T-2}(w)-\lambda_{T-1}(w))) + \ldots + (\lambda_1(w)-\lambda_2(w)). \tag{16}$$

Successive use of Proposition 2 and Eq. (16) yields the desired result. Proposition 3 breaks down the increase in search intensity over the life a temporary job into a sum of one period increases, about which Proposition 2 provides the lower bound for the relative increase later vs. earlier in the job.

With data on the discount factor and the transition probabilities out of temporary jobs, we can use Proposition 3 to form a lower bound for the portion of the increase in search intensity that occurs during the final period of a temporary job. While the ECHP data aren't fine enough to allow one to follow people within their temporary jobs, we can use the data to compare people with different total temporary contract durations or the same person in different jobs with different total durations. For example, a randomly chosen person with a 2 year contract will have on average more time remaining on his/her contract than a randomly chosen person on a one year contract, and so on. Thus, comparing people under different contracts will be similar to comparing people with different amounts of time remaining on their temporary job, as the model depicts. As shown below, in the ECHP data, the potential durations of temporary jobs are defined in four categories with enough observations on which to perform meaningful statistical analyses: under 6 months; 6 months to under a year; one year to under two years; and two years or more. This division of the data by the ECHP suggests considering a period to be 6 months and therefore that $T=5$. That is, period 1 is the less than 6 months category. Increasing by 6 month increments, we arrive at period 5, which is duration 2 to 2.5 years, or the highest category: period 2 is 6–12 months, period 3 is 12–18 months, period 4 is 18–24 months, and period 5 is 24+ months. This means that we need to use the following limit for the total increase in search intensity relative to the increase in the last period:

$$(\lambda_1(w)-\lambda_5(w))<(\lambda_1(w)-\lambda_2(w))\left(1+BE_2+B^2E_2E_3+B^3E_2E_3E_4\right). \quad (17)$$

However, periods 3 and 4 (12-18 months) and (18-24 months) are aggregated by the ECHP, so we must set $E_3=E_4$.

Appendix Table A1 shows transition rates from temporary jobs of various total duration levels. If we make the maintained hypothesis that differences in behavior across durations are the same as that for an individual as the time left in his/her temporary job falls, then we can use these transition rates to compute E_2, E_3, and E_4). The data in Table A1 imply that E_2 is 0.232 (1-.475-.293) and E_3 (and therefore E_4) is 0.344 (i.e., 1-.169-.487).[7] Using a 6 month discount factor B of 0.985 (i.e., a period is 6 months, and I am assuming an annual real discount rate of 3%), inequality (17) implies that at least 75% of the total increase in job search intensity going from the longest to the shortest duration temporary job should occur in the last period.[8] We therefore expect to see sharply increasing job search intensity as temporary job durations fall. This result does not assume myopia or hyperbolic discounting; however, we do predict very gradually rising search intensity throughout one's employment in a temporary job until the last period. If jobseekers are completely unresponsive to changes in the duration of their jobs, then we would conclude that they are myopic.

The search model's calibration implies that forward-looking workers on long duration temporary jobs will find it optimal to essentially wait to until the last period—the last 6 months–to substantially increase their search intensity, although there will be a small increase in search effort before then. Thus, shortening the potential duration of such jobs is likely to have the same kind of effect on jobseekers as limiting the duration of UI benefit receipt. The model predicts higher search intensity throughout a temporary job if its duration is shorter, and a sooner rapid increase in intensity after the job starts than if duration is allowed to be longer.

The model just outlined assumes that one's probability of a job offer given search effort is the same regardless of whether one is employed or unemployed. In reality, some temporary job contracts are renewed when they expire, and firms promote some workers from temporary jobs into permanent jobs. The search model presented above can be modified conceptually to accommodate the possibility of promotion into permanent jobs. Suppose, for example, that while one is employed in a temporary job with n periods left, the probability of finding a permanent job is:

$$(\lambda_n(w)+d),$$

with $d>0$ being the probability of promotion into a permanent job even if one doesn't search. If d is constant over time, then the first order condition for search intensity has the same form as before, since d doesn't interact with λ_n. Reservation wages in period 1 will still be the same as reservation wages while unemployed, but the value of employment will have risen due to the possibility of promotion at zero search cost. But as long as d is constant across time, then the model will still predict that the value of employment will fall as the end of the job approaches, leading to falling reservation wages and thus rising search intensity. On the other hand, if d falls over time, then this would be an additional reason for observing rising search intensity as the end of a temporary job approaches; conversely, if d rises as the end approaches (perhaps because the firm wants to retain the worker), then search intensity might not rise. Thus the likely impact of promotions on search behavior depends on the relationship between temporary contract duration and the probability of promotion. In the empirical analysis below, I investigate this issue.

It should be noted that it is theoretically possible for employees to affect the promotion probability by working harder, as in Givord and Wilner's (2009) analysis for France. While the authors didn't find evidence that high effort levels enhanced one's promotion probability, workers with very low effort levels (measured by work hours) had a lower promotion probability.[9] Additionally, Dolado and Stucchi (2008) find indirect evidence from Spain that a higher perceived probability of conversion of temporary to permanent jobs raises total factor productivity and thus, implicitly, worker effort. Allowing for the possibility of effort-based promotions would lead to a more complicated model that would include both search effort and work effort on the job. Since the ECHP doesn't have good measures of actual effort on the job, I must abstract from this additional issue. Further, if search is more efficient while unemployed, then this would reduce the gap between $V_1(w)$ and V_0, reducing the rate at which search intensity rises as time left in one's temporary job falls.

Finally, lowering search costs (e.g. the expansion of employment agencies or direct job search subsidies) raises the value of being unemployed and therefore reservation wages on all types of jobs. Moreover, lowering search costs raises the relative value of temporary employment compared to permanent employment, since an inherent feature of temporary jobs is the likelihood of needing to search for a new one. Therefore, reservation wages of temporary jobs will rise by less than those of permanent jobs, and such policies may raise the relative incidence of temporary employment.

3. Data and descriptive patterns

I use the ECHP data for 1995–2001 for the following countries to study the impact of temporary employment contracts on job search: Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal, and Spain.[10] This is a panel data base that follows individuals over the 1994–2001 period. The questions were harmonized as much as possible in order to produce a data base that would provide comparable information across countries.[11] Beginning in 1995

[7] The total transition rate is actually somewhat higher for the shortest duration temporary jobs than for the other categories, as the search model predicts. But the transition to permanent jobs is actually highest for those in the longest temporary jobs. It is possible that the respondents in the different duration temporary jobs differ in measurable or unmeasurable ways that could affect their transition probabilities. The empirical work below controls for measured factors as well as person-specific unmeasured factors that would affect search intensity.

[8] Note that the bound would become much more conservative, approaching 20% if the discount rate is zero and the staying probability near 1. One interpretation of this more conservative bound is that search effort will start earlier if workers value the future more and if the offer probability given search effort is low.

[9] This interpretation of the causal impact of low work hours depends of course on the assumption that the worker had control of his/her work time. It could also have been true that the firm assigned low productivity temporary workers to work low hours even if promotions weren't affected by work effort.

[10] Of the fifteen countries in the ECHP, these eleven are the only ones with data on both on the job search and contract type (i.e., permanent vs. temporary) and with repeated observations on the same person.

[11] For further description of the methods and sample characteristics of the ECHP, see the Eurostat web site: http://circa.europa.eu/irc/dsis/echpanel/info/data/information.html.

Table 1
Contract duration and on the job search effort. 1995-2001 (employed workers).

Employment contract type	Fraction of sample	Incidence of on the job search	Incidence of active search behavior	Sample size
A. Temporary Contract				
<6 months	0.022	0.276	0.217	5672
6 months to<1 year	0.039	0.175	0.127	9915
1 year to under 2 years	0.019	0.179	0.132	4721
2 years to under 5 years	0.010	0.140	0.106	2308
5 years or more	0.004	0.152	0.096	779
2 years or more	0.013	0.143	0.103	3087
B. Permanent Contract	0.908	0.074	0.048	213919
Total	1.000	0.085	0.057	237314

Source: ECHP data. Adjusted sampling weights used, where the raw weights are modified so that each country receives the same total weight. Sample is limited to those age 16-65 from the following countries: Austria, Belgium Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal, and Spain.

Table 2
Contract duration incidence by country 1995-2001.

	Temporary contracts with duration					
	<6 months	6 months to <1 year	1 year to <2 years	2 years or more	Permanent contracts	Sample size
Austria	0.010	0.014	0.013	0.014	0.949	18221
Belgium	0.015	0.035	0.017	0.016	0.917	14449
Denmark	0.012	0.017	0.015	0.013	0.943	15669
Finland	0.039	0.049	0.031	0.013	0.867	18676
France	0.028	0.030	0.014	0.014	0.914	30315
Greece	0.011	0.046	0.013	0.017	0.913	14746
Ireland	0.013	0.014	0.015	0.010	0.948	13642
Italy	0.021	0.028	0.013	0.011	0.927	32172
Netherlands	0.012	0.007	0.006	0.007	0.967	27638
Portugal	0.014	0.073	0.021	0.009	0.884	25914
Spain	0.066	0.109	0.051	0.021	0.753	25872
Total	0.022	0.038	0.019	0.013	0.908	237314

for all of these countries except Finland and in 1996 for Finland, the ECHP asked each employed wage and salary worker whether his/her job was characterized by a fixed term contract. Specifically, each employed respondent is asked: "What type of employment contract do you have in your main job?" The possible responses are: 1) permanent employment; 2) fixed-term or short-term contract; 3) casual work or no contract; 4) some other working arrangement. For the purposes of analyzing the determinants of temporary employment, I include only those with responses 1) or 2), that is, those that state they have a permanent or a temporary employment contract. Respondents with a temporary contract were asked how long the total duration of their contract was, with possible responses: less than 6 months, 6 months to less than a year, 1 year to under 2 years, 2 years to under 5 years, 5 years or more.

To gauge on the job search activity, I use two questions from the ECHP. First, I use responses to the question asking employed workers whether they are looking for a job. Second, the ECHP asks whether in the last four weeks, a respondent has taken active steps to find a job. Examples given by the survey include: "contacted a public employment office,…, applied to an employer, studied or replied to advertisements, contacted a private employment or vocational guidance agency, asked friends or contacts, or taken steps to start your own business" (ECHP codebook, p. 273). In the empirical work below, I examine responses to both questions. The second question (about taking active steps) is more closely related to search effort than the first one, although the results were very similar for either measure of on the job search activity. Unfortunately, the ECHP does not include further measures of search intensity such as the time devoted to job search or the number of search methods used.

Tables 1–4 provide some descriptive information about contract duration and search activity. All statistics are weighted using the ECHP's provided person weights, and these have been adjusted in the data pooled across countries so that each country receives the same weight. Included in the tables are all employed workers with complete data on the explanatory variables used below and who have either a known fixed contract duration or a permanent job. The age range is restricted to 16–65 years. Table 1 provides these data aggregated across the eleven countries listed above. About 9% of the sample has a temporary contract, and the most common duration is 6–12 months (about 42% of temporary jobs), followed by less than 6 months (24%), and 1–2 years (21%).[12] A very small fraction have 5 years or more duration (4%). The incidence of on the job search and active search behavior look at first blush to be consistent with the theoretical model outlined earlier. First looking at the figures for on the job search, the fraction of workers searching rises from 0.074 of those in permanent jobs to 0.143 for those with at least two years' duration on a temporary contract.[13] This figure rises again to 0.175-0.179 for those with 6 months to two years duration, and rises sharply to 0.276 for those with the shortest contract duration (under 6 months). In other words, the incidence of search activity rises by 20.2 percentage points between those with permanent jobs and those with the shortest temporary jobs, and 10.1 percentage points of this rise occurs between the 6–12 months duration and <6 months duration categories. Moreover, among those with temporary contracts, search incidence rises by 13.3 percentage points from the 2+ years category to the shortest category, with, as just noted, 10.1 percentage points or 76% of the rise occurring in between the two shortest duration categories.

Table 1's figures for search intensity (the incidence of active search behavior) are very similar to those for the incidence of job search. 4.8% of those on permanent contracts have engaged in active search behavior in the last four weeks (compared to 7.4% who said they were looking for a new job), a figure that rises to 10.3% for those with temporary jobs with at least two years' duration and finally to 21.7% of those on the shortest temporary contracts. Again, for those on temporary contracts, 79% of the increase in search intensity between the longest and the shortest temporary contracts occurs between the two shortest duration categories. But there is still a slight increase in search intensity from the 2 + years category to the 6–12 months duration category.

Overall, then, workers appear to be forward-looking in the sense that the shorter one's employment contract, the more likely one is to search and the more intensively one searches. But most of the increase in search activity occurs for those in the shortest duration category. This result is especially noteworthy because the difference in expected duration between the two shortest categories likely to be less than for the other categories; this is the case because the potential duration only increases by 6 months going from the shortest to the second shortest category, while the increase is at least a year between the other pairs of adjacent temporary job duration categories. This set of outcomes is precisely what is predicted by the search model outlined earlier. Tables 2–4 examine whether this pattern is common to each of the countries individually. Table 2 shows that for all of the countries except the Ireland and Netherlands, the 6–12 months duration category is the most common temporary duration, while the least common is usually the 2+ years duration jobs. Tables 3 and 4 show a remarkable consistency across countries in the incidence and intensity of job search

[12] Earlier work has shown that the ECHP data on the incidence of temporary employment contracts match up well with published sources such as the OECD. See Kahn (2010).

[13] In the empirical work, I will be aggregating the 2–5 years and 5 years plus categories because of the small numbers of cases in the later duration category. Table 1 shows that these categories have very a similar incidence of on the job search and search intensity.

Table 3
Incidence of on the job search by contract duration.

	Temporary contracts with duration				
	< 6 months	6 months to <1 year	1 year to <2 years	2 years or more	Permanent contracts
Austria	0.124	0.105	0.120	0.102	0.038
cell size	171	234	255	224	17337
Belgium	0.270	0.184	0.212	0.137	0.063
cell size	227	498	242	205	13277
Denmark	0.344	0.294	0.260	0.197	0.091
cell size	176	282	262	192	14757
Finland	0.254	0.203	0.179	0.168	0.092
cell size	757	926	595	257	16141
France	0.417	0.275	0.212	0.109	0.050
cell size	880	964	437	440	27594
Greece	0.274	0.180	0.211	0.115	0.055
cell size	170	793	200	267	13316
Ireland	0.157	0.142	0.210	0.093	0.065
cell size	189	219	218	160	12856
Italy	0.333	0.228	0.272	0.214	0.061
cell size	746	1006	441	403	29576
Netherlands	0.394	0.386	0.429	0.314	0.206
cell size	310	179	160	191	26798
Portugal	0.160	0.076	0.070	0.077	0.024
cell size	300	1950	604	183	22877
Spain	0.250	0.163	0.127	0.131	0.059
cell size	1746	2864	1307	565	19390
Total	0.276	0.175	0.179	0.143	0.074
cell size	5672	9915	4721	3087	213919

Table 4
Incidence of active on the job search measures by contract duration.

	Temporary Contracts with Duration				
	< 6 months	6 months to <1 year	1 year to <2 years	2 years or more	Permanent contracts
Austria	0.118	0.090	0.092	0.077	0.028
cell size	171	234	255	224	17337
Belgium	0.178	0.126	0.132	0.092	0.037
cell size	227	498	242	205	13277
Denmark	0.284	0.229	0.206	0.157	0.073
cell size	176	282	262	192	14757
Finland	0.217	0.163	0.137	0.150	0.068
cell size	757	926	595	257	16141
France	0.296	0.169	0.129	0.071	0.033
cell size	880	964	437	440	27594
Greece	0.170	0.107	0.142	0.074	0.036
cell size	170	793	200	267	13316
Ireland	0.124	0.108	0.169	0.075	0.049
cell size	189	219	218	160	12856
Italy	0.265	0.154	0.229	0.142	0.044
cell size	746	1006	441	403	29576
Netherlands	0.303	0.248	0.245	0.193	0.097
cell size	310	179	160	191	26798
Portugal	0.122	0.051	0.052	0.044	0.013
cell size	300	1950	604	183	22877
Spain	0.210	0.134	0.101	0.099	0.046
cell size	1746	2864	1307	565	19390
Total	0.217	0.127	0.132	0.103	0.048
cell size	5672	9915	4721	3087	213919

in the various employment contract duration categories. In each case, those in permanent jobs are least likely to search or have the smallest amount of search activity, while those in the shortest temporary jobs search the hardest, with the exception of Ireland (Table 4). In addition, in most cases the largest increase in search activity among the temporary job holders occurs between the 6–12 months and the less than 6 months duration categories.

3.1. Empirical procedures and regression results

Tables 3 and 4 show that search behavior is consistent with the model discussed earlier, which of course did not assume myopia or hyperbolic discounting. In the empirical work that follows, I test whether these patterns hold up controlling for worker human capital, pay or economic conditions, as well as individual worker fixed effects. For example, it is possible that the shortest duration jobs pay lower wages than longer duration temporary jobs, and these purportedly lower wages could in principle explain the patterns in Tables 1, 3 and 4.

The basic empirical setup for testing the job search model presented earlier is to estimate the intensity of search as a function of contract duration and control variables:

$$\text{Active Search} = f(\text{dur0}-6,\ \text{dur6}-12,\ \text{dur12}-24, \text{dur24}+,\ X,\ u), \tag{18}$$

where for each employed individual, Active Search is a dummy variable for having taken active measures to find a job in the last four weeks, dur0-6, dur6-12, dur12-14, dur24+ are dummy variables for being a temporary job with respectively, less than 6 months, at least 6 but less than 12 months, at least 12 but less than 24 months, and at least 24 months total duration, X is a vector of control variables to be discussed below, and u is a disturbance term.[14]

In Eq. (18), the dependent variable is the ECHP's proxy for search intensity, although I also estimated models with an employed search dummy variable as dependent variable, with very similar results to those presented below. The duration variables correspond to the categories in Tables 2–4, and the omitted category is those who have permanent jobs. While, as noted earlier, the duration variables refer to total contract length, the person's remaining duration will on average be less than this.[15] Therefore the duration dummy variable categories likely measure the remaining duration with some error. On the other hand, given the total potential duration of one's contract, the remaining duration is clearly endogenous with respect to one's search effort. One might therefore view the duration variables as instruments in a model where the first stage determines the remaining duration, and the second stage determines search intensity. One can thus interpret Eq. (18) as a reduced form of such a process.

The controls include age, age squared, dummy variables for low (ISCED levels 0–2) and middle levels (ISCED level 3) of schooling with high levels of schooling the omitted category (ISCED levels 5–7), the log of hourly earnings expressed in purchasing power parity units in 2001 US dollars, the regional unemployment rate, and year dummy variables.[16] Individual fixed effects models automatically control for country and gender level effects, although in some models I stratify by gender and country. While, as discussed above, hourly earnings may be endogenous with respect to search activity, it is a potentially important control, since 6 month jobs may differ in quality from, for example, 24 month jobs. And we would expect job match

[14] Because active search is defined to occur over the previous four weeks, it is theoretically possible that some jobs began after this reported search. If so, it would be inappropriate to call this kind of search employed search from the current job. To examine this possibility, I estimated the model on a subsample of people who had been on their job at least two months. For this group, the search activity variable definitely refers to the current job, and the results were virtually identical to those presented below for the full sample.

[15] Unfortunately, the ECHP doesn't include information on the contract type of any intervening jobs that may have occurred between interviews. Therefore, we can't observe changes for the same individual in time remaining on a given temporary job. While there is information on when the current job started, we can't know for sure what type of job contract it started with.

[16] The ECHP provides purchasing power parity rates for each country in each year, allowing one to transform the earnings data into US purchasing power units for that year. These transformed earnings variables were then corrected for US inflation by using the Personal Consumption Expenditures deflator for the US, taken from www.bea.gov. I excluded observations with hourly earnings less than $1 or greater than $300 in 2001 purchasing power parity units. These exclusions amounted to about 0.2% of the sample.

quality (as proxied by wages) to independently affect one's search propensity, both directly as in the model presented earlier and perhaps also as a signal that one is likely to be promoted to a permanent job. Including current wages, therefore, increases the likelihood that the key duration variables reflect actual time remaining on the job rather than the overall quality of the job.[17] The regional unemployment rate information was collected from the European Labour Force Survey and matched to the regional indicators in the ECHP data.[18] The unemployment rate and human capital controls account for likely wage offers relative to the current wage, which is also a control. Year dummies account for continent-wide economic factors, as well as for the value of the US dollar in purchasing power. The standard errors were clustered at the country level.

The basic model is linear and is estimated using fixed effects, where I take advantage of the longitudinal nature of the ECHP data. The standard errors are, as suggested above, robust to heteroskedasticity.[19] I also estimated some models using conditional fixed effects logit analysis, with very similar results to those presented for the linear model. In either the linear or the logit case, the basic variation in the duration of the contract comes from individuals who change contract type[20]; however the linear model uses the whole sample, while the fixed effects logit model only includes those who changed search intensity. Thus the linear model may provide more efficient estimates of the effects of the control variables. The fixed effects models account for possible spurious correlation between an individual's propensity to search and the type of job one has. For example, if most workers want a permanent job, then other things being equal, those who are most willing to look hard for work will be most likely to have permanent jobs. If this willingness is a fixed trait, then we may observe a spurious negative correlation between search intensity and the incidence of temporary work. Fixed effect models can account for this possibility. It is also possible that those with better prospects for landing a permanent job may be more likely to accept a short term job in the first place. Thus, any correlation between short jobs and search activity may reflect one's underlying job prospects rather than the search model presented earlier. Again, however, to the extent that these job prospects are due to person-specific effects, the fixed effects will account for such a possibility.

In addition to the basic Eq. (18), which constrains the effects of job duration to be same across the sample, I also estimated the basic model separately by gender and also separately by country. These specifications allow each country's laws and economic structure to have different effects on search intensity as well as for possible gender differences in search behavior. In particular, continued inclusion of time dummies in the models disaggregated by country allows each country to have a flexible trend in its job search intensity.

Before discussing the regression results, I note that in a fixed effects model, the parameters are identified from the behavior of those who change contract duration during the panel. Appendix Table A2 compares mean values for three samples: i) the entire panel ("Full Sample"); ii) those who changed contract duration—the key group on which identification is based; and iii) those currently on a temporary job. From Table 2, we know that only about 9% of the sample has a temporary job. Therefore, it is not surprising that the sample who changed contract duration is more similar to those who currently have a temporary job than to the entire sample. In the second two columns of Table A2, we see that compared to the full sample, duration changers and those on temporary contracts are more likely to be women, live in areas with higher unemployment, have lower schooling levels, earn less money, and are more likely to live in Spain, Portugal, or Finland, countries with high levels of temporary employment. Most of those who changed contract duration had permanent jobs at least once during the panel, so it is not surprising that this group had somewhat higher wages, a (slightly) lower female representation, was slightly better-educated, and less likely to live in Spain (the country with the highest incidence of temporary jobs) than those currently on temporary jobs.

Table 5 contains basic fixed effects regression results for the determinants of search intensity among employed workers. Looking first at the full sample results, we see that the increase in search intensity from permanent jobs to short duration temporary jobs is very similar to the raw means shown in Table 1. Active searching increases by a highly significant 2.1 percentage points going from permanent jobs to longest duration temporary jobs, all else equal. The effect further increases to 3.6% for jobs with 6–12 months duration and all the way to 11.6% for the shortest duration jobs. Among temporary jobs, 84% of the increase in search intensity that occurs between the longest and the shortest duration jobs occurs in the last period. Moreover, Table 5 shows that the effect for the 0–6 month category is significantly different at better than the 1% level from that of permanent jobs and from each of the other temporary job duration categories. The point estimates show gradually increasing search intensity between the 24+ months category and the 6–12 months category, although Table 5 indicates that these increases are not statistically significant. Nonetheless, the point estimates and the significantly positive effect for 24+ months duration (relative to permanent jobs, the omitted category) both suggest forward-looking behavior by workers.

Other results for the full sample are that older workers have lower search intensity (the negative main term outweighs the positive quadratic term for all ages up to 61.6 years), and more highly paid workers have lower search intensity. Table 5's results for men and

Table 5
Individual fixed effects regression results for search intensity, pooled across countries.

	Pooled	Men	Women
Age	−0.007**	−0.004*	−0.011**
	(0.002)	(0.002)	(0.002)
Age Squared	0.000*	0.000	0.000**
	(0.000)	(0.000)	(0.000)
Low Level Schooling	−0.002	0.008	−0.015
	(0.009)	(0.011)	(0.011)
Middle Level Schooling	0.005	0.011	−0.003
	(0.009)	(0.013)	(0.005)
Contract Duration <6 mos.	0.116**	0.125**	0.104**
	(0.012)	(0.016)	(0.017)
Contract Duration 6-12 mos.	0.036**	0.037*	0.035**
	(0.009)	(0.015)	(0.008)
Contract Duration 12-24 mos.	0.032**	0.021	0.042**
	(0.010)	(0.016)	(0.010)
Contract Duration 24+ mos.	0.021**	0.040*	−0.001
	(0.006)	(0.015)	(0.012)
Regional Unemployment rate	0.000	−0.000	0.000
	(0.000)	(0.000)	(0.001)
Log Real Hourly Earnings	−0.022**	−0.029**	−0.012
	(0.006)	(0.007)	(0.008)
Year Dummies?	yes	yes	yes
N	237314	136699	100615
P-values for tests of Duration Coeffs:			
<6 mos. vs. 6-12 mos.	0.0000	0.0015	0.0028
<6 mos vs. 12-24 mos.	0.0001	0.0007	0.0204
<6mosvs.24 + mos.	0.0000	0.0045	0.0006
6-12 mos. vs. 12-24mos.	0.6880	0.1076	0.6254
6-12 mos. vs. 24 + mos.	0.1989	0.8743	0.0102
12-24mos. vs. 24 + mos.	0.2972	0.3792	0.0240

Standard errors clustered at the country level. Permanent contract is the omitted duration category. + p<0.10, * p<0.05, ** p<0.01.

[17] As discussed above, due to the possible endogeneity of wages, I also estimated models with wages excluded.

[18] I am grateful to Alison Davies and Rhys Powell for their help in acquiring the European Labour Force Survey regional unemployment rate data. Since the ECHP did not collect regional information for Denmark or the Netherlands, I used the national unemployment rate for those countries.

[19] Moffitt (1999) argues that the linear probability model may be justified as a linear approximation to a more general data-generating process.

[20] Below, I examine the representativeness of the subsample of individuals who change contract type relative to the whole sample.

Table 6
Individual fixed effects regression results for search intensity, separately by country.

	Austria	Belgium	Denmark	Finland	France	Greece	Ireland	Italy	Netherlands	Portugal	Spain
Age	−0.008*	−0.007	−0.001	−0.016+	−0.005	−0.013*	−0.015+	−0.010**	−0.009	0.002	−0.010
	(0.004)	(0.006)	(0.006)	(0.008)	(0.004)	(0.005)	(0.008)	(0.004)	(0.006)	(0.003)	(0.007)
Age Squared	0.000	0.000	−0.000	0.000	0.000	0.000+	0.000+	0.000*	0.000+	−0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Low Level Schooling	0.037	−0.029	−0.019	0.019	−0.056	0.025	−0.029	−0.027	0.002	0.001	−0.036*
	(0.028)	(0.025)	(0.023)	(0.027)	(0.071)	(0.023)	(0.026)	(0.048)	(0.040)	(0.026)	(0.018)
Middle Level Schooling	0.050+	−0.030	0.004	0.043+	−0.059	0.001	−0.025	−0.027	0.041	0.009	−0.009
	(0.025)	(0.021)	(0.019)	(0.023)	(0.052)	(0.016)	(0.021)	(0.047)	(0.041)	(0.022)	(0.017)
Contract Duration <6mos.	0.064+	0.074+	0.169**	0.123**	0.174**	0.063	0.043	0.135**	0.177**	0.079**	0.099**
	(0.036)	(0.041)	(0.052)	(0.029)	(0.023)	(0.050)	(0.053)	(0.025)	(0.041)	(0.027)	(0.020)
Contract Duration 6-12 mos.	0.067*	0.001	0.118**	0.048*	0.055**	0.056*	0.005	0.028	0.062	0.012	0.020
	(0.026)	(0.030)	(0.040)	(0.024)	(0.020)	(0.023)	(0.025)	(0.018)	(0.044)	(0.012)	(0.016)
Contract Duration 12-24 mos.	0.041	−0.005	0.086*	0.004	0.045+	0.072+	0.081*	0.049+	0.016	0.031	0.010
	(0.026)	(0.036)	(0.037)	(0.025)	(0.027)	(0.039)	(0.039)	(0.028)	(0.051)	(0.021)	(0.016)
Contract Duration 24+ mos.	−0.003	0.035	0.021	0.008	−0.010	0.022	0.048	0.039	0.068	0.003	0.014
	(0.016)	(0.038)	(0.045)	(0.049)	(0.035)	(0.031)	(0.042)	(0.027)	(0.070)	(0.024)	(0.024)
Regional Unemployment rate	0.016*	0.001		−0.001	−0.001	0.002	−0.001	0.003+		0.005*	−0.002
	(0.006)	(0.001)		(0.001)	(0.003)	(0.003)	(0.004)	(0.002)		(0.002)	(0.002)
Log Real Hourly Earnings	0.002	−0.024	−0.029	−0.025	0.005	−0.055**	−0.033*	−0.048**	−0.013	−0.025**	−0.037+
	(0.010)	(0.017)	(0.022)	(0.017)	(0.007)	(0.013)	(0.014)	(0.012)	(0.011)	(0.010)	(0.021)
Year Dummies?	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
N	18221	14449	15669	18676	30315	14746	13642	32172	27638	25914	25872
P- values for tests of Duration Coeffs:											
<6mos. vs. 6-l2mos.	0.9451	0.0948	0.4251	0.0290	0.0000	0.8885	0.4909	0.0001	0.0445	0.0165	0.0003
<6 mos vs. 12-24 mos.	0.5852	0.1174	0.1862	0.0006	0.0001	0.8862	0.5337	0.0166	0.0070	0.1510	0.0001
<6 mos vs. 24+ mos.	0.0792	0.4943	0.0243	0.0303	0.0000	0.5014	0.9277	0.0079	0.1670	0.0336	0.0033
6-12 mos. vs. 12-24 mos.	0.4670	0.8777	0.5446	0.1016	0.7280	0.7106	0.0869	0.5257	0.4704	0.4066	0.5758
6-12 mos. vs. 24+ mos.	0.0345	0.4883	0.1001	0.4211	0.0687	0.3805	0.3408	0.7406	0.9398	0.7049	0.8063
12-24 mos. vs. 24+ mos.	0.1236	0.4965	0.2392	0.9378	0.2480	0.3528	0.5394	0.7828	0.5357	0.3487	0.8749

+ p<0.10, * p<0.05, ** p<0.01.
Permanent contract is the omitted duration category. Standard errors are heteroskedasticity robust.

women separately are very similar to those of the pooled sample, and together they confirm that the raw increase in search intensity observed in the overall means as duration falls is not simply a gender composition effect. The impact for the 0–6 month duration category remains significantly different from that of each of the other temporary duration levels as well as of course the level for permanent jobs. For both men and women, most of the increase in search intensity across temporary job duration categories occurs in the last period (104% for men and 66% for women). And there is evidence of forward-looking behavior for both men and women: for men, there is significantly more search activity in the longest temporary jobs than on permanent jobs, while for women, there are significant increases in search intensity going from 24+ month duration jobs to those with 12–24 months or 6–12 months duration. A further interesting result concerns the relative effects of wages for men and women. In both cases, the effect is negative, is statistically significant for men and is 1.5 times its standard error in absolute value for women. But the magnitude is more than twice as high for men as for women, and the gender difference in the effects of wages is nearly significant (1.6 times its standard error in absolute value). Since the sample mean search intensity is 0.055 for men and 0.059 for women, men's search elasticity with respect to wages is more than twice as high in absolute value as women's. This suggests a higher labour supply elasticity to the firm for men than women, a factor that could help explain part of the gender pay gap.[21]

Table 6 shows that these results for the pooled ECHP sample largely hold up within individual countries. First, for each country, search intensity is significantly greater for those in temporary jobs than on permanent jobs. While search intensity doesn't always monotonically increase as contract duration falls, it generally rises, and in nearly every case, it is much higher for the shortest duration contract than for longest duration temporary jobs.[22] Nearly all, or some cases, more than the full rise in search intensity across temporary job categories occurs in the shortest jobs in seven of the eleven countries, including Belgium, Finland, France, Italy, the Netherlands, Portugal and Spain, and the increase in search intensity in the last period is statistically significant in each of these countries. Searchers are forward looking as indicated by the presence of several significant positive coefficients in duration categories in the 6 month to 2 year range, and the shortest duration jobs have the most search intensity.

In addition to the results shown so far, I attempted some further, alternative specifications. First, I limited the sample to those age 19–60 to reduce the influence of school and retirement on the results.

Table 7
Regulation of temporary and regular employment contracts, late 1990s.

	Temporary contracts		Regular employment
			OECD protection strictness
	Maximum number	Maximum total duration	Index
Austria	1.5	None	2.9
Belgium	4	30 months	1.7
Denmark	1.5	30 months	1.5
Finland	1.5	None	2.3
France	2	18 months	2.3
Greece	2.5	None	2.3
Ireland	None	None	1.6
Italy	2	18 months	1.8
Netherlands	3	None	3.1
Portugal	3	30 months	4.3
Spain	3	24 months	2.6

Source: OECD (2004), pp. 113 and 117. Protection Strictness Index ranges from 0 (least strict) to 6 (the most restrictive laws).

[21] The evidence on the relative wage elasticity of male and female quitting is somewhat mixed. See, for example, Blau and Kahn (1981), Viscusi (1980), Barth and Dale-Olsen (2009), and Ransom and Oaxaca (2010).

[22] An exception to monotonicity is a seemingly anomalous rise to 8.1% for those in 12–24 month contracts in Ireland.

Table 8
Incidence of one year transition rates to permanent job from temporary job in the same firm, by temporary contract duration.

Duration of last year's temporary contract	All workers	Workers with no search effort
0-6 months	0.157	0.174
6-12 months	0.200	0.214
12-24 months	0.229	0.243
24+ plus	0.177	0.184
Total	0.192	0.206
Sample size	17691	15171

Sample includes those who had a temporary contract one year ago. The figures represent the fraction of this sample that was at the same firm at the current year's interview and had a permanent contract.

Table Al
Transition rates for jobseekers: from temporary jobs to permanent or new temporary jobs.

Temporary job duration	Transitions to		
	Permanent job	New temporary job	N
<6 months	0.474	0.330	796
6-12 months	0.475	0.293	867
12-24 months	0.487	0.169	398
24+ months	0.627	0.130	199
All	0.495	0.260	2260

Sample includes people in temporary jobs who were searching for a new job.

Second, I re-ran the models excluding current wages on the idea that these may be endogenously determined at the worker-firm match level. Third, in some models, I controlled for current tenure and its square, although roughly 5% of the sample has missing tenure data. Controlling for tenure brings the duration variables closer to measuring time remaining on the contract, although as mentioned earlier, we don't know when the current contract started, in contrast to when the employment relationship started. Since current tenure is potentially endogenous with respect to search activity, I don't include it in the basic model in Eq. (18). But in all three of these cases, the results were virtually identical to those shown above, giving one further confidence in the robustness of the basic findings. I also investigated whether the impact of contract duration varied with wages on the idea that incentives to start searching earlier would likely be stronger with lower current wages. However, interaction effects between wages and duration showed no consistent patterns. Finally, I estimated models separately by country-gender group. I obtained similar results to the more aggregated analyses both with respect to the countries with large increases in search intensity in the shortest jobs and the higher wage elasticity of search activity for men.

It is interesting to consider my results for the individual countries in the context of their systems of employment protection both with respect to temporary contracts and regular (permanent) employment. In this regard, Table 7 summarizes these regulations as of the late 1990s, the period when most of the ECHP data were collected. The Table shows the limits (if any) on the number and duration of temporary contracts as well as the OECD's index of strictness of protection mandates on regular jobs. In countries where firms can renew temporary contracts without limit, one might expect less incentive to search during temporary employment than otherwise. And where it is very difficult to fire workers from permanent jobs, one might expect less search from such jobs, both because the probability of keeping the job is higher than otherwise and because in such settings, job creation is suppressed due to anticipated firing costs. The reduced job creation will also likely reduce the return to search from temporary jobs as well. Fig. 1 indeed shows a negative relationship between the incidence of searching on the job from permanent jobs and the OECD's protection strictness indicator for such jobs, while Fig. 2 shows a similar relationship for the incidence of active search from very short duration jobs. These admittedly simple relationships suggest that reforms reducing the strictness of protection will generate additional search activity.

Table 9
Individual fixed effects results for the incidence of one year transition rates to permanent job from temporary job in the same firm, by temporary contract duration.

	All workers with temporary jobs at previous interview			Workers with temporary jobs at previous interview who had no search activity		
	Pooled	Men	Women	Pooled	Men	Women
Age	0.127**	0.123**	0.133**	0.132**	0.129**	O.133**
	(0.015)	(0.027)	(0.021)	(0.019)	(0.032)	(0.023)
Age Squared	−0.00l**	−0.001+	−0.001**	−0.00l**	−0.001+	−0.001*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Low Level Schooling	−0.277+	−0.252+	−0.241	−0.278+	−0.246	−0.251+
	(0.138)	(0.138)	(0.147)	(0.132)	(0.162)	(0.133)
Middle Level Schooling	−0.045	0.021	−0.112	−0.025	0.032	−0.087+
	(0.033)	(0.069)	(0.072)	(0.029)	(0.070)	(0.044)
Contract Duration 6-12 mos. (prev. yr)	0.014	0.038	−0.006	0.024	0.046	0.006
	(0.018)	(0.026)	(0.026)	(0.022)	(0.033)	(0.037)
Contract Duration 12-24 mos. (prev. yr)	0.016	0.014	0.016	0.015	−0.003	0.030
	(0.020)	(0.039)	(0.024)	(0.016)	(0.037)	(0.028)
Contract Duration 24+ mos. (prev. yr)	0.046	0.041	0.050	0.032	0.028	0.033
	(0.034)	(0.035)	(0.045)	(0.043)	(0.033)	(0.065)
Regional Unemployment rate	−0.009*	−0.014*	−0.004	−0.011*	−0.016*	−0.006
	(0.004)	(0.006)	(0.004)	(0.005)	(0.006)	(0.004)
Log Real Hourly Earnings (prev. yr)	0.014	0.010	0.021	0.013	0.021	0.011
	(0.031)	(0.048)	(0.032)	(0.023)	(0.045)	(0.026)
Year Dummies?	yes	yes	yes	yes	yes	yes
	17691	8866	8825	15171	7612	7559
P- values for tests of Duration Coeffs:						
6-l2mos = 12-24mos = 24 + mos = 0	0.5709	0.5029	0.1408	0.6583	0.5786	0.5162
6-12 mos. vs. 12-24 mos.	0.9154	0.5582	0.2795	0.7173	0.3584	0.4336
6-12 mos. vs. 24+ mos.	0.1991	0.9266	0.1833	0.8301	0.6292	0.6268
12-24mos. vs. 24 + mos.	0.3462	0.4957	0.5202	0.7297	0.5735	0.9598

+ $p<0.10$, * $p<0.05$, ** $p<0.01$.
Sample for columns 1-3 includes all who had a temporary job at the previous interview. Sample for columns 4-6 includes those with temporary jobs at the previous interview who had no search activity. Omitted duration category is 0-6 months.

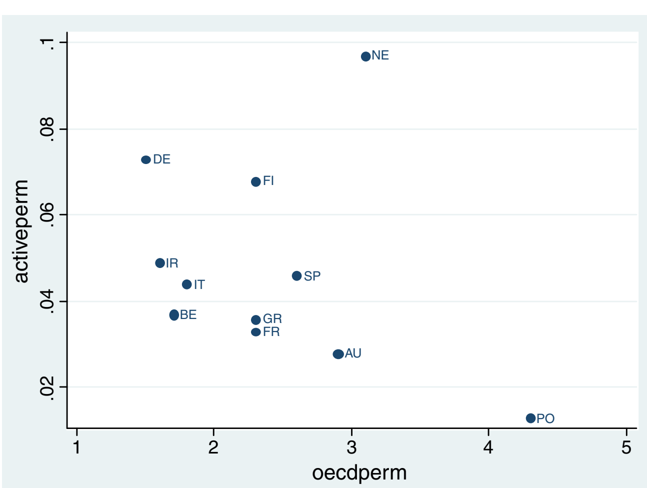
Table A2
Mean values for full sample, those who changed contract duration during the panel, and those currently in temporary jobs.

	Full sample	Changed contract duration	Currently on temporary job
Incidence of active search	0.057	0.106	0.146
Age	38.546	33.405	32.040
Age Squared	1597.227	1207.482	1125.964
Low Level Schooling	0.293	0.322	0.361
Middle Level Schooling	0.411	0.362	0.342
Contract Duration <6 mos.	0.022	0.092	0.236
Contract Duration 6-12 mos.	0.038	0.175	0.416
Contract Duration 12-24 mos.	0.019	0.098	0.205
Contract Duration 24+ mos.	0.013	0.064	0.142
Female	0.425	0.481	0.503
Regional Unemployment Rate	9.361	10.381	11.747
Log Real Hourly Earnings	1.926	1.737	1.649
Denmark	0.091	0.076	0.056
Netherlands	0.091	0.038	0.032
Belgium	0.091	0.092	0.081
France	0.091	0.071	0.085
Ireland	0.091	0.051	0.051
Italy	0.091	0.083	0.071
Greece	0.091	0.089	0.085
Spain	0.091	0.195	0.243
Portugal	0.091	0.123	0.114
Austria	0.091	0.068	0.050
Finland	0.091	0.112	0.131
Sample Size	237314	37396	23395

Sampling weights are used, where each country in the full sample receives the same total weight.

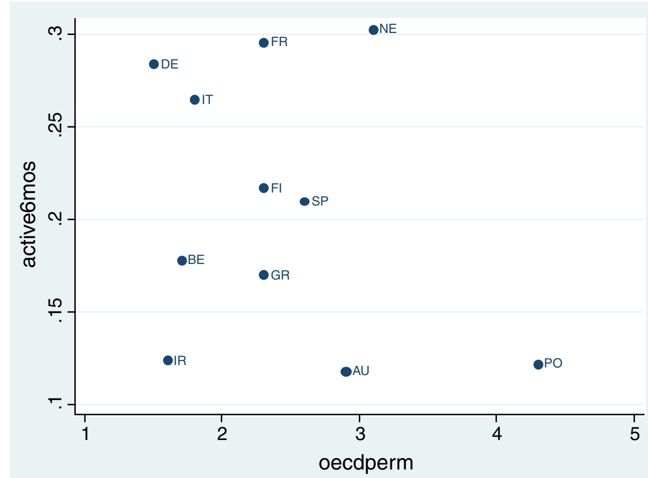
I investigated further the relationship between the OECD's permanent employment protection index and search intensity by re-estimating the basic pooled fixed effects models interacting the duration categories with the country's late 1990s permanent protection index. The interaction effects were each about -1 percentage point, were significantly different as a group from the omitted category (permanent jobs), but were not significantly different each other. Thus, there was some evidence that stricter employment protection lowers the gap in search intensity between temporary jobs and permanent jobs but not within the set of temporary jobs.

On the regulation of temporary contracts, Table 7 shows especially strict regulations for both France and Italy, with relatively few contracts allowed and tight limits on their duration. And for both countries, active search from very short jobs is especially high, as shown in both Table 4 and in the regression coefficients in Table 6. These results suggest that restricting firms' ability to offer temporary jobs will generate more active job search by workers on such jobs, although the Netherlands, with relatively lax regulations of temporary work also sees a high search intensity from such jobs (Tables 6 and 7). Estimating the basic fixed effects model adding the OECD's late 1990s index of temporary employment protection interacted with contract duration yielded a positive, small interaction effect for the shortest jobs relative to permanent jobs, as one might expect. However, this effect was insignificantly different from zero, as were the effects for the other duration categories. The effects of the duration interactions were also insignificant as a group and not significantly different from each other. But overall, the findings suggest the possibility that changing regulatory policy may generate the expected changes in search activity.[23]



Source: Table 4 and OECD (2004), p. 117.

Fig. 1. Incidence of active search measures from permanent jobs by OECD indicator of strictness of protection on permanent jobs.



Source: Table 4 and OECD (2004), p. 117.

Fig. 2. Incidence of active search measures from temporary jobs with under 6 months' duration by OECD indicator of strictness of protection on permanent jobs.

3.2. The role of within-firm promotions from temporary to permanent jobs

In the analysis of the theoretical model presented in Section II, the possible impact of promotions from temporary to permanent jobs was mentioned. If, for example, workers on temporary contracts anticipate being promoted to permanent jobs, then they may not search hard for new work. Moreover, as discussed, if the anticipated probability of promotion on temporary jobs with a long duration is greater than that on short-term jobs, then anticipated promotions could explain the pattern of results I have presented so far. To investigate this issue, one would ideally have access to information on whether and when workers on temporary contracts were offered permanent job status. Unfortunately, the ECHP doesn't have information on the existence of such offers. But we do know if a person who was on a temporary job in one year was in a permanent job with the same firm at the next wave of the survey. Unless such workers left their temporary job and were rehired into a permanent job at their original firm during the intervening year, we can infer that they were promoted into a permanent job. And we can observe the incidence of such events by using the

[23] In both Figs. 1 and 2, the relationship has a negative slope but is insignificant. As pointed out by Bentolila et al. (2010), a perhaps more relevant indicator of the effects of permanent and temporary employment regulation is to compute the difference in firing costs between the two types of jobs. However, the OECD's employment protection indicators as shown in Table 7 don't allow such a difference to be meaningfully computed.

Table A3
Individual fixed effects results for the incidence of one year transition rates to permanent job from temporary job in the same firm, by temporary contract duration for those who had no search activity, by country.

	Austria	Belgium	Denmark	Finland	France	Greece	Ireland	Italy	Netherlands	Portugal	Spain
Age	0.073	0.168*	−0.010	0.185**	0.142**	0.143*	0.461	0.131*	0.191*	0.090	0.108**
	(0.073)	(0.083)	(0.128)	(0.058)	(0.038)	(0.060)	(0.281)	(0.053)	(0.086)	(0.104)	(0.041)
Age Squared	−0.000	−0.001	0.001	−0.002*	−0.001*	−0.001	−0.004**	−0.001+	−0.002+	0.000	−0.000
	(0.001)	(0.001)	(0.002)	(0.001)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Low Level Schooling	0.094	0.032	−0.581	−0.211	0.014	0.447*	−0.634+	−0.243		−0.207	−0.219**
	(0.214)	(0.230)	(0.453)	(0.142)	(0.211)	(0.181)	(0.361)	(0.289)		(0.246)	(0.080)
Middle Level Schooling	−0.032	0.165*	−0.024	−0.130	−0.069	0.309**	0.078	0.007		0.057	−0.155*
	(0.147)	(0.073)	(0.279)	(0.120)	(0.205)	(0.094)	(0.320)	(0.212)		(0.124)	(0.061)
Contract Duration 6-12 mos. (prev yr)	0.018	0.131	0.297+	−0.030	−0.008	−0.061	−0.105	−0.067	−0.069	0.024	0.031
	(0.089)	(0.090)	(0.153)	(0.039)	(0.037)	(0.185)	(0.103)	(0.055)	(0.107)	(0.062)	(0.028)
Contract Duration 12-24 mos. (prev yr)	0.100	0.016	0.154	0.005	0.019	0.050	−0.045	−0.133	−0.006	0.101	0.035
	(0.108)	(0.103)	(0.119)	(0.044)	(0.052)	(0.198)	(0.099)	(0.112)	(0.090)	(0.097)	(0.042)
Contract Duration 24+ mos. (prev yr)	−0.114	0.170	0.241*	0.026	−0.001	−0.149	−0.085	−0.049	0.071	0.263*	0.088+
	(0.151)	(0.115)	(0.121)	(0.062)	(0.058)	(0.240)	(0.130)	(0.096)	(0.098)	(0.108)	(0.053)
Regional Unemployment rate	−0.011	0.038		−0.011	0.014	−0.056	0.025	−0.010		−0.014	−0.004
	(0.099)	(0.050)		(0.007)	(0.018)	(0.037)	(0.139)	(0.014)		(0.030)	(0.013)
Log Real Hourly Earnings (prev yr)	−0.109	0.002	0.052	0.009	−0.038	−0.111	0.050	0.111	0.055	0.118	0.071
	(0.113)	(0.114)	(0.134)	(0.065)	(0.048)	(0.079)	(0.130)	(0.079)	(0.166)	(0.124)	(0.046)
Year Dummies?	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
N	626	821	535	1515	1696	927	458	1565	456	2223	4349
P- values for tests of Duration Coeffs:											
6-l2mos = 12-24mos = 24 + mos = 0	0.2827	0.1372	0.1558	0.6271	0.9395	0.3446	0.7836	0.5459	0.7246	0.0552	0.3903
6-12 mos. vs. 12-24 mos.	0.3726	0.1350	0.3058	0.3570	0.5343	0.3071	0.5995	0.5312	0.6102	0.3060	0.9277
6-12 mos. vs. 24+ mos.	0.2902	0.6477	0.6796	0.2754	0.8826	0.4538	0.8668	0.8503	0.2609	0.0063	0.2435
12-24 mos. vs. 24+ mos.	0.0713	0.0571	0.3181	0.6579	0.7268	0.0788	0.7085	0.5009	0.4830	0.0682	0.3654

+ $p<0.10$, * $p<0.05$, ** $p<0.01$.
Omitted duration category is 0-6 months. Standard errors are heteroskedasticity robust.

longitudinal feature of the ECHP. However, we are unable to observe offers of promotion for those who left their firm. Nonetheless, with these caveats, one can obtain at least a partial assessment of the role of promotions in explaining my basic results.[24]

Table 8 shows mean values for the incidence of changes from temporary to permanent status at the same firm between consecutive waves of the ECHP for i) all workers with temporary contracts in the initial year ("All Workers") and ii) workers with temporary jobs in the initial year who were not actively searching for work that year. These mean values are based on an underlying variable which equals 1 if the worker in survey year t had a permanent job in the same firm at which he/she had a temporary job in survey year t-1. The variable equals 0 if at survey year t the worker either was not employed, employed at another firm, or employed at his/her period t-1 firm but still had a temporary contract (not necessarily the same contract that applied in period t-1). The overall incidence of such status changes is 19-21% on average and is roughly similar across the differing levels of initial contract duration. The incidence is slightly lower for the shortest jobs, although it is also lower for the longest jobs than for the 6–12 and 12–24 month categories. But to properly assess the role of promotions in explaining the search intensity results in Tables 5 and 6, one needs to compare my estimate of the promotion incidence across contract durations while controlling for other factors affecting such outcomes.

Tables 9 and A3 perform such an analysis by estimating individual fixed effects models of promotion to permanent jobs in equations that are similarly specified to our models of search intensity, except that temporary contract duration and real wages are measured as of the initial year.[25] Because of the individual fixed effects, the results in Tables 9 and A3 are identified from comparisons of individuals who were in temporary contracts of different duration levels at different survey dates. However, as noted, because of possible renewals of temporary contracts between surveys, one cannot observe the amount of time remaining on a temporary contract. In this paper, I have used the total potential duration as a reasonable proxy for time remaining. In Table 9, the results for both samples (i.e. all temporary job holders in period t-1 and all temporary job holders in period t-1 who weren't actively searching in period t-1) are very similar. Specifically, Table 9 shows that the duration variable coefficients are insignificant as a group and also not significantly different from each other in each model and sample. Moreover, the effects are quantitatively small relative to the mean incidence of promotions shown in Table 8. Thus at the aggregate level, there is no evidence to support the idea that search intensity rises dramatically in the 0–6 month duration category due to supposedly smaller perceived promotion probabilities.[26] Other results from the analysis of promotions in Table 9 are that older workers are more likely to be promoted (at least through ages 61–67), while those with low levels of schooling or living in high unemployment regions are less likely to be promoted.

[24] In principle, the same reasoning I have used for transitions to permanent jobs would apply to a worker who anticipates that his/her temporary contract will be renewed. Unfortunately, for the longer temporary job duration categories, it is impossible to determine from the year-to-year data whether one's contract has been renewed. Therefore, unlike the transitions we do observe to permanent work, we can't observe renewals of temporary contracts, since for many of those in temporary jobs, contract duration will be the same from one year to the next both for those with renewals and for those in their original contract. There thus would be no way to distinguish the renewals from the continuations, whereas for those who were promoted into permanent jobs, we do observe a change in contract type in the ECHP.

[25] I am calling the transition from last year's temporary job to this year's permanent job at the same firm a "promotion" even though, as mentioned, an unknown number of workers may have left their firm and returned with a permanent job during the months between consecutive waves of the ECHP. And of course, I also don't observe offers of promotion for those who left their firm and didn't return. I measure initial wages in order to avoid the simultaneity of current wages and promotion status.

[26] Even with a constant per period promotion likelihood, we would expect greater search intensity for those in the later stages of their temporary contract because they have not yet been promoted. The fact that they are still observed in their temporary contract means that they have likely been passed over for promotion. In the ECHP data, we have a snapshot for workers in a job with a given contract duration. On average, those workers in longer jobs have likely had more opportunities to be promoted but haven't been, compared to those observed in shorter jobs. Therefore, this selection phenomenon would lead us to expect more search intensity among those currently in long jobs than in short jobs. Of course we observe the opposite, as predicted by the search model.

The basic conclusion about the lack of impact of temporary contract duration on promotions holds within countries as well, as shown by the models in Table A3. The sample here is restricted to those in an initial temporary job who weren't actively searching on that initial job, although the findings are similar when I include everyone on an initial temporary contract. Table A3 shows that in each case, the contract duration variables are insignificant as a group and are not significantly different from each other in 31 of 33 possible comparisons. Moreover, Table 6 showed large increases in search intensity in the 0–6 month duration jobs for Belgium, Finland, France, Italy, the Netherlands, Portugal and Spain. Table A3 shows that the promotion probability was insignificantly different between the two shortest duration categories in each case. Additionally, in four of the seven countries with rapid final period increases in search intensity, the point estimate in Table A3 shows a *higher* promotion probability in the 0–6 month duration jobs than in the 6–12 month jobs, with three of these showing a lower promotion probability for the shortest jobs. And the point estimates among these seven countries are, with the exception of Belgium, small in absolute value. Again, promotion probabilities do not appear to explain our basic search intensity results, which are consistent with the calibrated search model presented in Section II.[27]

4. Welfare implications of reforms liberalizing the use of temporary contracts

The calibrated search model and supporting empirical results presented above together can be used to assess some welfare implications of one of the most popular reforms in European labour market regulation that have been enacted in recent years. Specifically, several countries have liberalized the use of temporary contracts, including increasing their maximum potential duration or allowing them to be renewed.[28] An intuitively plausible definition of welfare in this context is output net of search costs.[29] The goal of this section is to study the impact on welfare when the allowable duration of temporary contracts rises by one unit, say one six month period. Since the calibrated search model and empirical results discussed above imply little search effort until the last period, let us compare welfare when two period contracts are allowed with welfare when only one period contracts are allowed, in addition to of course permanent jobs.

To begin the analysis, I take a partial equilibrium approach and then discuss the possible general equilibrium responses to the policy liberalization. In such a framework, allowing a two period contract produces gains for both firms and workers because the worker always has the option of postponing search until the last period and can produce output in both periods, and the firm's options are also expanded. In a one period job, the worker must begin searching immediately, and it is not certain that there will be any output produced in the next period because he/she may be unemployed by then. As discussed in further detail below, this reasoning becomes more complicated in a general equilibrium framework: the reform may alter firms' probabilities of offering permanent jobs and therefore may affect output net of search costs by affecting the transition rate to permanent jobs as well as the likelihood of being unemployed.

Let y_t and y_p be, respectively, per period output on temporary jobs and permanent jobs. Recent research suggests that $y_p>y_t$ (Boeri 2011; Stancanelli 2002; and Booth et al., 2002). This is an intuitively plausible result, since firms won't offer permanent jobs unless there is a compensating gain to them, and the possibility of sharing in the productivity advantage of a worker on a permanent job can produce such a gain (Blanchard and Landier 2002). In addition, offering permanent jobs may help solve agency problems in the provision of firm specific training and thus may lead to higher productivity than temporary jobs would have (MacLeod and Nakavachara 2007). Let Y_{11} be the expected discounted lifetime output net of search costs of a worker on a temporary job in an economy where only permanent jobs and one period temporary jobs are allowed; this will necessarily be the last period of the job. Let Y_{22} be the expected discounted lifetime output net of search costs of a worker at the start of a two period temporary job (i.e. with two periods left) in an economy where all temporary jobs last two periods, and let Y_{12} be the corresponding quantity for a worker in the last period of a two period temporary job. Finally, let Y_p be the expected discounted lifetime productivity of a worker who takes a permanent job (in the simple model shown above, there is no search from permanent jobs, so output is the same as output net of search costs), and Y_0 be the expected discounted lifetime productivity net of search costs of a currently unemployed worker. The partial equilibrium welfare analysis suggested above compares Y_{11} and Y_{22}: when a worker starts a temporary job, what is the effect of increasing the duration to two periods from one period?[30]

Let λ_{11} be the search intensity the worker chooses when employed in a one period temporary job where the law only allows temporary jobs to last one period. Let P_{11} and T_{11} be respectively the endogenous probabilities that a given permanent or temporary job offer is acceptable to a worker in this one period job.[31] Then from the search model analyzed earlier, we can write the following expression for Y_{11}:

$$\begin{aligned} Y_{11} = {} & y_t + B\lambda_{11}P_{11}E(Y_p - Y_0 | \text{a given permanent offer is acceptable}) \\ & + Ba\lambda_{11}T_{11}E(Y_{11} - Y_0 | \text{ a given temporary offer is acceptable}) \\ & + BY_0 - .5(\lambda_{11})^2. \end{aligned} \tag{19}$$

Eq. (19) expresses the welfare associated with a worker on a temporary job where such jobs can only last one period. It is the sum of current output net of current search costs plus the discounted expected welfare associated with the three possible outcomes of current job search: i) an acceptable permanent job is found; ii) an acceptable temporary job is found; iii) the worker becomes unemployed.

To study the welfare associated with employment in a two period temporary job, define the search intensities in the two periods as λ_{22} (two periods left) and λ_{12} (one period left), and the probabilities that a permanent (P) or temporary (T) offer is acceptable as P_{22}, and T_{22}

[27] Table A3 shows a large, positive, marginally significant promotion effect of the 6–12 month jobs (relative to the omitted 0–6 month category) for Denmark. There is also a large, but insignificant coefficient for this category for Belgium. While these point estimates could explain the last period increase in search intensity in both countries, the promotion argument can't explain the monotonically increasing search intensity as duration falls for Denmark, since Table A3 shows for Denmark, falling, then rising promotion probabilities as duration falls. In the case of Belgium, the non-monotonic changes in search intensity in Table 6 match the direction of the changes in the coefficients in Table A3. But the only statistically significant search intensity coefficient for Belgium is that for the 0–6 month category (Table 6), while the comparison of the 12–24 and 24+ month categories is the only significant contract duration effect for promotion, as shown in Table A3.

[28] For example, between 1987 and 2003, the following OECD countries liberalized the use of temporary employment contracts at some time: Belgium, Denmark, Germany, Greece, Italy, Japan, the Netherlands, Norway, Poland, Portugal, the Slovak Republic, Spain, and the United Kingdom (OECD 2004, pp. 119–120).

[29] Cahuc and Postel-Vinay (2002), for example, define welfare as output net of firm recruiting costs, a similar concept to the one proposed here.

[30] Note that fixing contract length at two periods in the liberalized regime is analytically similar to giving firms the option of renewing a one period contract. In the latter scenario, the worker forms an estimate of the probability that the job will last more than one period, so the expected duration is higher than when contracts cannot be renewed. This example implicitly assumes that firms would optimally offer longer temporary contracts if they were allowed to do so.

[31] For notational convenience, I have suppressed the expressions showing the dependence of search intensity and reservation wages (and thus the probability that a given offer is acceptable) on current wages.

for offers with two periods left and P_{12} and T_{12} for offers with one period left, where temporary jobs last two periods. Then we have:

$$Y_{22} = y_t + B\lambda_{22}P_{22}E(Y_p - Y_{12}| \text{ a given permanent offer is acceptable}) + Ba\lambda_{22}T_{22}EY_{22} - Y_{12}| \text{ a given temporary offer is acceptable}) + BY_{12} - .5(\lambda_{22})^2. \tag{20}$$

Eq. (20) states that the welfare associated with a two period temporary job consists of current output net of current search costs plus the discounted expected welfare associated with the three possible search outcomes: an acceptable permanent job is found, an acceptable two period temporary job is found, or the worker continues in his/her current temporary job, which now is in its last period.

The basic theoretical and empirical result of the analysis of the earlier sections is that almost all job search on a temporary job takes place in the last period. If so, then λ_{22} will be much smaller than λ_{12}.[32] If in an extreme case λ_{22} is near zero, then we have:

$$Y_{22} \approx y_t + BY_{12} \text{ and} \tag{21}$$

$$Y_{22} - Y_{11} \approx y_t + BY_{12} - Y_{11}. \tag{22}$$

According to expression (22), the partial equilibrium gain to the reform lengthening the duration of temporary contracts is made up of two parts: i) the productivity of an extra period on the temporary job; and ii) the discounted value of being in the last period of a temporary job next period where temporary jobs have two periods relative to the current value of being in a temporary job in a labour market where such jobs last only one period. One can infer in a partial equilibrium setting that $(Y_{22} - Y_{11})$ is positive because the worker starting a two period temporary job always has the option of treating it as the last period of a one period temporary job, and Lemma 2 showed that a temporary job with two periods remaining is worth more than one with only one period left. We can also infer that in the partial equilibrium framework, $Y_{12} > Y_{11}$, since the worker at the end of a two period temporary job may land a new two year temporary job, which has higher productivity than a new one year temporary job. Therefore, for discount factors near 1, a lower bound for the welfare gain to the reform is roughly one period's output on a temporary job.

The simple framework of this welfare analysis assumed that the only way one could make a transition from a temporary to a permanent job is to actively search for one. As discussed earlier, it is also possible to be promoted into a permanent job with no job search activity, although the possibility of such promotions did not affect the basic conclusion that active search efforts would be largely postponed until the last period. To study the role of promotions in the welfare analysis of policies liberalizing the use of temporary contracts, let q_{22} be the probability of a promotion without any job search effort to a permanent job during the first period of a two period temporary job. Then the expected productivity net of search costs of a worker in this situation becomes:

$$Y_{22} = y_t + Bq_{22}E\left(Y_p - Y_{12}\right) + B\lambda_{22}P_{22}E(Y_p - Y_{12}| \text{ a given permanent offer is acceptable}) + Ba\lambda_{22}T_{22}E(Y_{22} - Y_{12}| \text{ a given temporary offer is acceptable}) + BY_{12} - .5(\lambda_{22})^2, \tag{20'}$$

assuming that promotion and obtaining a permanent or temporary offer through job search are mutually exclusive events. In addition, Eq. (20') assumes that the promotion offer is accepted, a reasonable assumption. If search effort is approximately zero in the first period of the two period temporary job, then the gain to the reform becomes:

$$Y_{22} - Y_{11} \approx y_t + Bq_{22}E\left(Y_p - Y_{12}\right) + BY_{12} - Y_{11}. \tag{22'}$$

With the possibility of promotion, the two period job gives the worker an additional chance to be promoted relative to the one period job, where in expression (22'), Y_{12} and Y_{11} both include terms referring to the probability of being promoted. The empirical analysis in Tables 8 and 9 suggests that the per period probability of being promoted is not affected by contract duration. This result implies that the lower bound for the welfare gain due to lengthening the duration of temporary jobs, as shown in expression (22'), is even larger than expression (22) implies. Specifically, the empirical estimates of Table 8 imply that q_{22} is about 0.2, implying that with a low discount rate the welfare gain is at least equal to roughly the one period output on a temporary job plus 0.2 times the lifetime productivity advantage of a permanent job relative to a temporary job.

Recent research on the wage effects of permanent vs. temporary employment sheds some light on the per period productivity differential between the two types of contract, although one would expect the lifetime productivity advantage of a permanent job to be greater than the per period effect because of its longer duration. For example, Stancanelli (2002) used ECHP micro data and an extensive set of controls to find hourly wage effects of permanent relative to temporary jobs across ten of the eleven countries studied here (all except Finland) averaging 0.116 for women and 0.121 for men. Boeri (2011) used ECHP and other European microdata and found monthly wage effects for the eleven countries in the current study averaging 20.9%, although his list of controls was far less extensive than Stancanelli's (2002), and his use of monthly rather than hourly earnings may have also helped lead to his larger estimate.[33] While these estimates are suggestive, they may be upward biased because workers on permanent jobs are likely to have higher levels of unmeasured productivity than workers on temporary jobs. Supporting this idea, Booth et al. (2002) used individual panel data for Britain and found that fixed effects estimates of the permanent job premium were smaller than cross-sectional estimates. For example, the cross-sectional effect for men was 0.171 log points but the fixed effects estimate was only 0.069; for women, the cross-sectional estimate was 0.144, and the fixed effects estimate was 0.109.

In light of these studies, I used the ECHP to estimate the impact of permanent jobs on hourly wages using both cross-sectional and fixed effects methods. I used the same controls as in the analysis of search intensity, except that contract type was measured as a single dummy variable indicating a permanent job. The cross-sectional estimate was 0.128 log points (which is much closer to Stancanelli's (2002) estimates than Boeri's (2011) results), while the fixed effects estimate was only 0.026 log points, and both effects were statistically significant. Thus, these estimates of the wage effects of permanent jobs averaged across European countries range from a low of 0.03 (my fixed effects estimate) to a high of 0.21 (Boeri's 2011 estimate), with a middle range of 0.12-0.13 (my cross-sectional estimate and Stancanelli's estimates). The smaller fixed effects estimates I found and that Booth et al. (2002) found suggest that an important portion of the cross-sectional estimate represents unmeasured individual heterogeneity rather than a true

[32] The possibility of a transition to a permanent job without job search (i.e. a promotion without search) will be discussed shortly.

[33] For example, Boeri (2011) controlled for education and tenure, while Stancanelli (2002) controlled for these as well as age, sector, occupation, and unemployment history.

effect of permanent jobs. Moreover, some of this wage premium may be also due to the worker's ability to appropriate some of the firing costs on permanent jobs and willingness to take a low wage on temporary jobs rather than a productivity differential.[34]

Using the range of estimates of the wage premium for a permanent job as estimates of the lifetime productivity advantage of such jobs adds an additional 0.6% to 4.2% of the productivity of a temporary job, as shown in the lower bound estimate in expression (22') for the welfare effects of replacing one period with two period temporary jobs (i.e. the promotion probability of 0.2 times the wage premium estimates). This implies a welfare gain of at least 1.006-1.042 times the one period output of a temporary job, since we expect $(Y_p - Y_{12})$ to be no less than the one period productivity effect.[35]

The foregoing analysis suggests positive effects of allowing firms to offer two period rather than one period temporary contracts. As noted, however, in a general equilibrium context, the effects may differ. The analyses of both Blanchard and Landier (2002) and Cahuc and Postel-Vinay (2002) both suggest that such jobs would be harder to come by after reforms making it easier to fire workers from temporary jobs (Blanchard and Landier 2002) or increasing the firm's ability to offer temporary jobs (Cahuc and Postel-Vinay 2002). The distortions induced by firing costs on permanent jobs produce a wedge between wages and productivity, possibly reducing their profitability. Allowing firms to offer two period temporary jobs may therefore lead them to offer fewer permanent jobs. Thus, the positive partial equilibrium effects might conceivably be undone by such induced changes. Moreover, due to higher turnover from temporary jobs, a higher vacancy rate may be required for a given unemployment rate (Wasmer 1999). This outcome is conceptually similar to a deterioration of the matching function. But overall, the gains due to keeping temporary workers employed for potentially another period must be weighed against these potential general equilibrium costs.

5. Conclusions

In this paper, I have examined the job search behavior of those employed in temporary jobs with a known duration level. A theoretical model of optimal search from a temporary job was constructed, and it predicts that workers employed in shorter duration temporary jobs would search harder than those in longer duration temporary jobs. Moreover, calibration of the model using observed transition rates to permanent work and to temporary work implies that at least 75% of the increase in search intensity over the life of a 2+ year temporary contract occurs in the last six months of the contract. I then used the ECHP data on employed workers for 1995–2001 from 11 countries to study the impact of contract duration on job search intensity. The countries were Austria, Belgium, Denmark, Finland, France, Greece, Ireland, Italy, the Netherlands, Portugal and Spain. In regression models that controlled for human capital, pay, regional unemployment, time and individual fixed effects, I found that workers on temporary jobs indeed search harder than those on permanent jobs. Moreover, search intensity increases as temporary job duration falls, and roughly 84% of this increase occurs on average in the shortest duration jobs. These results largely held up when I disaggregated by gender and also by country. In addition, an analysis of promotions from temporary to permanent jobs within one's firm showed that anticipated promotions cannot explain the basic findings. These empirical results are noteworthy, since it was not necessary to assume myopia or hyperbolic discounting in order to explain them, although the data clearly also do not rule out such explanations.

During the past 20 years, a number of European countries have enacted reforms of their employment protection systems in an attempt to stimulate job creation and increase the flexibility of their labour markets. The most common of these reforms have been new regulations allowing firms greater freedom to offer temporary employment contracts (OECD 2004), including increasing the number and duration of such contracts. I performed a partial equilibrium analysis of the impact of reforms lengthening the allowable duration of temporary contracts. Controlling for the supply of permanent jobs and wages, such reforms were shown to have positive welfare effects primarily due to the increased duration of temporary employment, including an increased opportunity for promotions to permanent jobs. However, theoretical and empirical research suggests that such reforms will reduce firms' relative supply of permanent jobs (Blanchard and Landier 2002; Kahn 2010). One must weigh the direct benefits to a worker of having a longer duration temporary contract with the reduction in the supply of permanent jobs.

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Appendix. Proof that the value of a temporary job falls and reservation wages fall as the end of a temporary job approaches

The reservation wages for temporary and permanent job offers conditional on being employed in a job with n periods left leave one indifferent between taking the job offer and remaining employed in the current job. Thus, if there are currently n periods left in one's temporary job, the value of alternative employment at the reservation wages is V_{n-1}. From my discussion of the final period on a temporary job, we have:

$$V_1(w) = V_0 + w - b > V_0 \tag{A1}$$

as long wages are strictly greater than unemployment benefits.

[34] While individual heterogeneity is a likely explanation for at least a portion of the smaller fixed effects estimates, it is possible that measurement error could also have contributed to the smaller fixed effects estimate. This is the case since the fixed effects models are more sensitive to measurement error in contract status than ordinary cross sectional models are when most of the sample is not changing contract status, as is the case here (see Table A2). Freeman (1984) makes a similar point about individual fixed effects estimates of the union wage impact. In the current case, the basic search intensity results for contract duration using fixed effects models were very similar to those using cross-sectional regressions. For example, Table 5 shows fixed effects coefficients on the duration variables ranging from 0.021 to 0.116, and the corresponding estimates from a cross-sectional model (with gender and country dummies added to the list of variables in the basic model in equation (18)) range from 0.041 to 0.150. This similarity suggests that measurement error is not an important cause of the difference between fixed effects and cross-sectional estimates of the wage effect of permanent contracts.

[35] While it is plausible that the one period productivity advantage of a permanent relative to a temporary job is less than the discounted lifetime advantage, the conclusion about lower bounds must be tempered by the possibility that the one period wage advantage of a permanent job overestimates the one period productivity advantage of such jobs. This is the case, since as mentioned, part of the wages on a permanent job reflect the worker's superior bargaining power relative to temporary jobs (Blanchard and Landier 2002).

To prove that reservation wages fall as the end of the temporary job approaches, it is sufficient to show that $V_n > V_{n-1}$ for all $n >= 2$. I prove this using induction. First, note that:

$$V_n(w) = w + B\lambda_n(w) \int_{Rn(w)}^{\infty} [W(x) - V_{n-1}]dF(x) + Ba\lambda_n(w) \int_{Tn(w)}^{\infty} [V_T(y) - V_{n-1}]dG(y) + BV_{n-1} - .5(\lambda_n(w))^2 \quad \text{(A2)}$$

Using (A2), this value when n = 2 is at least as large as the expected value of searching if one used the period 1 reservation wages and search intensity:

$$\begin{aligned} V_2(w) >&= w + B\lambda_1(w) \int_{R1(w)}^{\infty} [W(x) - V_1]dF(x) + Ba\lambda_1(w) \int_{T1(w)}^{\infty} [V_T(y) - V_1]dG(y) + BV_1 - .5(\lambda_1(w))^2 \\ &= V_1(w) + B(w - b)[1 - \lambda_1(w)(1 - F(R_1(w)) \\ &\quad - a\lambda_1(w)(1 - G(T_1(w))] > V_1(w) \end{aligned} \quad \text{(A3)}$$

as long as wages are strictly greater than UI benefits and the probability of receiving an acceptable wage is less than one.

Now assume that $V_n > V_{n-1}$. Using the same reasoning as in expression (A3), one finds that

$$\begin{aligned} V_{n+1}(w) >&= w + B\lambda_n(w) \int_{Rn(w)}^{\infty} [W(x) - V_n]dF(x) + Ba\lambda_n(w) \int_{Tn(w)}^{\infty} [V_T(y) - V_n]dG(y) + BV_n - .5(\lambda_n(w))^2 = V_n(w) \\ &\quad + B(V_n - V_{n-1})[1 - \lambda_n(w)(1 - F(R_n(w)) \\ &\quad - a\lambda_n(w)(1 - G(T_n(w))] > V_n(w) \end{aligned} \quad \text{(A4)}$$

by the induction hypothesis.

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