

Collegio Carlo Alberto

Do Larger Severance Payments Increase Individual Job Duration?

Pietro Garibaldi

Lia Pacelli

Working Paper No. 39

February 2007

www.carloalberto.org

Do Larger Severance Payments Increase Individual Job Duration?¹

Pietro Garibaldi
University of Torino and Collegio Carlo Alberto

Lia Pacelli
University of Torino, and Laboratorio R. Revelli - Collegio Carlo Alberto

February 2007²

¹We thank the editor and two anonymous referees for very helpful comments, and Giuseppe Bertola, Francesco Devicienti, Michelangelo Filippi, and Hugo Hopenhayn for detailed comments. We also thank seminar participants at Labor (Torino), Ecares (Brussels), Bocconi University of Milan, IZA Bonn, University of Rome La Sapienza, University of Brescia, CEPR Workshop in Bristol and at the joint World Bank-Ludwig Boltzmann Gesellschaft conference on Severance Payments and Individual Accounts. Pietro Garibaldi is also affiliated with IGIER, IZA and CEPR. Lia Pacelli acknowledges financial support by a grant from the Italian Ministry of Education, University and Research to the project “Evaluating the impact of public interventions: methods and case studies?”

²© 2007 by Pietro Garibaldi and Lia Pacelli. Any opinions expressed here are those of the authors and not those of the Collegio Carlo Alberto.

Abstract

This paper analyzes the effect of severance payments on the probability of separation at given tenure, wages and other individual and firm characteristics. It studies a mandatory deferred wage scheme of the Italian labour market (Trattamento di Fine Rapporto, TFR). Deferred wages increase job duration if two conditions hold: wages are rigidly set outside the employer-employee relationship, and past provisions are accumulated at interest rates that are below market rates. Under such circumstances, workers who withdraw from their accumulated stock of unpaid wages should experience, at given tenure, a subsequent increase in the probability of separation. This prediction appears empirically robust and quantitatively sizeable. A withdrawal of 60% of the TFR stock (the median observed withdrawal) increases the instantaneous hazard rate by almost 20%. In other words, an individual with at least ten years of tenure that experiences an early withdrawal increases his/her hazard rate from 10% to about 12%. The empirical result takes into account the existence of unobserved heterogeneity and a variety of further robustness tests.

JEL Classification: J0, J3, J6

Keywords: labor markets, severance payments, wage schemes, job tenure, job separation

1 Introduction

More stringent Employment Protection Legislation (EPL) should induce labour hoarding from the firm stand-point: other things equal, employer initiated separations should be lower for individuals with stricter EPL¹. While such theoretical prediction is unambiguous, little is known on the empirical links between EPL and separation rates at the job level². Indeed, two empirical regularities prevent straightforward identification of the theoretical prediction. On the one hand, it is well established that the probability of job termination declines markedly with tenure (Farber 1999), independently of the presence of severance payments. On the other hand, severance payments increase with job tenure. As a result, it is very difficult to identify the effect of severance payments on labour hoarding and employer initiated separation. This paper exploits an institutional feature of the Italian labour market that makes it possible to identify this effect at the micro level.

Traditional severance payments, defined as statutory firm worker transfers in case of firm initiated separation, do not exist in Italy (Brandolini and Torrini, 2002). Nevertheless, Italy features a mandatory deferred wage (*Trattamento Fine Rapporto*, *TFR* hereafter) that can be akin to a severance payment. It is paid to the worker at the end of the employment relationship, regardless of the reasons behind the separation. During the entire working relationship, the worker accumulates a credit vis-à-vis the firm, which is in turn obliged to recapitalize its debt to the worker at policy determined interest rates. Over the last twenty years, such policy determined interest rates have been traditionally very low, so that *TFR* resulted in subsidized financing of firm operations from the part of workers. We solve a simple partial equilibrium model of job destruction with deferred wages, and show that deferred wages are akin to a severance payment. As long as wages are rigidly set outside the employer-employee relationship, and deferred wages are accumulated below market rates (both conditions held in Italy during our observation period), a scheme like *TFR* increases the firm propensity to hoard labour. As a way to keep the subsidized financing on the part of workers, firms have incentives to delay job separation and to increase the average duration of jobs. The latter result suggests that *TFR* has the same effects of a severance payment with fixed wages, and appears equivalent to a piece of employment protection legislation.

Within the existing institutional setting workers are allowed to proceed to an advance withdrawal of their accumulated credit. Our theoretical model clearly predicts that following a random shock, workers who withdraw from their accumulated credit of unpaid wages increase the probability of a firm initiated separation. This suggests that we can empirically identify the effect of

¹See for example, Nickell, 1986; Bentolila and Bertola, 1990 and Bertola 1999; Hopenhayn-Rogerson, 1993; Ljungqvist, 2001.

²A possible example is Kugler (1999), who studies the effect of changes in firing costs in Colombia.

severance payments on the probability of separation, at given tenure, wages and other individual and firm characteristics. Our empirical analysis suggests that such effect is both sizeable and robust.

We rely on the Work Histories Italian Panel (WHIP), a longitudinal micro dataset drawn from the social security administration (INPS) archives and processed in a public-use file by Laboratorio Revelli³. Using information on about 20,000 employment spells initiated between 1985 and 1988 and followed up to 1999, we test the impact of advance withdrawals on the hazard rate, i.e. on the probability of firm initiated separation between t and $t + 1$ conditional on having a tenure of at least t , using a variety of survival models. Overall, we find that withdrawing from the *TFR* fund significantly increases the hazard rate, even when we control for wages, industry and occupational effects. Our results are also quantitatively non negligible. A withdrawal of 60% of the TFR stock (the median observed withdrawal) increases the instantaneous hazard rate by almost 20%. In other words, an individual with at least ten years of tenure that experiences an advance withdrawal increases his/her hazard rate from 10% to about 12%. We perform a variety of robustness checks, focussing mainly on unobserved heterogeneity and reverse causality, and we find our results robust.

Our results are also relevant in the policy debate. The social security reforms approved by the Italian Parliament in 2004 and 2006 envisage a two pillar system. The first pillar will be a national pension. The second pillar will be made by private pension schemes. The 2006 law considers the *TFR* as a base payment to fill up the private pension funds. Our results suggest that a shift of *TFR* funds into pension funds will increase labour turnover, so that implementing such reform will have a direct effect in the social security area, and an indirect effect on the labour market.

The paper proceeds as follows. Section 2 presents the Italian institutional setting, with particular emphasis on the existence of deferred wages. Section 3 presents a model of stochastic and dynamic job destruction with deferred wages, while section 4 characterizes a simple version of the model, and derives a key empirical implication. Section 5 presents the data and sample design. Section 6 describes our empirical methodology and econometric issues. Section 7 presents the empirical results while section 8 concludes.

2 Deferred Wages in the Italian Institutional Setting

In the literature on employment protection legislation, statutory severance payments are defined as mandatory payments (monetary transfers) to which a worker is entitled in case he or she is dismissed without fault of his or her own. In practice, beyond statutory payments, collective bargaining can

³www.laboratoriorevelli.it/whip

and do set additional severance payments to which workers are entitled in case of redundancy.

As was pointed out by Brandolini and Torrini (2002), Italy does not feature any mandatory severance payment⁴. In the Italian institutional setting, employment restrictions take the form of an-in-kind protection that forces firms to rehire unlawfully dismissed workers (reinstatement clause). Individual dismissal for economic reasons is contemplated by the law and it can be carried out at no cost. Yet, dismissed individuals have the right to appeal employer initiated separations. Ultimately, a court ruling decides whether a specific dismissal is admissible. Previous research has focused on such restriction. Ichino et al. (2003) argue that judges are biased by labour market conditions in their court ruling. Garibaldi et al. (2004) study the effect of a size threshold in the reinstatement clause, and find that there is some evidence of an increase in employment persistence around the threshold, but the size of the effect is quantitatively small.

This paper studies a deferred wage scheme, called *Trattamento di Fine Rapporto* (Remuneration upon Separation, *TFR* in the rest of the paper). *TFR* is an amount of money to which the worker is entitled at the end of the employment relationship, regardless of the cause beyond the job separation (quit, layoff, retirement, even firing for cause). The *TFR* is technically defined as a fraction of the yearly wage that is paid to all employees (including those under probation) with a time delay. It is a sum of yearly provisions that are explicitly included in the firm's balance sheet (and so they are part of the labor cost for accounting purposes) and it is periodically re-valuated. For each year of service, a provision of two twenty-seventh (or $1/13.5$) of the yearly gross salary is included in the individual *TFR* account, and is re-valuated yearly according to the following coefficients: 1.5% fixed plus 75% of the CPI inflation of the previous year. In legal terms, the *TFR* is a credit from the employee to the firm, and it is guaranteed by the social security administration in case the firm goes bankrupt.

Under normal circumstances, the *TFR* is paid to the worker at the time of job separation. Nevertheless, workers have the right to withdraw up to 70% of the *TFR* in advance if several conditions hold. First and foremost, they must have at least 8 years of tenure; second, they have to use the advance payment only for health related expenses, for buying a house, or for specific periods of unpaid leave (e.g. training). Finally, the advance withdrawal is a right on the part of the worker as long as less than 10% of the eligible workers in the firm apply, and as long as it corresponds to less than 4% of the total workforce; this implies that firms below 25 employees may be exempted from this obligation⁵. Failure to meet these conditions requires that the advance

⁴While there are no statutory severance payments, collective agreements do set severance payments when the parties bargain over layoffs. No official survey is available on the size of these transfers or on the frequency of these events. In this work we focus on a different, universal, institution that is akin to a severance payment and whose effects may be added to bargained severance payments when they exist.

⁵The legislation says that early withdrawal is a worker's right as long as it is requested by "less than 10 percent

withdrawal is approved by the employer. In fact, workers do withdraw their individual accounts through employer approval. We observe about 2% of employees withdrawing before the 8th year of tenure, while this share increases to 5-6% afterward.

From the firm standpoint, the policy determined interest rate implies a subsidized financing of firm operation. Between 1988 and 1999, the years on which we will base our empirical analysis, the best rate available in the banking system was approximately ten percentage points higher than the financial cost of the *TFR*. This is clearly visible in Table 1, where we report the implicit ex-post interest rate of the *TFR* as well as the prime rate available in the banking system. From the worker standpoint, the outside riskless option is a financial investment in long term government bonds. As shown in Table 1, between 1988 and 1999 the average long term rate on government bonds is some 8 percent higher than the rate of the *TFR*. This suggests that workers face a financial opportunity cost induced by the institutional setting.

Table 1: Prime Rates and Policy Determined Interest Rate for Recapitalizing TFR stocks

| year | Policy \tilde{r}^a | Market r^b | T-Bill r_{tb}^c |
|------|----------------------|--------------|-------------------|
| 1988 | 5.596 | 13.00 | 10.6 |
| 1989 | 6.387 | 14.00 | 10.9 |
| 1990 | 6.28 | 13.00 | 12.8 |
| 1991 | 6.032 | 12.50 | 13.5 |
| 1992 | 5.068 | 16.25 | 13.3 |
| 1993 | 4.491 | 10.38 | 11.2 |
| 1994 | 4.542 | 9.38 | 10.4 |
| 1995 | 5.851 | 11.50 | 12.2 |
| 1996 | 3.422 | 10.75 | 9.4 |
| 1997 | 2.643 | 9 | 6.9 |
| 1998 | 2.626 | 7.875 | 4.9 |

^a Policy Determined Yearly Capitalization rate for *TFR*
^b Market Determined Annual Prime Rate
^c Long Term Annual Rate on Government Bonds
Source: Authors' calculation, and Datastream.

3 Job Destruction with Deferred Wages

This section develops a simple model of deferred wages. It highlights the links between job duration and an institution like *TFR*. In particular, it helps understanding why we can identify the effect of severance payments at given tenure, wages and other characteristics. We make several simplifying assumptions, whose bearings on the empirical analysis will be discussed in section 6.

that does not exceed 4% of the workforce". Indeed, it is not entirely clear whether early withdrawal after 8 years of tenure is a workers' right in firms with less than 25 employees. On the one hand, there is a high court ruling that interprets the 4% threshold literally, and argues that firms with less than 25 employees are not obliged to obey to the early withdrawal legislation. On the other hand, various judges argue that the early withdrawal request apply as soon as "the cumulative yearly value of eligible workers does not reach 1" This implies that a firm of 5 employees should agree to one advance withdrawal every 5 years.

We consider a firm-worker pair that is engaged in a multi-period relationship that lasts T periods and produces a value of p_t at time t . In each period p_t is drawn from a continuous distribution $F(x) = \text{prob}(p \leq x)$ with finite upper support p^{\max} . Firms and workers are risk neutral and discount the future at the constant interest rate r . We do not model job creation, and we normalize the outside option of the firm and the worker to zero. The key decision we are modelling is job destruction conditional on the state of demand p_t , and we give full authority to the firm in such dimension. There are two possible ways in which a relationship ends. First, the firm decides to stop production at time t . Second, the job survives up to time T . The most natural interpretation of T is that of the retirement age, but it is also consistent with a temporary job. In this parsimonious setup the worker is fairly passive, no quits are considered. We assume that the per period wage is exogenously fixed at w throughout the employment relationship; however, we discuss how the results would change if wages were endogenous and flexible (Appendix 9.1).

Our focus is on the deferred wage. We assume that in each period a fraction δ of the wage is postponed and paid at the end of the employment relationship, regardless of the reasons behind the job termination. Firms are obliged to accumulate the unpaid wages at rate \tilde{r} where \tilde{r} is exogenously set. For most of our purposes, we consider the case in which $r > \tilde{r}$, even though the model is perfectly viable for a different assumption. This is fully consistent with the data reported in Table 1.

3.1 Value Functions

In what follows we indicate with Π_t the expected present discounted value of a job at time t and with R_t the value of accumulated unpaid past wages. The timing of the decision is as follows. At the beginning of each period the firm decides whether continuing production is optimal. The firm decision to continue production is based on the realization of the productivity shock. If production takes place, part of the wages are paid and the relationship moves to the next period. If production is interrupted, the entire stock of TFR is paid at the end of the period. For example, the sequence of events for a job that lasts 2 periods is as follows. In the first period, production takes place, and the relationship moves to the second period where job destruction can take place. If it does, R_2 is paid to the worker; if it does not, production takes place again and the job terminates when it reaches its natural end at $T = 2$. In this case R_3 , the deferred wage, is paid to the worker.

We focus on $(\Pi_t|p_t)$ which is the present discounted value of a firm who has *just* decided to continue production at time t , and its expression reads

$$(\Pi_t|p_t) = \begin{cases} p_t - (1 - \delta)w_t + \frac{1}{1+r} \left\{ \int \text{Max}[\Pi_{t+1}(z); -R_{t+1}]dF(z) \right\} & t = 1 \dots (T - 1) \\ p_t - (1 - \delta)w_t - \frac{1}{1+r}R_{t+1} & t = T \end{cases} \quad (1)$$

where the firm maximization decision over the job continuation is done in every period but the very last one. In equation (1), the per period operational profits are $p_t - (1 - \delta)w_t$ while the continuation value depends on whether job destruction is optimal. Note that if the job is destroyed TFR is still to be paid. Note further that the deferred wage R_{t+1} is paid to the worker regardless of the reason behind the job termination, so that the payment R_{t+1} is not a traditional severance payment. The value of R_t evolves according to the following rule

$$R_{t+1} = (1 + \tilde{r})[R_t + \delta w_t] \quad t = 1 \dots T,$$

with $R_1 = 0$.

Since the function Π_t is monotonic in p_t , the firm continuation decision is described by a reservation strategy p_t^* such that

$$(\Pi_t | p_t^*) = -R_t \quad (2)$$

whose value will be determined below. While the problem is dynamic and non-stationary, since it depends on the actual tenure of the worker, it can be easily solved by backward induction.

Definition 1 *The solution to the firm problem specified by equation 1 is a sequence of reservation productivity $\{p_t^*\}_{t=1}^T$ that solves equation 2.*

Note that at time $t = 1$, in light of the definition of deferred wages, the reservation productivity is akin to a job creation condition, and specifies that a firm is willing to open a job as long as its present discounted value is positive, or at least as large as the outside option of zero.

We also specify the worker's value function, even though his role is fairly passive. This is nevertheless useful for clearly specifying our assumptions. The present discounted value at time t of a job to a worker that is currently employed is

$$W_t = \begin{cases} (1 - \delta)w_t + \frac{1}{1+r} \{ [1 - F(p_{t+1}^*)] W_{t+1} + F(p_{t+1}^*) R_{t+1} \} & t = 1 \dots (T - 1) \\ (1 - \delta)w_t + \frac{1}{1+r} R_{t+1} & t = T \end{cases}$$

where it is clear that the worker takes as given the firm decision over the continuation policy, which is described by the probability of separation $F(p_t^*)$.

4 Characterizing The Model

4.1 One Period Model

To establish some very basic results we can start from a 1 period model ($T = 1$), where the only decision is a static reservation productivity, which is akin to a job creation decision. The firm profit

can be written as

$$\Pi|p = p - (1 - \delta)w - \frac{\delta w(1 + \tilde{r})}{1 + r}$$

Note that if wages are fully flexible and proportional to the realization of productivity, so that $w = \gamma p$ with $\gamma < 1$, firm profits are strictly proportional to productivity and can be written as

$$\begin{aligned} \Pi|p &= p\Delta; & w &= \gamma p \\ \Delta &= (1 - \gamma) + \delta\gamma \left[1 - \frac{1 + \tilde{r}}{1 + r} \right] > 0 \text{ if } \tilde{r} < r \end{aligned}$$

In the one period model with flexible wages the reservation productivity solves $(\Pi|p^*) = 0$, i.e. $p^* = 0$, and the firm job creation decision is $p > 0$ independent of deferred wages. This implies that mandatory deferred wages have no allocative impact. This feature should not be surprising, since it is just an application of the Lazear (1990) neutrality result: any mandatory transfer from the worker to the firm can be neutralized by wage flexibility, i.e. deferred wages are irrelevant.⁶

Remark 2 *With flexible wages, deferred wages are irrelevant*

To make the problem interesting, wages need to be not strictly proportional to productivity. An extreme assumption is assuming that wages are fixed and determined outside the relationship. As Devicienti et al. (2006) show, in the case of Italy such assumption turns out to be not so extreme. With constant wages, the reservation productivity reads

$$p^* - w = -\delta w \left[\frac{r - \tilde{r}}{1 + r} \right] \quad (3)$$

Equation 3 highlights an important result. When wages are fixed and $\tilde{r} < r$, the firm has an incentive to run a job with a marginal productivity that is lower than the wage. This mechanism represents the basic channel for the labour hoarding effect of deferred wages, a result that can be much more appreciated with a proper dynamic setting.

4.2 Dynamic Setting

In this section we analytically solve the model with $T = 2$, so that the solution of the model is described by the productivities p_1^* and p_2^* . It turns out that most of the properties that we want to emphasize apply to a two periods model.

⁶The neutrality result can also be seen in terms of total surplus of the job. The worker value function in the one period model is

$$W|p = (1 - \delta)w + \frac{\delta w(1 + \tilde{r})}{1 + r}$$

and the total surplus from the job is $S = W + \Pi = p$ which is positive as long as $p > 0$

We solve the model backward and obtain, sequentially, the reservation productivity at time $T = 2$ and $T = 1$. The equations for the two productivities solve $(\Pi_2|p_2^*) = -R_2$ where R_2 is the stock of TFR at the beginning of period 2 and $(\Pi_1|p_1^*) = -R_1$, where R_1 is the stock of TFR at the beginning of the relationship at $T = 1$. Their respective values solve

$$p_2^* - w = -\Gamma_2(r, \tilde{r}, \delta w) \quad (4)$$

$$p_1^* - w = -\frac{1}{1+r} \int_{p_2^*}^{p^{\max}} (z - w) dF(z) - \Gamma_1(r, \tilde{r}, \delta w, p_2^*) \quad (5)$$

where

$$\Gamma_2(\cdot) = \delta w \left[\frac{r - \tilde{r}}{1+r} \right] [1 + (1 + \tilde{r})] > 0$$

$$\Gamma_1(\cdot) = \delta w \left[F(p_2^*) \frac{r - \tilde{r}}{1+r} + (1 - F(p_2^*)) \frac{r^2 + 3(r - \tilde{r}) - \tilde{r}^2}{(1+r)^2} \right] > 0$$

The sign of the two Γ functions is positive as long as $r > \tilde{r}$, an assumption that we maintain throughout the work. The structure of the model, which is based on the assumption that the firm has an outside opportunity of 0, implies that the firm has an option value associated to hiring labour, even when deferred wages do not exist. To see this, one can solve the model when $\delta = 0$, and obtain

$$p_{2(\delta=0)}^* - w = 0$$

$$p_{1(\delta=0)}^* - w = -\frac{1}{1+r} \int_{p_{2(\delta=0)}^*}^{p^{\max}} (z - w) dF(z)$$

where it is clear that $p_1^*(\delta = 0) < w$ so that at time $t = 1$ the firm is willing to run a current loss in exchange of future profit gains. In the very last period, conversely, the problem is static and the firm hires only if labour productivity is as large as the wage, exactly as in a static textbook model of labor demand. As the next remark shows, we say that the firm hoards labour

Remark 3 *With an outside option of zero and constant wages, the firm hoards labour in every period but the very last one*

Over and beyond the fixed wage assumption, the firm propensity to hoard labour depends also on the structure of the productivity shock. In our current setting shocks are i.i.d. and the distribution faced by the firm is time invariant. If shocks were persistent and autoregressive, the firm propensity to hoard would fall, since a fast and large turnaround would be less likely. For analytical simplicity, we work only with i.i.d. shocks.

One of the main questions of this paper is whether deferred wages increase the firm propensity to hoard labour. To see this one needs to study the marginal impact of the deferred share δ on the reservation productivity. After simple algebra, the result reads

$$\frac{\partial p_2^*}{\partial \delta} = -w \left[\frac{r - \tilde{r}}{1 + r} \right] [1 + (1 + \tilde{r})] < 0. \quad (6)$$

$$\frac{\partial p_1^*}{\partial \delta} = -\frac{\Gamma_1(\cdot)}{\delta} < 0 \quad (7)$$

where the latter expression was obtained after a simple substitution from equation (4). Equations (6) and (7) are key equations, and show that the existence of *TFR* increases the firm propensity to hoard labour. Note also that the result requires not only $\delta > 0$, but also $r > \tilde{r}$. We can now state one of our key results.

Remark 4 *TFR increases the firm propensity to hoard labour. If $\tilde{r} < r$ and $\delta > 0$, the reservation productivity at time t falls with the size of *TFR*.*

An intuition of this result is as follows. *TFR* creates on the part of the firm an incentive to delay the time of separation, since the longer the average tenure, the lower the average labour cost. Thus, following a negative temporary shocks, the firm optimally holds on to current losses just to increase tenure and postpone the payment of the *TFR*.

While the presence of *TFR* unambiguously reduces the reservation productivity at time t , the dynamic evolution of the reservation productivity for given *TFR* is more complicated, since there are two labour hoarding effects that influence the value of p_t^* and p_{t+1}^* . One the one hand, the larger is t the larger is the accumulated stock of *TFR*, and the larger the firm incentive to hold on to the worker; this reduces the reservation productivity. On the other hand, the larger is t the lower is the future value of the firm rent. The net effect of these two forces is thus ambiguous, and one can not establish ex-ante the dynamic evolution of the productivity.

One can nevertheless establish that for a given forward looking time span, the labour hoarding effect of *TFR* increases with tenure. To see this we just consider two workers in the very last period of the relationship ($T = 2$ in the context of our model) but with different elapsed tenure. Equation 4 describes a firm employing a worker with elapsed tenure $\tau = 1$. Suppose the firm employs also a worker with elapsed tenure $\tau = 2$. Simple algebra shows that in this case⁷ $\hat{p}_2^* - w = -[\Gamma_2(r, \tilde{r}, \delta w) + \hat{\Gamma}_2(r, \tilde{r}, \delta w)]$, where $\hat{\Gamma}_2 = \delta w \frac{r - \tilde{r}}{1 + r} (1 + \tilde{r})^2 > 0$ if $\tilde{r} < r$ and $\delta > 0$, i.e. the wedge between productivity and wages increases with elapsed tenure. It is straightforward to generalize the result for a worker with elapsed tenure $\tau = K$. In this case $\bar{p}_2 - w = -\bar{\Gamma}(r, \tilde{r}, \delta w, K)$ where

⁷The reservation rule becomes $-\delta w(1 + \tilde{r}) - \delta w(1 + \tilde{r})^2 = \hat{p} - (1 - \delta)w - \frac{1}{1 + r}[\delta w(1 + \tilde{r}) + \delta w(1 + \tilde{r})^2 + \delta w(1 + \tilde{r})^3]$

$\bar{\Gamma}(r, \tilde{r}, \delta w, K) = \delta w \frac{r - \tilde{r}}{1 + \tilde{r}} \sum_{\tau=1}^{K+1} (1 + \tilde{r})^{\tau-1} > 0$ if $\tilde{r} < r$ and $\delta > 0$, increasing in K . As an example, $\bar{\Gamma}(r, \tilde{r}, \delta w, K)$ relative to w can be computed setting $\delta = 0.074$ (as stated by law), $r = 0.12$, $\tilde{r} = 0.05$, based on average values from table 1. $\bar{\Gamma}(r, \tilde{r}, \delta w, K)$ increases from 1% of w for $K = 1$ to 3.1% of w for $K = 5$ to 6.6% of w for $K = 10$; all but a negligible size.

Having characterized the firm reservation productivity, we can turn to the effects of *TFR* on firm profits. To study such effects, we can differentiate total expected profits at the beginning of the employment relationship with respect to δ . We can prove the following statement (see proof in Appendix 9.2):

Remark 5 *As long as $\delta > 0$ and $\tilde{r} < r$ TFR increases firms' present discounted profits*

The previous remark suggests that *TFR*, through its effect on profits, is likely to have an effect on job creation. Such effect would arise in an equilibrium model of the labour market, which is not within the scope of the current research. Models of this type have been extensively solved in the literature. See notably for a survey Bertola (1999), Ljungqvist (2001) for the effects of EPL in a variety of models and Garibaldi and Violante (2005) for a paper that distinguishes between various forms of EPL.

Turning to workers' behaviour, we spell out the key theoretical assumption on workers' behaviour:

We assume that for a given wage, workers enjoy the job security provisions determined by the TFR.

Technically, this assumption is linked to the effect of a larger *TFR* on workers' welfare. As we show in more details in appendix 9.3, the impact of larger *TFR* on workers welfare is made of two components, which have opposite effects, and we label them "income effect" and "labour hoarding effect". The *income effect* of TFR decreases workers' welfare. A larger *TFR* (i.e. an increase in the share δ of wages that are paid at the end of the relationship) induces a fall in workers' utility, since the worker is financing the firm at the interest rate \tilde{r} and reduces the present discounted value of its wage stream. The *labour hoarding effect of TFR* increases workers' welfare, since it grants more stability to the worker. Our assumption implies that for a given wage and firm profit maximization behaviour, workers are better off with *TFR*. This is consistent with the evidence that advance withdrawals are a rare event⁸.

⁸Average observed withdrawals among eligible workers (more than 8 years of tenure) working in large firms (above 25 employees) is about 6.5%; well below the maximum share stated by the law (10%).

4.3 The Effects of Advance Withdrawals

In this section, in order to obtain a key empirical implication, we look into another institutional dimension of the *TFR* legislation, namely the possibility that *TFR* is paid in advance to the worker. The last section explicitly assumed that, for given wages, workers enjoy the job security determined by the *TFR*. This assumption implies that in the baseline model the individual is better off with *TFR*, and if s/he had the option to withdraw the accumulated stock of *TFR* s/he would not exploit such option. Hence, since *TFR* is welfare improving at the given wage, an advance withdrawal has to be the result of an exogenous and random shock. We assume that such shock takes place at an exogenous rate μ . The interpretation we give to the shock μ is an i.i.d. liquidity shock at the individual level.

The advance withdrawal has an impact on firm behaviour. The existence of the withdrawing shock modifies the sequence of events within the model. At the beginning of the period a withdrawing shock is realized. Then the firm observes the realization of the productivity shock. If the worker did withdraw from his stock of *TFR* (at rate μ), the firm continues operation as long as $p_2 > \tilde{p}_2^*$. In case there is a firm initiated separation (i.e. p is below the threshold \tilde{p}_2^*), no *TFR* is due⁹. If the worker did not withdraw from his stock of *TFR* (at rate $1 - \mu$), the firm continues operation as long as $p > p_2^*$. In case there is a firm initiated separation (i.e. p is below the threshold p_2^*), the full *TFR* is due.

Clearly, the value of the continuation profit depends on whether an advance withdrawal shock hits the relationship. As the analysis in the appendix 9.4 makes clear, the value function for $\Pi_t(p_t)$ specifies a different reservation productivity conditional on the fact that a withdrawal will or will not take place. Characterizing the two reservation productivities p_2^* and \tilde{p}_2^* in the context of our two periods model, it is immediate to see that:

$$\tilde{p}_2^* > p_2^*$$

The previous expression implies that when a worker withdraws its *TFR*, the firm reduces its incentive to hold to marginal losses. The following empirical implication follows.

- **EMPIRICAL IMPLICATION.** *Other things equal, workers who withdraw their TFR, have a larger probability of experiencing a separation.*

Since in the model there are no quits, the separation is firm initiated. The intuition of the empirical implication is straightforward. Conditional on an advance withdrawal, the firm incentive

⁹For the sake of simplicity, we impose that 100% of the stock of *TFR* is withdrawn. The law imposes a maximum withdrawal of 70%; in the actual data the median share is about 60%.

to hold on to marginal losses disappears, and the labour hoarding dimension of TFR is no longer relevant. As a consequence, for given tenure, the firm will prefer to hold on to those individuals that did not experience an advance withdrawal.

The empirical implication highlighted above is the key prediction that we take to the data in the rest of the work. Before introducing the empirical analysis we present the dataset available to us, and some descriptive statistics

5 Data and descriptive statistics

We have access to a single-spell flow-sample of Italian employment spells. The data source is the Work Histories Italian Panel (WHIP) originated from the Italian Social Security Administration (INPS) archives and processed in a public-use file by LABORatorio Revelli. We have a 1:90 random sample of the entire archive of employees of private firms for the period 1985 to 1999. This is a longitudinal archive, as the same worker can be followed over time through possibly different employers. From such sample we select all employees that start a new job between February 1985 and December 1988¹⁰ and we follow those particular employment spells till they end or until December 1999. Ongoing spells at December 1999 are right censored¹¹. Notice that, in principle we can observe more than one employment spell for the same worker. However, as we need "long" spells to observe advance withdrawals, and as the observation period lasts between 11 and 14 years this possibility is actually unavailable. Finally, we select spells that last a minimum of four years, to be able to condition on lagged covariates (about 20,000 spells). This unavoidable selection excludes from the analysis very short employment spells, scarcely relevant for the purpose of the work. On the other hand, this selection turns out to be very important for the robustness checks on reverse causality discussed in section 6.3¹².

We observe the workers once a year, even though we know the exact month in which the hire and the (possible) separation took place. Starting time and censoring are clearly exogenous.

An important empirical remark concerns the cause of firm-worker separation. In the data there are three possible causes of separation: natural turnover (reaching the retirement age T ¹³), worker initiated separation (quit), firm initiated separation. We are interested in firm initiated separations

¹⁰For those working in January 1985 we cannot distinguish between new hirings and left censored ongoing employment spells.

¹¹For those working in December 1999 we cannot distinguish between separations occurring in December and right censored ongoing employment spells. This generates an - unavoidable - underestimation of the separation rate in 1999.

¹²The point is discussed further in section 6.3.

¹³There are no temporary contracts in our sample. In the years 1985-1989 temporary contracts could last a maximum of 2 years; hence imposing a minimum tenure of 4 years selects either permanent contract workers or workers hired with a temporary contract subsequently transformed into a permanent one.

only. However, our dataset reports separations but not their cause. Empirically, we distinguish between different causes of separation in the following way. First, we are very conservative vis-à-vis the quit for retirement, and we define "retirements" all separations that occur when the worker is 55 or over. More subtle and problematic is how to disentangle quits and layoffs. There is a well known grey area between worker initiated and firm initiated separations, both from a theoretical and from an empirical point of view. We address the issue using the observed duration of the subsequent non-employment spell; the idea being that on average a quit is "more likely" to lead to a new job "sooner" than a layoff. Hence, we label "firm initiated separation" those separations followed by a spell of non-employment of at least two months. Of course what we obtain is to increase the share of firm initiated separations, not to exclude quits altogether. In Appendix 9.5 we discuss the robustness of the econometric results to this definition of firm initiated separation.

We now turn to some descriptive statistics on *TFR* and advance withdrawals. The *WHIP* archive records the *TFR* stock at December of year t . It is then possible to compute the rate of accumulation of *TFR* with respect to the total annual gross wage. Figure 1 shows its distribution. There is some variability around the $1/13.5=0.074$ coefficient stated by the law. This is likely to reflect a number of unobservable events that may be the outcome of union - firms agreements with respect to the "relevant part of the wage" mentioned by the law on which the *TFR* yearly increase is computed. However, there is also some variability that cannot easily be explained (e.g. small positive and negative values, clearly visible in Figure 1), and we label it "measurement error". The existence of measurement error imposes a more precise definition of "withdrawal", since "negative changes in *TFR* stock" would overestimate the event of interest. In the rest of the work, we define a withdrawal as a negative change in the stock of *TFR* between $t - 1$ and t that (i) does not occur in the separation year (it would be the compulsory payment, not a withdrawal); (ii) involves at least 20% of the existing *TFR* stock; (iii) amounts at least at 500 euro in real terms. Sensitivity analysis confirms that the - necessarily arbitrary - choice of the above mentioned thresholds is non influential on the econometric estimates.

As we mention in Section 2, withdrawing from the stock of *TFR* is a legal right of the worker only under specific circumstances: eight years of tenure are elapsed, a few causes are met (health, housing, leave for training), firms are larger than 25 employees (even though among law scholars there is no agreement over whether such constraint binds) and less than 10 percent of the workforce applies for an advance withdrawal. As a results, it should not be surprising that the total number of advance withdrawals observed is not large, and it concerns less than 2 percent of the entire firm-worker pairs we observe (table 2). The share of advance withdrawals from the stock of *TFR* increases with tenure, and reaches some 6% of the workforce having 12 years of tenure. This very

low share is consistent with our hypothesis that workers value the labour hoarding effect of TFR and normally do not withdraw, even if they could. Information on the mean and median shares of advance withdrawals suggests that workers go for a large share of the *TFR* (about 60%). Such amount is quantitatively relevant. Since the *TFR* stock increases approximately by one month of salary for every year of tenure, withdrawing 60% of the stock at the 10th year of tenure is equivalent to receiving about half of the yearly salary.

Withdrawals before tenure 8 are by definition consensual. Since an early withdrawal tends to reduce profits, we believe that firms authorize such advance withdrawals only if business conditions are good (a sort of informal profit sharing agreement on this specific aspect only), while advance withdrawals are denied if business conditions are less favorable. While we do not observe firms' profit, Table 2 shows that up to tenure 8 advance withdrawals are constantly more likely in growing firms than in shrinking firms. The opposite becomes true after the 8th year of tenure, when withdrawals may be non consensual. Although not a formal test, this is consistent with our interpretation.

Table 2: Characteristics of advance withdrawals.

| elapsed tenure (years) | total number (a) | percentage of employees (b) | percentage in growing firms (c) | percentage in shrinking firms (d) | mean withdrawal (e) | median withdrawal (f) |
|------------------------|------------------|-----------------------------|---------------------------------|-----------------------------------|---------------------|-----------------------|
| 1 | 39 | 0.11 | 0.13 | 0.11 | 0.82 | 0.98 |
| 2 | 136 | 0.39 | 0.51 | 0.44 | 0.82 | 0.97 |
| 3 | 213 | 0.81 | 0.85 | 0.96 | 0.71 | 0.71 |
| 4 | 268 | 1.31 | 1.42 | 1.21 | 0.64 | 0.60 |
| 5 | 243 | 1.46 | 1.67 | 1.28 | 0.63 | 0.62 |
| 6 | 264 | 1.90 | 1.83 | 1.74 | 0.63 | 0.58 |
| 7 | 319 | 2.64 | 2.97 | 2.3 | 0.62 | 0.59 |
| 8 | 439 | 4.14 | 5.12 | 4.05 | 0.59 | 0.57 |
| 9 | 546 | 5.83 | 5.94 | 7.14 | 0.57 | 0.57 |
| 10 | 454 | 5.46 | 5.99 | 6.23 | 0.58 | 0.58 |
| 11 | 397 | 5.38 | 5.26 | 6.89 | 0.58 | 0.59 |
| 12 | 283 | 6.19 | 5.99 | 8.13 | 0.58 | 0.60 |
| 13 | 129 | 5.14 | 6.9 | 4.91 | 0.61 | 0.61 |
| 14 | 50 | 4.96 | | | 0.57 | 0.60 |
| Total | 3,780 | 1.59 | | | | |

(a) Absolute number of withdrawals observed in the sample
(b) Pct. of withdrawers over total no. of employees, for given tenure
(c) Pct. of withdrawers over total no. of employees in growing firms, for given tenure
(d) Pct. of withdrawers over total no. of employees in shrinking firms, for given tenure
(e) Mean computed on positive withdrawals only
(f) Median of positive withdrawals only

Table 3 compares the sample composition at tenure 10 for the entire sample of those who survived in the job for at least 10 years and for those individuals among them that experienced an advance withdrawal at tenure 9. Individuals that experience an advance withdrawal are more

likely to be males, in the mid thirties, non-manual workers, employed in larger firms (median firm size is 100 workers instead of about 50). We will discuss the link between firm size and withdrawals presenting the results.

Table 3: Sample composition at tenure=10 years.

| | | all workers at tenure=10 (a) | withdrawers at tenure=9 only (b) |
|--------------------------------------------------------------------------------------|--------|------------------------------------|----------------------------------------|
| daily real wage (c) | median | 67 | 74 |
| firm size | median | 48 | 103 |
| age at entry | median | 25 | 24 |
| manual | % | 0.569 | 0.533 |
| female | % | 0.333 | 0.267 |
| part time | % | 0.063 | 0.038 |
| geografic mover | % | 0.289 | 0.300 |
| industry= utilities | % | 0.023 | 0.026 |
| industry= chemical etc. | % | 0.073 | 0.097 |
| industry= mechanical | % | 0.239 | 0.239 |
| industry= textile, food etc. | % | 0.236 | 0.200 |
| industry= construction | % | 0.064 | 0.072 |
| industry= wholesale, retail | % | 0.180 | 0.127 |
| industry= transports | % | 0.061 | 0.139 |
| industry= banking, insurances | % | 0.124 | 0.101 |
| Observations | no. | 8320 | 546 |
| (a) Characteristics of all workers at 10 years of tenure | | | |
| (a) Characteristics of workers at 10 years of tenure that withdrew at tenure=9 years | | | |
| (c) Euro. (.000 lire in the estimates) | | | |

6 Empirical Analysis and Econometric Strategy

In this section we first present the basic empirical model; then we turn to other important empirical issues, namely unobserved heterogeneity and reverse causality.

6.1 The Basic Empirical Model

We highlight once more the key empirical implication that stemmed from our theoretical model: other things equal, workers that withdraw increase the probability of firm initiated separation. Empirically, we exploit the variability the withdrawal generates in the *TFR* stocks for given wage and tenure; such variability can be used to test our hypothesis. In words, we set a test to compare the probability of observing a firm initiated separation, at a given tenure and wage, for workers that did not withdraw their *TFR* stock and for workers who did. The natural setting for testing this empirical implication is the use of survivals models (Lancaster, 1990, Wooldridge, 2002). We estimate

$$h(t; D, X) = f(t; D, X)/[1 - F(t; D, X)] \quad (8)$$

where $h(t; D, X)$ is the hazard rate, i.e. the probability of ending the employment spell between t and $t + 1$ conditional on having "survived" on the job up to t , or the ratio between the density $f()$ and its cumulative function $1 - F()$, or survival function. It is expressed as a function of elapsed tenure t , of an eventual withdrawal D , and of a set of characteristics X including the wage.

Remark 6 *The goal of our empirical exercise is to test whether the impact of D on h is positive, i.e. whether the probability of a firm initiated separation between t and $t + 1$, conditional on an elapsed tenure t , is increased by withdrawing from TFR.*

Before discussing how to bring the theoretical model - and its simplifying assumptions - to the data, we present the econometric setup.

The job termination process is a continuous time process, even though we describe it in discrete time in our theoretical model for analytical simplicity. The empirical approach to the problem should take into account such property, even though we work with discrete time data. Our specification of the hazard does take care of this issue. Assuming that the continuous time process can be specified as a proportional hazard model, one has:

$$h(t; D, X) = \varkappa(D, X)h_0(t) \tag{9}$$

where \varkappa is a non-negative function of the covariates and $h_0(t)$ is the baseline hazard. The important assumption here is that the process is separable in X, D and t , i.e. that the baseline hazard is the same for all individuals and it shifts due to the effect of the covariates. This is a very convenient assumption because odds ratios are then constant for every X . If we assume - as it is standard in the literature - that $\varkappa(D, X)$ is an exponential function, the hazard reads

$$h(t; D, X) = \exp(X\alpha + \beta D)h_0(t)$$

The specification of $h_0(t)$ can be parametric or non parametric. The most flexible option is to use a set of dummies on each t ; we choose it to mitigate as much as possible the effect of unobserved heterogeneity on the estimates¹⁴. Notice that the coefficients α, β are semi-elasticities of the hazard with respect to the covariates.

Prentice and Gloeckler (1978) and Jenkins (1995) show that the discrete time counterpart of an underlying continuous time proportional hazard model is a complementary log-log function. This means that under the hypothesis that the true process in continuous time is the above one, and under the hypothesis that the econometrician observes the process only at discrete points in time, the resulting hazard is the following:

¹⁴More on this point below.

$$h(t; D, X) = 1 - \exp\{-\exp[h_0(t) + X\alpha + \beta D]\}$$

so that taking logs

$$\log[-\log(1 - h(t; D, X))] = h_0(t) + X\alpha + \beta D \quad (10)$$

In light of the proportional hazard specification of the underlying continuous process, obtaining instantaneous odds ratios is straightforward:

$$OR = \frac{h(t; D = d, X)}{h(t; D = 0, X)} = \exp[\beta(d - 0)]$$

where we compute the relative increase in the instantaneous hazard due to D being equal to d with respect to D being equal to the baseline case 0 (no withdrawal).

6.2 Unobserved Heterogeneity

In the model individuals are ex-ante identical. In the empirical analysis we include a set of controls X for observable worker and firm specific characteristics. However, equation (10) does not allow for unobserved heterogeneity. If unobserved heterogeneity is important, omitting it implies (i) a downward biased estimated duration dependence, (ii) a downward bias on the (absolute value of the) estimated parameters α and β , (iii) that the downward bias on the (absolute value of the) estimated parameters α and β increases with t ¹⁵. It must be stated up-front that the effects of unobserved heterogeneity are mitigated by the use of a flexible specification of the baseline hazard (see Jenkins, 2005). This is the reason why we choose a non parametric and totally flexible specification for $h_0(t)$, i.e. a set of dummies on each t . Furthermore, the effects of non including unobserved heterogeneity work against the empirical implication we aim at testing, so that if we obtain a significant and positive β we may have a lower bound of the true and larger effect of a withdrawal on the hazard rate¹⁶. Nevertheless, we estimate equation (10) also including unobserved heterogeneity, as a robustness check. We assume that unobserved heterogeneity is uncorrelated to observables and we estimate a random intercept model. Data limitation forbids to model general correlation between unobserved heterogeneity and the (time invariant as well as time varying)

¹⁵It can be shown that $[1 - F(t, D, X|v)] = S(t, D, X|v) = [S(t, D, X)]^v$, i.e. unobserved heterogeneity v , or frailty, scales the no-frailty component survival function (see Jenkins, 2005). It means that high- v individuals leave the job faster than low- v individuals, changing the sample composition over time and generating the effects mentioned in text when omitted from the model.

¹⁶An intuition could be the following. Workers with low (unobserved) propensity to change jobs will stay longer and therefore be more likely to withdraw, making it harder to obtain a positive effect of withdrawal on separation. We thank an anonymous referee for this comment.

covariates¹⁷. Defining v as a positive random variable with unit mean and finite variance σ^2 , distributed independently of t, X, D , equation (9) becomes

$$h(t; D, X, v) = \varkappa(D, X)h_0(t)v \quad (11)$$

and equation (10) becomes

$$\log[-\log(1 - h(t; D, X, v))] = h_0(t) + X\alpha + \beta D + v \quad (12)$$

To estimate equation (12) we must specify the distribution of v . Two options are available: a parametric and a non parametric one. In the first case the most common choice is the Gamma distribution, that can be easily integrated out of equation (12) providing a closed form of the unconditional hazard. The non parametric approach applies Heckman and Singer (1984) seminal work, fitting an arbitrary discrete distribution whose parameters are its mass points and the probabilities that individuals belong to one of them. To be more specific: suppose we have two mass points, i.e. two kinds of individuals; equation (12) becomes

$$\log[-\log(1 - h_1(t; D, X, v_1))] = h_0(t) + X\alpha + \beta D + v_1 \quad (13)$$

$$\log[-\log(1 - h_2(t; D, X, v_2))] = h_0(t) + X\alpha + \beta D + v_2 \quad (14)$$

and the contribution of each individual to the estimate will be the probability-weighted average of the two above hazards.

In terms of our model, a parametric Gamma distribution for v implies that every individual is different from the others depending on a random draw from that distribution. A non parametric discrete distribution implies that we have n groups of individuals, that every individual is identical to the other members of his group and different from members of other groups; it also implies that each person is allocated randomly to a group. We apply both assumptions; however the non parametric one is more flexible and it could have a more straightforward interpretation in terms of our model. If $n = 2$ we could label the groups "good health" and "poor health" people, or "shirkers" and "non shirkers", or "movers" and "stayers", and so on¹⁸.

¹⁷Horowitz and Lee (2004) provide a consistent estimator for survival models with unobserved heterogeneity possibly correlated to observables. Such estimator requires data with repeated spells for each individual. As already anticipated, such dataset is not currently available for the specific purpose of this analysis.

¹⁸The obvious limitation of both approaches is the randomness of v . As already stated, the only way of relaxing this assumption is to have access to a repeated spells sample.

6.3 Reverse Causality and Other Issues

The theoretical model excludes quits and imposes exogenous withdrawals. Our empirical strategy encompasses these two assumptions. In the real world, with endogenous quits, it may well be possible that individuals first decide to leave their current job, and subsequently withdraw from their accumulated stock of *TFR*. In other words, there may be a reverse causality effect, in the sense that an anticipated separation is followed by an advance withdrawal of the *TFR* stock and then by the separation itself. Clearly, individuals that are certain to leave their current job are only interested in the income effect of *TFR* and the job security effect vanishes. However, the possibility of reverse causality can be fairly easily ruled out by introducing lags between the withdrawal and the observed separation. We introduce up to three lags, regarding as totally unrealistic the hypothesis that workers can plan to quit 24 to 36 months in advance. This makes it necessary the selection of employment spells that last a minimum of four years. The drawback is limited. Short spells provide little information to our analysis: accumulated *TFR* stock is very low, advance withdrawals are extremely rare; hence, variability in *TFR* stock at given wage and tenure is very limited. There is a clear trade off between higher order lags and sample selection. Three lags proved to be the best compromise.

Finally, we assume that the per period wage is exogenously fixed at w throughout the employment relationship. This is certainly a reasonable assumption at the individual level in Italy, since most of the wages are negotiated at the industry or national level¹⁹; individual wage cuts below the collectively set wage are hardly possible²⁰.

7 Results

7.1 Unconditional Hazard

We begin by providing simple statistics on the hazard rate conditional on a withdrawal having or not occurred at $t - 3$. We choose the longest lag to avoid any reverse causality problem. These hazard rates are estimated non parametrically, using the Kaplan-Meier estimator of the probability of separation between t and $t + 1$ conditional on having been employed for t periods:

$$h(\mathbf{t}) = \frac{m(t)}{n(t)}$$

¹⁹Only about 10% of the total wage is bargained at the individual level, on average (Devicienti et al., 2006).

²⁰Notice that this does not deny the firm the possibility to consent to an advance withdrawal if the worker is hit by an adverse liquidity shock when s/he has no right to access the *TFR* fund. This consent makes the wage neither endogenous nor downward flexible.

where $m(t)$ is the number of spells terminated between t and $t + 1$ and $n(t)$ is the number of ongoing spells at t . We compute $h(t)$ separately for individuals that have withdrawn at $t - 3$ and for individuals that have not. Figure 2 plots $h(t)$ for $t = 5$ to 14. Remarkably, the hazard rate conditional on advance withdrawals having occurred lies clearly above the hazard rate for individuals that did not withdraw three years before. Such difference is very large up to tenure 11, while the gap closes for longer tenures. While this pattern may look puzzling, it is actually linked to firm size. Figure 3 shows that beyond tenure 7 the median firm size of workers that withdraw their stock of TFR is twice as large the average firm size at that particular tenure. In other words, the additional withdrawals after tenure 7 take place in large firms. Not surprisingly, if we replicate Figure 2 selecting only firms in the first quintile of the size distribution (by elapsed tenure) the puzzle disappears (Figure 4). This discussion suggests that firm size is an important determinant of the individual probability of withdrawing, as we further discuss in the multivariate analysis that follows (it will confirm that, excluding composition effects, a non consensual withdrawal is more likely to increase the hazard of firm initiated separation).

7.2 Empirical Specification

The regressor of interest can be specified in different ways. The straightforward one is to use the dummy D to signal that an advance withdrawal has taken place. However, there are better alternatives. First, the (absolute value of the) withdrawal rate if drawing occurred, zero otherwise, so that

$$TFRdrawR(t) = \begin{cases} \left| \frac{TFR_t - TFR_{t-1}}{TFR_{t-1}} \right| & \text{if } D(t) = 1 \\ 0 & \text{if } D(t) = 0 \end{cases} \quad (15)$$

Second, the log of the (absolute value of the) amount withdrawn if it occurred, zero otherwise, so that

$$TFRdrawA(t) = \begin{cases} \ln |TFR_t - TFR_{t-1}| & \text{if } D(t) = 1 \\ 0 & \text{if } D(t) = 0 \end{cases}$$

The advantage of using $TFRdrawR$ or $TFRdrawA$ instead of the indicator D is clear: we allow for larger withdrawals to have a more sizeable effect on the hazard. The advantage of using $TFRdrawR$ with respect to $TFRdrawA$ is in the ease to interpret its estimated coefficient. However, as we will see, all definitions provide coherent estimated results, as a further proof of robustness.

As anticipated, the advance withdrawal regressor always enters the analysis as a lagged value. This is coherent with our theoretical setting, where we show that the realization of a withdrawal leads to a subsequent increase in the probability of separation. The risk of reverse causality in the relationship between advance withdrawal and subsequent job separation is present when the

time lag between the two events is sufficiently short. Such risk, conversely, is not present when the time lag between the two events is large. So we include three lags in the regressor of interest, acknowledging that lag 1 might be affected by a reverse causality bias, but not higher order lags. As a robustness check we also estimate the model excluding lag 1 and lag 2 from the specification, providing further evidence on the direct effect of an advance withdrawal on the probability of separation.

Our specification includes time varying as well as time invariant controls, detailed in Table 4. $h_0(t)$ is specified as a set of dummy variables on t . Available time invariant characteristics are individual's gender, age at entry, occupation, type of contract (full or part time), whether s/he works in a province different from where s/he was born, and firm's industry. Among time varying covariates, observed once a year, we have daily average real wage (at 2003 prices, 000 lire) and firm size (the number of employees); both enter the specification in logs and lagged one period. Two covariates deserve a special remark. We have a dummy for growing and a dummy for shrinking firms, defined as positive (negative) yearly changes in the stock of employees (lagged one period). They aim at controlling for firms' profitability in general terms, as separation rates should naturally be different in expanding and contracting firms.

7.3 Bottom-line Results

The bottom-line results of our multivariate analysis (equation 10) are specified in table 4. Estimated standard errors are robust to heteroschedasticity. Duration dependence is negative, as expected, but not smooth, supporting the choice of a flexible specification. Lagged wages and firm size have both the expected negative impact on the hazard, while the manual occupation dummy, as well as the part time and female ones feature the expected positive sign. Age at entry has a positive impact on the hazard of separation up to 27 years, then negative. As expected, having reduced the workforce at the firm level in the past has a positive impact on the hazard while having increased it does not impact the hazard differently from having had a constant employment level.

The main empirical result concerns the sign and significance of the advance withdrawals, that are indicated in the table as "withdrawal at t-j". Remarkably, *withdrawal at t - 1* and *withdrawal at t - 3* are positive and significant, while it does not appear significant at lag 2²¹. This is so whatever the definition of withdrawal (share, log of absolute value, dummy on the event), as reported in table 5, specification A²². Quantitatively, the result is also sizeable. Taking at face value the coefficient

²¹We must remember that we are dealing with a few cases of advance withdrawal, so that statistical significance cannot always be achieved.

²²Only the coefficients of interest are reported. Coefficients on controls are unchanged with respect to table 4 and are not reported.

Table 4: Hazard rate estimate. Baseline specification.

| | β | s.e. | p value |
|------------------------------------------------------------------------------------------------------------------------------------|---------|-------|---------|
| withdrawal at t-1 | 0.332 | 0.101 | 0.001 |
| withdrawal at t-2 | -0.010 | 0.130 | 0.937 |
| withdrawal at t-3 | 0.274 | 0.131 | 0.036 |
| elapsed tenure=4 | | | |
| elapsed tenure=5 | -0.043 | 0.036 | 0.230 |
| elapsed tenure=6 | -0.231 | 0.040 | 0 |
| elapsed tenure=7 | -0.266 | 0.043 | 0 |
| elapsed tenure=8 | -0.306 | 0.046 | 0 |
| elapsed tenure=9 | -0.270 | 0.049 | 0 |
| elapsed tenure=10 | -0.185 | 0.050 | 0 |
| elapsed tenure=11 | -0.338 | 0.056 | 0 |
| elapsed tenure=12 | -0.237 | 0.067 | 0 |
| elapsed tenure=13 | -0.362 | 0.093 | 0 |
| elapsed tenure=14 | -0.830 | 0.179 | 0 |
| log real daily wage at t-1 | -0.423 | 0.042 | 0 |
| log firm size at t-1 | -0.129 | 0.008 | 0 |
| dummy growing firm at t-1 | 0.008 | 0.035 | 0.828 |
| dummy shrinking firm at t-1 | 0.346 | 0.031 | 0 |
| dummy part time | 0.347 | 0.049 | 0 |
| dummy female | 0.095 | 0.029 | 0.001 |
| age at entry | -0.219 | 0.010 | 0 |
| age at entry squared/100 | 0.388 | 0.016 | 0 |
| dummy manual occupation | 0.106 | 0.030 | 0.001 |
| dummy geographic mover | 0.021 | 0.027 | 0.441 |
| industry= utilities | -1.538 | 0.226 | 0 |
| industry= chemical etc. | -0.355 | 0.059 | 0 |
| industry= mechanical | -0.428 | 0.044 | 0 |
| industry= textile, food etc. | -0.323 | 0.043 | 0 |
| industry= construction | | | |
| industry= wholesale, retail | -0.254 | 0.044 | 0 |
| industry= transports | -0.588 | 0.075 | 0 |
| industry= banking, insurances | -0.498 | 0.057 | 0 |
| constant | 3.061 | 0.241 | 0 |
| Wald chi2 | 3113.09 | | |
| no. observations | 99765 | | |
| no. workers | 19554 | | |
| no. Separations | 7624 | | |
| no. Sep. and withdraw at t-1 | 226 | | |
| no. Sep. and withdraw at t-2 | 141 | | |
| no. Sep. and withdraw at t-3 | 130 | | |
| Probability of firm initiated separation. Complementary log log model. Robust s.e. Withdrawal is defined as withdrawal rate. | | | |

Table 5: Hazard rate estimate. Robustness to definition of withdrawal and to reverse causality.

| | Specification A | | | Specification B | | |
|------------------------------------------------------------------------------------|-----------------|-------|---------|-----------------|-------|---------|
| | β | s.e. | p value | β | s.e. | p value |
| 1. withdrawal rate | | | | | | |
| withdrawal at t-1 | 0.332 | 0.101 | 0.001 | | | |
| withdrawal at t-2 | -0.010 | 0.130 | 0.937 | | | |
| withdrawal at t-3 | 0.274 | 0.131 | 0.036 | 0.276 | 0.131 | 0.035 |
| 2. log absolute withdrawal | | | | | | |
| withdrawal at t-1 | 0.015 | 0.008 | 0.052 | | | |
| withdrawal at t-2 | -0.008 | 0.010 | 0.454 | | | |
| withdrawal at t-3 | 0.021 | 0.011 | 0.046 | 0.021 | 0.011 | 0.044 |
| 3. withdrawal dummy | | | | | | |
| withdrawal at t-1 | 0.128 | 0.069 | 0.063 | | | |
| withdrawal at t-2 | -0.067 | 0.086 | 0.438 | | | |
| withdrawal at t-3 | 0.176 | 0.090 | 0.051 | 0.179 | 0.090 | 0.048 |
| Probability of firm initiated separation. | | | | | | |
| Complementary log log model. Robust s.e. | | | | | | |
| Coefficients of interest only. Controls as in baseline specification, not reported | | | | | | |
| Specification A - baseline specification | | | | | | |
| Specification B - robustness to reverse causality (lag 3 only) | | | | | | |

on lag 3 in the first model (share of withdrawal), in Figure 5 we draw the instantaneous odds ratio of withdrawing a given share of *TFR* relative to not withdrawing, along with the 95% confidence interval. The implied increase in the instantaneous hazard is not negligible. An individual that withdraws 60% of the fund increases the instantaneous hazard rate by some 17-18% three years later. In other words, an individual with at least 10 years of tenure that withdraws the *TFR* fund increases his hazard rate from 10% to about 12% three years later.

7.4 Extensions and Robustness

The first key robustness test refers to unobserved heterogeneity. Working on our bottom line specification, we introduce unobserved heterogeneity both Gamma distributed and non parametric (with 2 mass points). Table 6 shows that, although significant, unobserved heterogeneity has a negligible effect on the parameters of interest²³. Hence we are confident in presenting estimates that impose no unobserved heterogeneity.

Table 5 specification B presents the second key robustness test. In order to fully avoid the risk of reverse causality, we disregard all advance withdrawals that take place in the 24 months before the actual separation date, since such advance withdrawals may capture the reverse causality mechanism. An advance withdrawal that takes place 3 years before the separation date features a positive and significant coefficient.

We now turn to the effect of firm size on the hazard of job termination. As we pointed out in

²³It has a negligible effect also on the other controls, not reported.

Table 6: Hazard rate estimate. Robustness to unobserved heterogeneity.

| | β | s.e. | p value |
|------------------------------------------------------------------------------------|--------------|--------------|--------------|
| 1. unobserved heterogeneity: no | | | |
| withdrawal at t-1 | 0.3324 | 0.101 | 0.001 |
| withdrawal at t-2 | -0.0104 | 0.130 | 0.937 |
| withdrawal at t-3 | 0.2741 | 0.131 | 0.036 |
| 2. unobserved heterogeneity: gamma | | | |
| withdrawal at t-1 | 0.3317 | 0.100 | 0.001 |
| withdrawal at t-2 | -0.0111 | 0.128 | 0.931 |
| withdrawal at t-3 | 0.2744 | 0.133 | 0.039 |
| <i>Gamma var.</i> | <i>0.207</i> | <i>0.113</i> | <i>0.068</i> |
| <i>LR test of gamma var =0: chibar2(01)=</i> | <i>3.589</i> | | <i>0.029</i> |
| 3. unobserved heterogeneity: discrete, 2 m.p. | | | |
| withdrawal at t-1 | 0.3342 | 0.102 | 0.001 |
| withdrawal at t-2 | -0.0057 | 0.130 | 0.965 |
| withdrawal at t-3 | 0.2753 | 0.135 | 0.041 |
| <i>v1</i> | <i>0.000</i> | | |
| <i>v2</i> | <i>2.066</i> | <i>0.221</i> | <i>0.000</i> |
| <i>prob. Type 1</i> | <i>0.911</i> | <i>0.027</i> | <i>0.000</i> |
| <i>prob. Type 2</i> | <i>0.089</i> | <i>0.027</i> | <i>0.001</i> |
| Probability of firm initiated separation. | | | |
| Complementary log log model. Robust s.e. | | | |
| Coefficients of interest only. Controls as in baseline specification, not reported | | | |
| Withdrawal is defined as withdrawal rate. | | | |

section 2, there is a discontinuity over the firm size distribution in the application of the legislation on *TFR*, namely firms below 25 employees seem to be exempted from the obligation to allow withdrawals from the fund. This implies that in these firms withdrawals should be consensual, whatever the elapsed tenure, and hence they should be less harmful for the career of those who withdraw. As Table 2 confirms, the small number of observed withdrawals does not allow us to perform a proper Regression Discontinuity Design; however, we can interact the rate of *withdrawal at t-3*²⁴ with firm size at $t-3$ (above or below 25 employees). Notice that this does not mean that we have only consensual withdrawals in small firms and only non consensual withdrawals in larger firms. In the first case, the jurisprudence is not unanimous in interpreting the law with respect to firms below 25 employees²⁵; in the second case, consensual withdrawals can take place before the 8th year of tenure and even afterward (when parameters set by the law are exceeded). What we can achieve is to increase the share of non consensual withdrawals when focussing on firms above 25 employees. Table 7 reports the coefficients of interest²⁶. Although less precisely estimated, we learn that the average estimated coefficient of .27 increases to .34 among large firms and decreases to .22, not significant, among small firms. This is what we expected. This also sheds some light on

²⁴Changing the definition of the regressor does not change the results.

²⁵See footnote 4.

²⁶Again, coefficients on controls are unchanged and not reported.

Table 7: Hazard rate estimate. Interaction with firm size.

| | β | s.e. | p value |
|------------------------------------------------------------------------------------|---------|-------|---------|
| withdrawal at t-1 | 0.332 | 0.101 | 0.001 |
| withdrawal at t-2 | -0.010 | 0.130 | 0.936 |
| withdrawal at t-3 in firms below 25 employees | 0.222 | 0.174 | 0.201 |
| withdrawal at t-3 in firms above 25 employees | 0.341 | 0.196 | 0.082 |
| Probability of firm initiated separation. | | | |
| Complementary log log model. Robust s.e. | | | |
| Coefficients of interest only. Controls as in baseline specification, not reported | | | |
| Withdrawal is defined as withdrawal rate. | | | |

the pattern by size of the Kaplan-Meyer estimator of the unconditional hazard we discussed at the beginning of the section. Excluding composition effects confirms that a non consensual withdrawal is more likely to increase the hazard of firm initiated separation.

The legislation introduces also a discontinuity over the elapsed tenure dimension (before or after the 8th year). Again, a Regression Discontinuity Design is beyond the possibility of the available data. Related to this is the estimated pattern of the duration dependence, that seems to show a slowdown in the hazard at year 8 or 9 (Table 4). Abbring and Van Der Berg (2003) propose an estimator that would account for a possible interaction between D and t in the data; however, that estimator is identified only if the "no anticipation" assumption is acceptable. If reverse causality is an issue in our model, and we so believe, than the estimator is not identified.²⁷ Nevertheless, from a pure econometric point of view, unobserved heterogeneity, misspecification and heteroschedasticity are indistinguishable, and our very flexible baseline hazard is meant to capture not only unobserved heterogeneity but also eventual missing interactions between t and the covariates. In our baseline specification we estimate the residual average effect of a withdrawal on the hazard, getting possibly a lower bound estimate of the true effect.

Finally, in Appendix 9.5 we check the effect of our definition of firm initiated separation, as well as the effect of sample selection with respect to the inclusion or exclusion of workers separating because of retirement or quit.

Overall all our results point consistently in favour of the Empirical Implication of our theoretical model: other things equal, workers that withdraw from their stock of TFR increase the probability of a firm initiated separation.

8 Conclusions

This paper has studied the severance payment dimension of a mandatory deferred wage payment (TFR) of the Italian labour market. Theoretically, deferred wages increase labor hoarding from

²⁷We thank an anonymous referee for bringing this work to our attention.

the firm standpoint if two conditions hold: wages are rigid, and deferred wages are accumulated at interest rates that are below market rates. Indeed, if wages were fully flexible, the problem would not even arise, since deferred wages would be completely irrelevant from the allocative standpoint. The second condition ensures that the longer the tenure, the lower the average wage cost. As long as both conditions are satisfied, as it certainly seems to be the case for Italy, firms hoard labour in bad times, holding on to marginal loss as a way to increase the average duration of jobs and hence decrease average labour cost.

Empirically, the paper shows that individuals with larger severance payments (in the form of larger TFR credit) at given tenure and wage, have lower hazard rates of job termination. Specifically, we have used information on about 20,000 employment spells to test a key empirical implication: workers who withdraw their accumulated stock of unpaid wages in advance, have a subsequent increase in the probability of being fired. This prediction seems to be empirically robust and quantitatively sizeable. A withdrawal of 60% of the TFR stock increases the instantaneous hazard rate of job termination by some 20%.

Our results are also relevant in the policy debate, which has always analyzed the impact of deferred wage payments in terms of the social security system. To the best of our knowledge, the effect of *TFR* on labour market and labour mobility is new. Further, the social security reform approved by the Italian Parliament in 2006 that will go into full effect in 2008 aims at using the *TFR* to boost the second pillar of the social security system. Our results suggest that a shift of *TFR* funds into pension funds will increase labour turnover, so that implementing such reform would have a direct effects in the social security area, and an indirect effect on the labour market. For example, when the 2006 reform will be implemented, an individual with at least ten years of tenure will divert from the *TFR* to the pension fund the annual *TFR* quota, equivalent to withdrawing less than 10% of the fund; his hazard would increase from 10% to about 10.4% in the first year (figure 5). A significant but small effect, that will obviously increase as time goes by and *TFR* is "withdrawn" every year.

Appendix

I. The Solution with Flexible Wages

If wages were fully flexible, and the firm and the worker shared the total surplus from the job, mandatory deferred wages would not have any allocative impact. To see this simply define the surplus from the job as

$$S_t(p_t) = \Pi_t(p_t) + W_t$$

and assume that the worker gets a fraction β of the total surplus in each and every period. The total surplus at time t reads

$$S_t(p_t) = \begin{cases} p_t + \frac{1}{1+r} \left[\int_{p_{t+1}^*}^{p^{\max}} S_{t+1}(z) dF(z) \right] & t = 1, (T-1) \\ p_t & t = T \end{cases} \quad (16)$$

which is an expression that is independent of wages and deferred wages. This is not surprising, since the total surplus is independent of a transfer between the two parts. The optimal separation policy in this case would be $S_t(\hat{p}_t) = 0$. Proceeding backward it is clear that the reservation productivity in this case would be $\hat{p}_t = 0$ and that the wage would simply be

$$w(p_t) = (1 - \delta)\beta p_t$$

so that the marginal wage would be zero ($w(\hat{p}_t) = 0$). In words, deferred wages are irrelevant.

II. TFR increases firms' profits

To prove it, as long as $\delta > 0$ and $r > \tilde{r}$, we just need to show that the derivative with respect to δ of the firm present discounted value is positive. In our two periods specification the firm present discounted value reads

$$\begin{aligned} \Pi_1(p_1) &= p_1 - (1 - \delta)w + \frac{1}{1+r} \left\{ \int_{p_2^*}^{p^{\max}} \Pi_2(z) dF(z) - F(p_2^*)\delta w(1 + \tilde{r}) \right\} \\ \Pi_1(p_1) &= p_1 - (1 - \delta)w + \frac{1}{1+r} \left\{ \int_{p_2^*}^{p^{\max}} \left[z - (1 - \delta)w - \frac{R_3}{1+r} \right] dF(z) - F(p_2^*)\delta w(1 + \tilde{r}) \right\} \end{aligned}$$

$$\begin{aligned} \frac{\partial \Pi_1}{\partial \delta} &= w + \frac{1}{1+r} \left\{ \int_{p_2^*}^{p^{\max}} \frac{\partial \Pi_2(z)}{\partial \delta} dF(z) - F(p_2^*)w(1 + \tilde{r}) \right\} - \frac{f(p_2^*)}{1+r} \Pi_2(p_2^*) \frac{\partial p_2^*}{\partial \delta} - \frac{f(p_2^*)}{1+r} \delta w(1 + \tilde{r}) \frac{\partial p_2^*}{\partial \delta} \\ \frac{\partial \Pi_1}{\partial \delta} &= w + \frac{w}{1+r} \left[1 - \frac{(1 + \tilde{r})}{1+r} - \frac{(1 + \tilde{r})^2}{1+r} \right] (1 - F(p_2^*)) - \frac{F(p_2^*)w(1 + \tilde{r})}{1+r} - \frac{f(p_2^*)}{1+r} \frac{\partial p_2^*}{\partial \delta} [\Pi_2(p_2^*) + \delta w(1 + \tilde{r})] \end{aligned}$$

Since the last expression in the brackets is zero by virtue of the reservation productivity $\Pi_2(p_2^*) = -\delta w(1 + \tilde{r})$ the previous expression reads

$$\frac{\partial \Pi_1}{\partial \delta} = w \left\{ 1 - \left[\frac{(1 + \tilde{r})^2}{(1+r)^2} (1 - F(p_2^*)) + F(p_2^*) \frac{(1 + \tilde{r})}{(1+r)} \right] \right\} + \frac{w}{1+r} \left[1 - \frac{(1 + \tilde{r})}{(1+r)} \right] (1 - F(p_2^*))$$

which is positive since

$$\frac{(1 + \tilde{r})^2}{(1+r)^2} (1 - F(p_2^*)) + F(p_2^*) \frac{(1 + \tilde{r})}{1+r} < 1$$

To see the latter result simply recall that $\frac{(1 + \tilde{r})^2}{(1+r)^2} < \frac{(1 + \tilde{r})}{(1+r)} < 1$

III. The Impact of TFR on Workers' Welfare

In the two periods version of the model, the welfare of the worker depends stochastically on the probability of being employed at time $t = 2$. The welfare of the worker at the beginning of the relationship is

$$W_1 = (1-\delta)w + \frac{1}{1+r} \left\{ [1 - F(p_2^*)] \left\{ (1-\delta)w + \frac{1}{1+r} [(1+\tilde{r})\delta w + (1+\tilde{r})^2\delta w] \right\} + F(p_2^*)(1+\tilde{r})\delta w \right\}$$

where the worker enjoys the current wage for certainty, while the value of the relationship at time $t = 2$ depends on the probability that the worker is not fired, which happens with probability $1 - F(p_2^*)$ where p_2^* is determined by the firm continuation policy. To study the impact of *TFR* we need to study $\frac{\partial W_1}{\partial \delta}$ which can be written as

$$\frac{\partial W_1}{\partial \delta} = \frac{\partial W_1}{\partial \delta} \Big|_{p_2^* = \bar{p}_2^*} + \frac{\partial W_1}{\partial p_2^*} \frac{\partial p_2^*}{\partial \delta}$$

The formal value of the two derivatives is

$$\begin{aligned} \frac{\partial W_1}{\partial \delta} \Big|_{p_2^* = \bar{p}_2^*} &= -\frac{w(1 - F(p_2^*))}{(1+r)} \left[\frac{r - \tilde{r}}{1+r} \right] - w \left\{ 1 - \frac{(1+\tilde{r})}{(1+r)} \left[(1 - F(p_2^*)) \frac{(1+\tilde{r})}{(1+r)} + F(p_2^*) \right] \right\} < 0 \\ \frac{\partial W_1}{\partial p_2^*} \frac{\partial p_2^*}{\partial \delta} &= -\frac{f(p_2^*)p_2^*}{1+r} \frac{\partial p_2^*}{\partial \delta} > 0 \end{aligned}$$

The former is negative as long as

$$\begin{aligned} (1 - F(p_2^*)) \frac{(1+\tilde{r})}{(1+r)} + F(p_2^*) &< 1 \\ (1 - F(p_2^*))(1+\tilde{r}) + F(p_2^*)(1+r) &< 1+r \\ \tilde{r} &< r \end{aligned}$$

Our assumption on the welfare effect of *TFR* implies assuming

$$\frac{\partial W_1}{\partial \delta} > 0 \quad \Rightarrow \quad \frac{\partial W_1}{\partial p_2^*} \frac{\partial p_2^*}{\partial \delta} > \left| \frac{\partial W_1}{\partial \delta} \Big|_{p_2^* = \bar{p}_2^*} \right|$$

IV. Withdrawing Shocks

Formally, the existence of a withdrawing shock modifies the firm problem. Assuming that withdrawing takes place at rate μ , the firm problem reads

$$\Pi_t(p_t) = \begin{cases} p_t - (1-\delta)w + \frac{1}{1+r} \left\{ \begin{aligned} &+\mu[-F(\tilde{p}_{t+1}^*)R_{t+1} + \int_{\tilde{p}_{t+1}^*}^{p^{\max}} \Pi_{t+1}(z)dF(z)] \\ &+(1-\mu) \left[\int_{p_{t+1}^*}^{p^{\max}} \Pi_{t+1}(z)dF(z) - F(p_{t+1}^*)R_{t+1} \right] \end{aligned} \right\} & t = 1 \dots (T-1) \\ p_t - (1-\delta)w - \frac{1}{1+r} R_{t+1} & t = T \end{cases} \quad (17)$$

while the value of R_t evolves according to rule

$$R_{t+1} = \begin{cases} (1 + \tilde{r})\delta w & \text{if worker withdraws at } t \\ (1 + \tilde{r})[R_t + \delta w] & \text{if worker does not withdraw at } t \end{cases}$$

with $R_1 = 0$. Let us consider a firm at the beginning of the second period that is employing a worker who has just withdrawn the stock of TFR. For such firm, the continuation policy is described by the following reservation productivity \tilde{p}_2^*

$$0 = \tilde{p}_2^* - (1 - \delta)w - \frac{1}{1 + r}[\delta w(1 + \tilde{r})]$$

Conversely, when a firm is hiring a worker who has not withdrawn in the previous period, the continuation policy would be as in eq. (4)

$$-R_2 = p_2^* - (1 - \delta)w - \frac{1}{1 + r}[\delta w(1 + \tilde{r})^2 + \delta w(1 + \tilde{r})]$$

so that

$$\tilde{p}_2^* - p_2^* = \delta w(1 + \tilde{r}) \left(1 - \frac{1 + \tilde{r}}{1 + r}\right) > 0$$

i.e.

$$\tilde{p}_2^* > p_2^*$$

from which it follows that when a worker withdraws the *TFR*, the firm reduces its incentive to hold to marginal losses.

V. Firm initiated separations and sample selection: robustness check

Our separation indicator S is defined as follows. We let U be the length of the non employment spell that follows the separation. The indicator S is

$$S = \begin{cases} 0 & - & \text{if the spell is right censored;} \\ 1 & \text{firm initiated separation,} & \text{if } U \geq n \text{ and age at separation } < 55; \\ 2 & \text{quit,} & \text{if } U < n \text{ and age at separation } < 55; \\ 3 & \text{retirement,} & \text{if age at separation } \geq 55. \end{cases}$$

The cloglog model always tests $S = 0$ vs $S = 1$.

The first robustness check concerns n , the subsequent non employment spell length that defines quits. We modify it from 0 months (all separations are firm initiated, provided that the worker is younger than 55) to 4 months. We do not claim to disentangle quits from layoffs perfectly; what we obtain, as we increase n , is to increase the share of pure firm initiated separations included in

$S = 1$. This procedure let emerge more clearly the effect of the withdrawal. Table 8, sample 2, shows that while the coefficient of *withdrawal at $t - 1$* is almost unaffected by n , the coefficient of *withdrawal at $t - 3$* increases with n ²⁸. As expected, the labour hoarding effect emerges stronger in this case. $n = 2$ is the benchmark case, used in the text.

The second robustness check concerns the sample definition. Provided that we test $S = 0$ vs $S = 1$ we can either include or exclude from the sample individuals for which $S = 2$ or $S = 3$. Under the assumption of independence in competing risks, the destination-specific hazards (quit, layoff, retirement, censored spells) can be estimated separately (sample 1 in Table 8). This is exactly true in continuous time models, where in every instant only one exit can be taken. In our case of interval-censored data this is only approximately true, accuracy depending on how small destination specific hazards are (Jenkins, 2005). The alternative is selecting the sample to exclude individuals quitting ($S = 2$) or retiring ($S = 3$), and to estimate the parameters specific to layoffs ($S = 1$) (sample 2 and 3 in Table 8). The drawback in the first case is the mentioned approximation; in the second case it is the possibly endogenous sample selection introduced. Table 8 shows that all this is non influential for our results. The coefficients of interest are modified only marginally by the sample used and their significance never changes.

We choose sample 2 as the benchmark case (excluding only retirements from the sample). The reason for this choice is clear from Table 9. To estimate all destination-specific hazards simultaneously we pretend just for this exercise to work with truly discrete time data, and we assume a convenient multinomial logistic specification of the hazard (thus abandoning the assumption of proportional hazard). The results for $S = 1$ are coherent with those presented in section 7, as a further robustness. The results for $S = 2$ confirm that lag 1 withdrawal suffers from reverse causality, while lag 3 does not. The results for $S = 3$ point to the fact that retirement time is known well in advance, so that also lag 3 could be influenced by reverse causality. Hence, to be very conservative we exclude retirements from the sample and in text we always use sample 2.

²⁸Share of withdrawal. Log absolute withdrawal and dummy on withdrawal show exactly the same behaviour. Results not reported.

Table 8: Hazard rate estimate. Robustness over quit definition and sample selection.

| | sample 1 | | | sample 2 (base) | | | sample 3 | | |
|------------------------------------------------------------------------------------------------|----------|-------|---------|-----------------|-----------------|-----------------|----------|-------|---------|
| | β | s.e. | p value | β | s.e. | p value | β | s.e. | p value |
| U length ≥ 0 months | | | | | | | | | |
| withdrawal at t-1 | 0.2898 | 0.073 | 0 | 0.3285 | 0.075 | 0 | | | |
| withdrawal at t-2 | 0.0312 | 0.090 | 0.728 | 0.0841 | 0.093 | 0.366 | | | |
| withdrawal at t-3 | 0.1437 | 0.098 | 0.141 | 0.2093 | 0.101 | 0.039 | | | |
| | | | | | | | no quits | | |
| U length ≥ 1 month | | | | | | | | | |
| withdrawal at t-1 | 0.3418 | 0.082 | 0 | 0.3354 | 0.083 | 0 | 0.3637 | 0.083 | 0 |
| withdrawal at t-2 | 0.0107 | 0.105 | 0.919 | 0.0075 | 0.106 | 0.944 | 0.0323 | 0.105 | 0.759 |
| withdrawal at t-3 | 0.1648 | 0.113 | 0.145 | 0.1806 | 0.113 | 0.109 | 0.1807 | 0.113 | 0.109 |
| U length ≥ 2 months (base) | | | | | | | | | |
| withdrawal at t-1 | 0.3424 | 0.101 | 0.001 | 0.3324 | 0.101 | 0.001 | 0.3684 | 0.101 | 0 |
| withdrawal at t-2 | -0.0069 | 0.130 | 0.958 | -0.0104 | 0.130 | 0.937 | 0.0121 | 0.130 | 0.926 |
| withdrawal at t-3 | 0.2618 | 0.131 | 0.046 | 0.2741 | 0.131 | 0.036 | 0.2762 | 0.131 | 0.035 |
| U length ≥ 3 months | | | | | | | | | |
| withdrawal at t-1 | 0.3548 | 0.103 | 0.001 | 0.3433 | 0.104 | 0.001 | 0.3954 | 0.104 | 0 |
| withdrawal at t-2 | -0.0878 | 0.137 | 0.522 | -0.0924 | 0.138 | 0.503 | -0.0620 | 0.137 | 0.652 |
| withdrawal at t-3 | 0.2876 | 0.133 | 0.031 | 0.3002 | 0.133 | 0.024 | 0.3030 | 0.133 | 0.023 |
| U length ≥ 4 months | | | | | | | | | |
| withdrawal at t-1 | 0.3226 | 0.106 | 0.002 | 0.3115 | 0.107 | 0.004 | 0.3655 | 0.107 | 0.001 |
| withdrawal at t-2 | -0.0867 | 0.140 | 0.535 | -0.0911 | 0.140 | 0.516 | -0.0641 | 0.140 | 0.647 |
| withdrawal at t-3 | 0.3277 | 0.134 | 0.015 | 0.3408 | 0.134 | 0.011 | 0.3364 | 0.134 | 0.012 |
| <i>Definition of separation (S) and sample:</i> | | | | <i>sample 1</i> | <i>sample 2</i> | <i>sample 3</i> | | | |
| <i>S=0 right censored spell</i> | | | | <i>in</i> | <i>in</i> | <i>in</i> | | | |
| <i>S=1 firing i.e. U length $\geq n$ months and age at sep < 55</i> | | | | <i>in</i> | <i>in</i> | <i>in</i> | | | |
| <i>S=2 quit i.e. U length $< n$ months and age at sep < 55</i> | | | | <i>in</i> | <i>in</i> | <i>out</i> | | | |
| <i>S=3 retirement i.e. age at separation ≥ 55</i> | | | | <i>in</i> | <i>out</i> | <i>out</i> | | | |
| <i>cloglog always tests separation=0 versus separation=1</i> | | | | | | | | | |
| Probability of firm initiated separation. Complementary log log model. Robust s.e. | | | | | | | | | |
| Coefficients of interest only. Controls as in baseline specification, not reported | | | | | | | | | |
| Withdrawal is defined as withdrawal rate. | | | | | | | | | |

Table 9: Multinomial logit estimate.

| | β | s.e. | p value |
|------------------------------------------------------------------------------------|---------|-------|---------|
| S=1 | | | |
| withdrawal at t-1 | 0.384 | 0.107 | 0 |
| withdrawal at t-2 | 0.014 | 0.136 | 0.918 |
| withdrawal at t-3 | 0.305 | 0.139 | 0.028 |
| S=2 | | | |
| withdrawal at t-1 | 0.316 | 0.117 | 0.007 |
| withdrawal at t-2 | 0.187 | 0.136 | 0.171 |
| withdrawal at t-3 | 0.122 | 0.165 | 0.459 |
| S=3 | | | |
| withdrawal at t-1 | -0.051 | 0.240 | 0.833 |
| withdrawal at t-2 | 0.209 | 0.246 | 0.395 |
| withdrawal at t-3 | 0.458 | 0.246 | 0.063 |
| Definition of separation (S): | | | |
| S=0 right censored spell; benchmark case | | | |
| S=1 firing | | | |
| S=2 quit | | | |
| S=3 retirement | | | |
| Multinomial logit model | | | |
| Coefficients of interest only. Controls as in baseline specification, not reported | | | |
| Withdrawal is defined as withdrawal rate. | | | |

References

- [1] Abbring J. and G. van den Berg (2003): "The Non Parametric Identification of Treatment Effects in Duration Models", *Econometrica*, 71, 1491-1517.
- [2] Bertola G. and S. Bentolila (1990) "Firing Costs and Labor Demand: How Bad is Euroclerosis?", *Review of Economic Studies*, 57, 381-402.
- [3] Bertola, G. (1999) "Microeconomic Perspective of Aggregate Labour Markets in Ashenfelter and Card (editors) *Handbook of Labor Economics*, Volume 3, Chapter 37
- [4] Brandolini, A. and R. Torrini (2002) "La Tutela Del Rapporto di Lavoro Subordinato Secondo Gli Indicatori dell'OCSE", Bank of Italy, mimeo.
- [5] Devicienti F., Maida A. and Pacelli L. (2006): "The Resurrection of the Italian Wage Curve", Laboratorio R. Revelli working paper no. 52.
- [6] Farber, H. (1999), "Mobility and Stability: The Dynamics of Job Change in Labor Markets", in Ashenfelter and Card (editors) *Handbook of Labor Economics*, Volume 3, Chapter 37, 2440-2480.
- [7] Garibaldi, P., Pacelli, L. and Borgarello, A. (2004) "Employment Protection and the Size of Firms" *Giornale Degli Economisti*, April.

- [8] Garibaldi, P. and Violante, G. (2005) "Firing Taxes and Severance Payments in Search Economies: A Comparison" *Economic Journal*, October.
- [9] Kugler, A. (1999) "The Impact of Firing Costs on Turnover and Unemployment: Evidence from the Colombian Labor Market Reform," *International Tax and Public Finance*, vol. 6(3), pp. 389-410.
- [10] Hopenhayn, H. and R. Rogerson (1993) "Job Turnover and Policy Evaluation: a General Equilibrium Analysis", *Journal of Political Economy* 101, 915-938.
- [11] Ichino, A, Polo, M. and Rettore, E. (2003) "Are Judges Biased by Labor Market Conditions?" *European Economic Review*.
- [12] Jenkins, S. P. (1995). "Easy estimation methods for discrete-time duration models". *Oxford Bulletin of Economics and Statistics* 57(1): 129-138.
- [13] Jenkins, S. P. (2005) *Survival Analysis*. Web mimeo
<http://www.iser.essex.ac.uk/teaching/degree/stephenj/ec968/pdfs/ec968lnotesv6.pdf>
- [14] Lancaster, 1990, *The Econometric Analysis of Transition Data*. Cambridge: Cambridge University Press.
- [15] Ljungqvist, L. (2001); "How Do Layoff Costs Affect Employment?", *Economic Journal*.
- [16] Lazear, E. (1990) "Job Security Provisions and Employment," *Quarterly Journal of Economics*, 105, 699-726.
- [17] Heckman J. J. and Singer B. (1984) "A method for minimizing the impact of distributional assumptions in econometric models of duration data"; *Econometrica*, 52, 271-320.
- [18] Horowitz J. and Lee S. (2004); "Semiparametric Estimation of a Panel Data Proportional Hazards Model with Fixed Effects", *Journal of Econometrics*, March.
- [19] Nickell, S. Labour Demand Under Uncertainty (1986) in in Ashenfelter and Layard (editors) *Handbook of Labor Economics*.
- [20] Prentice, R. and L. Gloeckler. (1978). "Regression analysis of grouped survival data with application to breast cancer data". *Biometrics* 34: 57-67.
- [21] Wooldridge, 2002 *Econometric Analysis of Cross Section and Panel Data*, MIT Press

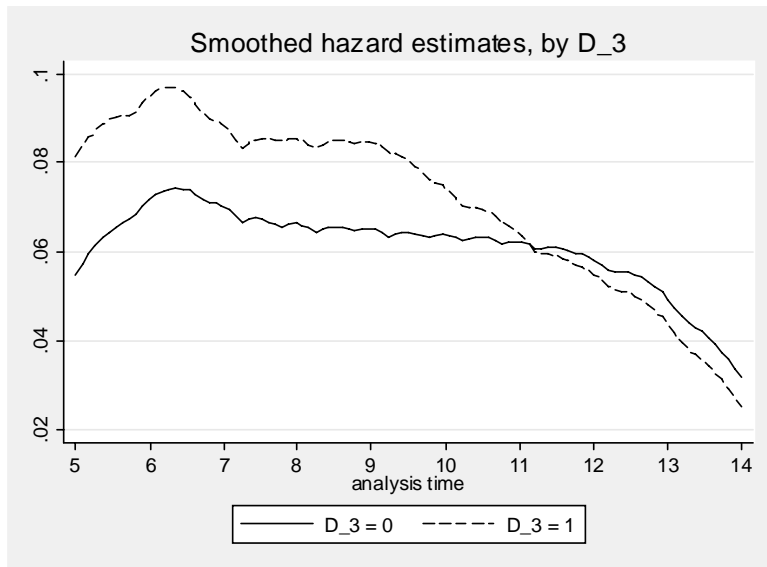


Figure 2: Kaplan-Meier hazard, withdrawers and non withdrawers at t-3

Figures

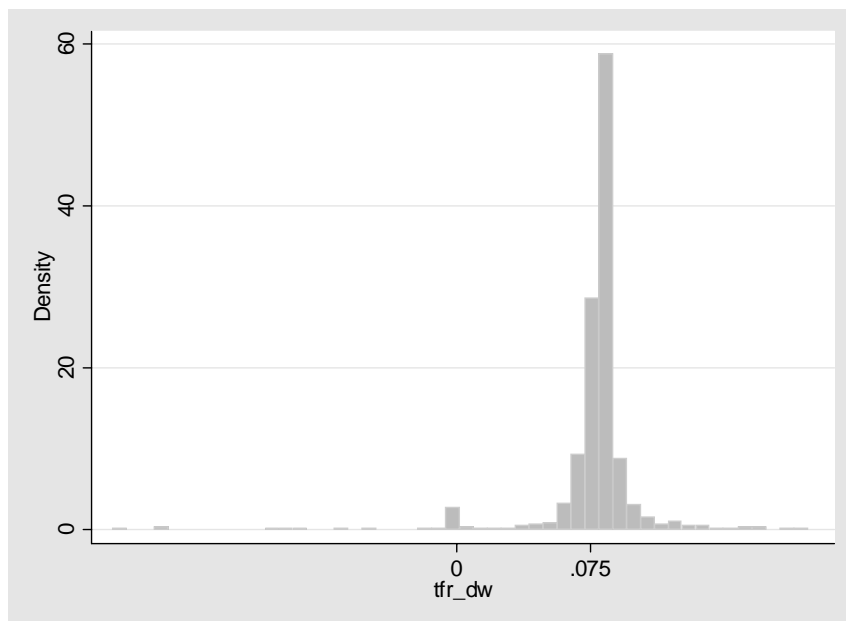


Figure 1: Distribution of TFR changes over annual wage, 1993

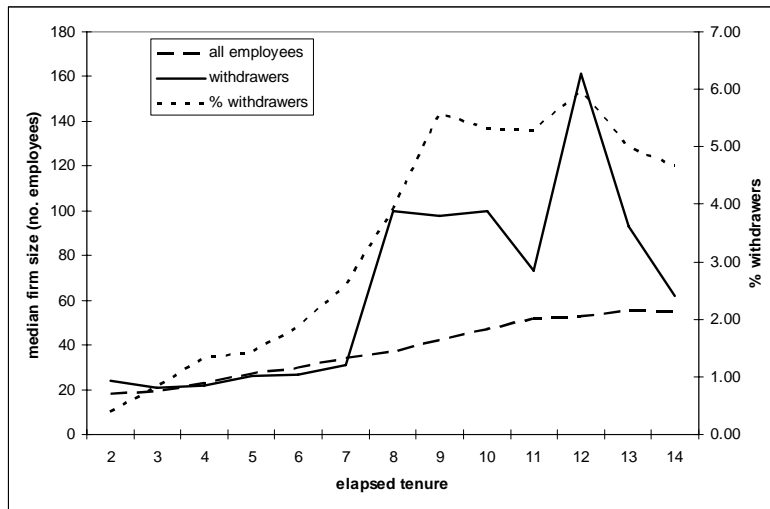


Figure 3: Median firm size and % withdrawers, by tenure

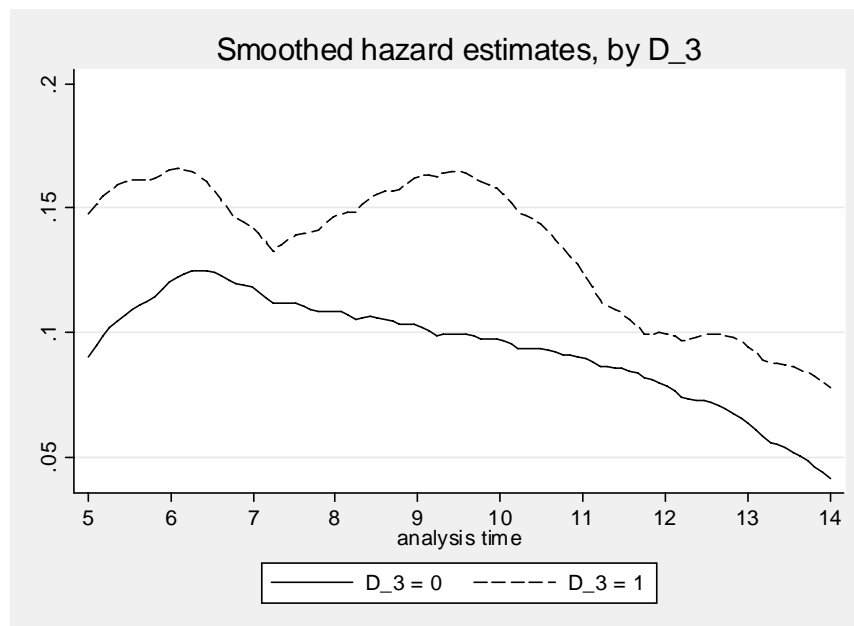


Figure 4: Kaplan-Meyer hazard, withdrawers and non withdrawers at t-3. First quintile of firm size distribution only

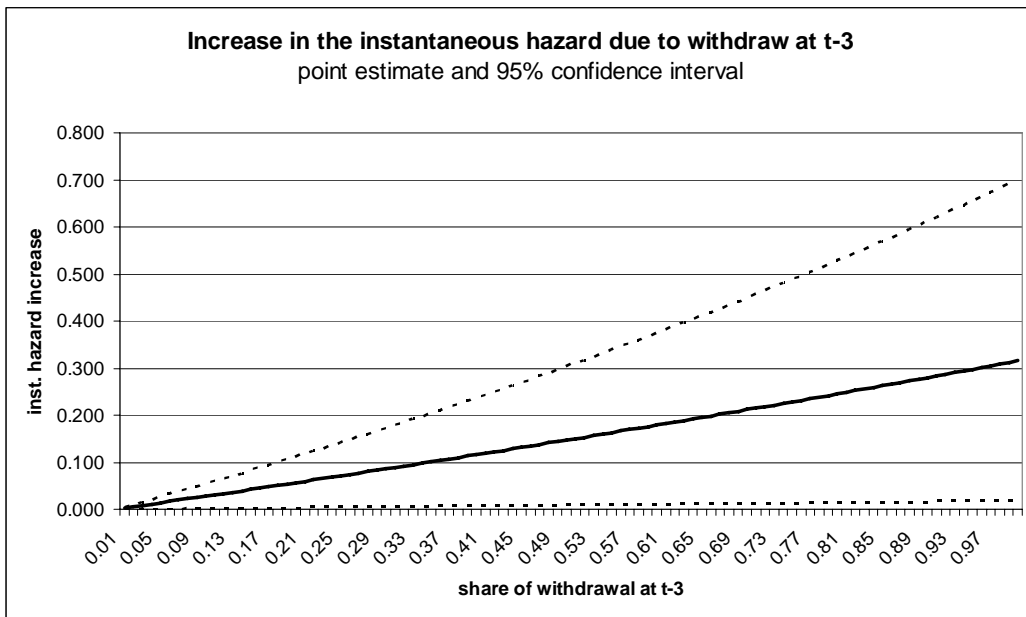


Figure 5: Odds ratios, bottom line specification

Fellows of the Fondazione Collegio Carlo Alberto Working Paper SERIES

1. Fabio Maccheroni, Massimo Marinacci and Aldo Rustichini. "Dynamic Variational Preferences", No.1, May 2006.
2. Maristella Botticini and Zvi Eckstein. "From Farmers to Merchants, Voluntary Conversions and Diaspora: A Human Capital Interpretation of Jewish History", No.2, May 2006.
3. Maristella Botticini and Zvi Eckstein. "Path Dependence and Occupations", No.3, May 2006.
4. Larry G. Epstein and Massimo Marinacci. "Coarse Contingencies", No.4, May 2006.
5. Massimo Marinacci and Luigi Montrucchio. "On Concavity and Supermodularity", No.5, May 2006.
6. Fabio Maccheroni, Massimo Marinacci, Aldo Rustichini and Marco Taboga. "Portfolio Selection with Monotone Mean-Variance Preferences ", No.6, May 2006.
7. Pietro Garibaldi. "Hiring Freeze and Bankruptcy in Unemployment Dynamics", No.7, May 2006.
8. Daniela Del Boca and Christopher J. Flinn. "Household Time Allocation and Models of Behavior: A Theory of Sorts", No.8, May 2006.
9. Luigi Montrucchio and Patrizia Semeraro. "Refinement Derivatives and Values of Games", No.9, May 2006.
10. Tito Boeri and Pietro Garibaldi. "Shadow Sorting", No.10, May 2006.
11. Diego García, Francesco Sangiorgi and Branko Urošević. "Overconfidence and Market Efficiency with Heterogeneous Agents", No.11, May 2006.
12. Fabio Maccheroni, Massimo Marinacci, Aldo Rustichini. "Ambiguity Aversion, Robustness, and the Variational Representation of Preferences ", No.12, June 2006.
13. Massimiliano Amarante, Fabio Maccheroni , Massimo Marinacci and Luigi Montrucchio. "Cores of Non-Atomic Market Games ", No.13, June 2006.
14. Daniela Del Boca and Christopher J. Flinn. "Modes of Spousal Interaction and the Labor Market Environment", No.14, June 2006.
15. Luigi Montrucchio and Marco Scarsini. "Large Newsvendor Games", No.15, June 2006
16. Andrea M. Buffa and Giovanna Nicodano. "Should Insider Trading be Prohibited When Share Repurchases are Allowed?", No.16, July 2006.

17. Peter Klibanoff , Massimo Marinacci and Sujoy Mukerji. "Recursive Smooth Ambiguity Preferences ", No.17, August 2006.
18. John Londregan and Andrea Vindigni. "Voting as a Credible Threat", No. 18, August 2006.
19. Larry G. Epstein and Massimo Marinacci. "Mutual Absolute Continuity of Multiple Priors ", No. 19, September 2006.
20. Daniela Del Boca and Robert M. Sauer. "Life Cycle Employment and Fertility Across Institutional Environments", No. 20, September 2006.
21. Matteo Triossi. "Reliability and Responsibility: A Theory of Endogenous Commitment", No. 21, September 2006.
22. Matteo Triossi and Antonio Romero-Medina. "Ramón y Cajal: Mediation and Meritocracy", No. 22, September 2006.
23. Matteo Triossi. "Application Costs in Sequential Admission Mechanisms", No. 23, September 2006.
24. Matteo Triossi and Luis Corchón. "Implementation with State Dependent Feasible Sets and Preferences: A Renegotiation Approach", No. 24, September 2006.
25. Giovanni Mastrobuoni. "The Social Security Earnings Test Removal. Money Saved or Money Spent by the Trust Fund?", No 25, September 2006.
26. Wioletta Dziuda and Giovanni Mastrobuoni. "The Euro Changeover and its Effects on Price Transparency and Inflation", No. 26, September 2006.
27. Peter Bossaerts, Serena Guarnaschelli, Paolo Ghirardato, and William Zame. "Ambiguity and Asset Prices: An Experimental Perspective", No. 27, September 2006.
28. Fabio C. Bagliano and Caludio Morana: "A New Approach to Factor Vector Autoregressive Estimation with an Application to Large-Scale Macroeconometric Modelling", No. 28, October 2006.
29. Elisa Luciano and Wim Schoutens: "A Multivariate Jump-Driven Financial Asset Model", No. 29, October 2006.
30. Elisa Luciano and Elena Vigna: "Non mean reverting affine processes for stochastic mortality", No. 30, October 2006.
31. Daniela Del Boca and Daniela Vuri: "The Mismatch between Employment and Child Care in Italy: the Impact of Rationing", No. 31, October 2006.

31. Daniela Del Boca and Daniela Vuri: "The Mismatch between Employment and Child Care in Italy: the Impact of Rationing", No. 31, October 2006.
32. Fabio C. Bagliano and Claudio Morana: "International Macroeconomic Dynamics: A Factor Vector Autoregressive Approach", No. 32, November 2006.
33. Giovanni Mastrobuoni: "Labor Supply Effects of the Recent Social Security Benefit Cuts: Empirical Estimates Using Cohort Discontinuities", No. 33, November 2006.
34. Daron Acemoglu, Davide Ticchi and Andrea Vindigni, "Emergence and Persistence of Inefficient States", No. 34, November 2006.
35. Kenn Ariga, Giorgio Brunello, Roki Iwahashi and Lorenzo Rocco, "On the Efficiency Costs of De-tracking Secondary Schools", No. 35, December 2006.
36. Giovanni Mastrobuoni and Matthew Weinberg, "Heterogeneity in Intra-Monthly Consumption Patterns, Self-Control, and Savings at Retirement", No. 36, February 2007.
37. Tito Boeri and Pietro Garibaldi, "Two Tier Reforms of Employment Protection: A Honeymoon Effect?", No. 37, February 2007.
38. Pietro Garibaldi, Francesco Giavazzi, Andrea Ichino ed Enrico Rettore, "College Cost and Time to Complete a Degree: Evidence from Tuition Discontinuities", No. 38, February 2007.
39. Pietro Garibaldi and Lia Pacelli, "Do Larger Severance Payments Increase Individual Job Duration?", No. 39, February 2007.

THE CARLO ALBERTO NOTEBOOKS