# Labour market transition in Italy: an empirical investigation

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## Abstract

In this paper I study the effects of the introduction of Treu package on the hazard rate of moving from a temporary job to a stable work in Italy. In order to enhance the flexibility of labour market and to stress the capabilities to create job, a progressive liberalization of flexible contracts has been adopted in Italy during the ‘90s, widening their scope and easing their adoption. The empirical exercise is carried out using an administrative dataset, Clap, extracted by the social security archives for the years 1992-2002. I apply a parametric and semi-parametric continuous time duration model to investigate whether the recent reform has changed the pattern of duration dependence in temporary employment duration. The main findings emphasize how the Treu package has substantially reduced the positive duration dependence in the hazard rate of moving from temporary to permanent employment. This result is robust to different specifications adopted for the baseline hazard. However, an even stronger conclusion concerns the temporary work agency (TWA) employment which exhibits a negative duration dependence in the post-reform period, probably indicating a negative relation between contract duration and the probability to moving into permanent employment for this contractual type. Finally I find large differences in terms of individual and job-related characteristics.

Keyword: temporary employment, mixed proportional hazard model, duration dependence

## 1. Introduction

Since the mid-1980s several European countries enhanced the flexibility of their labour market in order to cope with the high and persistent unemployment. During the last years, reforms aimed at increasing flexibility have been introduced in the Italian labour market. Generally the strategy used is consisted in implementation of the reform "at the margin", allowing for temporary employment without changing EPL for permanent jobs. Standard insider-outsider framework may provide a first explanation for the prevalence of this such an approach in which the traditional standard work has been exempted from labour market liberalization. As consequence, temporary jobs have been widely used and accounted for a growing share of the employed workforce in countries with rigid employment protection.

Temporary employment may potentially provide a path of entry into the stable jobs, but it is often associated with precarious labour condition and thus likely with lower job quality. There is an extensive debate on the extent to which such jobs help unemployed workers, providing them opportunities to gain work experience and acquire human capital, to deepen the attachment to the labour market.

Nevertheless the risk of experiencing precarious career paths can therefore be higher for people in such flexible contractual arrangements. There is an extensive literature that analyzes whether the existence of temporary contracts exacerbates the inherent precariousness in the labour market. Several studies have investigated the impact of fixed-term works on specific groups and the role of the unobserved heterogeneity and state dependence, to examine labour market transitions over time (Tattara and Valentini, 2006).

Indeed temporary work experience may be informative about the ability of the individuals. Some empirical studies evidence that the employers use flexible contracts as a way of screening for the stable works and to cope with demand fluctuations, in this way lowering adjustment costs in a traditional labour market characterized by a wide use of permanent contracts with costly firing regulations. Thus flexibility may be a choice on the demand's side, but it can worsen life's conditions for those workers passing through it, enhancing the level of precariousness in the labour market and raising concern among policy-makers (see Booth et al, 2002; D'Addio and Rosholm, 2005).

Although numerous attempts to investigate the consequence of fixed-term contracts, some consensus has formed among economist that the flexibility does not necessarily increase the overall employment level and produces segmentation in the labour market, creating a strong dualism between insiders and outsiders. Several theoretical investigations (Dolado et al., 2002;

Blanchard and Landier, 2002; Cahuc and Postel-Vinay, 2002) have been carried out in order to analyze the effects of flexible contracts and to isolate the mechanism through which temporary employment modify labour market equilibrium. Despite the large diffusion of two-tier labour marker reforms in Europe, macroeconomic studies in the context of matching models of the labour market have argued that the combination of deregulated temporary jobs and positive firing costs on stable employment may have a perverse impact in terms of aggregate unemployment and welfare. Boeri and Garibaldi (2007) investigate the transitional dynamics of such two tier reform strategies and they argued the presence of a link between the recent growth-less job creation and the partial labour market reforms introduced in the European labour market in the '90s.

Thus, more recently, the debate has focused upon the negative implications of a segmented labour market in which atypical workers operate in low-wage jobs and firms may be less inclined to invest in training, reducing the quality of the match between worker and firm (Güell and Petrongolo, 2007)

However the crucial question is the extent to which temporary jobs help individual workers to move from fixed-term to permanent employment, that is, the extent to which flexible contract acts as a stepping stone to stable job. Another important point concerns the pattern of promotion into indefinite duration jobs for temporary workers but mixed employment effects do not a provide clear signal that the atypical contracts are associated with good performances in terms of subsequent careers and wages.

This study aims at providing a description of the transition patterns of temporary workers towards a stable job in the Italian labour market and an identification of the duration dependence associated with less favourable working histories. In order to disentangle the duality of labour market I focus on exit margin, i.e. the transition after a temporary experience. In doing this, I estimate a duration model of temporary employment.

It has been often argued that the entry in temporary employment represents a first stepping stone into stable employment, increasing the probability of moving into permanent jobs with respect to the unemployed workers. As the extensive literature for the Spanish labour market has shown, the entry margin evidences little variation, while the main source of duality lies in the exit margin (Güell and Petrongolo, 2007).

The typical framework in the study of labour market transition is the job-search model where the duration of each working spell is specified as the conditional probability of leaving that spell, i.e. the hazard rate (Cahuc and Zylberg, 2004; D'Addio and Rosholm, 2005).

The empirical analysis in this study is based therefore on Clap (*Campione Longitudinale dei Lavoratori Attivi e dei Pensionati*) dataset, an administrative archive from the Social Security records (INPS), provided by Labour Minister. It consists of individual working histories and reports the sequence of daily contiguous spells. To evaluate whether the Treu package has modified the conversion pattern, I select a sub-sample that spans for the years 1993-2002. The modelling framework avoids the left-censoring issue in labour market histories. Then to assess the duration dependence in the transitions from temporary to permanent employment, I adopt a single risk mixed proportional hazard (PH) model, employing both parametric and semi-parametric baseline hazard. This study complements previous literature investigating whether the partial labour market reforms introduced in the Italian labour market during the '90s have modified the duration dependence pattern in the transitions from temporary to permanent employment. Moreover, I provide further evidence on the timing of conversion for different atypical contracts, such as fixed-term work and temporary work agency (TWA) employment .

The main empirical findings show that the introduction of Treu reform has substantially reduced the positive duration dependence in the probability of moving from a temporary work to a stable job. However, the results on positive duration dependence suggests the persistence of a human capital accumulation effect also in the post-reform period. Differently, the TWA employment reports a negative duration dependence in the post-reform period, which indicates a negative relation between contract duration and the hazard rate of moving into permanent employment, emphasizing the concerns about the inherent precariousness of this specific contractual form. Finally individual heterogeneity in terms of personal and job-related characteristics explains a large amount of differences in labour market performance.

The plan for the paper is as follows. Section 2 presents a literature review. Section 3 describes the dataset used and presents some preliminary descriptive evidence. Section 4 introduces the PH econometric model performed to investigate the impact of the Treu reform on the duration pattern in the transition from temporary to permanent employment. Section 5 reports the empirical findings and the sensitivity analysis to check the robustness of the results and section 6 draws the conclusion

## 2. Literature review on temporary employment

Previous research has extensively focused on several aspects of temporary employment to emphasize the impact of flexible contracts on labour market in OECD countries. In addition, an issue much debated has been at what extent the persistence of European unemployment has been exacerbated by the labour market institutions and in particular the discussion has stressed the role of employment protection legislation (EPL). Therefore, in order to reduce the high level of unemployment, several European countries have gradually enhanced the flexibility in the labour market, allowing the use of temporary contracts with low firing costs.

Thus a significant feature of most reforms is that they are reforms “at the margin”, which simply add further distortions without changing the strict dismissal costs for permanent contracts. The insiders-outsider models may provide an explanation for the prevalence of such an approach which could have undesirable consequence in terms of labour market segmentation.

However the introduction of fixed-term contracts has generally favoured a reduction of unemployment rates and raised the dynamicity of the European labour market. In addition, it has been often argued that employment spells in temporary jobs can increase workers’ human capital and provide the opportunities to acquire work experience and labour market contacts which might favour the subsequent transition to a stable job. Thus, whether a flexible contract can improve the labour market outcomes of the unemployed, representing a stepping stone towards a permanent job, has recently become the crucial question for the empirical studies.

In this contest Sciulli (2006) studies the impact of a partial labour market reform in Italy (the so-called Treu package) aimed at introducing a two-tier system and allowing temporary employment. Applying a proportional hazard model with competing risks, he compares the transition rate between work spells started before and after the exact day of introduction of the reform, exploiting a sample of young worker (aged 16 to 32). In the post-reform period, he finds an increase in negative duration dependence for non working state outflow, i.e. a worsening of the duality between the short-term and long-term unemployed. Moreover, having previous atypical job experiences reduces the probability of finding a permanent work for the unemployed, while there is no evidence that the probability of moving to a permanent job is higher for atypical workers than for the unemployed. However Sciulli (2006) finds a positive duration dependence for the transitions from temporary to permanent employment, suggesting the presence of a human capital accumulation effect. Finally the individuals’ heterogeneity in terms of geographic, gender and occupation variables accounts for a large amount of differences in labour market transitions.

Picchio (2007) tries to assess whether a temporary job experience represents a springboard towards a stable job or a trap of precariousness, comparing the probability of finding a permanent work for the temporary and unemployed workers. In order to provide the true stepping stone effect, he estimates a bivariate random effect probit model. The main finding suggest that a temporary job, rather than unemployment, increases the 2-year after probability of moving to a permanent work by 13-16 percentage points. Corsini and Guerrazzi (2007) provide a probabilistic evaluation of the transition from temporary to permanent employment in a regional contest, applying a multinomial nested logit estimation strategy, in order to account for the decisions made by the workers and those made by the employers. The model is based on two steps. In the first, the worker decides whether or not to participate in the labour market and whether as employee or self-employed. In the second, the firm decides whether or not to hire the worker and what type of contract offers. Their results show a stabilization probability equal to 61% but display significant heterogeneity across workers' characteristics. For instance, atypical contracts may represent a precariousness trap for women and older workers.

Berton et al. (2007) study the labour market transitions of young entrants and consider seven labour market states: permanent and temporary employment, apprenticeship, training programmes, self-employment, quasi subordinate jobs and unemployment, applying a dynamic multinomial logit model and controlling for individual fixed effects. The empirical analysis provides some interesting results. First of all, individual heterogeneity explains part of the observed transition rates and whatever the entry state, keeping the same contract is the most likely destination. This indicates the presence of true state dependence. Finally 1-year after all atypical contracts display a significant probability of moving to open-end contract. In this sense, the stepping stone hypothesis holds, in particular for the CFL. Berton (2008) and Alboni et al. (2008) provide further evidence on the stepping-stone effect using the timing of the events approach.[1]

While in Italy the introduction of partial labour market reform is a relatively recent issue, two-tier strategies have been applied since the mid-80s in Spain to foster job creation and to develop a more dynamic labour market. Alba-Ramirez (1998) and Amuedo-Dorantes (2000) have found that the probability of finding a permanent work reduces after some previous fixed-term contracts, i.e. it is not atypical contract per se but its intermittent nature that is detrimental in terms of career prospects. Garcia-Perez and Munoz-Bullon (2003) examine to what extent Temporary Help Agency (THA) intermediation affects workers' transitions into and out of employment, focusing upon the Spanish youth labour market and exploiting a discrete-time duration model. The findings on the success of THAs as a recruitment method are two-fold. On the one hand, the results show that those intermediaries help the unemployed workers to gain a new job.[2] On the other hand, job tenure in THA works is lower than for the other temporary contracts. Moreover, the probability of exiting from employment is negatively affected by job

[1] Gagliarducci (2005) also apply duration model to investigate the impact of repeated flexible contracts on the probability of moving into a stable job. He finds that good jobs are converted into permanent works after the initial screening while longer atypical jobs display a decreasing probability of being converted over time. Moreover, experiencing repeated temporary works (and in particular interruptions) are detrimental in order to find a stable job.

[2] In particular for short-term unemployed individuals.

tenure and increases with the duration of fixed-term contract. Finally individuals' heterogeneity explains a large amount of the difference in the hazard rate of finding a job.[3]

Güell and Petrongolo (2007) estimate a duration model with competing risk and flexible duration dependence in order to study the duration pattern of temporary contracts and their conversion into permanent employment versus alternative states. The results show that the conversion rate from fixed-term to permanent contract is rather low and it concentrates at three years of duration, close to the deadline of the temporary contract. However, there is another spike in the conversion rate around one year, suggesting in this case that atypical contracts are also used as a screening device.

Empirical research for other European countries provides evidence that temporary employment may indeed represent a stepping stone into a stable job. For instance, Hagen (2002) finds that escaping unemployment through a temporary job experience enhances the future employment chances for the unemployed workers both for temporary and permanent job in the West Germany labour market. Similar results are found in Booth et al. (2002) for the UK. Finally D'Addio and Rosholm (2005) focus upon the determinants of the transitions out of temporary employment in order to identify factors that are likely to increase the risk of precariousness for atypical workers and to understand whether fixed-term works provide a stepping stone effect for a stable job. They use a discrete-time duration model with competing risk in a multinomial logit framework.

## 3. Dataset

The data used in this study are extracted from the Clap (*Campione Longitudinale degli Attivi e dei Pensionati*), an administrative archive which spans the years 1985-2002. The original dataset collects the Social Security records in order to describe the functioning of the Italian labour market and to investigate the recent evolution of flexible contracts. Moreover, it includes a representative dynamic population of about 500.000 individuals.

The selection rule is based on a 1/90 random sample of employees born on the 10th of March, June, September and December of each year in the sample. The archives provide information on each worker that from the 1985 has paid some contributions to INPS, distinguishing between

[3] In addition, similar results in terms of mobility patterns after the 1997 reforms are reported also in Kugler et al. (2002).

permanent or temporary contracts[4] and individuals who are unemployed or retired. Therefore the sample provides a detailed description of the stock of workers, unemployed and retired under INPS regime, excluding the main part of public employees and worker employed with the Continuative and Coordinate Collaboration (Co.Co.Co.) contracts that are linked to other social security regime.[5]

This dataset collects individual longitudinal observations using social security records, showing directly monthly information about the duration of each spell. A peculiar feature of these archives, created by INPS to compute the retirement benefits, is that it follows individuals who move both into self-employment, the agricultural sector, as well as unemployment and retirement. However it is not possible to indirectly determine the non-working spells[6] because the transitions towards the public sector or the Co.co.co. are excluded, so I cannot distinguish between unemployment spell or exit from the sample.

I have information about worker's demographic characteristics, as age and gender, the type of occupation (white-collar, blue-collar, apprenticeship and other limited figures), wage, the initial and final month of job matches and crucial for this paper the type of contract (i.e. temporary or permanent) and the job duration (i.e. the tenure in months from the start of the contract). Moreover the dataset includes a firm identifier, that for dependent employment, makes possible to observe the change of firm for the on-the job transitions. However I do not assess whether they are quits or layoffs.

In this sense CLAP has been collected to evaluate labour policies, classifying each individual according to a specific working status and then identifying the entry and the exit from each category under observation. Therefore the administrative archives represent an important resource to study individual work history and to assess the impact of the Treu reform in the worker flows (Contini and Trivellato, 2005; Sciulli, 2006).

### 3.1 Sample selection and descriptive statistics

To investigate whether Treu package has modified the duration dependence related to the out-flows from temporary employment and whether the "flexibility at the margin" has increased the transitions towards a permanent work, I select a sub-sample using the information from 1992

[4] The information is inferred by the contribution paid to INPS. From the type of contributions it is possible to identify individuals employed with a fixed-term contract, a CFLs (work training contracts) and interim work (or temporary work agency (TWA) employment after the 1997 reform).

[5] Individuals employed with a *Co.Co.Co.* contract are classified in the Italian legislation as self-employed and so they are not included in the sample of employees in the INPS archives.

[6] As the difference between the end of the previous contract and the start of the subsequent one

and 2002 for individuals aged 16 to 55, who enters in the sample the first time after the January 1st 1992 and having at least one job spell in the years under observation. This makes possible to reconstruct the complete work history from the entry in the sample, excluding all individuals whose recorded spells began before 1992.

Constructing my sub-sample, if an individual presents at the same time more than one job relationship, I eliminate the shorter work relationship characterized by fewer day of effective work. Furthermore, when the second work starts before of the end of the first, but ends after it, I decide to censor at the left the second job spell, so assuming that the second job starts only when the first ends. I adopt this strategy in order to better reconstruct the individual working history and the passage from double work to a single one is considered as a job to job transition. Finally, I assume that this procedure and the elimination of multiple contemporary spells leave the sub-sample to be representative.

**Table 1: Average survival time in temporary employment by individual and job-related characteristics: a comparison of pre and post Treu refom period**

| | **Pre-reform** | **Post-reform** |
|---|---|---|
| Female | Median = 1276<br>n.obs = 7371 | Median = 1156<br>n.obs = 13626 |
| Male | Median = 1126<br>n.obs = 8387 | Median = 1095<br>n.obs = 14689 |
| North-west | Median = 944<br>n.obs = 3798 | Median = 854<br>n.obs = 7455 |
| North-east | Median = 1094<br>n.obs = 3391 | Median = 1187<br>n.obs = 6057 |
| Center | Median = 1090<br>n.obs = 2938 | Median = 1126<br>n.obs = 5789 |
| South | Median = 1507<br>n.obs = 4289 | Median = 1491<br>n.obs = 7661 |
| Blue-collar | Median = 760<br>n.obs = 2256 | Median = 729<br>n.obs = 8664 |
| White-collar | Median = 735<br>n.obs = 2010 | Median = 727<br>n.obs = 5680 |
| Training job | Median = 1094<br>n.obs = 6142 | Median = 1265<br>n.obs = 10217 |
| Age16-24 | Median = 1094<br>n.obs = 11077 | Median = 1217<br>n.obs = 18077 |
| Age25-33 | Median = 1415<br>n.obs = 3672 | Median = 758<br>n.obs = 8522 |
| Foreigners | Median = 1867<br>n.obs = 3189 | Median = 1460<br>n.obs = 7384 |

Source: elaboration on Clap data

As suggested in several empirical studies, I exclude from the sample all workers with one or more episodes of self-employment since their decision are driven also by more complex motivation. Then the selection on age is implemented to follow the standard statistical classification of the workforce and to avoid the inclusion of the workers that have exceeded the limit for retirement but keep themselves in the labour market. In particular the exclusion of older individuals (55-65 aged) is justified by the fact that the employment volatility close to retirement may modify the real impact of higher flexibility on duration dependence in the transition towards a stable job.

The sub-sample spans for the years 1992-2002 and it records 198,756 spells for 44,089 individuals (66,275 permanent contract, 126,986 temporary and 5,495 unemployed). Moreover, to perform the comparison between the pre-reform period and the post-reform one, I divide my sub-sample in two groups. In the first I put all the spells started and ended before the Treu law (June 24$^{th}$ 1997), while in the second I consider the work relationships started after the approval of the package. Further I define as right censored spells all the episodes starting before the reform and ending after it. This strategy leaves me 48,517 spells for 15,763 individuals (11,714 permanent contract, 35,704 temporary one and 1,099 unemployed) in the pre-reform period and 85,437 spells for 28,326 individuals (18,602 permanent contract, 65,382 temporary contract and 1,453 unemployed) in the post-reform one.

**Table 2: Sample characteristics of temporary workers**

| | **Complete sample** | **Pre-reform** | **Post-reform** |
|---|---|---|---|
| Female | 0.46 (0.49) | 0.48 (0.49) | 0.46 (0.49) |
| Age | 25.18 (6.66) | 23.50 (6.71) | 24.70 (6.57) |
| Temporary | 0.638 (0.48) | 0.73 (0.44) | 0.76 (0.42) |
| Unemployed | 0.027 (0.16) | 0.02 (0.14) | 0.02 (0.12) |
| Permanent | 0.333 (0.47) | 0.24 (0.42) | 0.22 (0.41) |
| Average daily tenure | 228.01 (130.85) | 242.48 (127.44) | 200.60 (130.23) |
| North-west | 0.30 (0.45) | 0.29 (0.44) | 0.30 (0.45) |
| North-east | 0.24 (0.43) | 0.24 (0.42) | 0.23 (0.42) |
| Center | 0.19 (0.39) | 0.19 (0.38) | 0.20 (0.40) |
| South | 0.25 (0.42) | 0.26 (0.43) | 0.26 (0.42) |
| Economic growth | 1.63 (1.13) | 1.64 (1.38) | 1.56 (1.07) |
| Year93 | 0.03 (0.15) | 0.09 (0.29) | - |
| Year94 | 0.03 (0.17) | 0.13 (0.34) | - |
| Year95 | 0.04 (0.20) | 0.18 (0.39) | - |
| Year96 | 0.06 (0.23) | 0.24 (0.43) | - |
| Year97 | 0.06 (0.25) | 0.26 (0.44) | 0.03 (0.11) |
| Year98 | 0.10 (0.28) | - | 0.08 (0.24) |
| Year99 | 0.11 (0.32) | - | 0.12 (0.32) |
| Year00 | 0.15 (0.35) | - | 0.19 (0.39) |
| Year01 | 0.18 (0.38) | - | 0.25 (0.44) |
| Year02 | 0.20 (0.40) | - | 0.33 (0.47) |
| Immigrant | 0.18 (0.39) | 0.16 (1.38) | 0.21 (0.40) |
| White-collar | 0.21 (0.40) | 0.17 (0.37) | 0.22 (0.41) |
| Blue-collar | 0.34 (0.47) | 0.25 (0.43) | 0.38 (0.48) |

Source: elaboration on Clap data, standard deviation is reported in parenthesis.

Table 1 reports the average survival time in temporary employment by individual and job-related characteristics. The results show that after the reform there is a reduction in the survival time in temporary job, larger for women. Looking at the regional variation, the assessment is not completely clear with huge differences between the different macro-regions. Moreover, the foreign workers report a lower average survival time in temporary job during the post-reform period. For what concerns the job-related characteristics, the blue-collar displays a larger reduction than the white collar, but the lack of educational attainment makes less clear these results.

Differently, table 2 shows the sample characteristics, comparing the pre and post-reform period. An interesting feature concerns the foreigners: in the second half of the '90s I can observe an increasing share of foreign workers in the employment dynamics, with a growth of five percentage points between the pre and post reform period (16% against 21%).

## 4. Econometric specification

The standard approach to modelling individual transitions in the labour market is based on job-search framework and on the event history analysis which tries to explain whether certain individuals are at a higher risk of experiencing the event of interest than others.

The key concept in this class of model is the hazard rate, defined as the product of the probability of receiving a job offer in a given period and the probability that such an offer will be accepted (Cameron and Trivedi, 2005). The hazard rate can be interpreted as the reduced-form of a standard job-search framework and modelled through the conditional probability of leaving a state. In practice, it depends on individual observed characteristics and other variables likely to influence the cost of the search behaviour as well as on the duration an individual has already spent in the starting state.

In other words, the hazard model concerns an approach in which the risk of experiencing an event at a certain time point is predicted with a series of covariates. This type of model presents several special features with respect to other standard regression model. First of all, it deals with the presence of censored observation, i.e. when the event of interest is not yet occurred during the time of observation. Then, in the duration model the covariates may change their value during the observation period, so introducing a sort of dynamic analysis. Finally, duration models concentrate on distributional issue regarding both the survival time and the unobserved heterogeneity in the transition probabilities.

Defining the hazard function as the instantaneous rate of occurrence of the event, it assumes the following form:

$$\lambda(t) = \lim_{dt \to 0} \frac{\Pr\{t < T \leq t + dt | T > t\}}{dt}$$

where the numerator of this equation is the conditional probability that the event will occur in the interval $(t, t + dt)$, given that it's not occurred before. The conditional probability in the numerator may be written as the ratio of the joint probability that T is in the interval $(t, t + dt)$ and $T > t$ to the probability of the condition $T > t$.

The hazard rate may be indicate also as follows:

$$\lambda(t) = \frac{f(t)}{S(t)}$$

i.e. the ratio between the probability density function of an event at $t$ ($f(t)$) and the probability of surviving until time $t$ without experiencing the event, which is defined as $S(t) = 1 - F(t)$.

The focus of this study is to estimate the determinants of worker's transitions from temporary to permanent employment. Thus I select a sub-sample of individuals who enters with a fixed-term work, observing the transition process out of temporary employment towards a permanent job (Guell and Petrongolo, 2003). The administrative data used in this empirical studies provide monthly information about the spell duration and allow to investigate the individual working history, identifying the initial condition and the type of transition whether the spell is uncensored.

In practice, the objective is to describe a sequence of $t_i = \{t_i^n\}_{n \in (1,....,N)}$ of contiguous spells in a well-defined state, where $t_i$ is the elapsed duration for each episodes and $n$ represents the *n*th spell of individual *i*. I apply a single risk mixed proportional hazard model (MPH) which allow for the presence of unobserved heterogeneity. As suggested in the literature (Lancaster, 1979), the proportional hazard model not controlling for the presence of unobserved heterogeneity may imply bias estimates and in particular it may lead to underestimation of the slope of hazard function. Further neglecting unobserved heterogeneity can also impact on the proportional impact of a change in a covariate on the hazard rate (Cameron and Trivedi, 2005).

The temporary work can finish with the conversion into a permanent contract or alternative destination states and in this case I should consider an independent single risk model that separates exit into permanent employment from exit into any other state. Therefore, exploiting some theoretical results, as explained in Guell and Petrongolo (2007), I can treat all temporary spells that do no end in a permanent contract as censored, without the need to distinguish the destination states.

The model is in continuous time where the elapsed duration is measured in days and the individual covariates have a fixed value at the beginning of each episode. Assuming the presence of *n* individuals, $i = 1,....,N$, I define the hazard rate denoting the probability of the transition out of the origin state (temporary work) as a continuous process towards destination state (permanent job)

$$\lambda_i(t_i^n \mid x, v, \beta) = \lambda_{0i}(t_i^k)\exp(x_i^{'}\beta\ )v_i$$

where $\lambda_{0i}$ defines the baseline hazard, which measures the effect of time spent in the origin state on the subsequent transition, i.e. what it is called the duration dependence of the process. The baseline hazard can be specified either parametrically or semi-parametrically. In the first case, a particular functional form is assumed and the Weibull is the parametric specification for the baseline hazard commonly used in the employment duration literature. For this reason I adopt a Weibull distribution for the baseline hazard, modelled as follow:

$$\lambda_{0ijk}(t) = pt_i^{p-1}.$$

In practice the duration dependence is measured by the parameter $p$ that implies positive duration dependence whether it is higher than one, negative duration dependence whether is less than one and no duration dependence whether equal to one. As customary in this literature, the multiplicative term $v_{ijk}$ takes into account the presence of unobserved heterogeneity, assumed to be distributed as a Gamma:

$$V \mid X \approx \Gamma(1,\theta)$$

The key assumptions used in the multiple-spells duration model are that the unobserved heterogeneity is independent of the observed covariates and usually enters the hazard function multiplicatively. Given these elements, the probability of surviving in the origin state for the elapsed duration $t$ is expressed by a survivor function defined as:

$$S(t \mid x, v) = \exp(-\int_0^t \sum_{j \neq k} \lambda_{jk}(s \mid x, v, \beta, \theta)\partial s^2$$

and substituting with the definition of the integrated hazard function, I can rewrite the above equation as follow:

$$S\ (t_i^n \mid x, v, \beta, \theta) = \exp\left\{-\Lambda\ (t_i^n \mid x, v, \beta, \theta)\right\}.$$

This expression defines the individual contribution to the likelihood function of a right censored spell, written in term of probability of surviving. Now I define the individual contributions to the likelihood function of a complete spell with duration $t_i^n$ in the origin state before transiting to stable work:

$$f_i(t_i^n \mid x, v, \beta, \theta) = S_i(t_i^n \mid x, v, \beta, \theta) \times \lambda_i(t_i^n \mid x, v, \beta, \theta)$$

Then, supposing a model without unobserved heterogeneity, I define the individual likelihood contribution with a sequence of spells $t_i^n$ as follows:

$$\ln L(\Omega \mid t_i^1, .., t_i^N, x_i) = \sum_{n=1}^{N}\left[d^N \ln\left(f_i(t_i^N \mid x_i; \Omega)\right) + r^N \ln(S_i(t_i^N \mid x_i, \Omega))\right]$$

where $d^N$ is a dummy indicator variable equals to one whether the individual transits from temporary to permanent employment in the $n^{th}$ spell and zero otherwise, while $r^N$ is a dummy variable, equal to one when the spell under consideration is censored and zero otherwise. The log-likelihood function is estimated both for the pre and the post-reform period. Moreover, I use robust standard error to correct the usual distortion due to the effects of clustered data.

Introducing the unobserved heterogeneity in the model becomes more complex, leading to the class of the mixture models. In fact, since the $v_i$ are unobservable, there is no possibilities to condition on them, but they are integrated over all possible values to obtain the so-called unconditional probabilities. As stated before, the multiplicative term $v_i$ are distributed as a Gamma and let us suppose that they are identically and independently distributed for all individuals.

Thus the individual likelihood function for the transition out to temporary employment is so defined:

$$L_i(\Omega \mid t_i^N, x_i, v) = \int_{-\infty}^{+\infty}\prod_{n=1}^{N} f_i(t_i^N \mid x_i, v_i, \Omega)^{d_i^n} \times \prod_{n=1}^{N} S_i(t_i^n \mid x_i, v_i, \Omega)^{r_i^n})\times d\Gamma(v_i)$$

## 5. Empirical results

In this section I first discuss the results of the PH model for the transitions out of temporary employment based on the Weibull specification which imposes a particular monotonic shape for the baseline hazard. The second set of results is from a piecewise constant baseline hazard specification adopted to check the robustness of the Weibull model once I have relaxed the parametric assumptions about the shape of baseline hazard.

### 5.1 Weibull model and duration dependence

Estimation results summarized in table 3 cover the 1992-2002 period in order to provide a preliminary evidence on the conversion pattern of temporary into permanent jobs during the '90s. Column (1) shows the results for the model not controlling for unobserved heterogeneity, while in column (2) I estimate a frailty PH model accounting for the unobserved heterogeneity among individuals. In order to do this, I introduce a multiplicative Gamma-distributed error term in the hazard function. Both specifications report the estimates for several individual characteristics that are consistent with the main indication of the empirical literature on job search model (Sciulli, 2006; Güell and Petrongolo, 2007; D'Addio and Rosholm, 2005). Moreover, columns (3) and (4) report the results for female and male workers. As can be seen from the results in column (2), the estimated coefficients in the PH model controlling for unobserved heterogeneity are substantially the same of those in column (1), nonetheless the variance of gamma error term is positive and statistically significant (0.121). Likelihood ratio test of zero unobserved heterogeneity is also rejected, indicating the frailty model is preferred to the reference non-frailty model. However, given that accounting for unobserved heterogeneity does not make a difference in the sense that leaves the estimated coefficients and the duration dependence parameter $p$ basically unchanged, I conclude that the other PH models may be accurately estimated without controlling for the unobserved heterogeneity.[7]

Referring to the estimated coefficients of the covariates included in the model emphasize how they affect the hazard of exiting out of temporary employment. Column (1) indicates that the probability of finding a permanent job is positively affected by age. This effect may capture a sort of job tenure impact which is increasing with age. In order to test this hypothesis, I estimate another PH specification that controls for on the job tenure, finding that longer market experience raises the hazard rate of transiting into permanent employment, probably capturing

[7] A similar conclusion has been found in on other microeconometric studies of employment duration model, as for instance Garcia-Perez and Munoz-Bullon (2005).

the impact of workers' skills on the future employment status. I decide not to report these estimates which are available from the author. In addition from a gender perspective the age effect is larger for men, as evidenced in columns (3) and (4). At a first glance, this gender difference is confirmed by the negative female coefficient (-0.162) in column (1) which remain basically unchanged once I control for the unobserved heterogeneity (-0.175).

**Table 3: Proportional Hazard model for transitions from temporary to permanent employment: 1992-2002, without and with unobserved heterogeneity**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | **All sample** | | **Female** | **Male** |
| Age | 0.107*** | 0.111*** | 0.080*** | 0.126*** |
| | (11.30) | (11.33) | (5.50) | (9.85) |
| Age^2 | -0.002*** | -0.002*** | -0.001*** | -0.002*** |
| | (-13.19) | (-13.20) | (-7.49) | (-10.24) |
| North-west | 0.292*** | 0.300*** | 0.311*** | 0.283*** |
| | (13.74) | (13.81) | (9.81) | (9.84) |
| North-east | 0.089** | 0.099*** | 0.115*** | 0.050* |
| | (3.65) | (4.18) | (3.39) | (1.66) |
| South | -0.215*** | -0.222*** | -0.355*** | -0.135*** |
| | (-9.09) | (-9.02) | (-9.35) | (-4.40) |
| Female | -0.162*** | -0.175*** | - | - |
| | (-10.32) | (-10.64) | | |
| Agriculture | 0.537*** | 0.073 | 0.655*** | 0.423** |
| | (5.28) | (1.36) | (4.62) | (3.00) |
| Mining industry | 0.553*** | 0.574*** | 0.682*** | 0.487*** |
| | (21.90) | (21.78) | (14.75) | (16.25) |
| Metal industry | 0.564*** | 0.589*** | 0.704*** | 0.447*** |
| | (20.35) | (20.18) | (16.02) | (12.56) |
| Mechanic industry | 0.533*** | 0.571*** | 0.578*** | 0.488*** |
| | (11.04) | (11.06) | (5.33) | (9.18) |
| Construction | 0.456*** | 0.506*** | 0.704*** | 0.291** |
| | (6.29) | (6.57) | (6.50) | (3.09) |
| Commerce | 0.548*** | 0.564*** | 0.627*** | 0.486*** |
| | (11.91) | (12.12) | (10.54) | (6.93) |
| Communication | 0.598*** | 0.633*** | 0.672*** | 0.507*** |
| | (26.61) | (26.24) | (19.69) | (16.36) |
| Credit | 0.546*** | 0.590*** | 0.682*** | 0.359*** |
| | (18.60) | (18.42) | (16.67) | (8.10) |
| Foreigners | 0.026 | 0.030 | -0.261*** | 0.145*** |
| | (1.05) | (1.15) | (-5.92) | (4.48) |
| Blue-collar | 0.556*** | 0.623*** | 0.615*** | 0.526*** |
| | (26.76) | (24.16) | (17.30) | (20.31) |
| White-collar | 0.599*** | 0.630*** | 0.627*** | 0.499*** |
| | (25.88) | (25.76) | (20.25) | (14.13) |
| Economic growth | 0.074*** | 0.079*** | 0.074*** | 0.071*** |
| | (11.61) | (11.62) | (7.80) | (8.18) |
| Gamma | | 0.121 | | |
| Likelihood ratio test | | 39.12 | | |
| | | p-value (0.000) | | |

Note: the table reports coefficient estimates of Weibull duration model and the test z are reported in parenthesis. All specifications control also for year dummies ***Significant at 1%, **significant at 5% and *significant at 10%.

The estimated coefficients also suggest strong regional variation in the pattern of exit from temporary employment towards a stable job. Generally, I find that to be resident in the Northern regions (mainly for those working in the North-West) affects positively the hazard rate of finding a permanent job and clearly the differences are larger for women than for men, as shown by the South coefficient in columns (3) and (4). The macro-region variables can be interpreted as a proxy for local labour market conditions that are usually captured using local unemployment rate and in this case, they confirm the dualistic structure of the Italian labour market. Focusing on the impact of economic growth, the estimated coefficients are positive and statistically significant, suggesting how the hazard rate of transiting from a temporary to a permanent job is positively affected by the economic cycle. In practice, this means better outside options for the workers and a higher incentive for the employer to retain the temporary workers at the end of the contract.

On the other hand, to be a foreign worker affects positively the hazard of exiting from temporary employment although this effect is not statistically significant in column (1). Quite different are the estimated coefficients if we look at gender differences. For women we detect a negative and significant coefficient (-0.261), indicating that non-Italians women have a lower probability of reaching a stable job, while for men the effect is positive (0.145) and this may be in line with expectation given the high diffusion of non-Italians workers in some sectors and in well-defined contractual type as interim work which may facilitate the transition to a permanent work after a probationary period.[8]

In the empirical studies on labour market transitions, the job-related characteristics are found to be an important factor in explaining the differences in the hazard of exit from the state of temporary workers. For this reason table 4 reports the same PH model but distinguishing between blue-collar and white-collar, in order to investigate whether the occupational status is a source of different labour transitions. At a first glance, the female dummy displays a positive and statistically significant effect for the white-collar in column (1), while the opposite is reported in column (2) for the blue-collar category, in line with the traditional results in the literature. Even though the nature of the data makes not possible to identify the source of this specific result, sample composition may provide a first explanation. Furthermore, adding more evidence to the previous results for the foreign workers, I detect a positive impact on the hazard rate of transiting form a temporary to a permanent job for non-Italians workers only in the blue-collar group, which represents the category of jobs in the industry sector.

[8] The idea is that in some sectors as manufacturing and mechanic the increasing diffusion of non-Italians workers has been emphasized by the introduction of new flexible contracts such as interim work which favours to employ the worker for a short period and to test his skills, so providing a possible path towards a stable jobs.

Concerning the regional variation, the estimated coefficients summarized in table 4 indicate a larger penalty in the Southern regions for blue-collar workers, in this sense emphasizing the traditional weakness of industry sector in these regions.

**Table 4: Proportional Hazard model for transitions from temporary to permanent employment: 1992-2002, by job-related characteristics**

| | (1) | (1) |
|---|---|---|
| | **White-collar** | **Blue-collar** |
| Age | 0.097** | 0.099*** |
| | (2.96) | (5.48) |
| Age^2 | -0.001** | -0.001*** |
| | (-2.06) | (-4.10) |
| North-west | 0.298*** | 0.272*** |
| | (7.10) | (6.48) |
| North-east | -0.067 | 0.030 |
| | (-1.31) | (0.69) |
| South | -0.027 | -0.079* |
| | (-0.56) | (-1.85) |
| Female | 0.072** | -0.062* |
| | (2.19) | (-1.91) |
| Agriculture | -0.031 | -0.005 |
| | (-0.20) | (-0.03) |
| Mining industry | 0.297*** | 0.056 |
| | (4.28) | (1.37) |
| Metal industry | 0.254 | -0.012 |
| | (3.22) | (-0.28) |
| Mechanic industry | 0.065 | 0.141 |
| | (0.56) | (1.58) |
| Construction | 0.058 | -0.191 |
| | (0.53) | (-1.61) |
| Commerce | 0.271*** | 1.22*** |
| | (3.63) | (3.95) |
| Communication | 0.211*** | -0.114*** |
| | (3.20) | (-2.61) |
| Credit | 0.116* | -0.097** |
| | (1.68) | (-2.01) |
| Foreigners | 0.018 | 0.140*** |
| | (0.19) | (3.47) |
| Economic growth | 0.093*** | 0.123*** |
| | (6.60) | (9.44) |

Note: the table reports the coefficient estimates of a Weibull duration model and the test z are reported in parenthesis. All specifications control also for year dummies. *** Significant at 1%, ** significant at 5% and * significant at 10%.

Estimation results summarized in table 5 display the results from a PH model with Weibull specification for the baseline hazard, comparing the pre and post-reform period in order to investigate whether the Treu package has modified the impact of the observable characteristics on the hazard rate of moving from temporary to a stable job and in what direction. With regard to individual characteristics, the age coefficient is larger for both women and men in the post reform period as evidenced in columns (3) and (4), probably capturing the positive impact of job experience on the hazard rate of exiting from temporary employment. Moreover, the estimated coefficients for the geographical dummies indicate that the Treu package has not substantially modified the traditional duality of the Italian labour market. However, the higher degree of flexibility and the introduction of new flexible contracts has reduced the negative effect of being resident in the South, as shown by coefficient in columns (3) and (4). Also the North-West area reports a positive effect on the hazard rate, significantly larger in the post-reform period.

Moreover, the gender differences increase in the post-reform period if we look at the female coefficient in table 5, as well as the economic growth coefficient shows a negative impact in the pre-reform period characterized by the 1992-93 recession and a positive and significant sign in the post-reform period, capturing the influence of the economic cycle on labour market transitions. A further interesting result concerns the foreign workers. In the pre-reform period being a non-Italians worker clearly reduces the hazard rate of moving into a stable job both for women and men. Differently the introduction of Treu package has modified this pattern and mainly for men the estimated coefficient for the post-reform period is positive (0.247). In practice the diffusion of temporary work agency (TWA) employment can have helped the entry in the Italian labour market of foreign workers, improving their human capital and providing them work experience, so increasing their chance of transiting into a permanent work, first of all in the industry sector.[9]

One popular hypothesis is that an individual’s permanent employment prospects improve with the duration of the temporary employment spell, which is emphasized as positive duration dependence. In practice, the idea is the existence of a tenure effect in the job transitions, i.e. the probability of finding a permanent job after a temporary work experience increases with the length of the temporary contract. For this reason changes in duration dependence have important implications for labour market policy. In this section I investigate whether the Treu package has modified the positive duration dependence pattern, in what direction and the role of individual heterogeneity in explaining the large amount of variability provided by the estimates.

Figures 1a and 1b display the hazard function in the pre-reform period by macro-regions and by gender. The general pattern shows that the level of estimated hazard rates differ widely between the Northern regions and the South as well as between women and men. Although the employment hazard increases for all groups considered with the duration of employment, the

[9] For a more detailed description of TWA employment see Labour and Social Policy Minister (MLPS, 2006).

differences remain remarkable. At the same way, figures 2a and 2b display the hazard function but for the post-reform period. A first interesting result concerns the reduction in the difference between the South and the North-East in hazard function as well as a lower gender gap in hazard rate at higher duration of temporary employment.

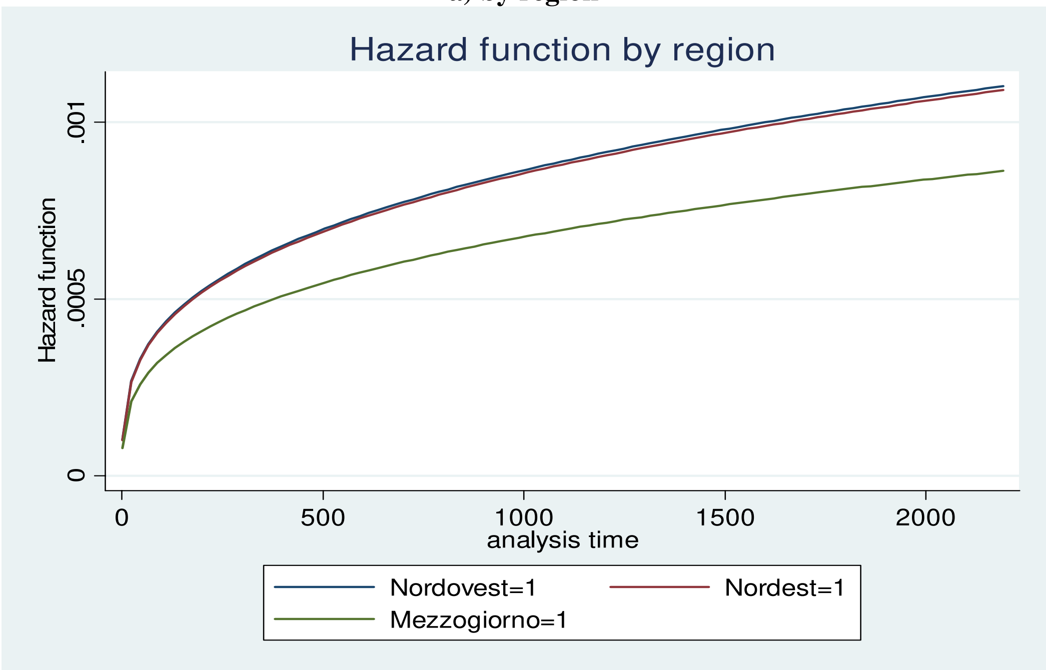
**Table 5: Proportional Hazard model for transitions from temporary to permanent employment: pre-reform period, 1992-1997**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | **Pre reform** | **Pre reform** | **Post reform** | **Post reform** |
| | **Female** | **Male** | **Female** | **Male** |
| Age | 0.051 | 0.095*** | 0.171*** | 0.186*** |
| | (1.42) | (3.44) | (8.11) | (10.05) |
| Age^2 | -0.002** | -0.002*** | -0.002*** | -0.002*** |
| | (-3.01) | (-4.02) | (-7.92) | (-8.70) |
| North-west | 0.197*** | 0.166*** | 0.381*** | 0.427*** |
| | (3.43) | (3.04) | (8.49) | (10.23) |
| North-east | 0.147** | 0.099* | 0.100** | 0.130** |
| | (2.33) | (1.67) | (2.03) | (2.89) |
| South | -0.526*** | -0.384*** | -0.232*** | -0.046 |
| | (-7.70) | (-6.39) | (-4.42) | (-1.07) |
| Agriculture | 0.604* | 0.481** | 0.574** | 0.645*** |
| | (1.81) | (2.11) | (2.86) | (3.61) |
| Mining industry | 0.706*** | 0.648*** | 0.816*** | 0.524*** |
| | (8.07) | (11.26) | (11.64) | (11.02) |
| Metal industry | 0.875*** | 0.524*** | 0.695*** | 0.484*** |
| | (11.66) | (7.79) | (9.51) | (8.52) |
| Mechanic industry | 0.677*** | 0.745*** | 0.726*** | 0.603*** |
| | (2.73) | (7.63) | (5.24) | (7.85) |
| Construction | 0.977*** | 0.363 | 0.724*** | 0.379*** |
| | (4.45) | (1.26) | (4.98) | (3.48) |
| Commerce | 0.623*** | 0.788*** | 0.725*** | 0.528*** |
| | (4.55) | (6.79) | (7.88) | (6.31) |
| Communication | 0.798*** | 0.742*** | 0.779*** | 0.527*** |
| | (12.94) | (12.66) | (13.60) | (10.91) |
| Credit | 0.842*** | 0.731*** | 0.828*** | 0.376*** |
| | (10.57) | (6.34) | (13.31) | (6.61) |
| Foreigners | -0.410*** | -0.111 | -0.047 | 0.247*** |
| | (-4.23) | (-1.63) | (-0.81) | (5.70) |
| Blue-collar | 0.572*** | 0.247*** | 0.838*** | 0.698*** |
| | (7.78) | (4.47) | (17.24) | (17.47) |
| White-collar | 0.390*** | 0.243*** | 0.698*** | 0.538*** |
| | (6.08) | (3.23) | (15.64) | (10.91) |
| Economic growth | -0.034** | -0.077*** | 0.208*** | 0.234*** |
| | (-2.18) | (-5.13) | (14.07) | (17.96) |

Note: the table reports the coefficient estimates of a Weibull duration model and the test z are reported in parenthesis. All specifications control also for year dummies. *** Significant at 1%, ** significant at 5% and * significant at 10%.

**Figure 1: Daily Hazard Function for transition to permanent employment, pre-reform period, 1992-1997**

**a) by region**




**b) by gender**

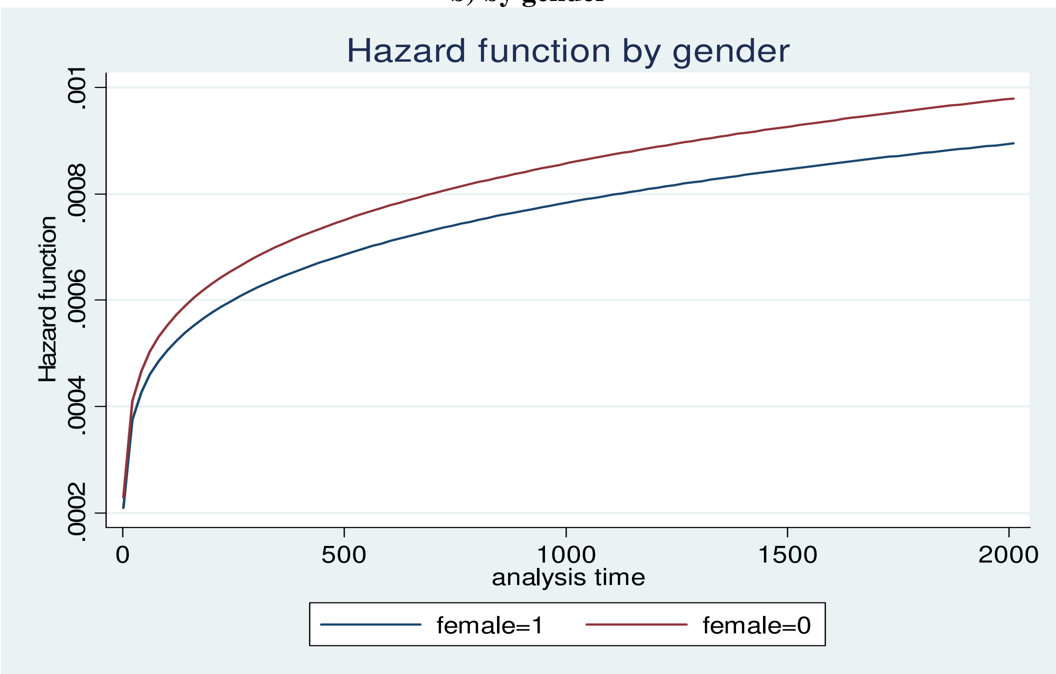
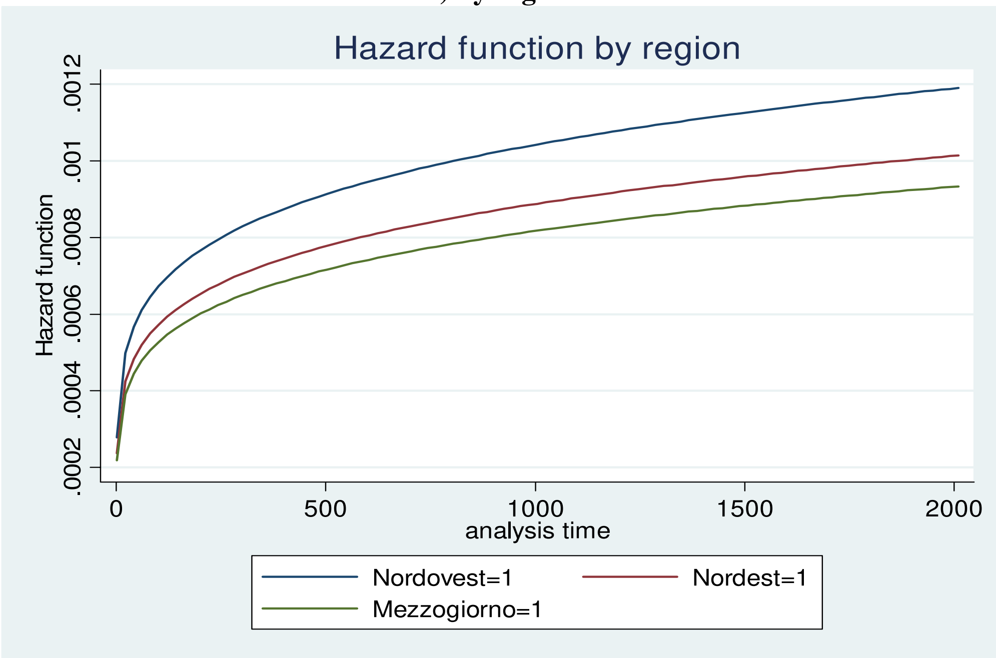
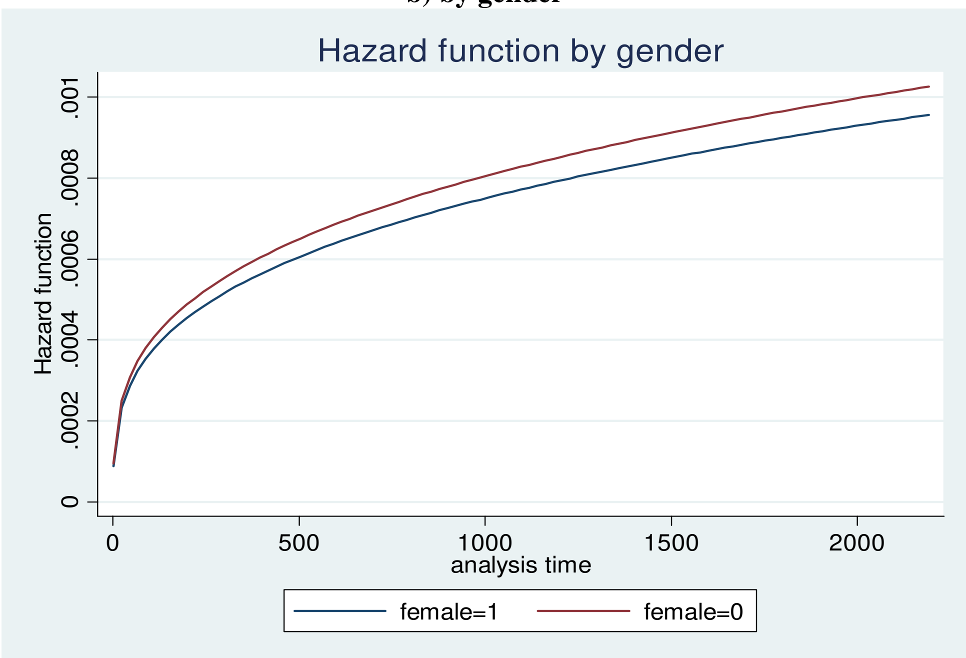


**Figure 2: Daily Hazard Function for transition to permanent employment, post-reform period, 1997-2002**

**a) by region**




**b) by gender**




The estimation results concerning the Weibull duration parameter $p$ for the pre and post-reform period are summarized in table 6. In addition, as previously discussed, accounting for unobserved heterogeneity does not make difference the estimated duration dependence parameters and given these results, I decide not to report the estimates for the model controlling for unobserved heterogeneity. At a first glance, the evolution of the parameter $p$ exhibits more or less a substantial reduction in the positive duration dependence in the post-reform period.[10] However, the persistence of positive duration dependence might indicate a sort of human capital accumulation effect, i.e. contract duration positively affects the hazard rate of moving out of temporary work towards a stable job. Even more pronounced is the reduction in duration dependence for women and white-collar, respectively from 1.37 to 1.25 and from 1.71 to 1.36.

Generally, after the introduction of the Treu package, the estimated parameters are lower for all groups considered. A possible explanation might be related to the introduction of TWA employment which allows for interrupted working spells, so reducing the hazard rate of exit to permanent employment (Gagliarducci, 2005 and Sciulli, 2006). More in detail, the estimated parameters suggest that TWA employment displays a negative duration dependence, given a parameter $p$ less than one (0.96). This result indicates that contract duration affects negatively the probability of moving into a stable jobs. A similar conclusion might be attributed to fixed-term works and to blue-collar workers and this could imply that declining duration dependence is an obvious result of the sorting process in these new contract types.

Figures 3 and 4 display the hazard function both for fixed-term jobs and TWA works, confirming the negative duration dependence summarized in table 6.

**Table 6: Duration dependence parameter by personal and job-related characteristics**

| | **Pre-reform** | **Post-reform** |
|---|---|---|
| | | |
| Female | 1.37<br>(0.026) | 1.25<br>(0.017) |
| Male | 1.28<br>(0.20) | 1.18<br>(0.014) |
| Blue-collar | 1.25<br>(0.034) | 0.99<br>(0.013) |
| White-collar | 1.71<br>(0.046) | 1.36<br>(0.025) |
| Fixed-term | - | 0.99<br>(0.013) |
| Temporary work | - | 0.96<br>(0.027) |

Baseline hazard: Weibull specification. Standard errors are reported in parenthesis.

[10] Sciulli (2006) also finds a similar reduction in duration dependence parameter in the post-reform period looking at the transition out a non-working state to permanent employment (negative duration dependence). This result confirms the theoretical predictions, suggesting that in a more flexible labour market the duality between short-term and long-term unemployed increases, i.e. to be a long-term unemployed may be perceived as a low productivity signal by firms.

**Figure 3: Daily Hazard Function for transition to permanent employment, post-reform period, by gender, fixed-term contracts**

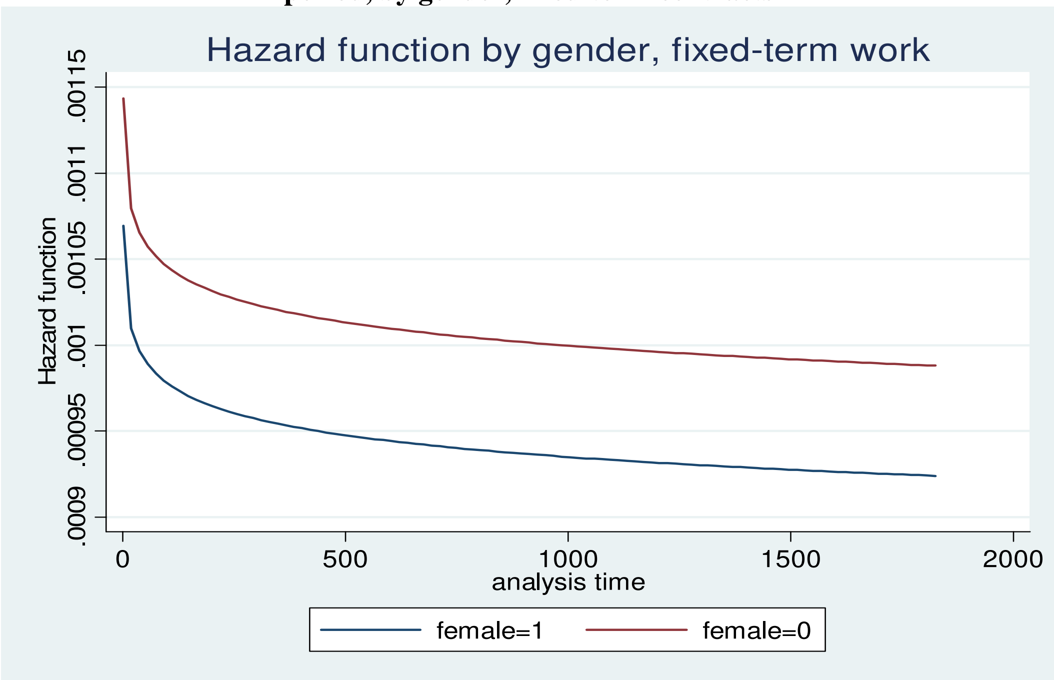
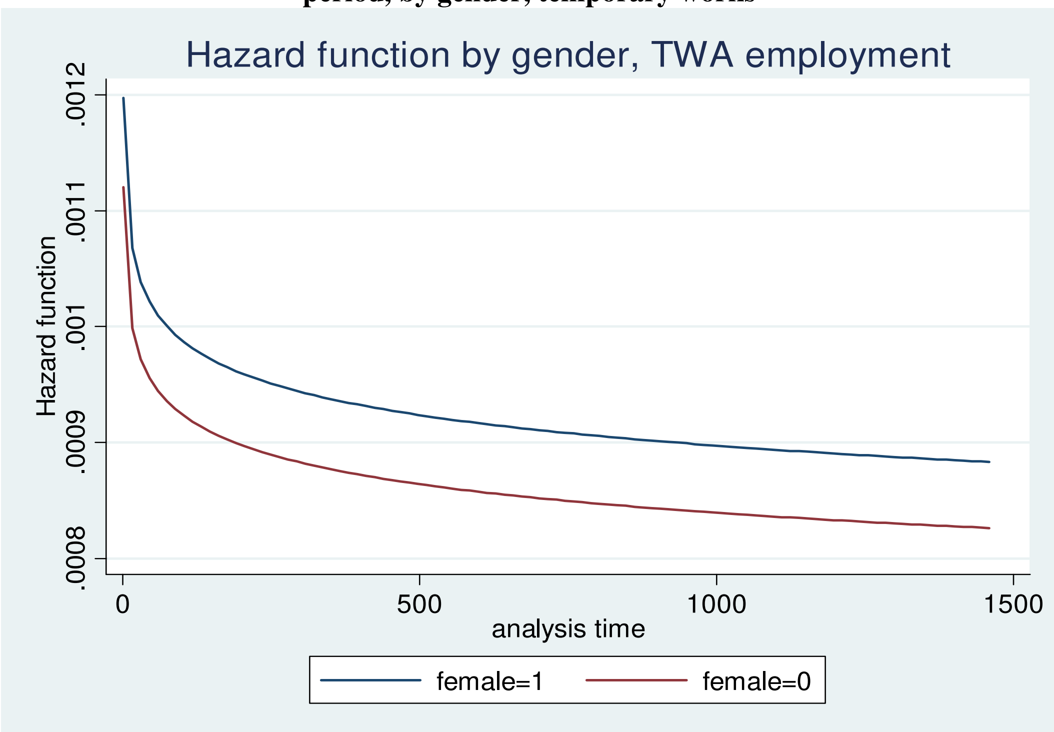



**Figure 4: Daily Hazard Function for transition to permanent employment, post-reform period, by gender, temporary works**



## 4.2 Sensitivity analysis

In this section I assess the robustness of the previous results, relaxing the strict parametric assumptions about the Weibull distribution of the baseline hazard. Indeed imposing a particular functional form for the baseline hazard may lead to a problem of misspecification. A way to solve this problem is first to adopt a semi-parametric piecewise constant specification and then to interact each piece of the baseline hazard with a reform dummy to investigate whether the piece is proportionally shifted by the Treu reform. Such specifications have the additional advantage in that estimated coefficients will be less sensitive to the distributional assumptions applied for unobserved heterogeneity, as discussed in the context of proportional hazard models (Han and Hausman, 19990). In other words, the piecewise constant hazard model is another example of the proportional hazard models and allows for period-specific differences in the risk of job transition, modelling the baseline hazard as a step function with $j$ interval so that

$$\lambda_0(t) = \lambda_j \qquad \tau_{j-1} < t < \tau_j$$

where $\tau_j$ specify the discrete changes in the baseline hazard. The $jth$ interval starts at time $\tau_{j-1}$ and ends at time $\tau_j$ but in the each interval the baseline hazard is constant and equal to $\lambda_j$. To investigate whether the duration dependence has increased or decreased proportionally over time, I further interact the reform dummy with each piece of the baseline hazard, namely $\lambda_j * r97$.

Estimation results for the proportional hazard model with a piecewise constant baseline hazard are summarized in table 7. To avoid a collinearity, as estimation strategy, I decide to include a constant term and to exclude the first piece of the baseline hazard which makes possible to look directly at how each interval differs from that of the interval corresponding to the excluded variable. In practice the estimated coefficient for the baseline dummies are to be interpreted relative to the base category, i.e. the first month of the temporary employment spell. I identify at most 11 baseline steps, so providing a specific hazard rate for the following time periods: month 1; month 2; month 3, months 4-6, months 7-9; months 10-12, months 13-24; months 25-36; months 37-48; months 49-60 and months>60. Since the purpose of this robustness check is to test for the changes in the duration dependence before and after the Treu reform, I will focus on the results for the baseline function and the relative impact of the reform and will only briefly concentrate on the estimates with respect to the other covariates of the traditional model, just discussed in the previous section.

**Tab 7: Proportional hazard model for transitions to permanent employment, piecewise baseline hazard**

| | Pre reform period (1992-1997) | | | Post reform period (1997-2002) | | |
|---|---|---|---|---|---|---|
| | **All sample** | **Female** | **Male** | **All sample** | **Female** | **Male** |
| Age | 0.122** | 0.225** | 0.079 | 0.261*** | 0.251*** | 0.265*** |
| | (2.16) | (2.59) | (1.21) | (16.22) | (10.52) | (11.97) |
| Age^2 | -0.001* | -0.004** | -0.001 | -0.003*** | -0.003 | -0.003*** |
| | (-1.69) | (-2.43) | (-0.67) | (12.21) | (-8.12) | (-8.80) |
| North-west | 0.137*** | 0.161** | 0.125** | 0.370*** | 0.352*** | 0.381*** |
| | (3.09) | (2.52) | (2.01) | (10.85) | (7.07) | (8.14) |
| North-east | 0.076 | 0.082 | 0.074 | 0.075** | 0.005 | 0.142*** |
| | (1.54) | (1.14) | (1.08) | (2.04) | (0.11) | (2.83) |
| South | -0.205*** | -0.262*** | -0.172** | -0.013 | -0.085 | 0.050 |
| | (-4.01) | (3.40) | (-2.49) | (-0.35) | (-1.45) | (1.02) |
| Female | 0.084** | | | -0.084*** | | |
| | (2.45) | | | (-3.37) | | |
| Agriculture | 0.315 | 0.312 | 0.406* | 0.441*** | 0.407** | 0.494** |
| | (1.52) | (0.78) | (1.83) | (3.09) | (1.98) | (2.46) |
| Mining industry | 0.559*** | 0.578*** | 0.565*** | 0.594*** | 0.652*** | 0.568*** |
| | (10.99) | (6.31) | (9.21) | (14.75) | (9.01) | (11.70) |
| Metal industry | 0.493*** | 0.604*** | 0.418*** | 0.535*** | 0.506*** | 0.559*** |
| | (9.17) | (7.42) | (5.79) | (11.76) | (6.84) | (9.73) |
| Mechanic industry | 0.658*** | 0.488* | 0.676*** | 0.484*** | 0.470*** | 0.479*** |
| | (6.69) | (1.75) | (6.43) | (6.62) | (3.13) | (5.79) |
| Construction | 0.613*** | 0.793*** | 0.474* | 0.461*** | 0.452*** | 0.473*** |
| | (3.16) | (3.39) | (1.72) | (4.76) | (2.75) | (4.03) |
| Commerce | 0.501*** | 0.320** | 0.677*** | 0.613*** | 0.609*** | 0.617*** |
| | (5.29) | (2.17) | (5.56) | (9.35) | (6.06) | (7.33) |
| Communication | 0.638*** | 0.612*** | 0.685*** | 0.515*** | 0.536*** | 0.507*** |
| | (13.79) | (9.01) | (10.64) | (13.79) | (9.35) | (9.99) |
| Credit | 0.636*** | 0.640*** | 0.660*** | 0.570*** | 0.632*** | 0.498*** |
| | (9.49) | (7.46) | (5.49) | (13.87) | (10.15) | (8.82) |
| **Baseline hazard** | | | | | | |
| e2 | 0.518*** | 0.467*** | 0.563*** | 0.321*** | 0.206** | 0.403*** |
| | (4.50) | (2.79) | (3.56) | (4.78) | (1.99) | (4.56) |
| e3 | 0.643*** | 0.508*** | 0.754*** | 0.290*** | 0.250** | 0.319*** |
| | (5.66) | (3.03) | (4.88) | (4.24) | (2.41) | (3.49) |
| e4 | -0.055 | -0.312* | 0.140 | -0.043 | -0.065 | -0.027 |
| | (-0.51) | (-1.93) | (0.97) | (0.71) | (-0.71) | (-0.34) |
| e5 | -0.022 | -0.148 | 0.085 | -0.363*** | -0.374*** | -0.355 |
| | (-0.21) | (-0.93) | (0.58) | (-5.42) | (3.72) | (-3.95) |
| e6 | 0.075 | 0.206 | -0.070 | -0.378*** | -0.362*** | -0.391*** |
| | (0.69) | (1.34) | (-0.45) | (5.48) | (-3.53) | (-4.18) |
| e7 | 0.566*** | 0.500*** | 0.628*** | -0.071 | -0.057 | -0.082 |
| | (5.89) | (3.60) | (4.71) | (-1.27) | (-0.69) | (-1.09) |
| e8 | 1.316*** | 1.470*** | 1.181*** | 0.969*** | 1.056*** | 0.886*** |
| | (13.26) | (10.28) | (8.56) | (16.83) | (12.39) | (11.33) |
| e9 | 1.219*** | 1.290*** | 1.187*** | 0.743*** | 0.664*** | 0.803*** |
| | (11.01) | (7.83) | (7.89) | (10.73) | (6.36) | (8.69) |
| e10 | 1.096*** | 0.835*** | 1.244*** | 0.876*** | 0.860*** | 0.886*** |
| | (8.15) | (3.06) | (7.14) | (8.65) | (5.77) | (6.41) |
| e11 | 1.094*** | 1.216*** | 1.033*** | 1.502*** | 1.649*** | 1.408*** |
| | (4.58) | (3.06) | (3.42) | (6.82) | (4.84) | (4.86) |

Note: the table reports the coefficient estimates of a piecewise constant baseline hazard model and the test z are reported in parenthesis. All specifications control also for foreign status, white-collar, blue-collar and economic growth. *** Significant at 1%, ** significant at 5% and * significant at 10%.

The first important result worthy of a note is that both for women and men the estimated parameters of the baseline hazard for temporary employment duration are substantially lower in the post reform period, in this sense confirming the hypothesis widely discussed in the previous section that the Treu package has reduced the degree of duration dependence in the hazard rate of moving out of temporary employment. Baseline hazard coefficients in both periods exhibit more or less a positive value, with some exceptions. In particular the baseline parameters are positive for the first two time period. After the second period, however, the baseline hazard estimates are found to be negative, mainly in the post reform period. This tendency lasts until the sixth-seventh time period, implying a negative duration dependence in this interval. Even more pronounced, after the seventh time interval, is the change in sign in the baseline hazard coefficients which become strongly positive and volatile, mainly for women.

The last robustness check is summarized in table 8 which compares two model both for men and women, the first is a traditional piecewise baseline model while in the second each piece of the baseline hazard is interacted with a reform dummy in order to test whether the Treu package has proportionally shifted the hazard rate. The estimated coefficients imply a rather strong negative effect of the reform on hazard rate equally distributed by gender, resulting in a well-defined and proportional reduction of each piece of the baseline hazard. In conclusion, results from semi-parametric baseline hazard tell a slightly different but more appealing story. Indeed, the hazard estimates are positive initially, then decline moving to negative duration dependence in the central time interval and finally increase substantially, reproducing the positive duration dependence already observed in the Weibull model.

In addition, the results with respect to other explanatory variable for both specifications (tables 7 and 8) are in accordance with one of the main implication of job-search theory and with the estimates discussed in the previous section.

## 6. Conclusion

The empirical evidence that the average hazard rate from temporary employment (towards a permanent work) rises with its duration is typically interpreted in literature as strong evidence for the hypothesis of positive structural duration dependence. This paper attempts to study the duration of temporary employment, focusing on the exit margin, and to investigate whether the introduction of Treu package in the 1997 has modified the duration dependence pattern in the probability of moving from a fixed-term work to a stable job.

**Tab 8: Proportional hazard model for transitions to permanent employment, piecewise baseline hazard**

| | Piecewise Baseline Hazard | | Piecewise Baseline Hazard Interacted with the reform dummy | |
|---|---|---|---|---|
| | Female | Male | Female | Male |
| Age | 0.101*** | 0.138*** | 0.147*** | 0.166*** |
| | (5.11) | (7.31) | (7.46) | (9.08) |
| Age^2 | -0.001*** | -0.001*** | -0.002*** | -0.002*** |
| | (-3.72) | (-4.95) | (-5.47) | (-6.41) |
| North-west | 0.251*** | 0.229*** | 0.258*** | 0.238*** |
| | (7.00) | (6.87) | (7.58) | (7.42) |
| North-east | 0.010 | 0.049 | 0.031 | 0.069** |
| | (0.26) | (1.36) | (0.87) | (1.99) |
| South | -0.130*** | 0.039 | -0.132*** | 0.039 |
| | (-3.00) | (1.10) | (-3.20) | (1.11) |
| Agriculture | 0.207 | 0.144 | 0.143 | 0.149 |
| | (1.32) | (0.90) | (0.96) | (0.98) |
| Mining industry | 0.364*** | 0.347*** | 0.308*** | 0.336*** |
| | (7.41) | (10.75) | (6.68) | (10.82) |
| Metal industry | 0.341*** | 0.306*** | 0.307*** | 0.295*** |
| | (7.43) | (7.9))3) | (7.07) | (7.94) |
| Mechanic industry | 0.270** | 0.313*** | 0.232** | 0.300*** |
| | (2.49) | (5.34) | (2.25) | (5.32) |
| Construction | 0.209* | 0.147 | 0.162 | 0.124 |
| | (1.68) | (1.39) | (1.36) | (1.22) |
| Commerce | 0.239*** | 0.353*** | 0.213*** | 0.352*** |
| | (3.63) | (5.10) | (3.57) | (5.49) |
| Communication | 0.295*** | 0.342*** | 0.241*** | 0.319*** |
| | (8.39) | (10.29) | (7.33) | (9.95) |
| Credit | 0.302*** | 0.218*** | 0.248*** | 0.194*** |
| | (7.03) | (4.56) | (6.17) | (4.20) |
| **Baseline hazard** | | | | |
| e2 | 0.273*** | 0.448*** | 0.200* | 0.399*** |
| | (3.17) | (6.02) | (1.94) | (4.50) |
| e3 | 0.311*** | 0.455*** | 0.238** | 0.307*** |
| | (3.62) | (6.07) | (2.29) | (3.36) |
| e4 | -0.151* | -0.041 | -0.087 | -0.055 |
| | (-1.94) | (0.60) | (-0.95) | (-0.67) |
| e5 | -0.314*** | -0.268*** | -0.404*** | -0.406*** |
| | (-3.83) | (-3.65) | (-4.03) | (-4.52) |
| e6 | -0.201** | -0.371*** | -0.407*** | -0.460*** |
| | (-2.46) | (-4.85) | (-4.00) | (-4.97) |
| e7 | 0.106 | 0.082 | -0.085 | -0.136* |
| | (1.54) | (1.31) | (-1.03) | (-1.83) |
| e8 | 1.114*** | 0.801*** | 0.966*** | 0.687*** |
| | (15.89) | (12.60) | (11.68) | (9.16) |
| e9 | 0.740*** | 0.624*** | 0.506*** | 0.429*** |
| | (9.40) | (9.00) | (5.52) | (5.29) |
| e10 | 0.584*** | 0.639*** | 0.395*** | 0.438*** |
| | (6.39) | (8.28) | (3.79) | (4.94) |
| e11 | 0.691*** | 0.742*** | 0.509*** | 0.559*** |
| | (7.16) | (9.42) | (4.85) | (6.40) |

Note: the table reports the coefficient estimates of a piecewise constant baseline hazard model and the test z are reported in parenthesis. All specifications control also for foreign status, white-collar, blue-collar, a reform dummy and economic growth. *** Significant at 1%, ** significant at 5% and * significant at 10%.

Indeed the partial labour market reform and the liberalization of flexible contracts in the Italian labour market have attracted the attention of many researchers and policy-makers. Empirical studies have tried to assess whether temporary jobs provide a stepping-stone towards a stable employment or an endless trap in the precariousness, but mixed employment effects do not provide a clear signal on the good performance of the flexible contracts in terms of subsequent careers and wage.

Parametric and semi-parametric continuous time duration model have been employed to study the duration of temporary work spells and to control for the unobserved heterogeneity, exploiting a large administrative dataset for the ‘90s. As already discussed, the results obtained indicate that accounting for unobserved heterogeneity does not matter and the estimated parameters with regards to duration dependence remains basically unchanged.

The main findings emphasize how the Treu reform has significantly reduced the duration dependence parameters in the probability of transiting from a temporary work to a permanent job and that this result is robust to different specifications adopted for the baseline hazard, both parametric and semi-parametric. Moreover, results on positive duration dependence might indicate the existence of a human capital accumulation effect, as suggested in other empirical studies. The reduction in duration dependence differs markedly with respect to observable characteristics and it is larger for women and white-collar in the post reform period.

An even stronger conclusion could be drawn, concerning the analysis for specific contract types: once I have controlled for different contractual arrangements, positive duration dependence in the employment hazard rate disappears or even becomes negative, at least for the temporary work agency (TWA) employment. This might indicate that for interim work the duration of the work spell affects negatively the hazard rate of moving into a permanent employment. Given the somewhat restrictive information of the dataset, more research to test for the robustness of the negative duration dependence in TWA employment seems warranted. Finally I conclude emphasizing the role of individuals’ heterogeneity in terms of personal and job-related characteristics in explaining a large amount of differences in hazard rates and, more in general, in labour market chances.

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