Flexibility Policies and Re-Employment Probabilities in Italy

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Flexibility Policies and Re-Employment Probabilities in Italy

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Abstract:

We analyze the effects of Italian labor market reforms "at the margin" on the probability of exiting from non-employment and entering permanent and temporary contracts, using WHIP data for the period 1985-2004. We find that the reforms have strengthened the duration dependence parameter, meaning a stronger labor market gap in employment opportunities between the short- and long-term non-employed. We suggest that in a flexible labor market, long-term unemployment is used by firms as a screening device to detect less productive workers. We also find evidence of greater differences in employment opportunities according to gender, and of reduced differences between regional labor markets.

Keywords: reforms at the margin, duration dependence, signaling hypothesis, duration models
JEL classification codes: J64, J08, C41

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1. Introduction

There is broad consensus among economists in considering close regulation of the labor market as the principal cause of high and persistent unemployment in Europe since the 1970s (e.g., Saint-Paul, 1997, and Nickell et al., 2005). Bentolila and Bertola (1990) – and more recently, the OECD (2013) – argued that a high level of employment protection legislation (EPL) tends to reduce movement from unemployment into employment and vice-versa because firms become more cautious about hiring and dismissals are more expensive as a result of high firing costs. Since the mid-1980s, policy makers in Europe seeking to reduce EPL and increase flexibility have reformed conditions 'at the margin'[1] by encouraging the use of pre-existing temporary contracts and introducing new atypical contractual forms that are characterized by fixed duration and lower firing costs. These gradual and/or partial reforms have been implemented throughout Europe and have led to changes in the institutional framework of the labor market (Boeri and Garibaldi, 2007). These changes have not affected European labor markets univocally in terms of unemployment duration (Blanchard and Landier, 2002), job destruction and job creation (Cahuc and Postal-Vinay, 2002, OECD, 2004), the desirability of temporary jobs in terms of wage, job satisfaction and training (Booth, Francesconi and Frank, 2002), and the probability of transiting into permanent employment (Amuedo-Dorantes, 2000, van den Berg et al, 2002, and Picchio, 2008).[2]

The Italian labor market, which is one of the most rigid in Europe, has been affected by the introduction of policies promoting flexibility. From the mid-1980s to the mid-2000s, this labor market was the site of important reforms, which essentially made it easier to utilize flexible, atypical, and part-time job contracts. There is no single date for these reforms; instead, they were implemented gradually. Legislative changes, indeed, were introduced throughout this entire period. The first legislative change dates to the end of the 1980s and was followed by a gradual implementation of the reforms and by additional relevant steps toward flexibility in the second half of the period. The first reform 'at the margin', indeed, was implemented in 1997 (Law No. 196/1997, the so-called 'Treu Package'); this reform included the introduction of temporary agency contracts and made apprenticeship contracts more widely applicable by aiming to make the Italian labor market more dynamic and to reduce the unemployment rate. Subsequent reforms, Legislative Decree No. 368/2001, Law No. 30/2003 ('Biagi's Law'), and Legislative Decree No. 276/2003,

[1] Because flexibility has typically been pursued by introducing new temporary contracts rather than by reducing the hiring and firing costs related to permanent contracts, it is a common claim that labor markets have been reformed 'at the margin' because only new entrants – and particularly young workers – have really been affected by flexibility policies.

[2] Other studies on the 'dead end' or 'springboard' effects of temporary contracts include those by D'Addio and Rosholm (2005) and Güell and Petrongolo (2007).

legitimized fixed-term contracts under general conditions and reformed apprenticeship contracts, thus further extending the use of flexible and atypical work arrangements.[3]

The initial low take-up of these contracts was because their introduction was conditioned on collective agreements. Furthermore, increased adoption of temporary contracts was constrained by the timing of the renewals of collective agreements and by the rigid resistance of the unions to the spread of new contractual arrangements. For these and other reasons, there is no single date for these reforms; instead, they were implemented gradually. This reforming process had various effects on the Italian labor market. Montanino and Sestito (2003), Ichino, Mealli and Nannicini (2005), Gagliarducci (2005) and Berton, Devicienti and Pacelli (2011) highlighted the effects of the wider use of temporary contracts on the job prospects of young Italian workers (including the existence of port-of-entry effects). The results indicate that taking on temporary employment (rather than being unemployed) has a substantial positive effect on the probability of transition into a stable job that is conditioned on the type of temporary contract entered into and on previous labor market history. Destefanis and Fonseca (2007) assessed the Treu Package in terms of matching efficiency and found an improvement in efficiency for the northern regions but also an increase in competition among skilled and unskilled workers, particularly in the south of Italy. Jimenez-Rodriguez and Russo (2012) found that partial labor market reforms have increased the response of aggregate employment to output shocks. Cappellari, Dell'Aringa and Leonardi (2012) evaluated the effects of the legislative changes of the 2000s and found that the reform of apprenticeship contracts had a positive impact on job turnover and productivity and that the reform of fixed-term contracts had a substantial negative impact.

Despite the increasing number of studies available in the literature, various aspects of the flexibility process have remained unexplored or merit further investigation. The primary contribution of this paper is to fill this gap in the empirical literature: the main question addressed herein is whether and to what extent employment probability rates for non-employed young people have changed in Italy in the face of the important labor market developments during the period examined. In addition, although several studies have assessed the impact of one or two subsequent labor market reforms, we offer an evaluation both of the overall reform process between 1985 and 2004, and of its step-by-step impact. The lengthy time period examined herein allows for a more precise evaluation of changes over time in the Italian labor market. Our analyses also allow us to investigate whether the legislative innovations that occurred between the mid-1980s and the mid-2000s affected structural factors that are central to the Italian labor market. In particular, we focus

[3]Until the second half of the nineties, the standard work arrangement in Italy was full-time, open-ended, and characterized by one of the strictest anti-dismissal employment protection schemes in the OECD (e.g., Lazear, 1990, Kugler and Pica, 2008, Leonardi and Pica, 2013).

our attention on gender-related and geographical differences/gaps in the probability of exiting from non-employment.

The present study uses data from the Work Histories Italian Panel (WHIP) on young workers in the 16-35 age range and for the 1985–2004 period[4]. Given the availability of interval-censored data and the possibility of identifying the types of exit contracts, we apply discrete time hazard models with competing risks. We also control for unobserved heterogeneity.

Our main finding is that the sample period was characterized by a significant and gradual increase in the gap in re-employment probabilities (and those involving permanent contracts, in particular) between the short-term and the long-term non-employed, which corresponded to an increase in negative duration dependence. Furthermore, we also show that the investigated period was characterized by an increase in gender-related differences in transitions to permanent contracts, while the geographical gaps decreased.

We provide alternative interpretations to the former results, as discussed in detail in Section 5. Certain explanations rely on a reinterpretation of the predictions of the employer screening model (Lockwood, 1991; Omori, 1997 for an application to non-employment spells) that is set in a dynamic context in which job opportunities increase because of flexibility policies or because of positive labor market conditions (see Kroft, Lange and Notowidigdo, 2013). A further interpretation refers to the changes in the Italian economy in terms of workforce participation and composition during the sample period under investigation.

The remainder of this paper is organized as follows: Section 2 describes the institutional background and the employment composition of the Italian labor market; Section 3 describes the data and the samples; Section 4 provides the empirical specifications for this study; the results of the econometric analyses and an in-depth discussion of our findings are provided in Section 5; and Section 6 concludes.

2. Institutional background and employment composition

Since the early 1980s, the Italian labor market has been the site of a significant process of reform that aimed at increasing the flexibility of that market. As in other European countries, flexibility has typically been increased by extending the use of certain types of pre-existing temporary contracts and by introducing new atypical contractual forms that are characterized by fixed duration limits and lower labor costs when compared with permanent contracts.

The flexibility process has been characterized by the introduction of a number of laws that have gradually liberalized the use of fixed-term contracts and on-the-job training contracts (i.e.,

[4] For details on the WHIP data, see http://www.laboratoriorevelli.it/whip/documentazione.

apprenticeship contracts and 'contratti di formazione lavoro'). These laws have also introduced collaboration contracts, temporary agency contracts and a plethora of other temporary contractual forms[5].

During the 1980s, the first slight increase in flexibility was pursued by extending the use of fixed-term contracts through Laws No. 56/87 and 223/91. However, the first significant step toward the liberalization of the Italian labor market was taken in the 1990s with the 'Treu Package' (Law No. 196/97). The Treu Package substantially increased flexibility by introducing temporary agency contracts, the regulation of collaboration contracts, the liberalization of on-the-job training contracts and by certain innovations regarding fixed-term contracts.

Finally, during the early 2000s, Decree Law No. 368/01 provided for the significant liberalization of fixed-term contracts, whereas the 'Biagi Law' (Law No. 30/03 and subsequent Legislative Decree No. 276/2003) introduced a number of new temporary and atypical contracts (e.g., job-on-call, job sharing, staff leasing, etc.) and modified legislation covering apprenticeship and collaboration contracts. The latter two laws became effective in 2005.

The effectiveness of this intervention in increasing the flexibility of the Italian labor market is confirmed by at least two sets of statistics that consider it both step by step (reform by reform) and in terms of its overall impact,. The first of these sets are indexes measuring the strictness of employment protection for temporary and permanent contracts. According to the OECD statistics, whereas the protection index for regular employment remained stable at a level of 2.76 from 1990 to 2004, the temporary employment index decreased continually, from 4.87 in 1990 to 2 in 2004, with major decreases after the introduction of the Treu Package (from 4.75 to 3.62) and following legislative changes at the beginning of the 2000s. The second stage refers to the changes in Italian employment composition during the period examined, 1985-2004. Figure 1 shows the employment compositions for the overall labor force and for young people (15-24 age range), respectively.

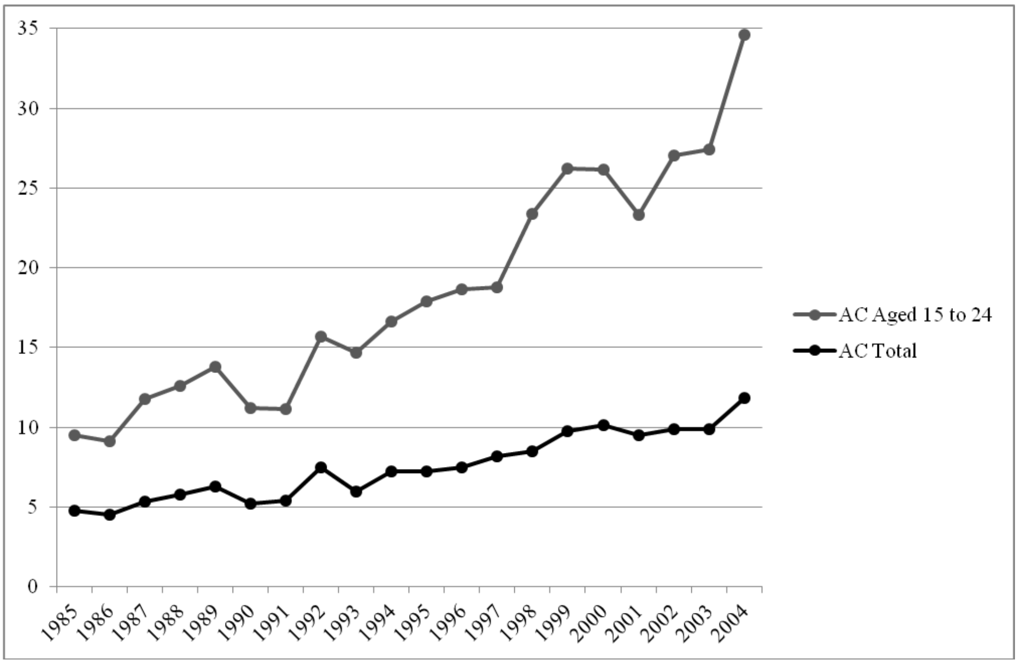
The share of temporary employment[6] with respect to total employment increased throughout the period, particularly for young workers in the age range of 15–24 years old. The percentage of

[5] Apprenticeship contracts and fixed-term contracts were originally introduced in 1955 and 1962, respectively, whereas the "contratti di formazione lavoro" were introduced in 1984. Since the 1980s, in addition to the laws listed above, a number of legislative changes to the regulation of labor market flexibility have been introduced. This intervention aimed at renewing the institutional framework by, for instance, introducing changes in incentives and obligations for employers, as well as changes in limitations on the applicability of temporary contracts (e.g. Law No. 108/90 and Law No. 451/94).

[6] The data for Figure 1 are available from the OECD statistics portal at: http://stats.oecd.org/. The information is only for employees. Temporary or atypical workers are defined under the following guidelines: Temporary employment is understood by both employer and employee as resulting when termination of the job is determined by objective conditions such as reaching a certain date, completion of an assignment or return of another employee who has been temporarily replaced. In the case of a work contract of limited duration, the condition for its termination is generally mentioned in the contract. These groups include the following: i) persons with a seasonal job; ii) persons engaged by an employment agency or business and hired out to a third party for the carrying out of a "work mission"; and iii) persons with specific training contracts.

the total labor force on temporary contracts increased from 4.8% in 1985 to 8.2% in 1997 (the year of the Treu Package). At the end of the period, it was approximately 11.8%. The corresponding share for the young in the 15-24 age range increased from approximately 9.5% in 1985 to approximately 18.8% in 1997, and at the end of the period (2004) more than one-third of the young workers (approximately 34.6%) were employed on temporary contracts.[7] As a consequence, the share of permanent employment decreased over the period, particularly for young people in the 15-24 age range (from approximately 90.5% in 1985 to approximately 65.4% in 2004). The statistics therefore confirm the need for both step-by-step and overall analyses of the impact of the Italian labor market reform/flexibility process for the 1985-2004 period.

Figure 1: Employment with Atypical Contracts, 1985-2004



Source: OECD statistics (2013)

3. Data

The Work Histories Italian Panel (WHIP) is a database of individual working histories based on the National Institute of Social Security (INPS) administrative archives and consists of a representative sample with a dynamic population of 370,000 individuals. The database provides full information for the 1985-2004 period, which was characterized by the introduction of a number of laws aimed at making the Italian labor market more flexible (see Section 2 for details).

[7] The indexes for the strictness of employment protection and the figures for employment composition are available online at http://stats.oecd.org.

The database permits the identification of job relationships on the basis of the social security contributions paid monthly to the INPS by employers and workers. This capability allows non-employment (NE) spells to be identified as periods characterized by the absence of paid social security contributions both by employers and workers.[8]

We study the effects of the labor market regulation on the hazard of exiting from non-employment to permanent and atypical contracts.

Because survival time occurs in continuous time but the spell lengths can only be observed at monthly intervals, the data are actually interval-censored. Although the data are available only up to 2004, the use of the WHIP dataset is nonetheless recommended for at least two reasons. First, it provides monthly information on private employment relationships, permitting the accurate estimation of transition times. Second, these data permit the effects of the gradual introduction of flexible employment contracts into the Italian labor market to be evaluated through several steps.

Based on the type of contribution rebates, it is possible to identify the contract forms held by individuals, i.e., permanent contracts (PCs) or atypical contracts (including on-the-job training contracts (OJTCs) and temporary agency contracts (TACs)), which indicates that a competing risks analysis can be performed.

From the original sample, we selected information for young individuals (aged 16 to 35) in the analyzed period because the flexibility policies were primarily targeted at young workers.

In detail, the waves (years) in which the individuals in the age range analyzed began a valid relationship with the INPS and become part of our dataset are displayed in the Appendix Table A1. For each year (wave) in our sample period (first column), this Table provides the number of individuals who began to be included in our dataset (Freq., second column). In addition, the third and fourth columns of this Table provide the percentage and cumulative percentage of individuals who were included through the 1985-2004 period.[9]

This selection resulted in a sub-sample of 44,114 individuals and 112,590 spells, which corresponded to 2,335,277 times at risk. This selection also allowed us to reconstruct complete individual working histories with accuracy and the impact of initial-condition problems is reduced because we can observe workers from the beginning of their careers.[10] The first month of a new

[8] For details on the WHIP data, see http://www.laboratoriorevelli.it/whip/documentazione. WHIP data do not present attrition problems (for details see Section 3.1); if either the worker or the firm is enrolled with INPS, they must provide INPS with all the relevant information (LABORatorio Revelli, 2009).

[9] This is important also for comparative purposes. For instance, an individual beginning his/her unemployment/inactivity experience in the 1980s and surviving until 2004 might be of a different type compared with those beginning unemployment/inactivity at a later date, particularly at the end of the period of observation. Needless to say, the increasing number of observations wave by wave (year by year) is the result of our selection, which includes individuals in the [16-35] age range.

[10] When constructing our sub-sample, if an individual was simultaneously in more than one work relationship, we eliminated the shorter job relationship; if they were of the same duration, we removed the part-time job or the work relationship characterized by fewer days of actual work. Finally, when the second job began before the end of the first

employment relationship permits identification of the time of exit from the state of non-employment, and the type of contract utilized in the new job makes it possible to identify the multiple failures that characterize the competing risks analysis. Because TACs represent only a small share of exit contracts, they are considered together with OJTCs indistinctly as atypical contracts (ACs).

Finally, because firms pay the same rate of social security contribution for permanent and fixed-term contracts,[11] i.e., a social contribution rate at 31% of gross earnings, and also because both PCs and FTCs do not provide any tax relief (e.g., Cappellari et al., 2012), FTCs are assimilated to PCs in the WHIP data.[12]

The WHIP data thus make available a set of individual and job-related variables. Specifically, information is provided on age, gender, working area, working conditions, firm-size, illness, wage, sector of economic activity and cumulative previous work experience in permanent and atypical contracts. With respect to working characteristics, these refer to the conditions held during previous work experience.

Finally, the business-cycle effect is controlled for by introducing the expected (next quarter) employment growth rate and by assuming rational expectations.[13] Descriptive statistics by time period for the most relevant variables are reported in Tables 1a and 1b in the usual manner.[14]

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job but ended after the first job, we censored the second work spell to the left and hypothesized that the second job started only when the first ended. In this way, the passage from a double job to a single job is understood as a transition from one job to another. This strategy is adopted to reconstruct non-employment duration spells with accuracy.

[11] A fixed-term contract of employment is defined as a contract of employment that has a definite start and end date, terminates automatically when a particular task is completed, or terminates after a specific event (other than retirement or summary dismissal). Prior to 2001, the law regulating fixed-term contracts provided a specific list of circumstances under which firms could use such contracts, such as during peaks in production or to replace workers on sick leave. Legislative Decree no. 368/2001 liberalized the use of fixed-term contracts to allow firms to use them to adapt quickly to changes in economic conditions.

[12] The WHIP dataset permits the identification of job relationships on the basis of the social security contributions paid to the INPS. For this reason (identification), given that firms pay the same (rate of) social security contributions for PCs and FTCs, the latter contracts are assimilated to PCs in the WHIP data. However, although we cannot exclude that this might affect the estimated hazard rates for the period 1992-Treu Package, it is plausible to believe that no significant estimation bias for the subsequent sub-periods arises from their non-identification, since the Treu Package, as explained in Section 2, resulted in only negligible changes to FTC legislation. Changes were indeed introduced in 2001 and became effective in 2003.

[13] Employment growth is measured with respect to the next quarter employment level using data from the 'Rilevazione sulle Forze di Lavoro', which is gathered by ISTAT.

[14] Descriptive statistics for the full set of covariates by time period are available upon request.

Table 1a. Descriptive statistics by time period, total sample.

	1985-1991		1992-Treu		Treu-2001		2002-2004	
	Mean	Std Dev.	Mean	Std Dev.	Mean	Std Dev.	Mean	Std Dev.
Age	18.574	1.734	21.631	2.824	24.187	3.988	26.095	4.859
Male	0.647	0.478	0.613	0.487	0.595	0.491	0.610	0.488
North-west	0.270	0.444	0.271	0.445	0.290	0.454	0.288	0.453
North-east	0.311	0.463	0.313	0.464	0.293	0.455	0.270	0.444
Center	0.171	0.376	0.180	0.384	0.194	0.396	0.194	0.395
South-Islands	0.248	0.432	0.236	0.424	0.223	0.416	0.248	0.432
Blue-collar	0.905	0.294	0.810	0.392	0.777	0.416	0.768	0.422
Manufacturing	0.402	0.490	0.344	0.475	0.295	0.456	0.238	0.426
Construction	0.145	0.353	0.135	0.342	0.111	0.315	0.130	0.336
Commerce	0.131	0.337	0.129	0.335	0.131	0.338	0.135	0.342
Tourism	0.178	0.382	0.159	0.365	0.149	0.356	0.157	0.363
Transport	0.016	0.127	0.030	0.170	0.044	0.206	0.045	0.208
Intermediation-Business	0.042	0.200	0.068	0.252	0.155	0.362	0.199	0.399
Other Sectors	0.086	0.281	0.135	0.342	0.113	0.317	0.096	0.294
Expected Employment Growth	-0.158	1.456	-0.395	1.482	0.797	1.407	1.253	1.454
Observations	13533		27075		39536		32446	

Source: our elaboration of WHIP data

Table 1.b Non-employment and employment spells: descriptive statistics.

	Mean	Std. Dev.
Average duration of non-employment	24.273	37.734
Average duration of non-employment (uncensored)	11.935	18.746
Average duration of permanent contracts	31.360	37.704
Average duration of atypical contracts	19.213	29.075
Average share of permanent contracts	79.87%	
Average share of atypical contracts	20.13%	

Source: our elaboration of WHIP data

3.1 Data Limitations

The Work Histories Italian Panel (WHIP), as described in Section 3, includes a representative sample of all the employee in the economic sectors of Industry, Construction and non-public Services that began a valid relation with the social security institute or INPS (Grand and Quaranta, 2011). Therefore, our data are a complete collection of all the information for employees working in the private sector.

As stated in Section 3, the use of the WHIP dataset for analyses of non-employment duration is recommended for a host of reasons. Nonetheless these data have certain limitations. In the following, we describe certain features of the data that must be borne in mind when interpreting the results of our analyses.

First, attrition, i.e., the fact that individuals are included in the data as long as they have a valid relationship with the INPS might be a relevant issue in these types of datasets and might reflect problems regarding data collection and management. However, as stated in the relevant literature/empirical evidence based on the WHIP data (e.g., Contini and Grand, 2010, Grand and Quaranta, 2011, Contini and Poggi, 2012) and by the specific documentation on those data (LABORatorio Revelli, 2009), the attrition that we observe is the product of perfectly explainable patterns of workforce utilization that have nothing to do with data collection.

Exits from the databases, as also explained by Contini and Poggi (2012), reflect transitions from employment to non-employment or unemployment. In principle, we do not have any attrition problems: once a certain group of individuals is selected (e.g., employee in the private sector), it is possible to follow them over the entire working life.

There are, however, a few exceptions to this principle: (i) movements from the private to the public sector go unrecorded when the job switch is accompanied by full tenure in the public sector, but this type of movement is rare for young people[15] in the period analyzed in our paper, 1985-2004, i.e., movements from the private to the public sector absorb approximately 2% of the young people (Dickson et al. 2012); (ii) shifts from private employment to self-employment, which are quite uncommon and involve only 0.8% of the young people employed in our sample; and (iii) movements from the private sector to the employment condition of 'collaborators',[16] but such movements absorb only 0.2% of the young people employed in our sample (see for instance Muehlberger and Pasqua, 2009). The case of no re-entry is also possible. Although exceptions (i), (ii) and (iii) involve a small number of individuals, the no re-entry includes movements from regular employment to the hidden or black market, which might be frequent particularly for low educated young workers, but they are undetectable by definition.

Finally, it must be borne in mind that the category 'non-employment' in the WHIP dataset includes individuals in different states of the labor market, i.e., unemployment and inactivity. As a consequence of the characteristics of the data explained above, we cannot capture transitions from states of unemployment and inactivity (together, "non-employment") to public sector and self-employment. However, the transitions between the two states, 'unemployment' and 'employment',

[15] Our sample, as explained above, includes young people from the age of 16 to the age of 35.

[16] 'Collaborators' are workers with a contract of continuous collaboration. Because of the ambiguity of the legal definition, the statistics available (mainly INPS and ISTAT) cannot give an accurate measure of the phenomenon. In any case, the WHIP dataset provides yearly-based information on collaborators, making it impossible to include them in our month-based duration analysis. For this reason we eliminated individuals who held at least one collaboration contract from our sample t, and this results in a reduction of about 2% of sampled individuals. Nonetheless this does not particularly affect estimation results.

account for a low share of total transitions. For instance, the transitions from unemployment to public sector employment and to self-employment absorb only 4.7% and 2.8% of the total labor force, respectively. In addition, movements from inactivity to both public sector and self-employment are also low during the sample period (Dickson et al. 2012).

The WHIP data, therefore, allow for following the employee in the private sector through their entire working life. Indeed, the exceptions and limitations explained above involve low shares of the labor market transitions of young people. In other words, the attrition that we observe is the product of perfectly explainable patterns of workforce utilization that have nothing to do with data collection.

4. Econometric specification

The duration analysis is developed using standard job search tools. Because the available data are interval-censored, discrete-time hazard models are estimated (Prentice and Gloecker, 1978). According to hazard models, the conditional probability that a transition to employment (either permanent or atypical) occurs in a given interval $[a_{j-1}, a_j)$ in the j_{th} period, conditional on the time already spent in non-employment, is defined as:

$$h_j \equiv \Pr\{T \in [a_{j-1}, a_j) \mid T \geq a_{j-1}\}. \qquad (1)$$

Assuming unit length intervals, the realization j of the discrete random variable T is the recorded spell duration.

A discrete-time hazard model requires that data are organized into a 'sequential binary form', which implies that data form an unbalanced panel of individuals with the i_{th} individual contributing to $j = 1, 2, \ldots\ldots t$ observations (where j is the number of periods at risk of the event).[17] Because some individuals transit to employment and possibly revert back to unemployment, multiple spells may be observed, $q = 1, 2, \ldots$ Q.

Models are estimated by assuming independent competing risks, which permits us to estimate models separately for each destination state (Narendranathan and Stewart, 1993). We adopt a cloglog specification, which consists of the discrete time representation of a continuous time proportional hazard model.

[17] Specifically, a binary dependent variable was created. If individual i's survival time is censored, then the dependent binary variable always takes the value of zero. If individual i's survival time is instead not censored, the dependent binary variable is zero in the first j-1 observation and one in the last observation.

In our analysis, we estimate a piecewise constant baseline hazard by using a non-parametric piecewise constant exponential specification, i.e., groups of months are assumed to have the same hazard rate, but the hazard may differ among groups. The total spell of non-employment is divided into specific sub-spells (D) for specific groups of months (e.g., D1_3 for the spells of non-employment with a duration from one to three months, D4_6 for a duration of four to six months, etc.).[18] The model is estimated by maximum likelihood, and the partial log-likelihood function for each destination, permanent contract (PC) or atypical contract (AC), is represented by the following:

$$\log L(\beta,\gamma)=\left[\sum_{i=1}^{N}\sum_{q=1}^{Q}\sum_{j=1}^{t}\left[y_{iqj}\log h_{iqj}+\left(1-y_{iqj}\right)\log\left(1-h_{iqj}\right)\right]\right]_{PC}+\left[\sum_{i=1}^{N}\sum_{q=1}^{Q}\sum_{j=1}^{t}\left[y_{iqj}\log h_{iqj}+\left(1-y_{iqj}\right)\log\left(1-h_{iqj}\right)\right]\right]_{TC} \quad (2)$$

where y_{ij} takes the value of one if the individual transition occurs in month j (i.e., the spell is uncensored) and zero otherwise. Because of the independence assumption, the total log-likelihood function logL(β,γ) is the sum of the partial log-likelihood function derived for the contract of destinations PC and AC.

The model presented above assumes that all the differences between individuals are captured by observed explanatory variables. However, it is well known that it may be relevant to use a model that allows for unobservable individual effects to prevent estimation bias that derive from omitted variables and/or measurement errors in the observables, for instance (Jenkins, 2005). Unobserved heterogeneity is modeled by assuming a Gaussian distribution defined at the individual level. We estimate random-effect cloglog models.[19] By avoiding any assumption about the functional form of the baseline hazard, i.e., by adopting the piecewise constant specification, estimation bias problems are reduced, and the estimation results may be considered reliable (Nicoletti and Rondinelli, 2010).

In addition, flexible employment contracts were introduced into the Italian labor market gradually, and the reforms proceeded in several steps. To account for the full reform process in detail, and also for the step-by-step impact, we split the overall period into four sub-periods. The first period, 1985-1991, predates the period when the changes in flexibility regulations became effective; the second period begins in 1992 (after the first changes had been introduced) and ends in 1997 (June), i.e., just before the Treu package. The third period runs from the introduction of the

[18] Table 2 contains the complete set of duration dependence estimates.
[19] We use STATA (ver 12.1) statistical software, which provides a command, xtcloglog, to estimate random-effects complementary log-log models. Random effects estimation might be biased if the omitted effect is correlated with explanatory variables, whereas fixed-effects estimation is not applicable in a complementary log-log framework. A second-best approach would require the application of a logistic model (that would be particularly indicated when survival times are intrinsically discrete), for which both random effects and fixed effects are allowed.

Treu package to the subsequent reform of 2001. Finally, the last period covers the years from 2002 to 2004.

A set of interaction dummy variables (*R*X*) is introduced to identify the effects of the labor market reforms on the duration dependence parameters and on other explanatory variables, *X*. α and β are vectors of unknown parameters, including intercepts to be estimated that refer to *X* and (*R*X*), respectively. The effect of the variables in the post-reform period is summarized by ($\alpha + \beta$). Finally, γ_j and $\left(R * \gamma_j\right)$ summarize the baseline hazard and the interaction of the baseline hazard with the dummy for the labor market reforms, respectively.

The hazard function assuming a complementary log-log specification with Gaussian unobserved heterogeneity is defined as:

$$h\left(j, X \mid \nu\right) = 1 - \exp\left\{-\exp\left[\alpha' X + \left[\sum_{a=1}^{3} \beta_a^{'}\left(R_a * X\right)\right] + \gamma_j + \left[\sum_{a=1}^{3}\left(R_a * \gamma_j\right)\right] + \log\left(\nu_i\right)\right]\right\} \qquad (3)$$

where $\log(\nu) \equiv u$ has a Normal distribution with zero mean and finite variance, and with a=1, 2, 3 for the three cutoff points of the overall period. Specifically, a=1 in (December) 1991, a=2 in (June) 1997 (Treu Package), and a=3 in (December) 2001.

To estimate this model, it is necessary that both the survival and the density function expressions that enter the likelihood function are not conditioned by unobserved effects. Therefore, the likelihood contributions are obtained by integrating out the random terms, as is appropriate for the Gaussian case because the integral does not have a simple closed form.

5. Results

Tables 2-4 show the duration dependence parameters and the coefficient estimates for the total sample examined over the 1985-2004 period. Specifically, Table 2 shows duration dependence parameters for each period of non-employment. We divided the total non-employment duration into nine periods, which correspond to specific numbers of months (e.g., the first period for a duration of from one to three months; the second period from four to six months, etc.).

Columns 1 and 5 of Table 2 report, for the NE-PC and NE-TC transitions, respectively, the estimated coefficients related to the baseline hazard specification for the base-category period (1985-1991). Table 2 also displays the estimated coefficients[20] of the baseline hazard specification

[20] Estimated coefficients related to the NE-PC transitions are often larger than those related to the NE-AC transitions, although the policy changes had more to do with atypical contracts than with permanent contracts. This result might have ensued because the reformatory process affected the entire hiring policies of firms and search behaviors of individuals, regardless that the application of legislative changes were limited to ACs.

interacted with dummy variables identifying the following sub-periods: 1992-June 1997 (Treu Package) (columns 2 and 6), July 1997 (Treu Package)-2001 (columns 3 and 7), 2002-2004 (columns 4 and 8). It follows that the estimated coefficients reported in columns 2-4 and 6-8 identify the exact change, with respect to the base-category period (1985-1991), of the baseline hazard parameters. For example, looking at Table 2, the estimated coefficient of the dummy variable 1-3 months of the NE-PC transitions is equal to 0.489 (γ) for the 1985-1991 period, whereas the estimated coefficient of the dummy variable for 1-3 months of the NE-PC transitions interacted with the 1992-June 1997 dummy variable is 0.358 ($R^*\gamma$). This finding means that the baseline hazard parameter for 1-3 months non-employment duration has increased by 0.358 from the 1985-1991 period to the 1992-June 1997 period and that the parameter for the overall 1992-June 1997 period is 0.847 (the sum of γ and $R^*\gamma$ coefficients). In other words, the $R^*\gamma$ coefficient identifies the significance and the magnitude of the change in a specific parameter for each period.

It should be noted that estimated coefficients related to the NE-PC transitions are often larger than those related to the NE-AC transitions, although the policy changes had more to do with atypical contracts than with permanent contracts. This result might have ensued because the reformatory process affected the entire hiring process of firms and search behavior of individuals, regardless that the application of legislative changes were limited to ACs.

Table 2 and Figures 1 and 2 reveal a negative relationship between the hazard rate and the time spent in non-employment (negative duration dependence).[21]

We also found that the probability rate has a non-monotonic pattern with two peaks for periods of one to three months' and 10 to 12 months' duration; however, after one year of unemployment the probability of finding a job (either permanent or atypical) falls quite sharply. It should be noted that the increase in re-employment probabilities for individuals experiencing 10-12 months' duration of non-employment might be a consequence of the design of standard unemployment benefits in Italy. In fact, individuals younger than 50 years old, who have been registered at least two years at the National Institute of National Security and who have paid at least 52 weeks of contributions have the right to eight months of unemployment benefits. It follows that individuals might possibly reduce their search efforts during the period in which they are protected by unemployment benefits, whereas these efforts sharply increase as the provision of unemployment benefits ends.

[21] Quite reassuringly, estimation results obtained from a random effects logistic model provide results that are strictly comparable with the random effects cloglog model. The fixed-effects specification, although providing estimations different in magnitude, provides similar results in statistical and economic meanings, i.e., the strengthening of the negative duration dependence.

Table 2. Duration dependence for exits to Permanent and Atypical Contracts, total sample 1985-2004.

	Permanent Contract				Atypical Contract			
	D	D*R1991	D*R1997	D*R2001	D	D*R1991	D*R1997	D*R2001
	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.
1-3 months	0.489***	0.358***	0.663***	0.833***	0.420***	0.089	0.505***	0.556***
4-6 months	-0.176***	0.387***	0.573***	0.773***	-0.157**	-0.042	0.372***	0.525***
7-9 months	-0.054	0.350***	0.545***	0.759***	0.016	-0.043	0.190*	0.334***
10-12 months	0.510***	0.101	0.283***	0.177***	0.712***	0.151	0.200**	0.086
13-18 months	base-category				base-category			
19-24 months	-0.270***	0.111	0.119	0.116	0.022	-0.228*	-0.198*	-0.131
25-36 months	-0.486***	0.081	0.110	0.098	-0.405***	-0.091	-0.206	-0.074
37-48 months	-0.629***	0.023	-0.035	-0.065	-0.638***	-0.262	-0.211	-0.116
more than 48 months	-0.649***	-0.249	-0.469***	-0.798***	-1.107**	-0.062	-0.379	-0.755*

Notes: We split the total period into four subperiods to explain the full reform process in detail. The first period, from 1985 to 1990 (D), predates the reforms. The second (D*R1991) and third (D*R1997) periods are before and after the Treu Package, i.e., from 1991 to the introduction of the Treu Package (June 1997) and from July 1997 to 2001, respectively. The last period (D*R2001) runs from 2002 to 2004.

* Significant at the 10% level; ** significant at the 5% level; and *** significant at the 1% level.

Source: our elaboration of WHIP data

Overall, the estimation results reveal that flexibility policies affected duration dependence parameters significantly, particularly after the introduction of the Treu Package (i.e., the third and fourth periods). In general, with few exceptions, we find positive interaction dummy coefficients for Short-Term Non-Employment (STNE) and smaller positive or negative interaction dummy coefficients for Long-Term Non-Employment (LTNE).[22] This finding indicates a strengthening of the negative duration dependence for LTNE (Table 2). In fact, as anticipated in the methodological section, the coefficients reported in columns 1 and 5 refer, for NE-PC and NE-TC transitions, respectively, to the baseline hazard dummy variables in the pre-reform period (1985-1991), whereas coefficients reported in columns 2 to 4 and in columns 6 to 8, refer to the baseline hazard dummy variables interacted with dummy variables identifying successive periods, i.e., 1992-June 1997, July 1997-2001, and 2002-2004, respectively. It follows that the coefficients reported in columns 2-4 and 6-8 identify the change (and the significance of that change) in the baseline hazard parameters due to the introduction of flexibility legislation. For example, referring to the 1-3 months of non-employment duration for NE-PC transitions, the estimated parameter is 0.489 for the 1985-1991 period, whereas it increases to 0.847 in the 1992-June 1997 period (i.e., the sum of 0.489 for the 1985-1991 period and 0.358 for the 1992-June 1997 period), to 1.152 in the July 1997-2001 period

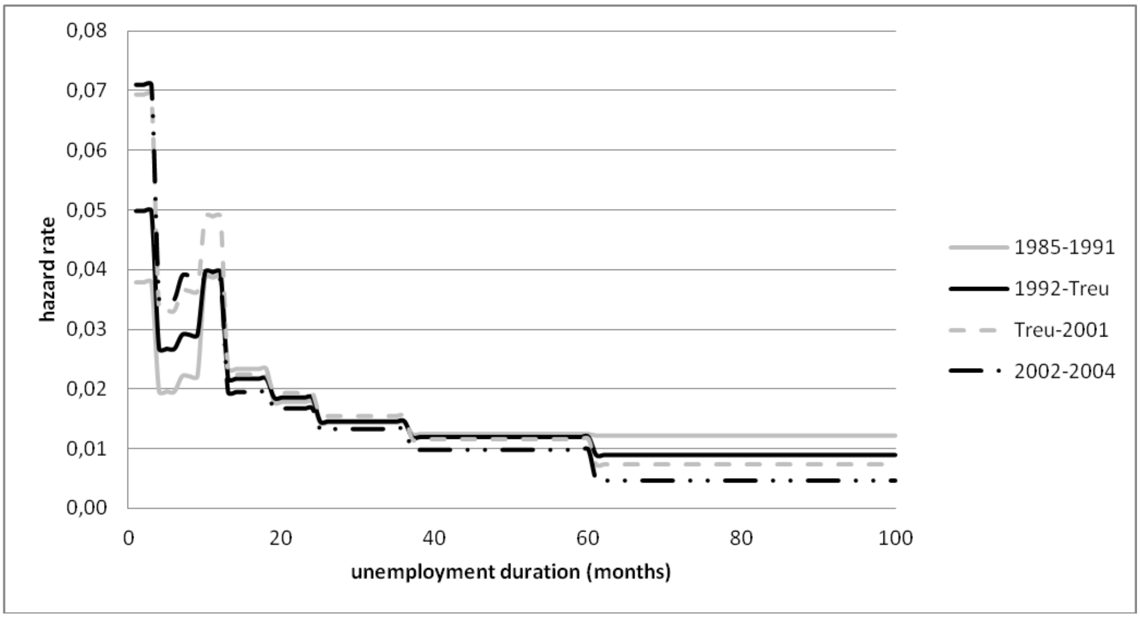
[22] STNEs are individuals in the state of non-employment for less than twelve months, whereas LTNEs are individuals in the state of non-employment for twelve months or more.

(i.e., 0.489 plus the estimated change in the parameter for the period July 1997-2001 taking value 0.663), and to 1.322 in the 2002-2004 period (i.e., 0.489 plus the estimated change in the parameter for the period 2002-2004 taking value 0.833).

A graphical analysis (Figures 2 and 3) helps clarify that, in addition to strengthening the negative duration dependence, the Treu Package also slightly increased the hazard rates for exits, particularly with respect to permanent contracts and from the second to the third period (from the Treu Package to 2001). This finding is possibly explained by the positive effects of both the business cycle, which improved in the post-reform period, and of greater job creation due to the implementation of flexibility policies.

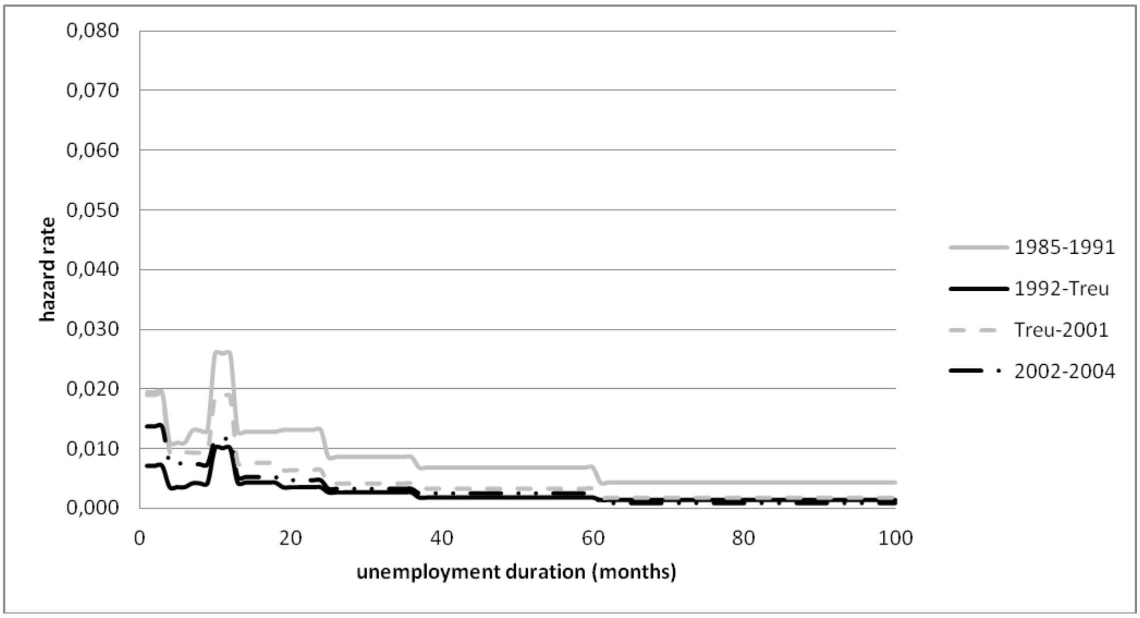
Finally, Figure 4 shows that for young people (age 16-24), the probability of exiting from non-employment and entering into atypical contracts more clearly increased with the introduction of the changes in labor market legislation (from introduction of the Treu Package to 2001 and also into the final period) compared with the probability for the total sample (Figure 3). This finding might suggest that the flexibility policies were more effective for young people.

Figure 2. Piecewise constant baseline hazard rate for permanent contracts by period, Total Sample, 1985-2004.



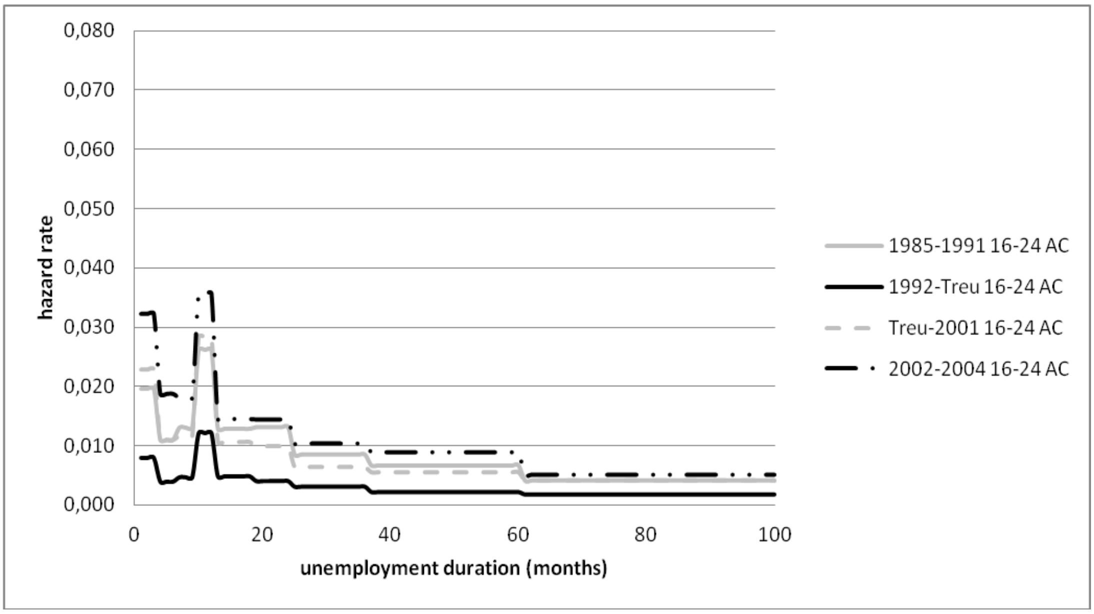
Source: our elaboration of WHIP data.

Figure 3. Piecewise constant baseline hazard rate for atypical contracts by period, Total Sample, 1985-2004.



Source: our elaboration of WHIP data.

Figure 4. Piecewise constant baseline hazard rate for atypical contracts by period, individuals aged 16-24, 1985-2004.



Source: our elaboration of WHIP data.

Estimation results are consistent with the hypothesis of a stronger STNE/LTNE gap in re-employment opportunities due to the introduction of labor market reforms, particularly the Treu Package. The stronger negative duration dependence parameter in the more flexible labor market can be understood as the empirical consequence of the use (by firms) of non-employment duration as a screening device in the hiring process.[23] This theory is reinforced by the fact that previous work experience is found to have less of an effect on re-employment probabilities after the legislative changes aimed at making the Italian labor market more flexible. Table 3 shows the coefficients for cumulative job experience and their positive but decreasing impact on re-employment probabilities through the period, particularly for permanent contracts.

Table 3. Cloglog coefficient estimates for previous job experience, total sample, 1985-2004.

	Permanent Contract											
	D			D*R1991			D*R1997			D*R2001		
	Coef.	s.e.		Coef.	s.e.		Coef.	s.e.		Coef.	s.e.	
Cumulated PC	0.048	0.004	***	-0.032	0.004	***	-0.045	0.004	***	-0.046	0.004	***
Cumulated PC square	-0.001	0.000	***	0.000	0.000	***	0.001	0.000	***	0.001	0.000	***
Cumulated AC	0.034	0.004	***	-0.017	0.004	***	-0.021	0.004	***	-0.023	0.004	***
Cumulated AC square	0.000	0.000	***	0.000	0.000	**	0.000	0.000	**	0.000	0.000	**
	Atypical Contract											
	D			D*R1991			D*R1997			D*R2001		
	Coef.	s.e.		Coef.	s.e.		Coef.	s.e.		Coef.	s.e.	
Cumulated PC	-0.011	0.008		0.008	0.008		0.000	0.008		0.002	0.008	
Cumulated PC square	0.000	0.000	**	0.000	0.000	*	0.000	0.000	*	0.000	0.000	*
Cumulated AC	0.071	0.006	***	-0.026	0.007	***	-0.032	0.007	***	-0.049	0.007	***
Cumulated AC square	-0.002	0.000	***	0.001	0.000	***	0.001	0.000	***	0.001	0.000	***

Source: our elaboration of WHIP data

This finding seems to suggest that previous job experience has a less significant signaling role in the post-reform period with respect to worker quality because the reforms have transformed non-employment duration into a more effective screening device.

[23] The use of unemployment duration as a screening device – a signaling hypothesis – is explained in Section 5.1, which offers a link to the literature and reformulates the hypothesis to explain our findings for Italy.

Table 4. Cloglog coefficient estimates, total sample, 1985-2004.

	Permanent Contract				Atypical Contract			
	D	D*R1991	D*R1997	D*R2001	D	D*R1991	D*R1997	D*R2001
	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.	Coef.
Age	-0.433**	1.766***	1.131***	0.927***	-0.132	-0.890***	0.342	0.137
Age square	0.014***	-0.042***	-0.027***	-0.022***	-0.007	0.024***	-0.001	0.004
Male	-0.167***	0.206***	0.302***	0.343***	-0.028	0.240***	0.057	0.031
North-West	0.389***	-0.050	-0.108**	-0.153***	0.276***	-0.245***	-0.062	0.175**
North-East	0.325***	0.055	-0.058	-0.080	0.297***	-0.030	0.046	0.106
Center								
South-Islands	-0.192***	-0.072	0.008	0.174***	-0.764***	0.017	0.272***	0.406***
Blue-collars	-0.252***	0.073	0.019	0.122**	0.175	0.026	-0.071	0.085
Manufacturing								
Construction	0.008	-0.129**	-0.106*	-0.018	0.066	-0.228***	-0.315***	-0.342***
Commerce	-0.163***	0.055	0.048	0.137**	-0.179***	0.140*	0.188**	0.133*
Tourism	0.107**	0.045	-0.144***	-0.058	-0.168***	-0.056	-0.260***	-0.279
Transport	0.035	0.136	0.115	0.183	-0.517**	0.426	0.325	0.279
Intermediation-Business	0.220***	-0.236***	-0.178**	-0.107	-0.093	-0.065	0.038	0.188
Other sectors	0.083	-0.157**	-0.352***	-0.217***	-0.351***	0.121	-0.086	0.046
Expected employment growth	0.107***	-0.002	-0.043***	-0.079***	0.208***	0.044***	-0.054***	-0.177***

Notes: All specifications also control for firm size, wage, part-time contract and health.

* Significant at the 10% level; ** significant at the 5% level; and *** significant at the 1% level.

Source: our elaboration of WHIP data.

Table 4 shows the effects of the other covariates used in our analysis. In addition to cumulative job experience, we also control for other individual, job-related and macroeconomic characteristics.[24] All the estimates must be interpreted as the relative effect with respect to the base category outcome, i.e., the state of non-employment.[25] The reforms have inverted the effect of age both on the NE-PC and on the NE-AC transitions. This latter effect possibly reflects both the greater range of applicability of ACs with reference to age as a result of the introduction of the Treu Package, and a greater possibility of taking on repeated temporary jobs based of this extension.

We add controls for gender and area of residence, which are two relevant and structural characteristics of the Italian labor market. The estimation results for the pre-reform period confirm the differences for both gender and geographical location. Nonetheless, to better assess whether differences between gender and geographical area of residence have changed in any way, we replicated our estimates for certain meaningful sample cuts (men versus women; North versus

[24] We also control for firm size, wage, part-time contract, and health. The full set of estimates is available upon request.
[25] It follows that an estimated coefficient with a positive sign indicates that the explanatory variable positively affects the re-employment probability rather than favoring permanence in the non-employment state. Moreover, because the non-employment state is the common base category, the sign and the magnitude of the same explanatory variable estimated for different transitions (NE-PC or NE-AC) define the differential effect (due to a specific covariate) on the probability of transition into alternative employment statuses.

Center and versus South). The graphs of the related hazard rates are provided in the Appendix, Figures A1-A4.[26]

We find that the changes in labor market legislation did increase gender-related differences, particularly with respect to the NE-PC transitions, whereas atypical job opportunities remained unchanged after the introduction of the Treu Package. In general, this finding suggests that an increase in the employment gender gap because males have increased – on a relative basis – (with respect to females) their opportunities of getting stable and more protected jobs, whereas females have similar probabilities of males of being employed with atypical contracts (that is typically characterized by instability and minor protections). Conversely, the legislative changes have been accompanied by a slight reduction in the geographical differences, particularly by increasing the employment probability under ACs in the South-Islands, which possibly indicates that, although ACs are valid instruments for avoiding non-employment, they may also represent a substitute for PCs.[27]

Seven dummy variables control for the role of economic sector specialization regarding re-employment probabilities, with manufacturing continuing as the reference category. With respect to transitions toward a PC, we find reduced re-employment probabilities, particularly for the sectors of Intermediation and Business. With respect to NE-AC transitions, previous experience in economic sectors other than manufacturing and commerce increases the length of non-employment spells.

Finally, we find that the expected employment growth variable has a positive and significant effect on re-employment probabilities before the reforms. Notably, a relatively greater impact is found for NE-AC transitions than for NE-PC transitions. Nonetheless, the effect has decreased since the introduction of the new legislation, and it is again greater (in magnitude) for ACs than PCs (except in the fourth sub-period). The reduced effect for exits from both types of contracts suggests that labor demand has progressively reduced the contents of its work intensity over the sample period.

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[26] In addition, we also replicate our estimates for young individuals (from 16 to 24 years of age) (see Figure 3 above). Nonetheless, we did not find any remarkable differences between the effects of the reforms on this age group and the overall sample (from 16 to 35 years). We therefore decided not to report these results. The estimates by gender, geographical area and for the age range 16-24 are available upon request.

[27] Figures A1 and A2 show that the gender difference increased after the reforms. The probability rate for women for a NE-PC transition fell sharply after the reforms, particularly for the STU category. The probability rate for STU decreased from 0.038 before the reform (dashed grey line) to 0.006 after the reform (dashed black line). Conversely, men had a greater probability of leaving NE for PC after the reforms (black line). NE-AC transitions do not show remarkable gender disparities either before or after the reforms. In terms of geographical differences, Figure A3 shows that the probability rates of NE-PC transition for the most relevant geographical areas (North, Center and South) were almost equally affected by the reforms. Indeed the gaps between the geographical probability rates remained almost stable after the introduction of the labor market reforms. Conversely, we find an increased probability of employment through AC (Figure A4) in all the areas after the introduction of the regulations.

5.1 Discussion

A major finding of this research is that the period under investigation has been characterized by a strengthening of the negative duration dependence, which means an increased gap in re-employment probabilities between STNE and LTNE. This result might be explained in various non-mutually exclusive ways. We now speculate on the possible explanations for our findings.

We suggest two main interpretations that are both based on the prediction of the employer screening model, according to which firms interpret unemployment duration as a 'private' signal of lower worker productivity (e.g., Lockwood 1991 and Omori 1997 for an application to non-employment). In this context, the exit rate from unemployment should be higher for the short-term unemployed than for the long-term unemployed, resulting in a negative duration dependence.[28]

Our explanations adapt the prediction of the employer screening model to a dynamic context in which job opportunities may increase because of flexibility policies or because of positive labor market conditions (e.g., economic cycle). In both cases, the increased job opportunities would improve the effectiveness of the signal that firms draw from non-employment duration, which finally results in an increased gap in re-employment probabilities between STNE and LTNE.

The first explanation relies on the fact that firms should be inclined to hire more STNE and less LTNE in a flexible labor market than in a rigid market for labor. It is indeed plausible that non-employment duration would be scarcely indicative of lower worker productivity in a rigid labor market because experiencing longer non-employment spells might be a consequence of fewer job opportunities instead of suspected lower productivity *per se*. Conversely, in a more flexible labor market that is characterized by more job opportunities, being a long-term non-employed might indeed be interpreted as an effective negative signal of worker productivity. It follows that firms operating in a more flexible labor market should be more prone to hire STNE with respect to LTNE because of the greater reliability of non-employment duration as a screening device. According to this view, flexibility policies should progressively contribute to increase the re-employment probability gap between STNE and LTNE as labor market flexibility increases.

In addition, this explanation is perhaps further reinforced by the finding of a smaller positive effect for previous work experience on re-employment probabilities, after the introduction of flexibility policies in the Italian labor market. This finding might be a consequence of the more effective role of non-employment duration as a screening device, which has partially replaced the signaling role of previous work experience in firms' hiring policies.

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[28] Negative duration dependence associated with unemployment spells is also consistent with the theoretical predictions derived from human capital models (Acemoglu, 1995), ranking models (e.g. Blanchard and Diamond, 1994) and search behavior models (e.g. Coles and Smith, 1998).

The second explanation relies on the fact that the effectiveness of the signal has changed as labor market conditions have changed (Kroft, Lange and Notowidigdo, 2013). According to this view, the difference in re-employment probabilities between STNE and LTNE should decrease in slack labor markets because there are fewer job opportunities; thus, the duration of unemployment would be less indicative of the unobservable characteristics of the workers. This explanation holds in a long-run perspective for the Italian labor market over the 1985-2004 period because this period has been characterized by a general increase of job opportunities and strengthening of negative duration dependence.

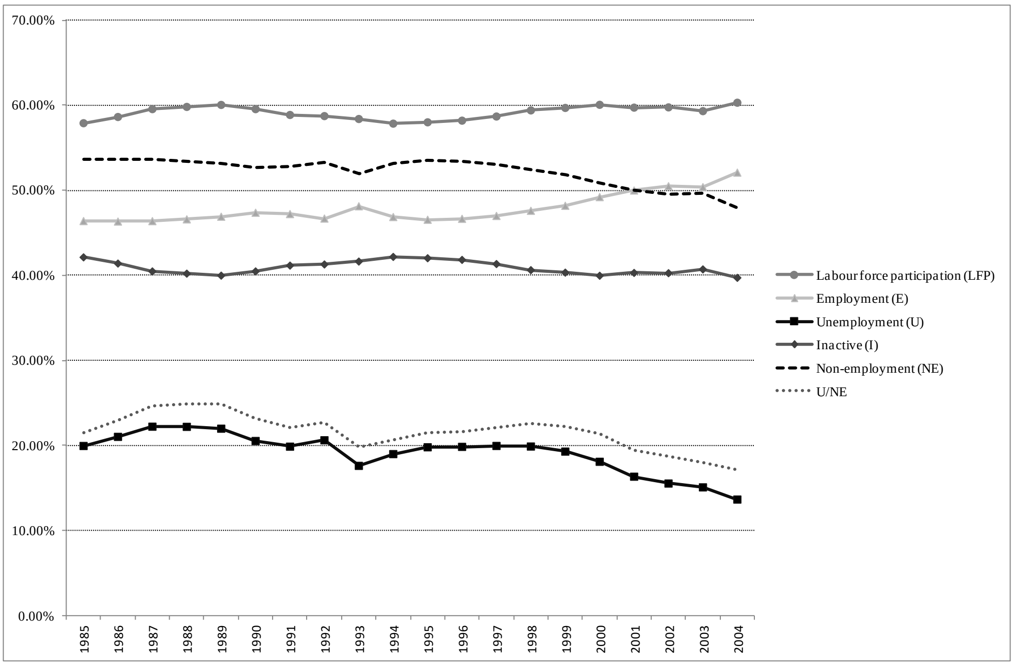
Finally, alternative explanations include changes in the Italian economy in terms of workforce participation and composition during the sample period (Figure 5). At the beginning of the 1990s (and particularly in 1992 and 1993),[29] the Italian economy sank into recession and workforce participation and composition rates changed: there was reduced unemployment among young people and a simultaneous increase in inactivity. These developments might be partly the result of the reforms described above. With respect to the former, the incidence of unemployment as part of non-employment (unemployment and inactivity) of young people in the 15-35 age range decreased from approximately 24.4% in 1985 to approximately 21.4% in 1997 (the year of the Treu Package) and was approximately four percentage points lower in 2004 (approximately 17.9%). As a consequence, the incidence of inactivity as part of non-employment increased among young people, rising from approximately 76.6% in 1985 to 78.6% in 1997 and approximately 82% in 2004.[30] These changes in the composition of non-employment and the previously discussed labor market reforms might have affected young workers' probability of leaving non-employment. In addition, the impact of the reforms might differ for the short- and long-term non-employed (depending on non-employment duration).

However, these speculations must be qualified because other forces might be at work. Thus, even if our findings about an increased gap in re-employment probabilities between STNE and LTNE are correct, we cannot exclude that other factors might explain this trend.

[29] Nonetheless, in Italy average GDP annual growth rate was approximately 1.9% during the two decades examined (Eurostat). Economic growth was indeed positive throughout the period, with the partial exceptions of 1993 (as a result of the early 1990s recession that witnessed a negative GDP growth rate of approximately -0.83%) and 2003 (GDP growth rate slightly negative at approximately -0.05%).

[30] These figures are available on the Internet at http://stats.oecd.org.

Figure 5. Labor force composition 1985-2004



Source: OECD statistics 2013.

6. Conclusions

This paper investigates the overall and step-by-step impact of Italian labor market reforms 'at the margin' on the probability of exiting from non-employment to permanent and atypical contracts. We estimate discrete time hazard models with competing risks and unobserved heterogeneity for a sub-sample of young workers in the 16-35 age range. The empirical analysis is based on 1985-2004 WHIP data.

From the mid-1980s to the mid-2000s the Italian labor market went through important reforms, which essentially made it easier to create flexible, atypical, and part-time job contracts. Our main finding is that the legislative changes of the period were associated with a significant and gradual increase in the gap in re-employment probabilities between the short- and long-term non-employed, particularly with respect to exits to permanent contracts, which provides evidence of negative duration dependence. We also find an increase in gender-related differences in the transitions to permanent contracts and a decrease in geographical differences. We offer some interpretations of our findings, which do not complete the picture of possible causes because other forces might be at work in causing the observed changes in probability during the period under investigation.

Our explanations run in four directions. First, long-term non-employment spells might be interpreted by firms as a negative signal of the unobserved productivity of workers. Second, there might be a link between the quality of the signal and labor market conditions, i.e., the signal might be stronger in a labor market with more job opportunities. Third, the flexibility process should result in an increase in job opportunities and therefore in a greater effectiveness of the signal. This result explains the gradual increase in the gap in re-employment opportunities for the short- and long-term non-employed (to the significant disadvantage of the latter) through the flexibility process characterizing the sample period. Fourth, changes in the composition of the labor force (particularly in terms of unemployed/inactive ratios and/or demographic issues) during the 20-year sample period might also help explain our findings.

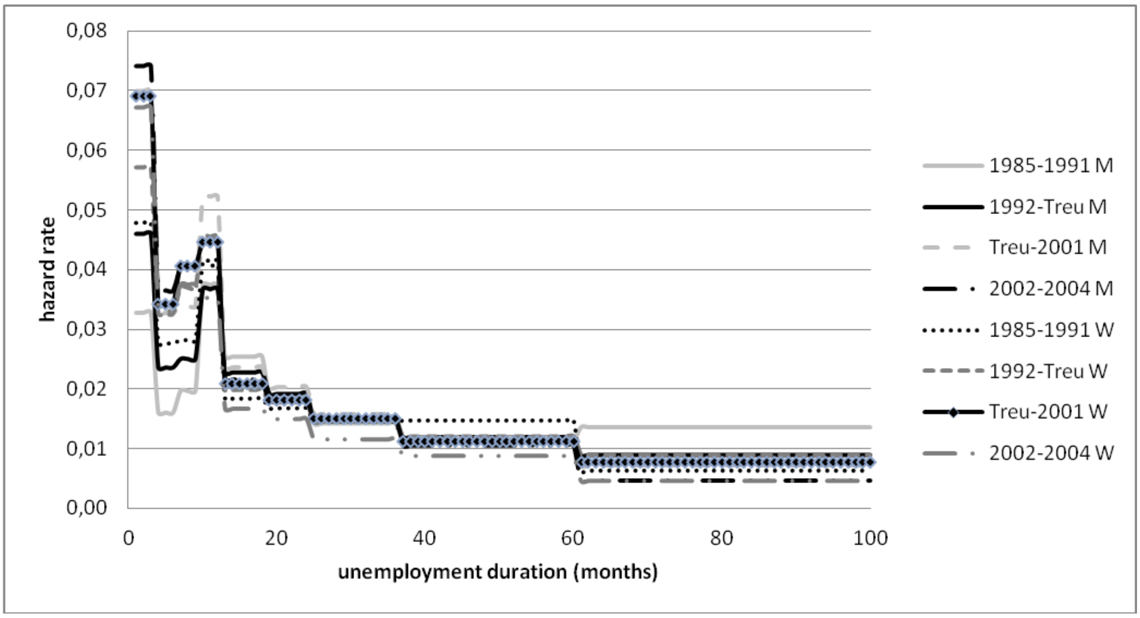
Appendix

Table A1. Number of individuals (starting of the first spell)

year	Freq	Percent	Cum
1985	262	0.59	0.59
1986	467	1.06	1.65
1987	864	1.96	3.61
1988	1,244	2.82	6.43
1989	1,577	3.57	10.01
1990	1,921	4.35	14.36
1991	2,017	4.57	18.93
1992	2,143	4.86	23.79
1993	1,854	4.2	27.99
1994	1,934	4.38	32.38
1995	2,268	5.14	37.52
1996	2,445	5.54	43.06
1997	2,609	5.91	48.98
1998	2,607	5.91	54.89
1999	3,044	6.9	61.79
2000	3,393	7.69	69.48
2001	3,450	7.82	77.3
2002	3,289	7.46	84.75
2003	3,715	8.42	93.17
2004	3,011	6.83	100
Total	44,114	100	

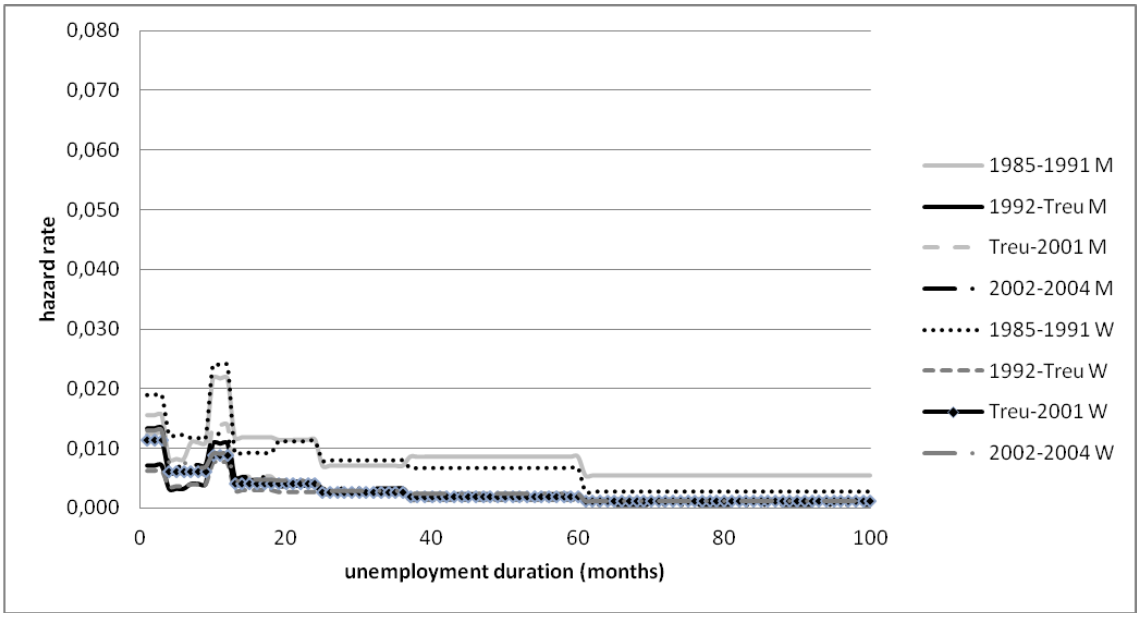
Source: our elaboration of WHIP data

Figure A1. Piecewise constant baseline hazard rate by Gender and Time Period for PCs. 1985-2004



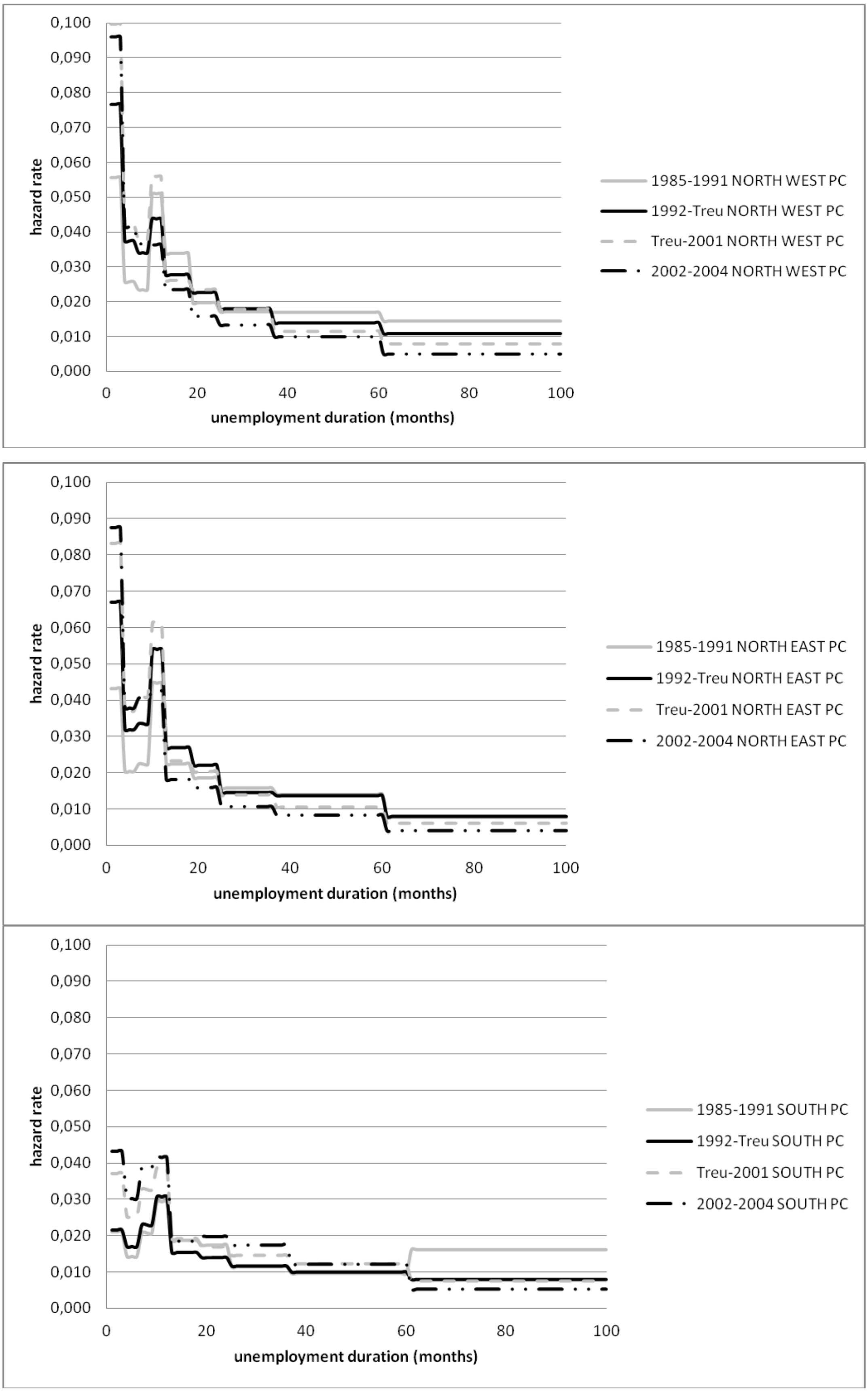
Source: our elaboration of WHIP data

Figure A2. Piecewise constant baseline hazard rate by Gender and Time Period for ACs, 1985-2004



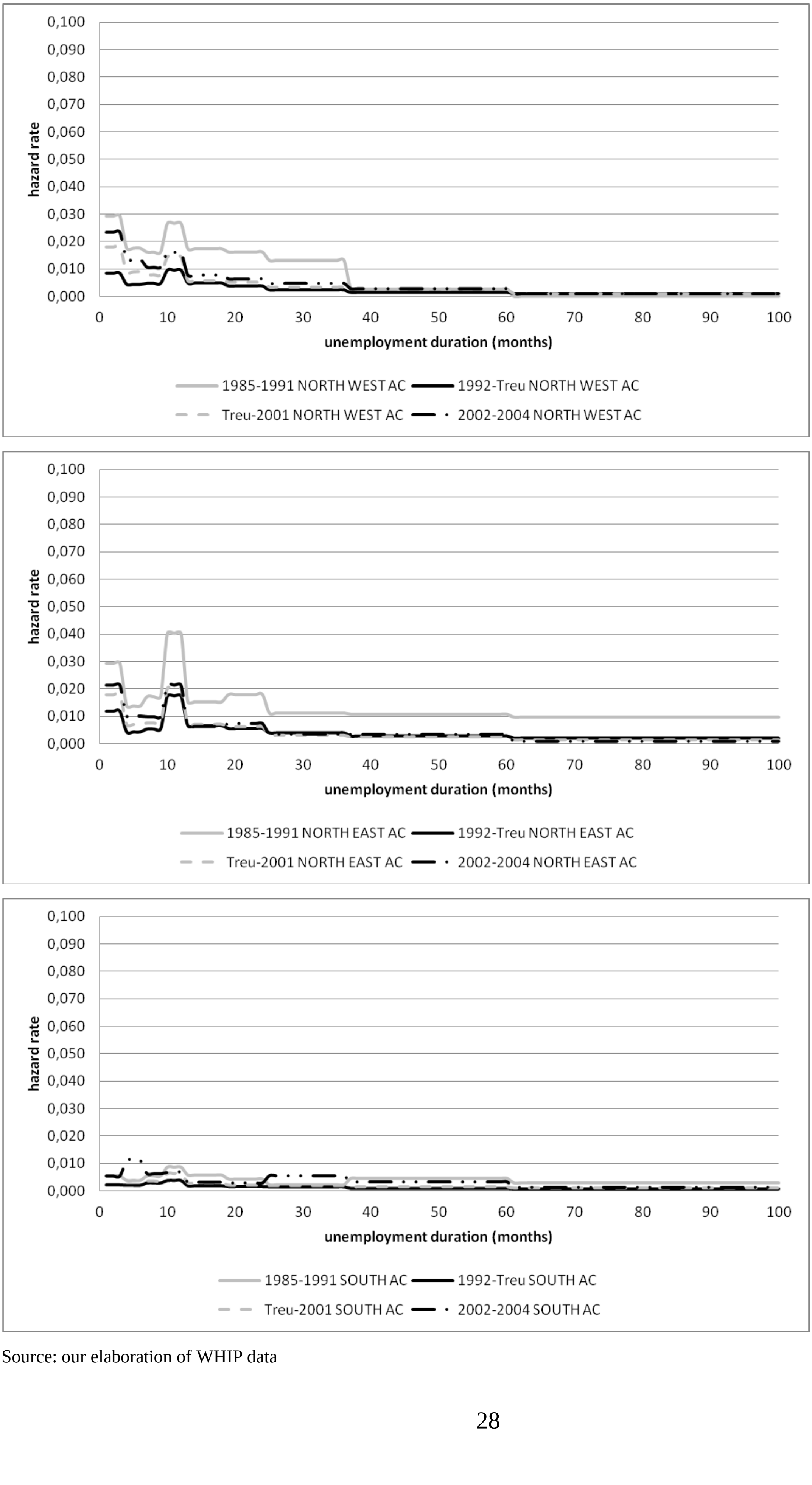
Source: our elaboration of WHIP data

Figure A3. Piecewise constant baseline hazard rate to PCs by Geographical Area and Time Period, 1985-2004



Source: our elaboration of WHIP data

Figure A4. Piecewise constant baseline hazard rate to ACs by Geographical Area and Time Period, 1985-2004



Source: our elaboration of WHIP data

References

Acemoglu D. (1995) 'Public policy in a model of long-term unemployment', *Economica*, 62 : 161-178.

Amuedo-Dorantes C. (2000) 'Work transitions into and out of involuntary temporary employment in a segmented market: evidence from Spain', *Industrial and Labor Relations Review*, 53(2): 309-325.

Bentolila S. and Bertola G. (1990) 'Firing Costs and Labour Demand: How bad is Eurosclerosis', *The Review of Economic Studies,* 57(3): 381-402.

Berton F., Devicienti F., Pacelli L. (2011), 'Are temporary jobs a port of entry into permament employment?: Evidence from matched employer-employee'. *International Journal of Manpower,* 32(8): 879-899.

Blanchard O.J. and Diamond P. (1994) 'Ranking, unemployment duration, and wages', *Review of Economic Studies* , 61: 417-434.

Blanchard O.J. and Landier A., 2002. 'The Perverse Effects of Partial Labour Market Reform: fixed--Term Contracts in France', *The Economic Journal*, 112(480): F214-F244.

Boeri T. and Garibaldi, P. (2007) 'Two tier reforms of employment protection: a honeymoon effect', *The Economic Journal*, 117: F357-F385.

Booth A.L., Francesconi M. and Frank J. (2002) 'Temporary jobs: stepping stones or dead ends?', *The Economic Journal*, 112: F189-F213.

Cahuc P. and Postel-Vinay F. (2002) 'Temporary jobs, employment protection and labor market performance', *Labour Economics,* 9(1): 63-91.

Cappellari L., Dell'Aringa C. and Leonardi M. (2012) 'Temporary Employment, Job Flows and Productivity: A Tale of Two Reforms', *The Economic Journal,* 122(562): F188-F215, 08.

Coles M.G., Smith E. (1998) 'Marketplaces and matching', *International Economic Review*, 39(1): 239-254.

Contini B. and Grand E. (2010) ‘Disposable Workforce in Italy’, Institute for the Study of Labor (IZA) Discussion Paper no. 4724, Bonn, Germany.

Contini B. and Poggi A. (2012) ‘Employability of Young Italian Men after a Jobless Period, 1989–98’, LABOUR, 26(1): 66-89.

D’Addio A.C. and Rosholm M. (2005) ‘Exits from temporary jobs in Europe: a competing risks analysis’, *Labour Economics*, 12(4): 449-468.

Destefanis S. and Fonseca R. (2007) ‘Matching efficiency and labor market reform in Italy: A Macroeconomic Assessment’, *LABOUR*, 21(1): 57-84.

Dickson M., Postel-Vinay F. and Turon H. (2012) ‘The lifetime earnings premium in the public sector: the view from Europe”, University of Bristol, unpublished manuscript.

Gagliarducci S. (2005) ‘The dynamics of repeated temporary jobs’, *Labour Economics*, 12(4): 429-448.

Grand E. and Quaranta M. (2011) ‘Completamento delle carriere lavorative WHIP con i dati del Casellario degli Attivi INPS’, WHIP Technical Report no. 3/2011, Labor Laboratorio Riccardo Revelli, Centre for Employment Studies.

Güell M. and Petrongolo B (2007) ‘How binding are legal limits? Transitions from temporary to permanent work in Spain’, *Labour Economics*, 14(2): 153-183.

Ichino A., Mealli F. and Nannicini T. (2005) ‘Temporary work agencies in Italy: A springboard to permanent employment?’, *Giornale degli Economisti e Annali di Economia*, 64(1): 1-27.

Jenkins S. (2005) *Survival analysis*. Unpublished manuscript, University of Essex.

Jiménez.Rodrìguez R. and Russo G. (2012), ‘Aggregate Employment Dynamics and (Partial) Labour Market Reforms’, *Bulletin of Economic Research*, 64(3): 430-448.

Kroft K., Lange F., Notowidgdo M.J. (2013) ‘Duration dependence and labor market conditions: Evidence from a field experiment’, *Quarterly Journal of Economics*, 128(3): 1123-1167.

Kugler, A. and G. Pica (2008) ‘Effects of Employment Protection on Worker and Job Flows: Evidence from the 1990 Italian Reform’, *Labour Economics*, 15 (1): 78-95.

LABORatorio Revelli (2009) WHIP Data House. www.laboratoriorevelli.it/whip/.

Lazear, E.P., (1990) ‘Job Security Provisions and Employment,’ *Quarterly Journal of Economics*, 105 (3): 699-726.

Leonardi, M. and Pica G. (2013) ‘Who pays for it? The heterogeneous wage effects of Employment Protection Legislation’, *Economic Journal*, 23: 1236-1278.

Lockwood, B. (1991). ‘Information Externalities in the Labor Market and the Duration of Unemployment’, *Review of Economic Studies*, 58(4), 733-753.

Miller R.B. and Hollist C.S. (2007) ‘Attrition Bias’, WHIP Technical Report no. 3/2011, Encyclopedia of Measurement and Statistics, Ed. Neil Salkind, Vol. 1: 57-60.

Montanino A. and Sestito P. (2003) ‘Le molte funzioni del lavoro interinale in Italia: da strumento di flessibilità a contratto di prova’, *Rivista di Politica Economica*, 93(3-4): 115-148.

Muehlberger U. and Pasqua S. (2009) ‘Workers on the border between employment and self-employment”, *Review of Social Economy*, 67(2): 201-228.

Narendranathan W. and Stewart M. (1993) ‘Modelling the probability of leaving unemployment: competing risks models with flexible baseline hazards’, *Applied Statistics*, 42, 63-83

Nickell et al. (2005), ‘Unemployment in the OECD Since the 1960s. What Do We Know?’, *The Economic Journal*, 115(500): 1-27.

Nicoletti C. and Rondinelli C. (2010) ‘The (mis)specification of discrete time duration models with unobserved heterogeneity: A Monte Carlo study’, *Journal of Econometrics*, 159(1): 1-13.

OECD (2004), ‘Employment Protection Regulation and Labour Market Performance’, in OECD Employment Outlook 20o4, OECD Publishing, http://www.oecd.org/els/emp/34846856.pdf.

OECD (2013), 'Protecting jobs, enhancing flexibility: A new look at employment protection legislation', in OECD Employment Outlook 2013, OECD Publishing, http://dx.doi.org/10.1787/empl_outlook-2013-6-en.

Omori, Y. (1997), 'Stigma Effects of Nonemployment', *Economic Inquiry*, 35(2), 394-416.

Picchio M. (2008) 'Temporary Contracts and Transitions to stable Jobs in Italy', Labour, 22 s1(06), Special Issue:147-174.

Prentice R.L. and Gloecker L.A. (1978), 'Regression Analysis of Grouped Survival Data with Application to Breast Cancer Data', *Biometrics*, 34 (1): 57-67.

Saint-Paul G. (1997), 'The Rise and Persistence of Rigidities', *The American Economic Review*, 87(2): 290-294.

van den Berg G.J., Holm A. and van Ours J.C. (2002) 'Do stepping-stone jobs exist? Early career paths in the medical profession', *Journal of Population Economics*, 15: 647-665.