

# 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 at a local labor market level, to allow for job-to-job flows. For our empirical study based on Italian matched employer-employee data, we use matching methods adjusting our observations with the generalized propensity score. We show that for the period 1988-1989 in Italy, the higher the decrease in the exit rate of older workers the lower the hiring rate of young workers, in a given local labor market. Then, we explain our findings using a theoretical matching model with heterogenous 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.

**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, the most of 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. First, Bozio (2006) argues 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 age<sup>1</sup>, 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 market<sup>2</sup>, 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 older worker and whose job requires less professional experience. Therefore, after

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<sup>1</sup>See also Chéron et al. (2007) for this question

<sup>2</sup>A 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

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.

Second, Layard, et al. (1991) argue that the idea of a negative relation between youth employment and participation of older workers may stem from 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. Consequently, to predict the effect of delaying retirement on youth employment, we have to better grasp the impact of this reform on job creation over the life cycle. Using a theoretical matching model with endogenous destruction "à la" Mortensen-Pissarides (1994), Chéron et al. (2007) point 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.

The relation between the labor force participation of older workers and the youth employment has also been considered empirically in a macroeconomic perspective. Such a study raises an empirical issue. Indeed, the employment rate among older workers has been severely affected by early retirement schemes, that have been implemented in a setting of mass youth unemployment. Consequently, the labor force participation of older workers is an endogenous variable that presents a simultaneity bias. To address this empirical issue, some recent studies use an instrumental variables approach. First, they compute a measure of the inducement to retire, in accordance with the previous literature on individual retirement incentives (Gruber and Wise, 1998). They determine the Social Security Wealth, the flows of discounted pension benefits expected by an agent who retires at a certain age. They compute this measure for three main exit paths, unemployment compensation early retirement and normal retirement age and for each age and year cells and they aggregate these measures into one indicator of global incentive towards early retirement. Then in a second step, after having checked whether the inducement to retire has a strong negative effect on elderly labor force participation or not, they estimate the impact of this measure on youth employment.

Using French data, Bozio et al. (2008) show that the inducement to

retire has a negative causal impact on employment and unemployment rates among workers aged 20-29 years. They explain their findings showing that a larger inducement to retire translates into a high schooling rate of the young. Using the same approach in the Belgium case, Jousten et al. (2009) provide empirical evidence of a negative causal effect of the incentives towards early retirement on youth employment rate and on schooling rate. Nevertheless, it has no causal effect on youth unemployment.

While the relation between elderly labor force participation and job opportunities for young workers has been studied at a country-level, too few papers deal with this issue 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 examine 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 also raises an empirical issue due to the endogeneity of the variable of interest. To address this issue, Portugal et al. (2009) investigate the effect of a decrease in the exit rate of elderly, 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 treatment<sup>3</sup>, and we control for the 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.

In a second part of the paper, we explain theoretically these results using a matching model with heterogeneous jobs and workers, following the literature (Dolado et al., 1996, Gautier 2002, Mortensen and Pissarides, 1999, Albrecht and Vroman, 1999 ) 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. We show in this framework 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.

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<sup>3</sup>Actually, 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

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

## 2 The empirical study

### 2.1 The data

For our empirical study, we use matched employer-employee data from the Worker Histories Italian Panel (WHIP) built from administrative data of the Italian Social Security Institute (INPS) by researchers of the laboratory R.Revelli of Turin. 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 aggregation level between having the finest grid and having a sufficient number of firms

in each observation unit. Agregating our data by province and by sector of activity, we consider a sample of 517 synthetic firms.

Furthermore, studying the mobility of workers at a synthetic-firm-level allows us to account for the job-to-job flows that count for the major part of the worker flows observed in Italy. Indeed, 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, by studying the mobility of workers at a local labor market level, we 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 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)<sup>4</sup> 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

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<sup>4</sup>The CFL is a fixed-term contract with a maximum duration of 24 months not renewable. This contract was introduced in 1984 to easy the entrance of youngsters aged between 15 and 29 into the labour market, providing fiscal benefits to firms hiring young people.

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 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 declining<sup>5</sup>. We provide some descriptive statistics in the table 2.

## 2.2 A difference-in-difference approach

### 2.2.1 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 adress this issue, we can determine the average treatment effect on the treated group using a difference-in-difference estimator denoted by  $\alpha_{DD}$  defined in the following way:

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<sup>5</sup>It 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

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)

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

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:

$$Y_{it} = \beta X_{it} + \lambda D + \psi \tau_t + \alpha_{DD} D \tau_t + c_i + \varepsilon_{it} \text{ with } \varepsilon_{it} \sim N(0, \sigma^2) \quad (1)$$

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  $\varepsilon_{it}$  is the error term of our regression. In our model, we include a calendar effect adding a binary variable  $\tau_t$  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  $\alpha_{DD}$  to determine the effect of the interaction between the variables  $D$  and  $\tau_t$  on the hiring rate of young workers. In this regression, the parameter  $\alpha_{DD}$  is the difference-in-difference estimator.

### 2.2.2 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  $\alpha_{DD}$  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 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.

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.

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)

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

## 2.3 Generalized propensity score matching method

### 2.3.1 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:

$$E(Y|T = \gamma, R = r) = b_0 + b_1T + b_2R + b_3TR \quad (2)$$

where  $\gamma$  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:

$$\mu(\gamma) = E[E(Y(\gamma)|r(\gamma, X))] \quad (3)$$

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

$$\hat{\mu}(\gamma) = \frac{1}{N} \sum_{i=1}^N \hat{b}_0 + \hat{b}_1 \gamma + \hat{b}_2 \hat{R} + \hat{b}_3 \hat{R} \gamma \quad (4)$$

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

### 2.3.2 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  $\beta_0$ ,  $\beta_1$  and  $\sigma_v^2$  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:

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

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.

We represent the dose-response function in the figure 1 .

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

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

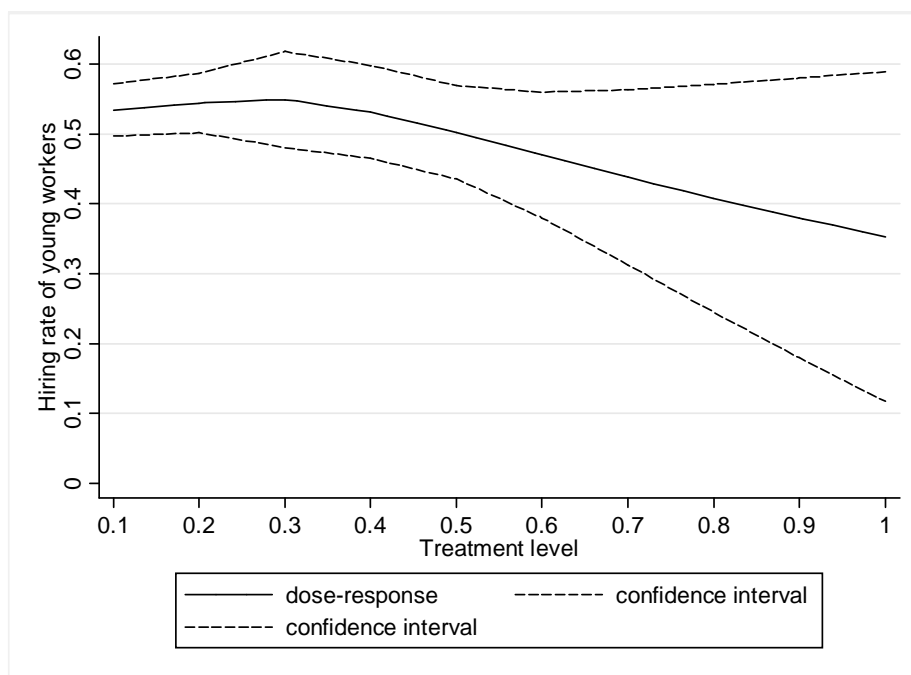


figure 1 that for a treatment level that ranges from 0.1 to 0.4, the treatment affects slightly the hiring rate of young workers. However, for a dose of treatment higher than 0.4, we remark that a slowdown in the exits of older workers has a negative effect on hirings of young workers.

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. Nevertheless, we have to take this results 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.

### **3 The theoretical model: A labor market with heterogenous 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  $\eta_0$  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  $\lambda$  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 wage<sup>6</sup>.

Both types of jobs may be broken up in the case of an idiosyncratic shock, according to a Poisson process with an arrival rate  $\delta$ . 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)<sup>7</sup>.

Following the Behaghel's specification (2007), we consider a guillotine retirement age, introducing a finite horizon in the career of worker<sup>8</sup>. We consider that a worker can retire as soon as he accumulates a sufficient experience level to apply for complex jobs. Let  $\eta$  the Poisson arrival rate at which an experienced worker retires. The horizon  $H$  of an experienced worker is defined in the following way:

$$H = \int_0^{\infty} t\eta e^{-\eta t} dt = \frac{1}{\eta} \quad (5)$$

Consequently, a decrease in  $\eta$  lengthens the career of an experienced worker. In the remainder of the paper, we assimilate a decrease in  $\eta$  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  $\eta$ , on hirings of young workers.

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<sup>6</sup>See also Belan et al. (2007) for a similar specification

<sup>7</sup>Contrarily to their paper, we define the skill level of a worker by his experience level

<sup>8</sup>The 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)

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 their hirings to a special type of workers<sup>9</sup>. 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  $\theta^S$ , 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  $\theta^S$ .

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  $\theta^C$ , 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  $\theta^C$ .

The figure 2 represents the worker flows in our model. The dashed arrow indicates job-to-job flows. Computing flow equations at steady-state, we determine the labor force structure at the local labor market level.

We find the six following steady-state flow equations:

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<sup>9</sup>However, 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.

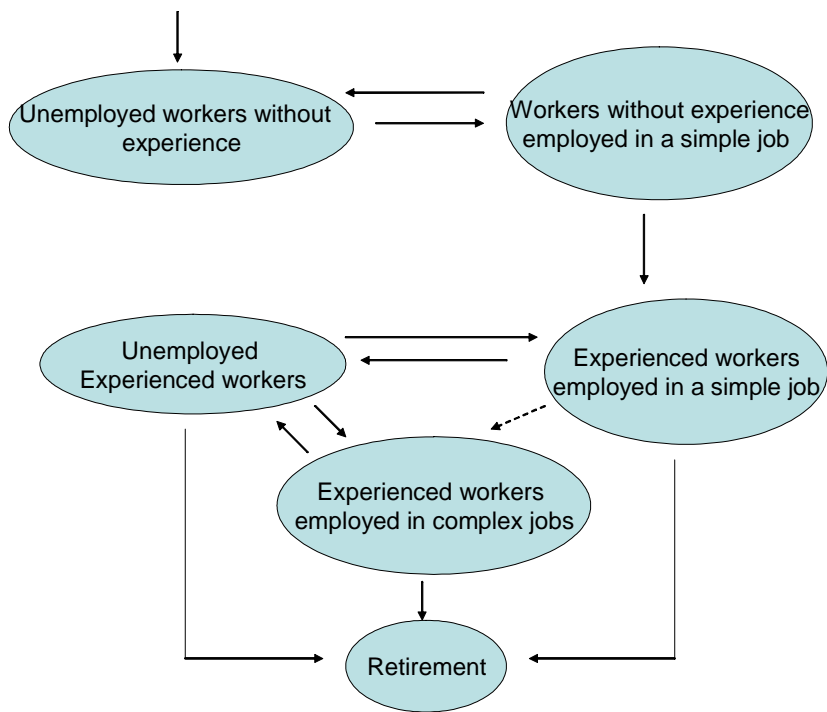


Figure 2: Worker flows in a local labor market

$$(u_e + o_e)p(\theta^C) = E^C(\delta + \eta) \quad (6)$$

$$(E^C + o_e)\delta = u_e[\eta + p(\theta^S) + p(\theta^C)] \quad (7)$$

$$u_e p(\theta^S) + E_{ne}^S \lambda = o_e[\delta + \eta + p(\theta^C)] \quad (8)$$

$$u_{ne} p(\theta^S) = E_{ne}^S(\delta + \lambda) \quad (9)$$

$$E_{ne}^S \delta + \eta_0 = u_{ne} p(\theta^S) \quad (10)$$

$$\eta_0 = \eta(u_e + o_e + E_e^C) \quad (11)$$

where  $u_e$  and  $o_e$  represent respectively the number of experienced job-seekers unemployed and employed,  $E_{ne}^S$  and  $E_e^C$  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  $\eta_0$  represent the number of young workers who arrive in the labor market.

Combining equations (9) and (10) we obtain the expression of the number of unemployed workers with no experience:

$$u_{ne} = \frac{\eta_0(\delta + \lambda)}{\lambda p(\theta^S)} \quad (12)$$

Furthermore, combining the equations (7) and (11) we deduce the expression of unemployment for experienced workers:

$$u_e = \frac{\delta \eta_0}{\eta[\eta + p(\theta^S) + p(\theta^C) + \delta]} \quad (13)$$

The equation (12) 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  $\theta^S$ , 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  $\eta$  has an impact on hirings of young workers with no experience. In the next subsection, we represent the job creation decision of firms.

### 3.1 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  $\gamma$  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:

$$rV^S = -c^S + q(\theta^S)[\gamma(J_{ne}^S - V^S) + (1 - \gamma)(J_e^S - V^S)] \quad (14)$$

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 :

$$rV^C = -c^C + q(\theta^C)(J^C - V^C) \quad (15)$$

where  $J^C$  is the flow value to an employer of a filled complex job. We assume that firms are free to entry 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 :

$$V^C = V^S = 0 \quad (16)$$

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  $\eta$ , 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:

$$rJ_{ne}^S = y^S - w + \delta(V^S - J_{ne}^S) + \lambda(J_e^S - J_{ne}^S) \quad (17)$$

And :

$$rJ_e^S = y^S - w + [\delta + \eta + p(\theta^C)](V^S - J_e^S) \quad (18)$$

Furthermore, when an experienced job seeker finds a complex job, he starts producing an output  $y^C > y^S$  and he receives a wage  $w^C > w$ . His job may break up for two reasons. First, it is hit by an idiosyncratic shock at a Poisson arrival rate  $\delta$  and the worker gets into unemployment. Second, the worker retires at an arrival rate  $\eta$ . So, the flow value to a firm from a filled complex job satisfies the following Bellman equation:

$$rJ^C = y^C - w^C + (\eta + \delta)(V^C - J^C) \quad (19)$$

After the description of value functions for firms, let us determine the value functions for workers. In our model, we assume that only the wage  $w^C$  may result from a negotiation between the employer and the employee. So, we focus only on Bellman equations for experienced workers. We define first the flow value  $U_e$  of unemployment to an experienced worker. Recall that a worker of this type may be matched with a simple or a complex job at respective arrival rates  $p(\theta^S)$  and  $p(\theta^C)$ . Furthermore, he may retire at an Poisson arrival rate  $\eta$ . In that case, he receives a pension  $P$  discounted over an infinite time. Let  $b$  the amount of unemployment benefits, so the value  $U_e$  satisfies the following Bellman equation:

$$rU_e = b + p(\theta^S)(W_e^S - U_e) + p(\theta^C)(W^C - U_e) + \eta\left(\frac{P}{r} - U_e\right) \quad (20)$$

When an experienced worker occupies a complex job, he receives a wage  $w^C$ . He can either lose his job at a rate  $\delta$  or retire at a rate  $\eta$ . Consequently, the value  $W^C$  to an experienced worker from being employed in a complex job is defined by the following Bellman equation:

$$rW^C = w^C + \delta(U_e - W^C) + \eta\left(\frac{P}{r} - W^C\right) \quad (21)$$

When an experienced worker occupies a simple job, he receives a minimum wage  $w$ , he may lose his job at an arrival rate  $\delta$ , he may be matched with a complex job at an arrival rate  $p(\theta^C)$  and he may retire at a rate  $\eta$ . So, the flow value  $W_e^S$  to an experienced worker from occupying a simple job satisfies the following equation:

$$rW_e^S = w + \delta(U_e - W_e^S) + p(\theta^C)(W^C - W_e^S) + \eta\left(\frac{P}{r} - W_e^S\right) \quad (22)$$

### 3.2 Wage posting equilibrium

We determine first the steady-state equilibrium when the wage received by an experienced worker occupying a complex job is not negotiated. Combining the Bellman equation (14) and the free-entry condition we obtain the following expression:

$$\gamma J_{ne}^S + (1 - \gamma)J_e^S = \frac{c^S}{q(\theta^S)} \quad (23)$$

Substituting the equations (17) and (18) into the expression (23) we get the following relation between  $\theta^S$  et  $\theta^C$ :

$$\frac{\gamma}{(r + \delta + \lambda)} [(y^S - w) + \frac{\lambda[y^S - w]}{[r + \delta + p(\theta^C) + \eta]}] + \frac{(1 - \gamma)(y^S - w)}{[r + \delta + p(\theta^C) + \eta]} = \frac{c^S}{q(\theta^S)} \quad (24)$$

Furthermore, substituting the Bellman equation (19) in the expression (15) and using the free-entry condition, we obtain the following expression:

$$\frac{c^C}{q(\theta^C)} = \frac{(y^C - w^C)}{(r + \eta + \delta)} \quad (25)$$

To determine the effect of a decrease in  $\eta$  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:

$$h^S(u_{ne} + u_e, v^S) = (u_{ne} + u_e)^{\frac{1}{2}}(v^S)^{\frac{1}{2}} \quad (26)$$

$$h^C(u_e + o_e, v^C) = (u_e + o_e)^{\frac{1}{2}}(v^C)^{\frac{1}{2}} \quad (27)$$

Substituting (27) into (25), we obtain the following expression:

$$(\theta^C) = \left[ \frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right]^2 \quad (28)$$

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  $\eta$  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  $\eta$  on the creation of simple jobs, that is on  $\theta^S$ .

Substituting (28) into (24), we find that at steady-state equilibrium,  $\theta^S$  satisfies the following equation:

$$\begin{aligned} L(\theta^S, \eta) &= \frac{\gamma}{(r + \delta + \lambda)} [(y^S - w) + \frac{\lambda[y^S - w]}{[r + \delta + \left[ \frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta]}] + \frac{(1 - \gamma)(y^S - w)}{[r + \delta + \left[ \frac{(y^C - w^C)}{c^C(r + \eta + \delta)} \right] + \eta]} \\ - \frac{c^S}{q(\theta^S)} &= 0 \end{aligned} \quad (29)$$

In the wage posting equilibrium, we find two main offsetting effects of a decrease in the exit rate among older workers denoted by  $\eta$  on the tightness for the submarket of simple jobs  $\theta^S$ . On the one hand, a decrease in  $\eta$  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  $\eta$  has a negative effect on  $\theta^S$  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  $\eta$ , 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  $\gamma$  is given by our flow equations at steady-state and therefore it is a function of both  $\theta^S$  and  $\theta^C$ . Indeed, combining equations (12) and (13), we write  $\gamma$  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  $\gamma$  is quite complicated and depends both on  $\theta^S$  and  $\theta^C$ . So it is not so straightforward to examine the sign of the derivative of  $\theta^S$  with respect to  $\eta$ .

**Proposition 1** *In the wage posting equilibrium, a decrease in  $\eta$  reduces  $\theta^S$ , 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  $\theta^S$  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  $\eta$  raises unemployment among young workers with no experience.

However, in this subsection we assumed that the wage received by a worker occupying a complex job is not negotiated between the employer and the worker. We relax this assumption in the next subsection.

### 3.3 Steady-state equilibrium with wage negotiation

Now, we aim at determining the steady-state equilibrium assuming that the wage  $w^C$  results from a bargaining process between the employer and the worker. To describe this process, we follow the previous papers of Shimer (2005) and of Hall and Milgrom (2008). With respect to the standard Nash bargaining process, the process that they describe has a different threat point for workers. Indeed, in a negotiation "à la" Nash the threat point of workers is the value of being unemployed while in a model "à la" Shimer, the threat point is the so-called "delay position" of the worker.

According to Shimer (2005), this strategic wage bargaining process allows to reproduce the extreme volatility of the ratio of the number of vacancies over the number of unemployed observed in the facts. The underlying idea is that in the case when the worker meets once a job offer, so as long as the job is not broken up, the employer can always make counter offers until the negotiation succeeds. Nevertheless, an experienced job seeker may retire at an arrival rate  $\eta$  and in that case, the employer can not make counter offers anymore. Consequently,  $D_e$ , that is the threat point of an experienced job seeker satisfies the following Bellman equation:

$$rD_e = b + \eta\left(\frac{P}{r} - D_e\right) + \delta(U_e - D_e) \quad (30)$$

The wage  $w^C$  received by a worker occupying a complex job results from a maximisation of the product of net surplus of both parties, weighted by their respective bargaining power. Consequently:

$$w^C = \arg \max\{(W^C - D_e)^\beta (J^C - V^C)^{(1-\beta)}\}$$

where  $\beta$  is the bargaining power of the worker. At steady-state equilibrium, given the free-entry condition (16), solving the maximisation program yields the following sharing rule:

$$W^C - D_e = \beta(J^C + W^C - D_e) \quad (31)$$

Rearranging terms, we can formulate this expression in the following way:

$$-(1 - \beta)(r + \eta + \delta)D_e = (r + \eta + \delta)[\beta(J^C + W^C) - W^C]$$

Substituting expressions (19), (21) and (30) in the previous expression, we deduce the following wage equation:

$$w^C = (1 - \beta)b + \beta y^C \quad (32)$$

Replacing the wage  $w^C$  by its expression in the equation (19), and given the free-entry condition we obtain:

$$\frac{c^C}{q(\theta^C)} = \frac{(1 - \beta)(y^C - b)}{(r + \eta + \delta)} \quad (33)$$

So endogenizing the wage bargaining for complex jobs and keeping the same functional forms than in the previous subsection, we show that at steady-state equilibrium  $\theta^S$  satisfies the following equation:

$$\begin{aligned} L(\theta^S, \eta) &= \frac{\gamma}{(r + \delta + \lambda)} [(y^S - w) + \frac{\lambda[y^S - w]}{[r + \delta + \left[\frac{(1-\beta)(y^C-b)}{c^C(r+\eta+\delta)}\right] + \eta]}] + \frac{(1 - \gamma)(y^S - w)}{[r + \delta + \left[\frac{(1-\beta)(y^C-b)}{c^C(r+\eta+\delta)}\right] + \eta]} \\ - \frac{c^S}{q(\theta^S)} &= 0 \end{aligned} \quad (34)$$

Consequently, even though we account for a wage bargaining process in the submarket of complex jobs, a decrease in the exit rate among older workers reduces hirings of young workers and raises unemployment among young workers with no experience if the value for an employer of a filled complex job is sufficiently high such that:

$$(1 - \beta)(y^C - b) > c^C(r + \eta)^2 \quad (35)$$

We remark that the condition (35) accounts for the bargaining power of the worker denoted by  $\beta$  and for the amount of unemployment benefits  $b$ , contrarily to the condition obtained in the wage posting case. Nevertheless,

these two conditions are very similar so we can conclude that endogenizing the wage received by a worker occupying a complex job does not change our theoretical results when we consider a wage negotiation process "a la" Hall and Milgrom (2008).

Consequently, through our theoretical study, we can explain our previous empirical findings obtained in the Italian case. Indeed, if the exit rate among older workers falls sufficiently, it encourages employers to open more complex jobs, which will reinforce job-to-job flows from simple jobs to complex jobs but only for experienced workers. This intensification of job-to-job flows reduces creation of simple jobs and therefore, it is harmful to young workers who arrive in the labor market and who experience a fall in their probability to be hired on a simple job.

## 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 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 significant 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.

In the second 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 experienced 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.

However, one extension of this theoretical model may come to mind, endogenizing the minimal experience level which allows a worker occupying a simple job to apply for a complex job. 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:

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

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  $L_1 = \partial L(\theta^S, \eta) / \partial \theta^S$

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

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  $\theta^S$ , we deduce that  $\gamma$  is decreasing with  $\theta^S$ . 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:

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

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  $\eta$ , we obtain:

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

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  $\eta$ , 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:

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

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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