Working Paper no. **76**

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# are temporary jobs the shortest way?

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April 1, 2008

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# The (long) run out of unemployment: are temporary jobs the shortest way?*

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**Abstract**

A higher job creation is a common result by many theoretical approaches trying to model marginal labor market reforms. In the framework proposed by Berton and Garibaldi [2006], in particular, the equilibrium arrival rate of temporary job offers is expected to be higher than the arrival rate of permanent ones. In this paper I use a sample of prime aged male workers from WHIP in a competing risks framework in order to compare the duration of unemployment spells terminated by jobs that only differ in their formal duration. I find that the arrival rate of fixed term jobs is actually larger than the arrival rate of permanent ones; this result is robust to the main sources of unobserved heterogeneity. However, the average duration of unemployment in Italy is still very high and the liberalization of flexible contracts as a policy to reduce it did not completely solve the problem.

**Key words:** temporary jobs, unemployment duration, competing risks

## 1 Introduction

In the last decade the Italian labor market has been progressively liberalized through a long sequence of reforms. The introduction at the margin of new labor contracts with very limited or even no employment protection has been the leading feature of basically all of them. The purpose of this kind of strategy was to improve the employment opportunities as long as to reduce the unemployment rate. Following Bertola [1990], lower firing costs lead to a higher workers' turnover - and therefore to a higher expected arrival rate of job offers, a higher destruction rate and a lower unemployment duration - but have no effect on the average employment level. More recent approaches try to face directly the issue of marginal reforms. The results are quite homogeneous. In the model

*I thank Dalit Contini, Lia Pacelli, Roberto Quaranta and the seminar participants to the Phd Workshop held in Moncalieri for comments and suggestions. The CRT Foundation with "Progetto Alfieri" is kindly acknowledged for financial support. This research was partly developed while visiting the Center for Labor Economics, University of California, Berkeley.

by Blanchard and Landier [2002] firms only post temporary vacancies and after a period they decide whether to retain the worker on a permanent basis or to post another new vacancy; the destruction rate of temporary jobs increases with the EPL gap between permanent and temporary workers and simulations show that also the arrival rate of temporary job offers increases after the reforms; an excessive workers' turnover, however, may even result in a higher equilibrium unemployment rate. Cahuc and Postel-Vinay [2002] instead assume that only permanent contracts exist; however, an exogenous share of them may be interrupted at firm's will and at no cost after the first period. A higher share of contracts with no firing costs - a way to model the marginal reforms - increases both the job creation and the job destruction, thus leading to a higher workers' turnover. Finally, Berton and Garibaldi [2006] provide a model in which firms face a trade-off between the ex-ante job filling rate and the ex-post flexibility: temporary contracts prevent inefficient workforce retention, but the waiting time to fill a temporary vacancy is longer. From the workers' standpoint this means that temporary jobs imply both a higher arrival rate of job offers and a higher destruction rate. The average unemployment duration decreases, but - as in the paper by Blanchard and Landier - its equilibrium level may even increase.

A higher arrival rate of job offers after the reforms - and thus a lower average unemployment duration - is a common result of all the theoretical approaches briefly depicted so far. The model by Berton and Garibaldi, in particular, shows that in equilibrium the arrival rate of temporary job offers is higher than the arrival rate of permanent ones. The purpose of this paper is to check whether this result holds in the Italian labor market. Using a sample from WHIP of male unemployed workers strongly attached to the formal labor market in a competing risks framework, I find that the hazard rate of exit to temporary jobs is higher than to permanent ones. This result is stronger among the workers who receive the unemployment benefit, which is in line with the theoretical model. However, the unemployment duration is in absolute terms still very long, even for the workers moving to a non-standard job. As long as the liberalization of flexible jobs was expected to soundly reduce the unemployment duration, therefore, it seems that much work is still to be done.

The existing empirical literature supports these conclusions. Bover and Gomez [2004] use the Spanish labor force survey and show that the exit rates to temporary jobs are ten times larger; Van Ours and Vodopivec [2006], with Slovenian administrative data, find a positive correlation between the unemployment duration and the probability of finding a permanent job. Finally, Blanchard and Landier [2002] find that in France the higher arrival rate of job offers had almost no effect on the unemployment duration.

This paper proceeds as follows: section 2 describes the data and the sampling strategy, section 3 presents the econometric model and its specification, section 4 includes the results and section 5 the discussion. Section 6 concludes.

## 2 The data

WHIP is a dataset of individual work histories built up by LABORatoro R. Revelli from the social security administration (INPS) archives[1]. The reference population includes all the employees of the private sector, temporary workers from the public administration, craftsmen, traders, collaborators, professionals without an autonomous security fund and benefit recipients (unemployment and collective dismissals). WHIP also takes into account retirement but is not able to distinguish among unsupported unemployment, non-participation and unobserved employment spells (basically, permanent civil servants and irregular workers). Careers are observed monthly and the observed series cover the period 1985 - 1999.

Since my purpose is to measure the duration of unemeployment, I sample all the separations occurred from January to November 1998. The month of December has been excluded because the firm data series end in 1998, and possible (unobserved) merges and buyouts occurred between 1998 and 1999 may be the origin of spurious separations. The distance between each individual separation and the beginning of a new job is the object of my analysis; through direct access to raw individual data, the censoring time is at December 2002.

In order to identify proper unemployment spells I also exclude the individuals who have a high probability of non-participation or of unobservable employment: women (who may decide to stop working in order to take care of the children; see Berton and Pacelli [2007]), people working in the South (where the black economy is widely diffused), seasonal workers (who have a sort of "natural" employment cycle), temporary employees of the public sector (who are more likely to move to a permanent job in the public administration) and individuals aged under 20, over 40, retired or moving to retirement. Job-to-job transitions (a possible signal of voluntary separations[2]) and non-employees (whose data are poorer and in some cases less reliable) have been dropped too. Finally, when more than one separation for the same individual is observed, only the last one enters the sample[3]. It includes 4095 individuals (3560 uncensored unemployment spells) and 75727 person-period observations (see below).

### 2.1 Some descriptive statistics

Table 1 shows that this prime-aged sample is young on average (67% was at most 30 years old at separation) and works mainly in northern regions. Most of the workers saparated from a full time permanent job, thus confirming that open end contracts do not prevent workers from losing their jobs (see Berton *et al.* [2007]); the share of full time fixed term arrangements, however, is all but

[1] See www.laboratoriorevelli.it/whip for further details.

[2] I define a job-to-job transition as a new job that begins within the first month after the separation. Almost 67% of the job-to-job movers find a full time permanent job, while this share is only 42% in my sample.

[3] Since seasonal workers have been dropped and in 1998 there is no change in the business cycle (the Italian GDP has been growing from 1996 to 2001), the calendar time of the separation shouldn't be relevant for the probability of finding a new job.

Table 1: Who's in the sample?

| Age class | No. of obs. | Col. perc. | Contract at separation | No. of obs. | Col. perc. |
|---|---|---|---|---|---|
| 20 - 25 | 1608 | 39.3 | Full time open ended | 2452 | 59.9 |
| 26 - 30 | 1129 | 27.6 | Part time open ended | 259 | 6.3 |
| 31 - 35 | 769 | 18.8 | Full time fixed term | 579 | 14.1 |
| 36 - 40 | 589 | 14.3 | Part time fixed term | 70 | 1.7 |
| | | | Full time CFL | 276 | 6.7 |
| Work area | No. of obs. | Col. perc. | Part time CFL | 13 | 0.3 |
| | | | Full time apprenticeship | 338 | 8.3 |
| North-west | 1614 | 39.4 | Part time apprenticeship | 6 | 0.1 |
| North-east | 1366 | 33.4 | Other (supported, TWA) | 102 | 2.6 |
| Center | 1115 | 27.2 | | | |

Source: analysis on WHIP data

negligible. Finally, 18% of the sampled workers (728 individual observations) receive an unemployment benefit after the separation; this information will turn out to be relevant later in the paper.

The average unemployment duration (Table 2) is long - at least one year - even when censored observations are dropped from the sample (column 2). It increases with age, but once censored observations are dropped, the trend becomes U-shaped, possibly meaning that more mature individuals either find a job quickly or prefer/are forced to stay unemployed longer. As far as the labor contract at separation is concerned, then, two aspects are worth a comment: part time workers display longer duration, possibly meaning that they are less attached to the labor market[4]; on the contrary, CFLs enjoy the lowest waiting time.

The main concern of the paper, however, is on the contract at destination. Columns 3 and 4 compare two contracts that only differ with respect to their formal duration, i.e. full time permanent and full time fixed term (see the specification strategy); as expected, irrespective of the age and the work area, non-employment duration is longer for workers moving to a permanent job, the only exception being for individuals aged between 31 and 35. The waiting time in the temporary submarket, anyway, is all but short.

# 3 The econometric strategy

## 3.1 The Model

The problem is that of a single duration $T$ (unemployment) which is terminated by exit to one out of an exhaustive set of mutually exclusive possible destinations (the different types of jobs) described by the subscript $m = 1, ..., M$. One way of formulating a duration model with multiple destinations is to postulate the existence of $M$ independent latent durations $T_1, ..., T_M$ and to assume that unemployed workers enter the shortest of them, so that what one observes is $T =$

[4]For most of them, working part time is a matter of choice [ISTAT 2005]. Maybe, they prefer non-participation or unemployment to full time jobs.

Table 2: Average unemployment durations (months)

| | Overall mean | No censored obs. | To FT perm. | To FT fixed term |
|---|---|---|---|---|
| Age class | | | | |
| 20 - 25 | 17.8 | 13.4 | 14.1 | 12.6 |
| 26 - 30 | 18.4 | 12.8 | 13.0 | 12.3 |
| 31 - 35 | 18.2 | 12.6 | 12.3 | 13.3 |
| 36 - 40 | 20.8 | 13.1 | 12.4 | 11.2 |
| Work area | | | | |
| North-west | 18.2 | 13.2 | 13.4 | 12.6 |
| North-east | 18.2 | 12.2 | 12.1 | 11.7 |
| Center | 19.3 | 13.9 | 13.8 | 13.2 |
| Contract at sep. | | | | |
| Full time CFL | 16.3 | 11.8 | 13.9 | 7.8 |
| Part time CFL | 20.0 | 14.2 | 15.8 | 13.0 |
| Full time temp. | 18.2 | 13.2 | 12.9 | 12.1 |
| Part time temp. | 18.1 | 14.1 | 11.5 | 14.0 |
| Full time perm. | 18.6 | 12.9 | 12.8 | 12.7 |
| Part time perm. | 21.1 | 14.9 | 16.7 | 15.0 |
| Full time app. | 16.6 | 12.3 | 12.5 | 12.4 |
| Part time app. | 36.5 | 27.0 | 24.0 | 32.5 |
| Source: analysis on WHIP data | | | | |

$Min(T_1, ..., T_M, T_C)$, where $T_C$ is the latent time for right-censored observations. Lancaster [1990] shows that in this framework the transition intensities

$$h_l(t) = \lim_{dt \to 0} \frac{\Pr\{t \leq T \leq t + dt, D_l = 1 | T \geq t\}}{dt}$$

i.e. the probability of departure to state $l$ in the short interval $(t, t + dt)$ conditional upon survival to $t$, can be given the interpretation of a hazard rate.

In the following I will use this *competing risks* approach. I assume that the data are intrinsecally discrete, i.e. that exits to employment occurs in monthly cycles described by the index $k$. As shown by Jenkins [2005], the contributions to the likelihood funtion for an individual surviving $j$ cycles read

$$L_l = h_l(j)S(j-1) = \left[ \frac{h_l(j)}{1 - \sum\limits_{m=1}^{M} h_m(j)} \right] S(j)$$

for uncensored spells terminated by an exit to state $l$, and

$$L_C = S(j)$$

for censored observations, where $S(j)$ is the probability of surviving for $j$ cycles. The overall individual likelihood contribution for a survival lenght of $j$ months therefore is

$$
\begin{aligned}
L &= \left\{ (L_C)^{1-\sum\limits_{m=1}^{M}\delta_m} \right\} \times \left\{ \prod_{m=1}^{M} (L_m)^{\delta_m} \right\} = \\
&= \left\{ \prod_{k=1}^{j} \left[ 1 - \sum_{m=1}^{M} h_m(k) \right] \right\} \times \left\{ \prod_{m=1}^{M} \left[ \frac{h_m(j)}{1 - \sum\limits_{m=1}^{M} h_m(j)} \right]^{\delta_m} \right\}
\end{aligned}
$$

where $\delta_l$ is an indicator variable taking the value of one if the individual exits to the $l$th state.

Allison [1982] suggests a straightforward method for the estimation of the discrete time competing risks model depicted above. He assumes a logistic specification for the destination-specific transition intensities

$$
h_l(j) = \frac{\exp(\beta_l' X)}{1 + \sum\limits_{m=1}^{M} \exp(\beta_m' X)} \tag{1}
$$

and shows that the resulting likelihood function takes the same form of the likelihood of a multinomial logit applied to a person-period reorganised dataset. Since the multinomial logit model implies independence across the destinations, $h_l(j)$ reads as a hazard rate.

## 3.2 Specification issues

In the observed period plenty of different temporary contracts were available on the Italian labor market [Tronti *et al.* 2003]. In order to compare arrangements that only differ with respect to their formal duration and to avoid possible cost-related confounding effects, I focus my attention upon unsupported full time open end and unsupported full time fixed term contracts[5]. The probability in (1) is therefore specified with $M = 3$, including "other labor contracts" as the residual alternative. As for the individual characteristics, Jenkins [1995] suggests to include in matrix $X$

- a set of individual covariates (age, unitary wage, actual previous experience, local unemployment rate, work area, firm size, industry and dummies for welfare provisions); in my specification $X$ is time-invariant and captures the individual characteristics at separation;
- a set of time-dummies such that $d_t = 1[T = t]$, in order to identify the shape of the destination-specific baseline hazard in a non-parametric way.

What I expect is that the baseline hazard of exit to a fixed term job is higher than the one to a permanent. Since in a multinomial logit framework each

[5] In the following simply "fixed term" and "open end" or "permanent".

coefficient read as the partial effect of the covariate on the log-odds ratio between destination $l$ and the reference exit, I expect the exponentiated coefficients of the time-dummies to be larger than one for destination $l =$ *fixed term* when full time open end contracts are the normalizing state. Formally:

$$\exp(\beta_{d_t, l|l=fixed\ term}) > 1$$

The specificaton depicted so far suffers from two major limitations: unobserved heterogeneity is not controlled for and the alternative destination states are assumed to be independent. Both of them can be tackled using a mixed logit model, so that the hazard rate in (1) becomes

$$h_l(j) = \frac{\exp(\beta_l' X + \alpha_l)}{1 + \sum_{m=1}^{M} \exp(\beta_m' X + \alpha_m)} \tag{2}$$

where $\alpha$ is an individual and destination specific unobserved heterogeneity component. Following Haan and Uhlendorff [2006], $\alpha$ is independent of the set of observables $X$[6], is identically and independently distributed across the individuals and follows a trivariate normal distribution with mean $a$ and unrestricted variance covariance matrix $W$

$$\alpha \sim N(a, W) \text{ where } W = \begin{bmatrix} var_1 & cov_{12} & cov_{13} \\ cov_{12} & var_2 & cov_{23} \\ cov_{13} & cov_{23} & var_3 \end{bmatrix}$$

which allows for correlation among the destination states.

In order to estimate model (2) - with proper individual specific effects - a panel of repeated unemployment spells should be observed. Through WHIP data at the moment this is not possible; employment spells, defined as a continuative relationship between a worker and an employer, can be identified only up to 1998, since firm data are no more available from 1999 onwards. In addition, the information about the legal duration of the contract (open end, fixed term or seasonal) appears only from 1998[7].

This notwithstanding, the person-period structure of the data still allows to have repeated observations for the same individual and thereofore to identify the mixed logit model. The estimation procedure, however, is computationally burdensome even when using Halton sequences as suggested by Haan and Uhlendorff themselves. For this reason, and for the moment, I will present only the results from model (1), and I will furtherly discuss the main sources of unobserved heterogeneity in section 5.

[6] To the best of my knowledge, removing the hypothesis of orthogonality between $X$ and $\alpha$ requires to assume a simpler dynamic structure, such as a Markov chain [Magnac 2000]. Since the purpose of my analysis is measuring a duration, I won't follow this strategy.

[7] This is the reason why I chose to sample the flow of 1998 separations.

# 4 Results

## 4.1 The effect of the covariates

Table 3 reports the partial effects of the individual characteristics as they are observed at the moment of separation. Previous actual experience decreases the log-odds ratio of exit to a fixed term job; the local unemployment rate, working in a North-eastern region and holding a temporary contract before the separation have in turn a positive effect. None of these results is unexpected: after a sufficient experience in the labor market, in fact, one is expected to find her own personal stable job; this is coherent with Gagliarducci [2005] who finds that past experience - whatever the labor contract - increases the probability of finding a permanent job. Many reasons, then, stand behind a positive correlation between the unemployment rate and temporary jobs: for instance, a higher unemployment rate may be a signal of bad business, so that firms are less capable or willing to invest in a long run employment relationship; in addition, it also means that more individuals are seeking for a job (the labor market tightness is lower), and that firms use more temporary arrangements in order to screen them. As for the work area, Berton and Pacelli [2007] observe a larger share of temporary jobs in North-eastern regions than in the North-west, where permanent contracts are instead more diffused. Finally, Berton *et al.* [2007] and Picchio [2006] show that state dependence in the labor contracts affects Italian working careers, which explains the strongly positive effect of "fixed term contract" and "trainees or apprentices" respectively on the log-odds of exit to fixed term jobs and to other contracts[8].

The effects of the other significant determinants are more puzzling. The workers who separated from a small firm have a lower probability to get a fixed term contract. As long as they move to another small firm, two facts may be relevant: first, open end contracts are less binding for small firms and their opportunity cost is smaller[9]. In addition, small firms seem to prefer other types of temporary arrangements, such as collaborations [ISTAT 2004]; the coefficient for the exit to other jobs is in fact significant and larger than one. Unfortunately, the firm data are no more available after 1998, so that I'm not able to control for the size at destination. Also workers from the constructions sector display low odds ratios of exit to flexible jobs. Since most of the transitions occur within the same sector, I can assume that they move to another job in that sector, where many temporary positions are informal and therefore unobservable.

On the contrary, the unemployment benefit positively affects the odds ratios of exit to a fixed term job. This is somehow surprising, since the unemployment benefit should allow the unemployed workers to wait for better job opportunities; one should therefore expect longer unemployment spells and exits to stable

[8] The residual destination "other jobs" includes both CFL contracts and apprenticeship.

[9] Small (meaning with at most 15 employees) and large firms in Italy face different rules in case of unfair layoffs (i.e. without the "just cause"). Small firms simply have to pay the worker a compensation (from twice to ten times the last wage if the worker's tenure was larger than ten years), while large ones - in addition to the compensation - are forced to re-hire the worker.

Table 3: The effect of the covariates

| No. of obs.: 75727<br>No. of individ.: 4095 | | | | | | LR chi2 (108): 1299.79<br>Pr > chi2: 0 |
|---|---|---|---|---|---|---|
| | Exit to fixed term jobs | | Exit to other jobs | | No exit (cens. obs.) | |
| | $\exp(\beta)$ | Pr > $z$ | $\exp(\beta)$ | Pr > $z$ | $\exp(\beta)$ | Pr > $z$ |
| Part time | 0.940 | 0.773 | 1.716* | 0.000 | 1.363* | 0.004 |
| Age (years in 1998) | 0.996 | 0.697 | 0.950* | 0.000 | 0.999 | 0.885 |
| Montlhy wage | 1.000 | 0.983 | 1.000 | 0.331 | 1.000 | 0.666 |
| Previous exp. (months) | 0.995* | 0.000 | 0.999 | 0.507 | 0.996* | 0.000 |
| Collaborator in 1997 | 2.058 | 0.147 | 1.528 | 0.293 | 1.345 | 0.359 |
| Collaborator in 1998 | 1.587 | 0.430 | 1.429 | 0.438 | 1.205 | 0.617 |
| Unemp. rate in 1998 | 1.050* | 0.037 | 1.047* | 0.005 | 1.048* | 0.000 |
| North east | 1.449* | 0.006 | 1.322* | 0.002 | 1.257* | 0.000 |
| Centre | 1.271 | 0.098 | 1.158 | 0.146 | 1.128 | 0.100 |
| Small firm (< 20) | 0.779* | 0.026 | 1.238* | 0.006 | 1.036 | 0.520 |
| Constructions | 0.537* | 0.000 | 0.726* | 0.002 | 0.762* | 0.000 |
| Services | 0.984 | 0.896 | 0.927 | 0.379 | 0.933 | 0.272 |
| White collars | 0.991 | 0.956 | 1.596* | 0.000 | 1.365* | 0.001 |
| Managers | 0.664 | 0.711 | 0.390 | 0.249 | 1.712 | 0.166 |
| Fixed term contract | 2.664* | 0.000 | 1.262* | 0.036 | 1.214* | 0.018 |
| Trainees or apprent. | 1.867* | 0.000 | 2.424* | 0.000 | 1.551* | 0.000 |
| Sickness benefit | 1.030 | 0.854 | 0.948 | 0.639 | 0.979 | 0.789 |
| Coll. dismissal benefit | 1.657 | 0.551 | 5.786* | 0.000 | 1.599 | 0.269 |
| Severance payment | 1.674 | 0.219 | 1.292 | 0.402 | 1.288 | 0.269 |
| Ord. unemp. benefit | 1.849* | 0.001 | 1.218 | 0.130 | 0.947 | 0.559 |
| Red. unemp. benefit | 2.338* | 0.000 | 0.993 | 0.958 | 0.832 | 0.062 |

* significant at 95% level
Note: relative risk ratios; exit to permanent contracts is the baseline outcome

positions. This is not the case in Italy. Possibly due to the low amount of the benefit (30% of the wage for at most six months) and to the administrative controls that force to accept the first suitable job opportunity[10], the recipients display shorter durations [Berton and Pacelli 2007] and a higher probability to move to fixed term jobs.

Finally, some comments about the exit to the residual destination are in order. The state dependence in the labor contracts explains the positive effect of the "part time" variable and age is likely to capture di effect of experience. Tax and social fees rebates, then, induce the firms to hire collectively dismissed workers, who therefore move to some form of supported jobs.

## 4.2 The time dummies

Table 4 shows that in any moment during an unemployment spell the arrival rate of a fixed term job is higher than the arrival rate of a permanent one, meaning that whatever the amount of time one has already spent in unemployment, the expected waiting time for a fixed term job is lower than for a permanent job. This result is coherent with the theoretical priors discussed in the introduction. However, only three out of fifteen time-dummies coefficients are statistically significant.

[10]Following the model by Berton and Garibaldi, this should indeed be a temporary one.

Table 4: The duration dependence structure

| No. of obs.: 75727<br>No. of individ.: 4095 | | | | | | LR chi2 (108): 1299.79<br>Pr > chi2: 0 |
|---|---|---|---|---|---|---|
| Duration of unemp. (months) | Exit to fixed term jobs | | Exit to other jobs | | No exit (censored obs.) | |
| | exp($\beta$) | Pr > $z$ | exp($\beta$) | Pr > $z$ | exp($\beta$) | Pr > $z$ |
| 2 | 1.049 | 0.864 | 0.996 | 0.979 | 0.470* | 0.000 |
| 3 | 1.477 | 0.191 | 1.337 | 0.112 | 0.370* | 0.000 |
| 4 | 1.197 | 0.529 | 0.594* | 0.005 | 0.247* | 0.000 |
| 5 | 1.527 | 0.138 | 0.644* | 0.023 | 0.277* | 0.000 |
| 6 | 1.331 | 0.344 | 0.863 | 0.443 | 0.292* | 0.000 |
| 7 | 1.522 | 0.173 | 0.719 | 0.114 | 0.325* | 0.000 |
| 8 | 2.299* | 0.004 | 0.854 | 0.439 | 0.301* | 0.000 |
| 9 | 1.142 | 0.692 | 0.770 | 0.214 | 0.299* | 0.000 |
| 10 | 1.101 | 0.763 | 0.807 | 0.273 | 0.225* | 0.000 |
| 11 | 2.293* | 0.006 | 1.113 | 0.604 | 0.278* | 0.000 |
| 12 | 1.532 | 0.213 | 1.046 | 0.838 | 0.307* | 0.000 |
| 13 - 15 | 1.400 | 0.234 | 0.732 | 0.087 | 0.395* | 0.000 |
| 16 - 18 | 1.184 | 0.571 | 0.986 | 0.939 | 0.370* | 0.000 |
| 19 - 24 | 1.447 | 0.156 | 0.784 | 0.138 | 0.369* | 0.000 |
| 25 - 30 | 1.854* | 0.044 | 0.843 | 0.415 | 0.698* | 0.016 |
| * significant at 95% level<br>Note: relative risk ratios; exit to permanent contracts is the baseline outcome | | | | | | |

The firm size at destination and the unemployment benefit after the separation are likely to affect these estimates. As already mentioned before, open end contracts are less binding for small firms, since firm-initiated separations are not prevented. Small firms have therefore a lower incentive to use fixed term contracts; as a consequence, the expected arrival rate of permanent jobs could be overestimated by the presence of open end contracts that do not involve a long-run committment. WHIP data, anyway, do not allow at the moment to control for the firm size at destination.

Another major source of possibly confounding effects is the unemployment benefit. Benefit recipients are subject to administrative controls that force them to accept the first suitable formal job opportunity; this is actually in line with the assumptions of the model in the previous chapter, where any job is always preferred to unemployment. Non-recipients, on the contrary, are more free to refuse a job, to tune the search effort, to move to non-participation or even to work in the informal market. When estimating destintion-specific transition intensities one should therefore control for such differences. Table 5 reports the coefficients of the time dummies as interacted with a control for the unemployment benefit. As expected, in the recipients subgroup all (but one) the coefficients for exit to fixed term jobs are significant and far larger than one, while among non-recipients the estimated coefficients are instead larger than one but in general not statistically significant.

Table 6 shows that subsample selection does not explain these results. The unemployment benefit recipients display higher age, experience, monthly wage and a larger share of full time permanent workers; the local unemployment rate, the firm size and the share of workers coming from North-eastern regions are instead roughly the same. The individual characteristics therefore induce a

Table 5: The duration dependence structure: benefit recipients vs. non-recipients

| No. of obs.: 75727 | | | | | | LR chi2 (108): 1311.50 |
|---|---|---|---|---|---|---|
| No. of individ.: 4095 | | | | | | Pr > chi2: 0 |
| Duration of unemp. | Exit to fixed term jobs | | Exit to other jobs | | No exit (censored obs.) | |
| (months) | exp($\beta$) | Pr > $z$ | exp($\beta$) | Pr > $z$ | exp($\beta$) | Pr > $z$ |
| | | | Recipients | | | |
| 2 - 3 | 2.715* | 0.011 | 1.280 | 0.367 | 0.614* | 0.018 |
| 4 - 6 | 2.704* | 0.001 | 0.626* | 0.038 | 0.195* | 0.000 |
| 7 - 9 | 3.087* | 0.000 | 0.999 | 0.998 | 0.174* | 0.000 |
| 10 - 12 | 3.574* | 0.001 | 1.102 | 0.737 | 0.262* | 0.000 |
| 13 -18 | 3.590* | 0.001 | 1.252 | 0.428 | 0.400* | 0.000 |
| 19 - 24 | 4.186* | 0.002 | 0.957 | 0.909 | 0.510* | 0.013 |
| 25 - 30 | 3.232* | 0.034 | 0.520 | 0.205 | 0.534* | 0.041 |
| Beyond | 2.169 | 0.117 | 0.905 | 0.768 | 0.799 | 0.338 |
| | | | Non-recipients | | | |
| 2 - 3 | 1.332 | 0.309 | 1.070 | 0.669 | 0.388* | 0.000 |
| 4 - 6 | 1.357 | 0.269 | 0.703* | 0.028 | 0.281* | 0.000 |
| 7 - 9 | 1.676 | 0.070 | 0.720 | 0.057 | 0.345* | 0.000 |
| 10 - 12 | 1.622 | 0.086 | 0.932 | 0.667 | 0.254* | 0.000 |
| 13 -18 | 1.244 | 0.447 | 0.786 | 0.144 | 0.366* | 0.000 |
| 19 - 24 | 1.441 | 0.214 | 0.749 | 0.098 | 0.340* | 0.000 |
| 25 - 30 | 1.970* | 0.050 | 0.891 | 0.607 | 0.706* | 0.033 |

* significant at 95% level
Note: relative risk ratios; exit to permanent contracts is the baseline outcome

higher transition rate of the unemployment benefit recipients towards permanent jobs, so that the result of a higher fixed term jobs arrival rate is even reinforced.

## 4.3 Uneasy vs. stable workers

In this section I simulate the probability time pattern of two hypothetical workers, using the results from Table 5. I compare a typical uneasy profile - a trainee aged 20 with a small experience and small wage - with a more stable one, a thirty years old white collar who separated from a standard full time job after a quite long tenure (Table 7 provides a full description of the two simulated workers).

Provided that one has been unemployed up to time $t$ and conditional to be an unemployment benefit recipient or not, the uneasy worker always displays a lower probability to persist in unemployment (Figure 1). This result is consistent with the model from the previous chapter. The stable worker is likely to have a higher reservation outside utility and places a lower value upon finding a new job more quickly; in other words, he can wait longer, for a more suitable job opportunity. The gap disappears for very long durations.

With respect to the unemployment benefit, the differences gather at durations between three and twelve months. The non-recipients profile is flatter, while recipients display a more typical U-shaped time dependence, possibly meaning that either non-recipients are actually less attached to the labor market and many of them unobservably exit to non-participation, or that they are

Table 6: Recipients vs. non-recipients: the individual characteristics

| Characteristics at separation | | Recipients | Non-recipients |
|---|---|---|---|
| Mean age (years) | | 30.2 | 27.7 |
| Mean actual experience (months) | | 60 | 46.4 |
| Monthly wage (1998 Euros) | | 1123.5 | 947.7 |
| Local unemployment rate (%) | | 6.9 | 6.9 |
| Work area (%) | North-west | 33.1 | 40.8 |
| | North-east | 34.8 | 33.1 |
| | Centre | 32.1 | 26.1 |
| Firm size (workers) | | 484 | 469.8 |
| Labor contract (%) | Full time open end | 63.7 | 59.0 |
| | Full time fixed term | 19.0 | 13.1 |
| | Full time CFL | 5.8 | 7.0 |
| | Part time open end | 5.4 | 6.5 |
| | Ful time apprent. | 1.2 | 9.6 |
| | Other | 4.9 | 4.8 |

Table 7: Two profiles: uneasy vs. stable workers

| Characteristics at separation | Uneasy worker | Stable worker |
|---|---|---|
| Age | 20 | 30 |
| Previous experience (months) | 24 | 72 |
| Worked as collaborators | No | No |
| Local unemployment rate (%) | 7.0 | 7.0 |
| Work area | North-west | North-west |
| Firm size | $> 20$ | $> 20$ |
| Contract at separation | Trainee | Open ended |
| Part time at separation | No | No |
| Sector at separation | Manifacturing | Manifacturing |
| Occupation | Blue collar | White collar |

Note: no severance payments or collective dismissal benefits

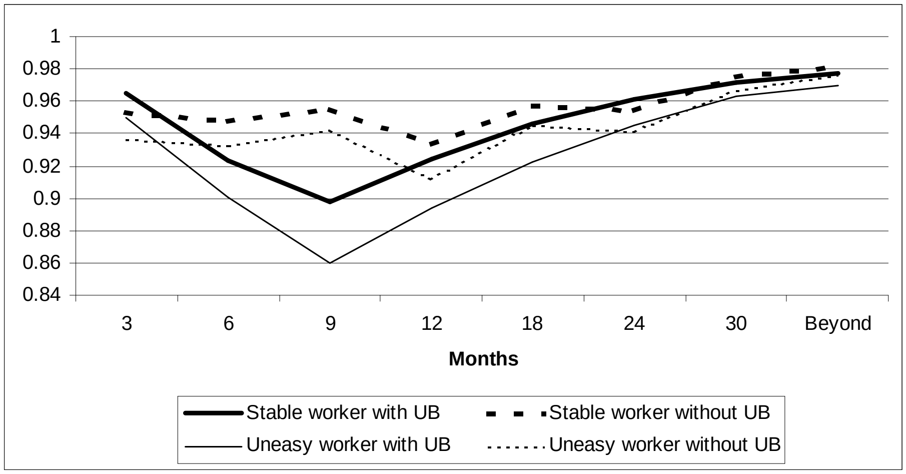


Figure 1: Probability of unemployment persistence

instead more likely to exit to unobservable jobs, for instance in the black economy. On the contrary, unemployment benefit recipients may be more attached to the formal labor market: they start searching for a formal job right after the separation, so that their exit rate initially increases with time; as time goes by, the share of less endowed individuals in the unemployment pool increases, and the exit rate starts decreasing. In addition, search effort is likely to be maximized at the end of the benefit, which explains the peak at six-to-nine months durations.

Anyway, the most outstanding fact from Figure 1 is the dramatically high persistence in unemployment. Averaging over the four profiles, after twelve months less than 30% of the simulated population finds a job, of any type; even within the observed prime-aged male sample I used for the estimates, this statistics is less than 55%. This means that any strategy aimed at reducing the unemployment duration - as the liberalization of temporary contracts was actually intended - would be of primary relavance from the labor market policy standpoint.

For the uneasy worker (Figures 2 and 3) the exit rates to a non-standard job are actually larger: on a population of 100 of them with (without) the unemployment benefit, after 12 months 8 (4) have joined a fixed term job, 9 (7) a permanent one and 20 (16) the residual category. This last destination includes both the apprenticeship and CFL contracts, and the result is thus consistent with the state dependence described by Berton *et al.* [2007]. On one hand it means that the simple higher arrival rate of fixed term job offers is not sufficient to overcome the effect of the individual characterisitics, of the state dependence, of possible port-of-entry contracts (as CFLs actually seem to have been: see again Berton *et al.*) and, eventually, of the incentive that firms


Figure 2: Probability of exit to a new job: uneasy worker with UB

have in using less costly arrangements. On the other, that the uneasy worker, in order to have a higher probability - but not decisive for a sound reduction of the unemployment duration - to find a job, must accept a less protected job and run the trap-risk.

For the stable worker (Figures 4 and 5) permanent jobs are slightly more likely and exits to fixed term contracts are almost negligible: out of a simulated population of one hunderd of them with (without) the unemployment benefit, after one year 12 (10) have gone to an open end contract, 13 (9) to the residual category and only 4 (1) to a fixed term job. In this case both persistence, individual characteristics and the higher reservation outside utility play a role, but, again, the exit rate from unemployment seems to be really low.

# 5 Discussion

The main issue in discussing the results I got so far is unobserved heterogeneity. In order to formally take it into account a mixed multinomial logit model with an individual random effect should be estimated. As discussed earlier, however, this procedure is computationally burdensome; the estimation is currently running, but the results are not yet ready.

This notwithstanding, unobserved heterogeneity can still be discussed through other instruments. First of all, Dolton and Van der Klaauw [1999] find that the specification with the time-dummies is not only robust to the misspecification of the duration dependence shape, but also to unobserved heterogeneity. Second, since the main source of unobserved heterogeneity is probably individual ability, one can argue that the arrival rate of fixed term job offers is even likely


Figure 3: Probability of exit to a new job: uneasy worker without UB


Figure 4: Probability of exit to a new job: stable worker with UB


Figure 5: Probability of exit to a new job: stable worker without UB

to be underestimated. Indeed, the most endowed individuals are more likely to find a permanent job, and to find it more quickly[11]; as time goes by, therefore, the sample is left with the less endowed individuals that instead have a higher probability to exit to a temporary job. Other relevant unobserved sources of individual heterogeneity are non-labor income and personal wealth. On one hand rich individuals can afford longer waiting times in order to find the job they like; as long as they look for a stable position, the arrival rate of fixed term jobs would be overestimated. On the other, personal wealth and non-labor income are positively correlated with individual ability; from this respect, the arrival rate of fixed term offers would therefore be underestimated. Underestimation, finally, occurs also if small firms are more prone to offer permanent positions that can be interrupted through a severance payment.

In addition, one has to consider that formal econometric approaches only allow to control for random effects; thus, the estimation of a mixed multinomial logit model does not allow to control for ability or individual wealth, which in turn are very likely to be correlated to the observed covariates. In sum, I therefore expect unobserved heterogeneity to affect the estimates but not the claim of a higher transition intensity to fixed term jobs.

Anyway a lower expected waiting time for a fixed term job also occurs when firms only open temporary vacancies in order to reach their optimal mix of permanent and temporary workers. This is probably what happened in Spain during the nineties, when 90% of new hires have been signed under a temporary arrangement [Dolado *et al.* 2002]. Analogously, Berton and Pacelli [2007] find that in Italy in 1998 most of the entrants (about 50%) found a temporary job;

[11] ISTAT [2006] find that education - the usual proxy for ability - is positively correlated with the probability to find a permanent job and negatively with the duration of unemployment.

however, only less than 12% signed a fixed term contract (including both part time and supported ones) while 27% is the share of new unsupported permanent workers. In the sample I used for the present paper the distribution is even more disproportionate: 12% of the observations exit to a full time fixed term job, and 37% to a full time open end one. Neither labor demand considerations, therefore, seem to affect the main conclusions of this paper.

From the labor market policies point of view a higher arrival rate of temporary job offers is of primary relevance. The reforms that took place since the mid-Nineties were intended - among the others - to create more job opportunities for the young workers and to reduce the unemployment duration, a problem affecting Italy from decades. This paper does not include a counterfactual analysis, so that I'm not in a position to properly evaluate the reforms. However, the simple descriptive statistics at the beginnig of the paper and the simulations I run show that after the reforms persistence in unemployment is still very high, a result found also by Blanchard and Landier [2002] for France.

# 6 Concluding remarks

In the last decade the Italian labor market has been liberalized through the introduction of more flexible labor contracts. The new contracts deeply loosen the employment protection legislation upon new hires. The theory shows that from the workers' standpoint a reduction at the margin of firing costs should result in a higher turnover, a higher job arrival rate and a lower unemployment duration, which actually are some of the goals of the reforms. The model proposed by Berton and Garibaldi [2006] - in particular - shows that after the liberalization a higher arrival rate of temporary jobs should be expected.

The purpose of this paper is to check this last implication. I use a sample of prime-aged male workers from WHIP in a discrete time competing risks framework and I compare the arrival rate of job offers that only differs with respect to their formal duration. The estimates show that at any moment during an unemployment spell the arrival rate of fixed term job offers is larger than the arrival rate of permanent ones, meaning that the expected waiting time for a fixed term job is actually lower. This result turns out to be robust to the major sources of unobserved heterogeneity. However, descriptive statistics and simple simulations show that workers in the temporary market still face very long unemployment spells, meaning that from the labor market policies point of view much work is still to be done.

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