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Young-in Old-out: a new evaluation^{*}

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Abstract

In this paper we examine the consequences of a decrease in exits of older workers on youth employment, allowing for job-to-job flows. We distinguish between young workers with no experience and experienced workers and we consider jobs with different experience requirements. Using a theoretical matching model with heterogeneous jobs and workers, we point out that a fall in the exit rate of older workers increases job-to-job flows, which makes firms more reluctant to open jobs for young workers. Then, we provide some empirical evidence for the Italian case.

Keywords : exit age; youth employment; propensity score; matching model; on-the-job search

JEL Classification : J23; J62; J63

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1 Introduction

Over the last three decades, most European countries experienced an ageing of their population and a fall in the labor force participation of older workers. In the face of steadily increasing dependency rates, some countries raised their legal retirement age to balance the budget of their pay-as-you-go retirement schemes. However, the efficiency of such a policy option depends strongly on its impact on the hiring rate and employment among every age group of workers and particularly the young workers.

At first sight, we can think that an increase in legal retirement age, encouraging firms to maintain their ageing workforce, could slowdown the dynamics of new hirings, if firms were expected to squeeze out their older workers and to replace them by new young workers. However, this simplistic idea, that led to the creation of early retirement schemes, has been severely criticized over the past ten years. Following the idea of Bozio (2006) we may argue that young and older workers may be imperfectly substitutable in the labor market. Consequently, once an older worker retires, his job can not be filled by a young worker. We can discuss about what makes these two cohorts of workers imperfectly substitutable. Cadiou et al. (2002) highlight the fact that the low job finding rates among young workers and among older workers are not of the same nature. While the low hiring rate of senior workers is due to the fact that they are close to the retirement age¹, the low job finding rate for young workers stems from their lack of experience. Consequently, they can not fill a job requiring some professional experience to replace an older worker retired.

However, even though the two sorts of jobs for the young and the old generation of workers differ greatly in the set of skill and experience requirements, a reduction in exits of older workers after an increase in the retirement age may have a negative impact on the hiring rate of young workers through a vacancy-chain effect (Contini and Revelli, 1997). The idea is the following: if we consider within a local labor market², that a vacancy opens up when an older worker leaves his job, there is a non-negative probability that it will be filled by an employed job-seeker (attracted by better opportunities or higher wage), who accumulated a sufficient level of experience to replace the

¹See also Chéron et al. (2007) for this question

²A local labour market is a group of firms belonging to the same sector of activity, located in the same area and whose jobs require the same educational level

older worker and whose job requires less professional experience. Therefore, after this job-to job flow, a new vacancy opens up and may be filled by another young employed job-seeker, involving again job-to-job flows, that may result finally in hiring a young worker from out of the local labor market. So, holding job creation constant and considering a vacancy-chain pattern, retaining older workers in the labor force leads to fewer complex vacant jobs, i.e. vacant jobs requiring some professional experience, which reduces job to job flows and decreases the number of vacant jobs requiring no experience. As a result, job finding rates of young workers are negatively affected.

For this idea to be relevant, two conditions are needed. First, the economy has to be characterised by a large share of job-to-job transitions in the total number of separations. Second, job creation has to be uncorrelated with a variation in the exit rates of older workers. Regarding the first condition, recent empirical studies highlight the major role of job-to-job flows to explain the total labour market turnover. In the case of the US labor market, Nagypal (2003) reports that job to job transitions account for more than 50% of the total separations of workers with a college degree and for more than 30% of separations of workers without a college degree. Regarding the European case, Theodossiou and Zangelis (2007) report cross-country evidence on job to job transitions for 6 European countries and they find that job to job flows account for up to 17% of total separations in the different European economies considered. So it appears that workers in Europe are less mobile than in the US labor market. Nevertheless, in Italy, Contini and Rapiti (1994) show that job to job transitions are about 25% of total separations. The Italian labor market can therefore be characterized by a large share of job to job transitions.

The second condition requires job creation to be exogenous. As argued by Layard, et al. (1991), it seems unrealistic and they refer to this idea as a lump-of-labor fallacy. Indeed, job creation is expected to be affected by a change in the exit rate of older workers, so we can not view the number of jobs in the economy as a fixed stock. One first goal of this paper is to predict the effect of delaying retirement on youth employment using the vacancy-chain setting and allowing for endogenous job creation. Using a theoretical matching model "à la" Mortensen-Pissarides (1994), Chéron et al. (2007) already pointed out that there exists a positive relation between the distance from retirement age and the hiring rate. This result implies that postponing retirement has a positive horizon effect on the hiring rate of each group of

workers, including the young workers. Nevertheless, while it is clear that the distance from retirement strongly affects job finding rates of older workers, it is less straightforward that it plays a significative role in the determination of job opportunities for young workers.

Our theoretical contribution is to investigate the effect of a decrease in the exit rate of older workers on the hiring rate among young workers, using a matching model with heterogeneous jobs and workers³ and considering two types of workers, workers with experience and young workers with no experience and two types of jobs, jobs requiring experience and jobs requiring no experience. Through this theoretical model, we represent an overlapping labour markets economy in the sense of Sattinger (2006), where new young entrants into a local labor market can not be matched with complex jobs requiring experience. However, as we allow for job-to-job flows, some experienced workers may be mismatched with jobs requiring no experience. These workers are therefore looking for a complex job due to better wage offers. In this framework we show that a decrease in the exit rate of older workers leads to an intensification of job-to-job flows from simple to complex jobs and reduces the flow value of a simple job to an employer. Consequently, it reduces the creation of simple jobs and exerts a negative effect on the job finding rate of young workers with no experience.

Then, in the second part of the paper, we provide some evidence of negative effects of retaining older workers in the labor force on the hiring rate of young workers at a local labor market level, examining the Italian case given that in this country job to job flows account for a large share of total worker flows. While the relation between elderly labor force participation and youth employment has been widely studied at a country-level (Bozio et al., 2008, Jousten et al., 2009), too few empirical studies have been carried out in a microeconomic perspective. However, as worker and job flows are strongly affected by firm-specific variables (Davis and Haltiwanger, 1999), it is interesting to investigate the effect of a slowdown in exits of older workers on the hiring rate of young workers at the firm-level, controlling for all regressors that may affect job creation.

However, such a study raises an empirical issue due to the endogeneity of the variation in exit rates of older workers. To address this issue, Portugal et al. (2009) investigate the effect of a decrease in the exit rate of elderly,

³For similar specification see also Dolado et al., 2009, Gautier 2002, Mortensen and Pissarides, 1999, Albrecht and Vroman, 1999

due to an increase in the legal retirement age for women, on worker flows of different age groups, using detailed matched employer-employee Portuguese data. Their empirical methodology lies on a quasi-experiment given the fact that only women have been affected by the retirement reform. Using a difference-in-difference matching method (Heckman, 1997), they show that firms employing women treated by the reform hire one to two fewer workers for each senior retained after the increase in legal retirement age and this decrease in the hiring rate is particularly strong for younger workers.

Our paper is close to the spirit of their study. Indeed, our goal is to examine the effect of a fall in the exit rate of older workers on the hiring rate of young workers, using Italian matched employer-employee data from the Worker Histories Italian Panel and considering the period 1988-1989. Nevertheless, our paper differs from the study of Portugal et al. for two reasons: first, we do not consider firms but rather local labor markets, using a synthetic-firm approach to account for job-to-job flows, that represent the starting point of the vacancy-chain theory. Second we do not exploit a quasi-experiment. We examine the period 1988-1989 given the fact that after 1990, the Italian labor market has been severely affected by changes in employment legislation (Kugler and Pica, 2008), then it has been hit by a severe recession and after 1997, job creation has been strongly influenced by the introduction of temporary contracts and of other atypical contracts by the Treu reform, making the Italian labor market more flexible (Sciulli, 2006). So, we choose to study the period 1988-1989 to be able to isolate the effect on hirings of young workers due to a fall in exits of older workers from effects due to other institutional effects or to economic fluctuations. However, we are faced with a selection bias, in the sense that local labor markets with different exit rates of elderly may also differ in their characteristics likely to affect the accession rate of young workers.

In the empirical part of the paper, we proceed in two steps. In a first step, we view a fall in the exit rate of older workers as a binary treatment and we examine the average effect of the treatment on the treated, using a difference-in-difference estimator to provide some preliminary results that are not robust to the selection bias. In a second step, we view a fall in the exit rate of older workers as a continuous treatment⁴, and we control for the

⁴Actually, we consider the absolute value of the decrease in the exit rate of older workers. Consequently, an high value of the treatment means that the observation unit experienced a severe fall in the exit rate of older workers

selection bias given observables using propensity score matching methods, generalized in a setting with a continuous treatment (Hirano and Imbens, 2004). After correcting for the selection bias we find that the higher the fall in exits of older workers, the less hirings of young workers.

The remainder of the paper is structured in the following way: in section 2, we develop our theoretical model and we present our theoretical findings. Then, in section 3, we present our empirical study, describing the type of data used, explaining our methodology and presenting our empirical results. Section 4 concludes.

2 The theoretical model: A labor market with heterogeneous jobs and workers

Following the wide literature about skilled and unskilled workers, we consider a theoretical model with two sorts of workers and two kinds of jobs. But in the contrary of the standard literature, we distinguish workers according to their professional experience and not according to their educational level. To allow for job-to-job flows, we focus on a local labor market where all firms belong to the same sector of activity, are located in the same area and whose jobs require the same educational level. In our model, we consider the young workers, that we denote by ne as subscript, who entry into the labor market with no professional experience and the workers with experience, denoted by e as subscript. We assume that a number η_0 of young workers exogenously arrive in the labor market as unemployed. They can only apply for jobs requiring no experience. In the remainder of the paper, we denote this type of jobs by S (simple job) as upscript. In our model, once a young is hired, he accumulates professional experience and can reach a sufficient experience level at a Poisson arrival rate λ to apply for complex jobs, denoted by the upscript C . Workers have interest to apply for complex jobs to have better wage opportunities, given that in our model, the wage received by a worker occupying a simple job is assumed to be the minimum wage⁵.

⁵See also Belan et al. (2007) for a similar specification

Both types of jobs may be broken up in the case of an idiosyncratic shock, according to a Poisson process with an arrival rate δ . In that case, a worker who occupied the job becomes unemployed but keeps his experience level. Indeed, whatever the unemployment duration of a worker, his working experience remains the same over time. Furthermore, as we introduce on-the-job search, workers with experience may always have interest to search for a simple job, implying a cross-skill equilibrium contrarily to the paper of Albrecht and Vroman (2002)⁶.

Following the Behaghel's specification (2007), we consider a guillotine retirement age, introducing a finite horizon in the career of worker⁷. We consider that a worker can retire as soon as he accumulates a sufficient experience level to apply for complex jobs. Let η the Poisson arrival rate at which an experienced worker retires. The horizon H of an experienced worker is defined in the following way:

$$H = \int_0^{\infty} t\eta e^{-\eta t} dt = \frac{1}{\eta} \quad (1)$$

Consequently, a decrease in η lengthens the career of an experienced worker. In the remainder of the paper, we assimilate a decrease in η to an increase in the retirement age.

Our specification is twofold: first, it allows us to reproduce the life cycle of a worker, who arrives in the labor market with no experience and who has to accumulate a sufficient level of experience occupying simple jobs to apply for complex jobs. Second, we build an economy with overlapping labor markets (Sattinger, 2006) in which young workers can not be matched with jobs requiring experience. Through this theoretical framework, we introduce an imperfect substitution between a worker occupying a complex job who retires and a young worker. Our aim is to study the effect of a decrease in the exit rate of older workers, that is η , on hirings of young workers.

In our model we assume that search is undirected, so workers have perfect information about the type of new vacancies but employers can not target

⁶Contrarily to their paper, we define the skill level of a worker by his experience level

⁷The implications of a finite horizon in a matching model on job creation and job destruction have been already investigated by Cheron et al. (2006 and 2007)

their hirings to a special type of workers⁸. In our economy, vacancies of type S and the two types of workers (with or without experience) meet each other according to a matching function denoted by $h^S(u_{ne} + u_e, v^S)$, that determines the number of hirings for this kind of jobs as a function of the number u_i of unemployed of type i , with $i \in \{ne, e\}$, and of the number of vacancies v^S of type S . We assume that the matching function is increasing, concave and linear homogenous.

An employer opening a S-type vacancy may fill his job according to a Poisson process at an arrival rate $h^S(\frac{1}{\theta^S}, 1) = q(\theta^S)$. $q(\theta^S)$ is decreasing with the tightness on the simple job market, denoted by θ^S , equal to the number of S-type vacancies over the number of job-seekers. We deduce therefore that an unemployed worker (whatever his type) may be matched with a simple job according to a Poisson process at an arrival rate $p(\theta^S)$, such that $p(\theta^S) = \theta^S q(\theta^S)$. $p(\theta^S)$ is therefore increasing with θ^S .

In a similar way, vacancies of type C meet job seekers with experience according to a matching function denoted by $h^C(u_e + o_e, v^C)$, that determines the number of hirings for this kind of jobs as a function of the number u_e of unemployed workers with experience, of the number o_e of employed job seekers and of the number v^C of C-type vacancies. Consequently, a C-type vacancy may be matched with a job seeker according to a Poisson process at an arrival rate $h^C(\frac{1}{\theta^C}, 1) = q(\theta^C)$. This arrival rate is decreasing with the tightness in the complex jobs market θ^C , equal to the number of C-type vacancies over the number of job seekers with experience. We deduce therefore that an experienced job-seeker may find a complex job according to a Poisson process at an arrival rate $p(\theta^C) = \theta^C q(\theta^C)$ increasing with θ^C .

The figure 1 represents the worker flows in our model. The dashed arrow indicates job-to-job flows. Flow equations at steady-state allow us to determine the labor force structure at the local labor market level.

We find the six following steady-state flow equations:

$$(u_e + o_e)p(\theta^C) = E^C(\delta + \eta) \tag{2}$$

⁸However, in the case of complex jobs, employers are sure to be matched with only workers with experience, given that youngsters can not apply for this type of jobs.

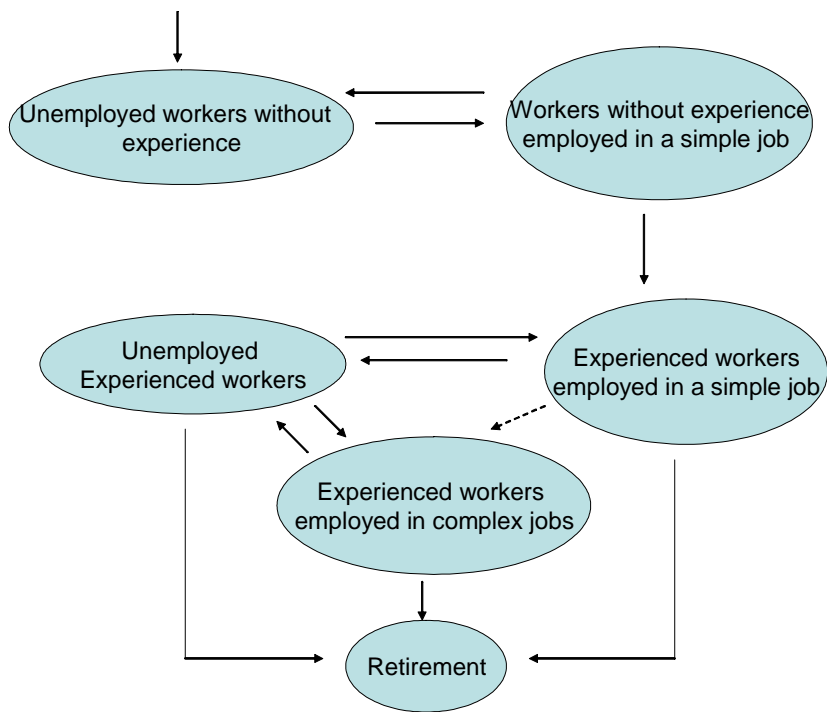


Figure 1: Worker flows in a local labor market

$$(E^C + o_e)\delta = u_e[\eta + p(\theta^S) + p(\theta^C)] \quad (3)$$

$$u_e p(\theta^S) + E_{ne}^S \lambda = o_e[\delta + \eta + p(\theta^C)] \quad (4)$$

$$u_{ne} p(\theta^S) = E_{ne}^S (\delta + \lambda) \quad (5)$$

$$E_{ne}^S \delta + \eta_0 = u_{ne} p(\theta^S) \quad (6)$$

$$\eta_0 = \eta(u_e + o_e + E_e^C) \quad (7)$$

where u_e and o_e represent respectively the number of experienced job-seekers unemployed and employed, E_{ne}^S and E_e refer respectively to the number of workers with no experience occupying a simple job and to the number of experienced workers occupying a complex job and η_0 represent the number of young workers who arrive in the labor market.

Combining equations (5) and (6) we obtain the expression of the number of unemployed workers with no experience:

$$u_{ne} = \frac{\eta_0(\delta + \lambda)}{\lambda p(\theta^S)} \quad (8)$$

Furthermore, combining the equations (3) and (7) we deduce the expression of unemployment for experienced workers:

$$u_e = \frac{\delta \eta_0}{\eta[\eta + p(\theta^S) + p(\theta^C) + \delta]} \quad (9)$$

The equation (8) is quite intuitive. The youth unemployment increases with the number of new entrants into the labor market and decreases with their probability to find a simple job denoted by $p(\theta^S)$. Our theoretical model allows us to endogenize this probability as a function of the tightness θ^S , depending on the simple-job creation behaviour of employers. As this behaviour may be strongly influenced by the horizon of workers, our approach allows us to investigate theoretically the channels through which a decrease in the exit rate η has an impact on hirings of young workers with no experience. In the next subsection, we represent the job creation decision of firms.

2.1 The firms' behaviour

In our model, for a sake of simplicity, each firm can not open more than one job. As long as the job is vacant, the firm pays a search cost c^i that depends on the type of job opened ($i \in \{S, C\}$). When a firm decides to open a S-type vacancy, it may be filled by a young worker with no experience according to a Poisson process at an arrival rate $\gamma q(\theta^S)$, where γ is the share of unemployed workers with no experience over the whole unemployment, namely $\gamma = u_{ne}/(u_{ne} + u_e)$. In a similar way, a S-type vacancy may be filled by an experienced job seeker according to a Poisson process at an arrival rate $(1 - \gamma)q(\theta^S)$. At the steady-state equilibrium, the flow value V^S to an employer from opening a S-type vacancy satisfies the following Bellman equation:

$$rV^S = -c^S + q(\theta^S)[\gamma(J_{ne}^S - V^S) + (1 - \gamma)(J_e^S - V^S)] \quad (10)$$

where J_i^S is the flow value to a firm of a simple job filled by a worker of type i , with $i \in \{ne, e\}$.

Furthermore, the flow value V^C to a firm which opens a C-type vacancy is defined by the following Bellman equation :

$$rV^C = -c^C + q(\theta^C)(J^C - V^C) \quad (11)$$

where J^C is the flow value to an employer of a filled complex job. We assume that firms are free to enter the market. Consequently, at steady-state equilibrium all profits from opening a new vacancy are exhausted for the two sorts of jobs, which implies the following free-entry condition :

$$V^C = V^S = 0 \quad (12)$$

In the submarket of simple jobs, a firm opening a S-type job vacancy may either recruit a unexperienced unemployed worker or an experienced one. Let J_i^S be the present flow value from hiring a worker of type i , with $i \in \{ne, e\}$. We assume that both types of workers have the same productivity on the same type of job. However, an experienced worker can quit his job to find a complex job at a Poisson arrival rate $p(\theta^C)$ and can also retire at a Poisson arrival rate η , contrarily to workers with no experience.

Borrowing the idea of Belan et al. (2007), we assume that each worker occupying a simple job receives a minimum wage w . Consequently, the present values of expected profit to a firm from a simple job filled either with a worker with no experience or with an experienced worker satisfy the following Bellman equations:

$$rJ_{ne}^S = y^S - w + \delta(V^S - J_{ne}^S) + \lambda(J_e^S - J_{ne}^S) \quad (13)$$

And :

$$rJ_e^S = y^S - w + [\delta + \eta + p(\theta^C)](V^S - J_e^S) \quad (14)$$

Furthermore, when an experienced job seeker finds a complex job, he starts producing an output $y^C > y^S$ and he receives a wage $w^C > w$. This part aims at providing a theoretical intuition, so we focus on the wage-posting equilibrium. His job may break up for two reasons. First, it is hit by an idiosyncratic shock at a Poisson arrival rate δ and the worker gets into unemployment. Second, the worker retires at an arrival rate η . So, the flow value to a firm from a filled complex job satisfies the following Bellman equation:

$$rJ^C = y^C - w^C + (\eta + \delta)(V^C - J^C) \quad (15)$$

2.2 Wage posting equilibrium

We determine the steady-state equilibrium. Combining the Bellman equation (10) and the free-entry condition we obtain the following expression:

$$\gamma J_{ne}^S + (1 - \gamma)J_e^S = \frac{c^S}{q(\theta^S)} \quad (16)$$

Substituting the equations (13) and (14) into the expression (16) we get the following relation between θ^S et θ^C :

$$\frac{\gamma}{(r + \delta + \lambda)} [(y^S - w) + \frac{\lambda[y^S - w]}{[r + \delta + p(\theta^C) + \eta]}] + \frac{(1 - \gamma)(y^S - w)}{[r + \delta + p(\theta^C) + \eta]} = \frac{c^S}{q(\theta^S)} \quad (17)$$

Furthermore, substituting the Bellman equation (15) in the expression (11) and using the free-entry condition, we obtain the following expression:

$$\frac{c^C}{q(\theta^C)} = \frac{(y^C - w^C)}{(r + \eta + \delta)} \quad (18)$$

To determine the effect of a decrease in η on job creation for each type of job, we use specific functional form to describe the two matching processes. We assume as in the standard literature that each matching function is Cobb-Douglas with an elasticity equal to $\alpha = 0,5$, such that:

$$h^S(u_{ne} + u_e, v^S) = (u_{ne} + u_e)^{\frac{1}{2}}(v^S)^{\frac{1}{2}} \quad (19)$$

$$h^C(u_e + o_e, v^C) = (u_e + o_e)^{\frac{1}{2}}(v^C)^{\frac{1}{2}} \quad (20)$$

Substituting (20) into (18), we obtain the following expression:

$$(\theta^C) = \left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right]^2 \quad (21)$$

Consequently, through our theoretical model, we can write the tightness for the submarket of complex jobs as a function of only exogenous parameters. This expression shows us clearly that a decrease in the exit rate among older workers denoted by η fosters the creation of complex jobs. Our result is in accordance with findings of Chéron et al. (2007) that highlight a positive horizon effect of an increase in the retirement age on job creation. However, in this paper, we investigate the effect of a decrease in the exit rate among older workers on hirings of young workers. We mention previously that young workers can not apply for complex jobs so we have to determine the effect of a decrease in η on the creation of simple jobs, that is on θ^S .

Substituting (21) into (17), we find that at steady-state equilibrium, θ^S satisfies the following equation:

$$\begin{aligned} L(\theta^S, \eta) &= \frac{\gamma}{(r + \delta + \lambda)} [(y^S - w) + \frac{\lambda[y^S - w]}{[r + \delta + \left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta]}] + \frac{(1 - \gamma)(y^S - w)}{[r + \delta + \left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta]} \\ - \frac{c^S}{q(\theta^S)} &= 0 \end{aligned} \quad (22)$$

In the wage posting equilibrium, we find two main offsetting effects of a decrease in the exit rate among older workers denoted by η on the tightness for the submarket of simple jobs θ^S . On the one hand, a decrease in η raises the horizon of experienced workers and therefore increases the average duration of simple jobs, which encourages employers to open more simple jobs. On the other hand, a decrease in η has a negative effect on θ^S through job-to-job flows. This is the key effect that we highlight in our model. Indeed, given that the creation of complex jobs is fostered by a decrease in η , the probability for an experimented worker occupying a simple job to be matched with a complex job is higher. Consequently, the average duration of simple jobs decreases and firms are more reluctant to open such jobs, reducing hiring of young workers.

Nevertheless, recall that the share of unemployed workers without experience denoted by γ is given by our flow equations at steady-state and therefore it is a function of both θ^S and θ^C . Indeed, combining equations (8) and (9), we write γ in the following way:

$$\gamma = \frac{(\delta + \lambda)\eta[\eta + p(\theta^S) + p(\theta^C) + \delta]}{\eta[\eta + p(\theta^S) + p(\theta^C) + \delta](\delta + \lambda) + \delta\lambda p(\theta^S)}$$

We see that the expression of γ is quite complicated and depends both on θ^S and θ^C . So it is not so straightforward to examine the sign of the derivative of θ^S with respect to η .

Proposition 1 *In the wage posting equilibrium, a decrease in η reduces θ^S , i.e. $\frac{\partial \theta^S}{\partial \eta} > 0$ only if $y^C - w^C > c^C(r + \delta + \eta)^2$*

Proof. See Appendix ■

We have shown theoretically that a decrease in the exit rate among older workers within a local labor market may reduce the tightness θ^S and therefore it has a negative effect on the hiring rate $p(\theta^S)$ of young unemployed workers, provided that the flow value to an employer of a filled complex job is sufficiently high. Thus, when this condition holds, the horizon effect is offset by the indirect effect due to job-to-job flows. Given the expression of u_{ne} we deduce that a fall in the exit rate η raises unemployment among young workers with no experience.

3 The empirical study

3.1 The data

For our empirical study, we use matched employer-employee data from the Worker Histories Italian Panel (WHIP) a database built by the University of Turin in cooperation with the Italian Social Security Administration (INPS). The original data set collects social security forms of a 1/90 random sample of employees every year. It provides accurate information about the employment spells of each worker. The data also includes also longitudinal records for firms employing the randomly selected workers in the sample. However, the data from the WHIP does not provide information about all individuals who work in a firm, so we can not observe the distribution of individual characteristics for each firm of the sample. To adress this issue, we carry out our study at the local labor market level, given that each observation unit is made up of a group of firms belonging to the same sector of activity, located in the same area and whose jobs require the same educational level.

We can discuss our synthetic-firm approach. Indeed, it leads us to underestimate the variance of each firm characteristic between firms given that each observation unit is a group of firms. To test the relevance of this approach, Contini et al. (2008) investigate the wage dynamics in the province of Veneto, confronting the results obtained through the synthetic-firm method with those obtained through an analysis at the firm-level. They conclude that both approaches lead to similar results when the number of firms in each local labor market is quite small (between 50 and 95 firms per observation unit). Consequently, there is a trade-off when choosing the agregation level between having the finest grid and having a sufficient number of firms in each observation unit.

Furthermore, studying the mobility of workers at a synthetic-firm-level allows us to account for the job-to-job flows that account for the major part of the worker flows observed in Italy. Using data from the WHIP, Leombruni and Quaranta (2005) show that job-to-job flows in the Italian labor market are observed within a sector, and within a geographical area. Consequently, we choose to agregate our data by province and by sector of activity, and we end up with a sample of 517 synthetic firms. Studying the mobility of workers at a local labor market level, allows us to account for the vacancy-

chain pattern and we can expect that exits of older workers may imply hirings of young workers, holding job creation constant.

Our empirical analysis aims at investigating the effect of a decrease in the exit rate among older workers on the hiring rate of the young, during the period 1988-1989. Here, the exit rate of older workers is defined as the number of exits of older workers observed in a year over the stock of older workers employed in the start of the year. In addition, the hiring rate of young workers is defined as the number of hirings of young workers observed in a year over the stock of young workers employed in the end of this year. We choose this period given that since 1990, employment protection legislation has been severely changed and has had a strong effect on the Italian labour market (Kugler and Pica, 2008). Moreover, since 1993, the Italian labour market has been strongly affected by the economic recession. At last, since 1997, the Treu reform liberalized the use of temporary contracts and introduced new atypical contracts, involving a strong effect on Italian worker flows (Sciulli, 2006).

Consequently, we study the period 1988-1989 to isolate the effect on youth employment due to a slowdown in the exit rate among older workers from other effects due to institutional changes or to economic fluctuations. However, even during the period 1988-1989, job creation among young workers has been increased by the introduction of CFL (Contratto Formazione e Lavoro)⁹ contracts (Tattara and Valentini, 2005). To control for this effect, we include in our set of regressors the share of CFL contracts observed in each local labour market.

As mentioned in the introduction, we carry out our empirical study in two steps. In the first step, we consider a binary variable equal to one if the observation unit experienced a reduction in the exit rate of workers aged 50 or more between 1988 and 1989 and we study the effect of this dummy on the hiring rate of young workers aged 30 or less in 1989. In a second step, we include in our sample only the treated observations and we examine to what extent a fall in the exit rate among older workers affect hirings of young workers. In this step, we view the absolute value of the decrease in the exit rate among older workers between 1988 and 1989 as the continuous treatment. For some descriptive statistics, we group our treated observations

⁹The CFL is a fixed-term contract with a maximum duration of 24 months not renewable. This contract was introduced in 1984 to ease the entrance of youngsters aged between 15 and 29 into the labour market, providing fiscal benefits to firms hiring young people.

into two categories: the group 1 is made up of synthetic firms, in which the value of the continuous treatment is lower than the median level, that is 13.5% and the group 2 is composed of synthetic firms, in which the value of the treatment exceeds 13.5%. We observe in the table 1 that the average hiring rate of young workers is higher in the group 1 (around 54,58%) than in the group 2 (around 51,57%). So a strong slowdown in the exit rate among older workers may reduce new hirings of young workers.

Table 1: Average hiring rate among workers aged 30 or less for different values of the continuous treatment.

	Group 1	Group 2	Total
Average hiring rate among young workers	54.58%	51.57%	53.07%
Number of observations	117	117	234

Note : The group 1 is composed of synthetic firms that experienced a decrease in the exit rate among older workers between 1988 and 1989 lower than 13.5% in absolute value.

The group 2 is composed of synthetic firms that experienced a decrease in the exit rate among older workers between 1988 and 1989 higher than 13.5% in absolute value.

Source : WHIP (waves 1988-1989)

To control for potential composition effects likely to influence the hiring rate of young workers or the exit rate of older workers observed within a local labor market, we use the set of information provided by our panel matched employer-employee data. We include first the age structure of each synthetic firms controlling for the share of younger workers and of older workers in each local labor market. We add also the share of women to test for gender-specific effects. We also introduce the average ratio between the average wage of young workers and the average wage of older workers in each synthetic firm. Indeed, this ratio may have a strong effect on the substitutability between young and older workers. We include also the share of CFL contracts signed in each local labor market, following empirical results supporting evidence of the strong positive impact of these contracts on youth employment (Tattara and Valentini, 2005).

Furthermore, empirical study of Contini et al. (2002) show that the age or the size of the firm affect the workers' mobility. Consequently, in our set of regressors, we include the share of young firms (less than 8 years) and the share of small firms (less than 9 workers) in each local labor market. To control for a potential effect due to economic trends in each synthetic firm, we add in our set of explanatory variables the unemployment rate and a dummy equal to one if the local labor market is declining¹⁰. We provide some descriptive statistics in the table 2.

Table 2: Descriptive statistics for the years 1988 and 1989

Variable	Year 1988	Year 1989
Hiring rate of young workers	50.37%	49.34%
Exit rate of older workers	21.62%	20.5%
Share of young workers	38.81%	38.91%
Share of older workers	14.89%	14.94%
Share of women	24.18%	24.6%
Share of CFL contracts	22.4%	21.35%
Share of young firms (less than 8 years)	33.96%	32.92%
Share of small firms (less than 9 employees)	40%	40.1%
Ratio between the average wage of youngsters and that of older workers	85.63%	86.97%
Unemployment rate	11.11%	10.84%
Dummy equal to one if the synthetic firms is declining	23.21%	26.31%
Number of observations	517	517

Source : WHIP (waves 1988 and 1989)

3.2 A difference-in-difference approach

¹⁰It means that in this synthetic firm, the number of jobs in the end of the year is lower than the number of jobs observed in the beginning of the year

3.2.1 Identification strategy

Let Y_{it}^D , the hiring rate of young workers observed at time t in a local labor market i in the state D , with $D \in \{0, 1\}$ and $t \in \{1988, 1989\}$. The binary variable D is equal to one if the synthetic firm experienced a decrease in the exit rate among older workers between 1988 and 1989. In that case, the synthetic firm is viewed as a treated observation. We have to determine the effect of the binary treatment on our dependent variable Y_{it} , in other words we have to estimate the difference $Y_{it}^1 - Y_{it}^0$. However, the estimation of this difference raises an issue of missing data. Indeed, at time t an observation i is either treated or not treated and we can not observe simultaneously the value of the dependent variable for one observation i in the two states $D = 0$ and $D = 1$.

To address this issue, we can determine the average treatment effect on the treated group using a difference-in-difference estimator denoted by α_{DD} defined in the following way:

$$\alpha_{DD} = \{E(Y_{it}|D = 1) - E(Y_{it}|D = 0)\} - \{E(Y_{it-1}|D = 1) - E(Y_{it-1}|D = 0)\}$$

The idea is to estimate the gap between the hiring rates of young workers observed respectively in treated and in control synthetic firms at time $t - 1$ and then to evaluate to what extent this gap has been affected at time t . However, this method holds only if the time invariance assumption is checked. This assumption implies that the hiring rate of young workers for the control group of observations would have evolved between $t - 1$ and t in the same fashion that the hiring rate for the treated group of observations, in the event that they had not been exposed to the treatment. The time invariance assumption may be formulated in the following way:

$$E(Y_{it}^0 - Y_{it-1}^0|D = 0) = E(Y_{it}^0 - Y_{it-1}^0|D = 1)$$

This assumption is particularly strong, given that treated synthetic firms may differ widely of not treated synthetic firms regarding some characteristics likely to affect hirings of young workers (a known issue that we refer to as the Ashenfelter's dip, 1978). To attenuate this bias, our strategy consists

in including a set of regressors when computing the difference-in-difference estimator, defining the following model:

$$Y_{it} = \beta X_{it} + \lambda D + \psi \tau_t + \alpha_{DD} D \tau_t + c_i + \varepsilon_{it} \text{ with } \varepsilon_{it} \sim N(0, \sigma^2) \quad (23)$$

where X_{it} is a vector composed of a set of regressors observed in a synthetic firm i at time t , c_i is an individual-specific effect, that allows us to control for unobserved heterogeneity and ε_{it} is the error term of our regression. In our model, we include a calendar effect adding a binary variable τ_t equal to one if the time t corresponds to the period when we consider our treatment, that is the year 1989. We also introduce a binary variable denoted by D equal to one if the synthetic firm i experienced a decrease in the exit rate among older workers between 1988 and 1989. Our aim is to estimate the parameter α_{DD} to determine the effect of the interaction between the variables D and τ_t on the hiring rate of young workers. In this regression, the parameter α_{DD} is the difference-in-difference estimator.

3.2.2 Results

We present our results in the table 3. We observe first that the Pearson's coefficient is equal to 0.3952. Consequently, our random-effect model is more relevant than a pooled-OLS regression, given that it allows to control for unobserved heterogeneity. We also see that the hiring rate of young workers at time t is strongly correlated with a lot of regressors like the job declining dummy, the share of young or small firms, the share of young workers or the share of CFL signed. It implies that the time invariance assumption is not a priori checked and that we need to control for the characteristics of each local labour market to examine the effect of the treatment on the treated group. We find that the difference-in-difference estimator α_{DD} is not significant.

We can discuss the relevance of our results given that treated synthetic firms are definitely not comparable with the control group. For instance, implementing a simple t-test of difference in means between these two groups for the job declining dummy, we find in the table 4 a t-stat equal to -3.3114 , quite higher in absolute value than 1.96. So the probability for the treated

Table 3: A difference-in-difference estimation of the effect of a decrease in the exit rate among older workers on the hiring rate among youngsters on the period 1988-1989

Variable	Coefficient	Standard error
Binary variable $\tau_t = 1$ if $t = 1989$	-0.013	0.013
Binary variable $D = 1$ if the synthetic firm is treated	0.024	0.016
Interaction term $D\tau_t$	0.007	0.019
Share of young workers	0.3634***	0.108
Share of older workers	0.266*	0.14
Share of women	-0.088*	0.046
Share of CFL contracts	0.273***	0.05
Share of young firms (less than 8 years)	0.173***	0.08
Share of small firms (less than 9 employees)	0,2***	0,058
Ratio between the average wage of young and that of older workers	0.07*	0.037
Unemployment rate	0.156	0.128
Dummy equal to one if the synthetic firms is declining	-0.089***	0.013
\bar{R}^2		0.3091
Number of observations		1034
Pearson's coefficient		0.3952

Note : Coefficients have been estimated through a random-effect model. Standard errors are corrected from the heteroscedasticity bias through the method of White.

Significativity thresholds : * : 10% ** : 5% *** : 1%

Source : WHIP (waves 1988 and 1989)

synthetic firms to experience job destruction in 1988 is quite higher than this probability for synthetic firms of the control group. It may be explained by the fact that declining sector tend to reduce their hirings, maintaining their ageing workforce. Consequently, we can not draw clear conclusions of a simple difference-in difference model due to the heterogeneity of the two groups regarding their covariates.

Table 4: Difference in means of the job declining dummy in 1988 between the treated and the control group

Group	Observations	Mean	Standard error
Control	283	0.177	0.023
Treated	234	0.299	0.029
difference		-0.0122	t-stat = -3.3114

Source : WHIP wave 1988

Furthermore, in this subsection, we viewed the decrease in the exit rate among older workers as a binary treatment. It could be interesting to estimate the effect of this decrease on the hirings of young workers, viewing this decrease as a continuous treatment. On the one hand, this strategy forces us to keep in our sample only the treated observations. On the other hand, this approach allows us to separate observations with a low level of treatment from those with a high level of treatment and to see how the hiring rate of young workers varies according to the level of the treatment.

Consequently, we use in the next subsection a propensity score matching procedure and we implement it to a continuous treatment, that is the decrease in the exit rate among older workers between 1988 and 1989 in absolute value.

3.3 Generalized propensity score matching method

3.3.1 Identification strategy

As mentioned previously, our strategy does not consist in regressing the hiring rate of young workers on absolute value of the decrease in the exit rate

among older workers. Indeed, even though we include in our model all the information provided by our database, we are in the face of a selection bias, in the sense that our explanatory variable of interest may be affected by a set of observables, likely to be correlated with the error term of our regression. To control for the selection bias through observables, we use a propensity score matching procedure generalized to a continuous treatment. As in the binary treatment context, propensity score methods in a setting with continuous treatment rely heavily on the key assumption that adjusting for pre-treatment differences solve the problem of drawing causal inference. Formally, we make the weak unconfoundedness assumption, introduced by Hirano and Imbens (2004), that requires that the treatment assignment mechanism is conditional independent of each potential outcome given the pre-treatment variables.

Let us consider a set of N synthetic firms characterized by a vector X_{it} of observables, by a number T_i equal to the absolute value of the decrease in the exit rate among older workers between 1988 and 1989 and by a number $Y_i(\gamma)$ equal to the hiring rate of young workers, associated with the treatment level $T_i = \gamma$. T_i is our continuous treatment and Y_i is our dependent variable. The unconfoundedness assumption may be formulated in the following way:

$$Y_i(\gamma) \perp T_i | X_i$$

Given unconfoundedness, we can apply matching methods based on the Generalized Propensity Score (GPS) with continuous treatment introduced by Imbens and Hirano (2004). The GPS is defined as the conditional density of the actual treatment given the observed covariates. Formally, let $r(\gamma, x) = f_{T|X}(\gamma|x)$ be the conditional density of the treatment given the covariates. Then, the GPS is $R_i = r(\gamma_i, X_i)$. The GPS is a balancing score, that is, within strata with the same value of $r(\gamma, x)$, the probability that $T = t$ does not depend on the value of X . In combination with the weak unconfoundedness assumption, this balancing property implies that:

$$f_T(\gamma | r(\gamma, X_i), Y_i(\gamma)) = f_T(\gamma | r(\gamma, X_i))$$

As a result, the GPS can be used to eliminate any biases associated with differences in the covariates. Formally, if the assignment to the treatment is

weakly unconfounded, given pre-treatment variables X_i , then we can define in a second step a specification regressing the conditional mean of our dependent variable on the value of the treatment T_i and on the value of the generalized propensity score R_i . In our analysis, we choose a simple linear model defined in the following way:

$$E(Y|T = \gamma, R = r) = b_0 + b_1T + b_2R + b_3TR \quad (24)$$

where γ and r are specific values of respectively the treatment and the propensity score.

At last, we have to determine the average causality effect or dose-response function to evaluate the effect of several values of the treatment on the hiring rate of young workers. The dose-response function denoted by $\mu(\gamma)$ may be formulated in the following way:

$$\mu(\gamma) = E[E(Y(\gamma)|r(\gamma, X))] \quad (25)$$

In other words we have to estimate $\hat{\mu}(\gamma)$ such that:

$$\hat{\mu}(\gamma) = \frac{1}{N} \sum_{i=1}^N \hat{b}_0 + \hat{b}_1\gamma + \hat{b}_2\hat{R} + \hat{b}_3\hat{R}\gamma \quad (26)$$

To obtain the dose-response function, we have to calculate $\hat{\mu}(\gamma)$ for different values of treatment, given that for each value γ , standard errors are computed through bootstrapping methods.

3.3.2 Results

The first step consists in matching our observations calculating the Generalized Propensity Score (GPS in the remainder of the paper). The GPS methods are implemented in our application using the `gpscore`, dose-response STATA package (Bia & Mattei, 2008). A key assumption in the STATA implemented version of the GPS methods is the normality of the treatment variable conditional on the pre-treatment covariates. In our study we assume that the Box – Cox transformation of the treatment has a Normal

distribution, given the covariates. Formally, let $BoxCox(T_i)$ denote the Box-Cox transformation of the treatment variable, we define $BoxCox(T_i)$ in the following way:

$$BoxCox(T_i) = \begin{cases} \frac{T_i^\lambda - 1}{\lambda} & \text{if } \lambda \neq 0 \\ \log(T_i) & \text{if } \lambda = 0 \end{cases}$$

Using the Shapiro-Wilk test, we check that $BoxCox(T_i)|X_i \sim N(\beta_0 + \beta_1 X_i, \sigma_v^2)$, where β_0 , β_1 and σ_v^2 are parameters to be estimated through a maximum likelihood procedure. So for each observation, the generalized propensity score is defined in the following way:

$$\hat{R}_i = \frac{1}{2\sqrt{\Pi\sigma^2}} \exp\left(-\frac{1}{2\sigma^2} [BoxCox(T_i) - \beta_0 + \beta_1 X_i]^2\right)$$

Then, according to the distribution of the values of the treatment, that is the absolute value of the decrease in the exit rate among older workers between 1988 and 1989, we divide our sample into different groups. In our study, we choose to break down our treatment into four intervals : $[0;0.05]$, $[0.05;0.15]$, $[0.15;0.3]$ and $[0.3;1]$. Therefore, at each time t , we divide our sample into quartiles according to the value of the treatment.

In each of these groups, we calculate the GPS for a treatment level T_i corresponding to the median treatment level for each group. We can determine the GPS of each observation i belonging to the group j , with $j = \{1, 2, 3, 4\}$ in the following way:

$$\hat{R}_i = \frac{1}{2\sqrt{\Pi\sigma^2}} \exp\left(-\frac{1}{2\sigma^2} [BoxCox(\bar{T}^j) - \beta_0 + \beta_1 X_i]^2\right)$$

where \bar{T}^j is the median treatment level for each group j . According to the distribution of the GPS in each of these groups, we divide them into K blocks. Here, we consider 5 blocks. For each block k^j , with $k = \{1, 2, 3, 4, 5\}$, composed of synthetic firms presenting a specific treatment level corresponding to their group j and a specific GPS corresponding to their block k , we compare the values of observables with those of observations belonging to the block $k^{j'}$ avec $j' \neq j$, namely synthetic firms with a similar GPS but a different treatment level.

Prior to evaluate the dose-response function, we have to ensure that the balancing property is checked. In the table 5, we report the values of the t-statistic when we carry out a test of difference in means for each explanatory variable and for different treatment levels after having adjusted by the GPS. The idea is to test for each group of observations, defined by a specific treatment interval, whether the observables are similar with those of synthetic firms of the other groups, after matching. We use the bold case values to indicate that the t-stat is higher than 1.96. In that case, balancing property does not hold. We remark in the table 5 that no t-stat values are higher than 1.96. We can conclude that our GPS matching procedure allows us to check the balancing property.

Table 5: Testing difference between means for different treatment interval after adjustment by the GPS

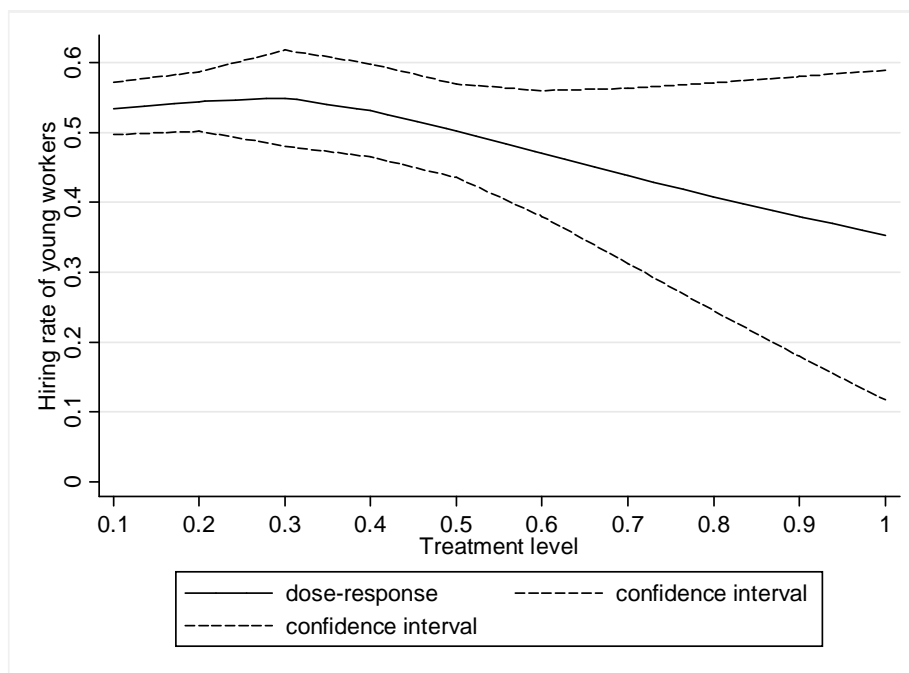
Variables for the year 1988	t-stat for different treatment intervals			
	[0;0.05]	[0.05;0.15]	[0.15;0.3]	[0.3;1]
Share of young workers	-0.18	0.34	-1.11	0.781
Share of older workers	0.369	-0.38	0.524	-0.918
Share of women	-1.15	1.26	0.01	-0.791
Share of CFL contracts	-0.389	0.78	-0.103	0.039
Share of young firms (less than 8 years)	-0.11	-0.243	-0.42	0.577
Share of small firms (less than 9 employees)	0.625	-0.705	-0.235	0.438
Ratio between the average wage of young and that of older workers	0.364	-1.073	0.494	0.592
Unemployment rate	0.941	-0.77	0.148	-0.252
Dummy equal to one if the synthetic firm is declining	-0.218	0.287	0.205	-0.304
Hiring rate of young workers	-0.011	-0.451	0.448	-0.505

Note : the treatment corresponds to the absolute value of the decrease in the exit rate among older workers between 1988 and 1989 for each synthetic firm.

Bold case values indicate that the t-statistic is higher than 1.96 in absolute value.

Source : WHIP (waves 1988 and 1989)

Figure 2: Dose-response function of the treatment on the hiring rate of young workers



We represent the dose-response function in the figure 2 .

This function describes to what extent a decrease in the exit rate among older workers has an effect on the hiring rate of young workers. Given that the treatment is equal to the absolute value of the decrease, a high treatment level corresponds to a strong fall in exits of older workers. We observe in the figure 2 that for a treatment level that ranges from 0.1 to 0.4, the treatment affects slightly the hiring rate of young workers. However, for a dose of treatment higher than 0.4, we remark that a slowdown in the exits of older workers has a negative effect on hirings of young workers.

Thus we show empirically that a strong decrease in exits of older workers reduces sharply hirings of young workers at a local labor market level, which

provides some empirical evidence of our theoretical results. Nevertheless, we have to take these empirical findings with caution. Indeed, we matched our observations given a set of observables which allows us to check the balancing property only for the variables that we include. Therefore we did not correct for a selection bias on unobservables.

Furthermore, recall that our study has been carried out through a synthetic-firm approach. It implies that our observations has been built grouping in a same cell firms belonging to the same sector of activity and located in the same province. Our approach is therefore strongly influenced by the aggregation level chosen to build each synthetic firm.

4 Concluding remarks

In this paper, we investigated the effect of delaying retirement age on hirings of young workers. Our starting point was to consider an imperfect substitutability between young and older workers. We explained the vacancy-chain pattern, through which an exit of an older worker may involve a new hiring of a young worker, even though these two types of workers are not perfectly substitutable for an employer. This vacancy-chain results from job-to-job flows observed within a local labor market composed of firms belonging to the same sector of activity, located in the same area and whose jobs require the same educational level.

In the first part of the paper, we highlighted the theoretical channels through which exits of older workers affect hirings of young workers within a local labor market. In this part, we represented a vacancy-chain in a matching model with heterogeneous jobs and workers, considering that workers are either experimented or not and that workers with no experience (i.e. the young workers) can not apply for a complex job requiring experience. In this model, to represent a vacancy-chain, we allowed for job-to-job flows from simple jobs to complex jobs, but only if workers have a sufficient level of professional experience to apply for complex jobs.

We pointed out two main effects of a fall in the exit rate among older workers on the hiring rate of young workers. First, this decrease lengthens the horizon of an experienced worker and encourages job creation for both types of jobs. Second, the increase in complex job creation raises the probability for an experienced employed job seeker to find a complex job. Therefore, it

reduces the average duration of a simple job making firms more reluctant to open simple jobs and reducing hirings of young workers with no experience. We show that the net effect of a decrease in the exit rate among older workers on the hirings of young workers is negative if the flow value to an employer of a filled complex job is sufficiently high.

Then, in the second part of the paper, we have studied empirically the effect of a slowdown in the exit rate among older workers on the hiring rate of young workers within a local labor market. We use matched employer-employee data drawn from the Worker Histories Italian Panel. We considered a sample of synthetic firms, each of these being composed of firms belonging to the same sector and located in the same area. We grouped first our observations into two categories: synthetic firms that experienced a decrease in the exit rate among older workers between 1988 and 1989 and those that did not experience such a decrease during this period. Using a difference-in-difference approach and controlling for characteristics of each synthetic firm, we showed that a slowdown in the exits of older workers does not exert a significative effect on hirings of young workers.

Nevertheless, we pointed out that some variables differ greatly among the two groups and therefore a simple difference-in-difference analysis has no sense, because it implies a strong selection bias. Consequently, keeping in our sample only synthetic firms that experienced a decrease in the exit rate among older workers between 1988 and 1989 and matching observations each other adjusting by the generalized propensity score to correct for the selection bias on observables, we found that a strong decrease in the exit rate among older workers reduces significantly the hiring rate of young workers within a local labor market. However, we have to take our results with caution given that our matching method does not allow us to correct for a selection bias on unobservables. Furthermore, our findings may depend on the aggregation level chosen to build our synthetic firms. To test the robustness of our results, we should reproduce this study on very big firms assimilating job-to-job flows to promotions within the firm.

One extension of this paper may come to mind. We could allow for both types of worker mobility, within a local labor market, through job-to-job flows and within a firm, through internal promotions. This specification could allow us to better grasp channels through which a decrease in the exit rate among older workers affects the hiring rate among young workers. We leave this idea for further investigation.

Appendix

Rearranging terms of $L(\theta^S, \eta)$, we obtain the following equation:

$$L(\theta^S, \eta) = \frac{\gamma(y^S - w) \left[\left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta \right] + (r + \delta + \lambda)(y^S - w)}{\left[r + \delta + \left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta \right] (r + \delta + \lambda)} - \frac{c^S}{q(\theta^S)} = 0$$

Differentiating $L(\theta^S, \eta)$ we get:

$$L_1 d\theta^S + L_2 d\eta = 0 \Leftrightarrow \frac{d\theta^S}{d\eta} = -\frac{L_2}{L_1}$$

where L_i is the partial derivative of $L(\theta^S, \eta)$ with respect to its i th element, with $i \in \{1, 2\}$. We determine $L_1 = \partial L(\theta^S, \eta) / \partial \theta^S$

$$L_1 = \frac{\left[(y^S - w) \left[\left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta \right] \frac{\partial \gamma}{\partial \theta^S} \right]}{\left[r + \delta + \left[\frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta \right] (r + \delta + \lambda)} - c^S \alpha (\theta^S)^{\alpha-1}$$

We have to determine the sign of $\partial \gamma / \partial \theta^S$. We remark that:

$$\frac{1}{\gamma} = 1 + \frac{\delta \lambda p(\theta^S)}{\eta [\eta + p(\theta^S) + p(\theta^C) + \delta] (\delta + \lambda)}$$

As the second term on the right-hand side of the equation is increasing with θ^S , we deduce that γ is decreasing with θ^S . Consequently, $L_1 < 0$.

Furthermore, let $E(J^S)$ be the the esperance of the expected income flows to an employer from a filled simple job such that:

$$E(J^S) = \frac{\gamma(y^S - w) [p(\theta^C) + \eta] + (r + \delta + \lambda)(y^S - w)}{\left[r + \delta + p(\theta^C) + \eta \right] (r + \delta + \lambda)}$$

where $p(\theta^C) = \frac{(y^C - w^C)}{c^C(r + \eta + \delta)}$. Consequently, L_2 has the same sign as $\partial E(J^S) / \partial \eta$. We consider the logarithm of the expression to simplify calculation so:

$$\begin{aligned}\ln[E(J^S)] &= \ln[\gamma(y^S - w)[p(\theta^C) + \eta] + (r + \delta + \lambda)(y^S - w)] \\ &\quad - \ln[r + \delta + p(\theta^C) + \eta] - \ln(r + \delta + \lambda)\end{aligned}$$

So, differentiating this expression with respect to η , we obtain:

$$\begin{aligned}\frac{\partial \ln E(J^S)}{\partial \eta} &= \frac{\gamma[\frac{\partial p(\theta^C)}{\partial \eta} + 1] + \frac{\partial \gamma}{\partial \eta}[p(\theta^C) + \eta]}{\gamma[p(\theta^C) + \eta] + (r + \delta + \lambda)} - \frac{[\frac{\partial p(\theta^C)}{\partial \eta} + 1]}{[r + \delta + p(\theta^C) + \eta]} \\ &= \left[\frac{\partial p(\theta^C)}{\partial \eta} + 1 \right] \left[\frac{1}{[p(\theta^C) + \eta] + \frac{(r + \delta + \lambda)}{\gamma}} - \frac{1}{p(\theta^C) + \eta + r + \delta} \right] \\ &\quad + \frac{\frac{\partial \gamma}{\partial \eta}[p(\theta^C) + \eta]}{\gamma[p(\theta^C) + \eta] + (r + \delta + \lambda)}\end{aligned}$$

Recall that:

$$\frac{1}{\gamma} = 1 + \frac{\delta \lambda p(\theta^S)}{\eta[\eta + p(\theta^S) + p(\theta^C) + \delta](\delta + \lambda)}$$

Differentiating this expression with respect to η , we obtain:

$$\frac{\partial(\frac{1}{\gamma})}{\partial \eta} = -\frac{\delta \lambda p(\theta^S)}{(\delta + \lambda)} \frac{(\eta + p(\theta^S) + p(\theta^C) + \delta + \eta[1 + \frac{\partial p(\theta^C)}{\partial \eta}])}{(\eta[\eta + p(\theta^S) + p(\theta^C) + \delta])^2}$$

We observe that:

$$p(\theta^C) + \eta \frac{\partial p(\theta^C)}{\partial \eta} = \frac{(y^C - w^C)}{c^C(r + \eta + \delta)} - \frac{\eta(y^C - w^C)}{c^C(r + \eta + \delta)^2}$$

Substituting this expression into the previous one, we obtain:

$$\frac{\partial(\frac{1}{\gamma})}{\partial \eta} = -\frac{\delta \lambda p(\theta^S)}{(\delta + \lambda)} \frac{(2\eta + p(\theta^S) + \delta + \frac{(y^C - w^C)}{c^C(r + \eta + \delta)}[1 - \frac{\eta}{(r + \eta + \delta)}])}{(\eta[\eta + p(\theta^S) + p(\theta^C) + \delta])^2} < 0$$

In addition, given that $\gamma < 1 + \frac{\lambda}{(r+\delta)}$, we deduce that:

$$\left[\frac{1}{[p(\theta^C) + \eta] + \frac{(r+\delta+\lambda)}{\gamma}} - \frac{1}{p(\theta^C) + \eta + r + \delta} \right] < 0$$

Consequently, $\partial \ln(E(J^S))/\partial \eta > 0$ only if

$$\left[\frac{\partial p(\theta^C)}{\partial \eta} + 1 \right] < 0 \Leftrightarrow (y^C - w^C) > c^C(r + \eta + \delta)^2$$

If this condition holds, it ensures that $L_2 > 0$ and therefore $\partial \theta^S / \partial \eta > 0$.

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