The Detrimental Effect of Job Protection on Employment: Evidence from France[1]

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August 2018 - PRELIMINARY DRAFT

[1]We are particularly indebted to Bérengère Patault for providing information on severance payments. We would also like to thank Giuseppe Bertola, Bruno Bouchard, Laurent Davezies, Grégory Jolivet, Xavier d'Haultfoeuille, Hélène Turon, Etienne Wasmer and Josef Zweimuller for their insightful comments. We are grateful to seminar participants at the Universities of Bristol, Paris I Sorbonne, Oxford, VU Amsterdam, Essex, ASSA meeting in San Francisco, TEPP winter school, EUROEMP 2018 meeting, SAM meetings in Barcelona, Cambridge, Marseille and Istanbul. We acknowledge the support of the Investissements d'Avenir grant (ANR-11-IDEX-0003/Labex Ecodec/ANR-11-LABX-0047) and of the Chaire de sécurisation des parcours professionnels.

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Abstract

According to French law, employers have to pay at least six monthly salary to employees whose seniority exceeds two years in case of unfair dismissal. We show, relying on data, that this regulation entails a hike in severance payments at two-year seniority which induces a significant rise in the job separation rate before the two-year threshold and a drop just after. The layoff costs and its procedural component are evaluated thanks to the estimation of a search and matching model which reproduces the shape of the job separation rate. We find that total layoff costs increase with seniority and are about four times higher than the expected severance payments at two years of seniority. Counterfactual exercises show that the fragility of low-seniority jobs implies that layoff costs reduce the average job duration and increase unemployment for a wide set of empirically relevant parameters.

Keywords: Employment protection legislations, Dismissal costs, Unemployment.
JEL classification: J65, J63, J32.

1 Introduction

Standard economic models predict that employment protection legislations reduce job destruction and job creation, hence suggesting that they have an ambiguous effect on unemployment.[1] The negative impact on job creation arises from the anticipation of separation costs, which reduce expected profits. Although this mechanism is clear in theory, it has very little empirical support.[2] In this paper, we show that the *anticipation effect* does play a key role in reducing job creation, but also in increasing job destruction, thus inducing significant employment losses. To establish this finding, we use a discontinuity in the relation between seniority and severance payments in case of unfair dismissal.

Until September 2017, French employers had to pay at least six months' salary to their employees whose seniority exceeded two years in case of unfair dismissal. We show, relying on data, that this regulation entails a hike in severance payments which induces a noticeable rise in the job separation rate just before the two-year threshold, followed by a drop of even greater magnitude. We devise a search and matching model which accurately reproduces the shape of the separation rate, and use the discontinuity at two years to identify its structural parameters. Estimating the model allows us to quantify the hike in layoff costs at the two-year threshold, and to separate the procedural from the severance components. We find that total layoff costs increase with seniority and are about four times larger than the expected severance payments of laid-off workers at two years of seniority. Procedural costs are therefore very large due to the extreme complexity of rules and procedures defined by the French labor code.

Once these costs have been quantified, we run counterfactual experiments to evaluate their impact on job tenure, labor market flows and unemployment.[3] In our empirical context, job protection *reduces* expected job duration for low skilled workers, but increases it for high skilled workers. The negative impact of job protection on job duration is driven by the anticipation effect: as firms forecast that they will have to pay higher costs in the future in the event of a separation, they find it optimal to anticipate the separation decision. The anticipation effect, which raises the destruction of jobs with short tenure, is counterbalanced by the well-known labor hoarding effect which lowers the separation rate for jobs with long tenure. The anticipation effect dominates for low skilled workers because they usually land jobs with small initial surpluses. The presence of the minimum wage amplifies this effect because it prevents downwards wage adjustments that could dampen the

[1] See for instance Bentolila and Bertola (1990) or Mortensen and Pissarides (2011).

[2] Scarpetta (2014).

[3] Although we focus on the French employment legislation, our approach could be applied to a broader set of countries because discontinuities in severance packages after a seniority threshold are fairly common, as shown by the OECD job protection database: http://www.oecd-ilibrary.org/employment/data/employment-protection-legislation_lfs-epl-data-en

impact of severance payments on profitability.

These findings rely on an empirically relevant property of the model consistent with the strongly decreasing profile of separation rates for jobs with short tenure. The fragility of low-seniority jobs, mostly overlooked by the literature on employment protection so far, implies that the anticipation effect is empirically important, especially in the presence of wage floors. In our empirical context, this suggests that job protection does not achieve its goal since its makes job retention less likely. However, in the absence of wage floors, or when the surplus of starting jobs is sufficiently high, the anticipation effect can be dominated by the labor hoarding effect. This is why we find that job protection raises the job duration of skilled workers, although it decreases that of unskilled workers.

We also take into account the impact of job protection on job creation to simulate its effect on unemployment. We find that job protection significantly raises unemployment for all categories of workers. This result is driven to a large extent by the fragility of low-tenured jobs, which implies that job protection raises the destruction of these jobs and significant lowers their rate of creation. The unemployment impact of job protection is larger for the unskilled, because the surplus of their jobs is lower and the minimum wage is more likely to be binding, but it is also strong and significant for skilled workers.

Related literature. Our paper is connected to several strands of literature. A large literature relying on theoretical models indicates that firing costs have countervailing effects on the firing and hiring margins. Bertola and Bentolila (1990) find in their seminal paper that employment protection legislations have a larger impact on firms' propensity to fire than to hire, thus suggesting that stringent regulations actually increase average employment in the long-run. Their conclusion has been qualified by subsequent research. For instance, Sargent and Ljunqvist (2008) underline that the overall impact of firing costs depends on the volatility of the economic environment. They argue that microeconomic turbulences which emerged in the 1980s turned firing costs from beneficial to detrimental.

Contributions which analyze the effect of job protection increasing with tenure confirm that job protection has an ambiguous impact on employment (Blanchard and Landier, 2002; Cahuc and Postel-Vinay, 2002; Güell and Rodriguez-Mora, 2017; Lalé et al., 2016). These contributions also stress that job protection is detrimental to workers who are unemployed or employed with a short tenure while it is beneficial to insiders. One limitation of this literature is that it does not precisely evaluate how seniority affects job separation. In order to properly evaluate the impact of job protection, we rely on contributions which explain the negative relation between seniority and the likelihood of separation (Jovanovic, 1979; Shimer, 1999; Moscarini, 2005; Pries, 2004; Prat, 2006). These contributions show that memory-less stochastic processes for productivity, like the Poisson process widely used in search and matching models, cannot explain why job fragility decreases with seniority. We add to this literature by introducing job protection, which depends on seniority, in a search and matching model where idiosyncratic job productivity follows a geometric Brownian motion.[4] It is well known that this type of model can account for the negative relation between seniority and the job separation rate.[5] We show that this setup also explains the rise in the job separation rate before the seniority at which the layoff costs suddenly increase, as well as the drop in the separation rate just after that seniority. Our analysis complements the contributions which have outlined that the anticipation of layoff costs increases the turnover of jobs with short tenure, by quantifying this mechanism and showing that it can be empirically so large that it dominates the usual labor hoarding effect, thus yielding a negative relation between average employment duration and the stringency of the legislation. We find that this detrimental effect is more important for low skilled workers, especially in the presence of binding wage floors.

A second literature has evaluated the structure of costs that firms face when adjusting their employment using surveys about hiring and separation costs (Abowd and Kramarz, 2003; Kramarz and Michaud, 2006; Garibaldi and Violante, 2005; Boeri et al., 2017). This literature reports information on several components of separation costs. However, it does not analyze the empirical relation between provisions of employment protection legislations and the layoff costs to the employers. This issue is nevertheless very important as it is clear that the impact of employment protection legislations on layoff costs depends, among other elements, on the interpretation by labor courts, on the length, the cost and the uncertainty of the legal process (Bertola et al., 1999; Ichino et al., 2003; Galdon Sanchez and Guell, 2003; Garibaldi and Violante, 2005; Besancenot and Vranceanu, 2009; Goerke and Neugart, 2015). We complement this literature by using original data about severance payments gathered from more than 27,000 court rulings and by providing a framework that links provisions of the employment protection legislations to the actual layoff costs. Information from court rulings allows us to compute the expected severance payments of laid-off employees. Then, we identify the procedural component of layoff costs at two-year seniority from the change in the job separation rate and in the expected severance payment. Our framework, although applied to a specific case, can be used in different contexts insofar as most provisions of employment protection legislations depend on job seniority. Its main advantage is that it makes possible to empirically

[4]Vindigni et al. (2015) have introduced job protection in a model where productivity shocks are described by a geometric Brownian process. As they assume that job protection does not depend on seniority, they do not account for the anticipation effect stressed in this paper.

[5]Besides its capacity to fit the empirical hazard rate of job separations, the assumption that output per worker follows a geometric Brownian motion is consistent with standard Mincerian regressions as well as more structural approaches, such as Prat (2011) or Buhai and Teulings (2014), which show that the distribution of wages among workers with the same seniority is accurately approximated by log-linear models with normal shocks when wages are set by surplus sharing rules.

distinguish the procedural from the severance components. Our evaluation indicates that procedural costs are significantly larger than severance payments. These findings are consistent with the literature in law which stresses that the extreme complexity of the French employment protection legislation is very difficult to manage for employers.[6]

Moreover, our framework allows us to look at the impact of layoff costs on labor flows and unemployment. From this perspective, we complement the large empirical literature which has developed empirical strategies to identify the impact of job protection by relying on reduced form models. Our results are in line with empirical findings based on natural experiments which consistently suggest that more stringent job protection increases the turnover of jobs with short tenure, generally temporary jobs, but has either very weak or no significant impact on the turnover of jobs with long tenure, thereby reducing rather than increasing average worker security (Schivardi and Torrini, 2008; Centeno and Novo, 2012; Hijzen et al., 2017).

Finally, our approach is also related to the contribution of Garicano, Lelarge and van Reenen (2016) who use firm size-contingent laws to identify the equilibrium and welfare effects of labor regulation. Their study relies on a static model and applies it to France, where many labor laws start to bind when firms have 50 or more employees. They also structurally estimate the key parameters of their model to run counterfactual analysis. We show that a similar approach can be used in a dynamic setup to analyze the consequence of employment protection legislations on labor market outcomes. Accounting for equilibrium effects is particularly important for the analysis of job protection insofar as job protection has effects on job destruction, job creation and wages.

The rest of the paper is organized as follows. Section 2 is devoted to the description of the institutional background. The model is laid out in section 3. Section 4 presents the estimation and the implications of the model regarding job stability in a partial equilibrium setting where the arrival rate of job offers is exogenous. Section 5 explores the consequences of job protection when the arrival rate of job offers is endogenous. Section 6 concludes. All proofs are relegated to the Appendix.

2 Institutional background

It is well known that employment protection legislations tend to decrease the job separation rate for workers with long tenure, but can have opposing effects for workers with little seniority (Blanchard and Landier, 2002; Cahuc and Postel-Vinay, 2002). This is particularly true in France, where job protection becomes really stringent after two years of seniority. After this threshold, employers have to pay at least six months' salary to their employees in case of unfair dismissal on a perma-

[6] Ray (2017).

nent job.[7] Before this threshold, no minimum amount is required. To avoid the cost of breach of permanent contracts, employers make an extensive use of temporary contracts. In principle, temporary contracts may be used in special circumstances only: to replace an employee who is absent, to cover changes in business activity or for seasonal work. Nevertheless, about 90% of hires are on temporary contracts.[8] In the period covered by our study, temporary contracts can be renewed only once, with a maximum total duration (including the two subsequent contracts) of 18 months. However, if the renewal of the contract is associated with changes in tasks, employers can renew temporary contracts more than once and so keep the same worker on a temporary position for more than 18 months. This strategy is widely used by French employers to avoid permanent contracts. But it becomes less profitable when the tenure of the employee exceeds two years because the employee can go to court if her contract is not renewed, and ask for a requalification of her temporary contract into a permanent contract, arguing that the circumstances required to use a temporary contract are not fulfilled. If the request of the employee is successful, the job separation induced by the non-renewal of the temporary contract is interpreted as a layoff by the court. Then the employer has to pay a severance of at least six months' salary whenever tenure is beyond two years.

In order to evaluate the impact of this legislation on actual compensation for wrongful contract breach, we rely on data from the Ministry of Justice. In the French system, labor Courts (Prud'hommes councils) are in charge of litigations about labor contract breach. However, the decision of the Prud'hommes councils, which are mainly composed of non-professional judges, are appealed in most of the cases: the appeal rates are, according to Guilloneau and Serverin (2015), between 60% and 67% in the 2004-2013 period. Because of such high appeal rates, compensations for wrongful contract breach decided at the Appeal Court level are a better measure of the compensation paid by firms. Our information about compensations for wrongful contract breach relies on the analysis of more than 27,000 French Appeal Court's rulings from 2003 to 2012.[9] Figure 1, which reports the average compensation by quarter of job seniority, clearly shows that the compensation jumps upwards at the two-year threshold.

The analysis of job separations shows that their likelihood[10] changes dramatically around the two-year threshold, thus mirroring the discontinuity in the compensation for wrongful termination. The job separation rate is estimated from the French Labor Force Survey over the period 2003-

[7]This rule has been changed by a law issued on 22 September 2017.

[8]Milin (2018).

[9]These data are presented in Appendix D.

[10]The job separation probability presented here includes transitions from employment to unemployment for consistency with the theoretical model. The hike before the two-year threshold is also apparent when all separations are taken into account.

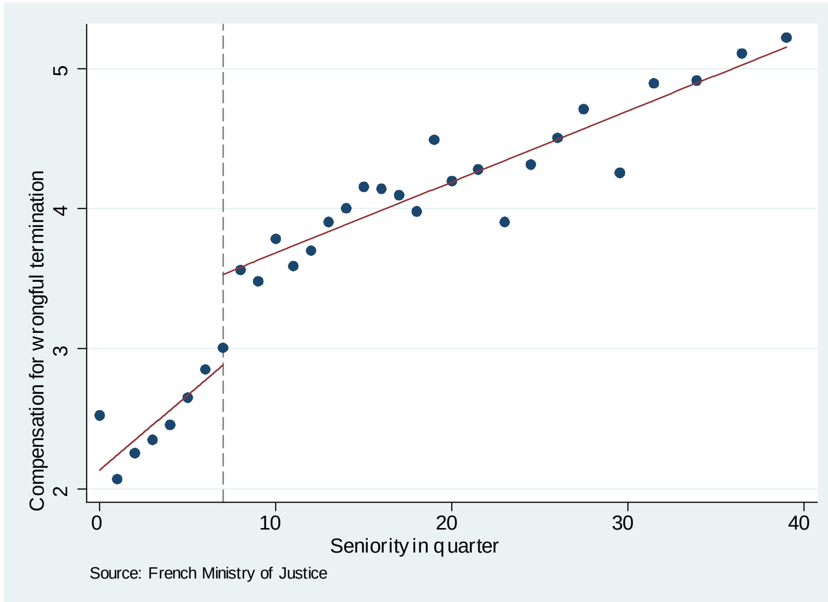

Figure 1: Average compensation in case of appeal for wrongful contract termination (in monthly wage) and seniority (in quarter).

Note: The graph represents a binscatter which groups the variable on the horizontal axis into equal-sized bins, computes the mean of the variables on the horizontal and vertical axes within each bin, and creates a scatterplot of these data points using a sample of 27,936 French Appeal Courts rulings over 2003-2012. The compensation for wrongful termination reported on the vertical axis is the average compensation for all litigations for contract breach, including the cases where the employee gets no compensation (39.9% of litigations).

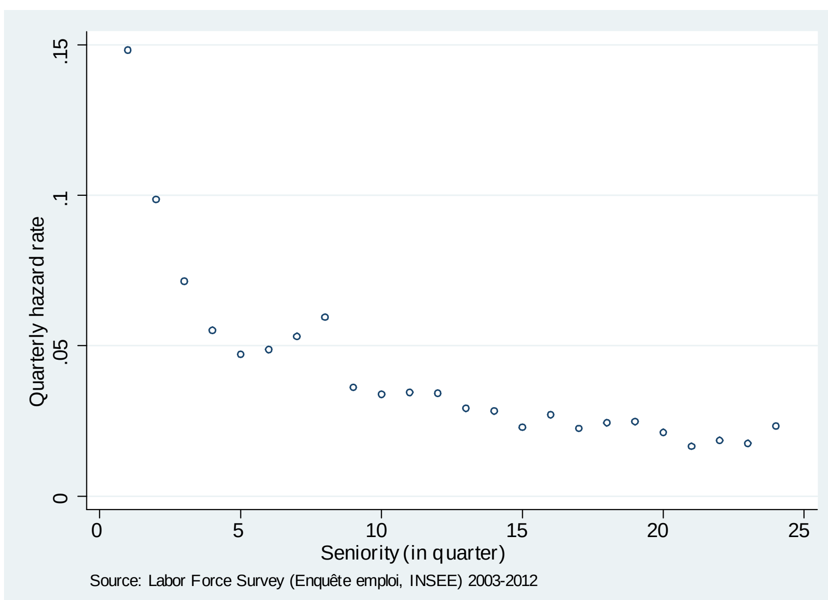

Figure 2: The relation between seniority and quarterly hazard rate from employment to unemployment

Note: Individuals working in the private sector, aged 15 to 54, with at most high school degree. Subsidized jobs are excluded.

2012. The Labor Force Survey provides information about the seniority of workers within the firm (in month) and a range of individual characteristics. Figure 2 displays the relation between the job separation rate and job seniority of workers aged from 15 to 54, who have completed at most their high school degree and work in the private sector.[11] It is clear that the job separation rate strongly decreases with seniority until the fifth quarter. Then, the job separation rate increases until the eighth quarter. Between quarter eight and nine, there is a huge drop, and from quarter nine the hazard rate slowly decreases with seniority.

In the rest of the paper, we provide and estimate a model that explains this non-monotonic relation between seniority and job separation. According to our model, the increase in the job separation probability from quarter five to quarter eight, and the drop from quarter eight to quarter nine, are induced by the increase in layoff costs at the two-year threshold. Moreover, we can use the size of the drop to identify the layoff costs which are not available in survey data. Once the layoff costs are estimated, together with the other parameters of the model, it becomes possible to evaluate their impact on labor market flows and on unemployment.

3 The model

Set-up. We consider an overlapping generations model in continuous time where people are born and die at rate χ. There are two goods: output, which is the numéraire, and labor. Individuals are risk neutral and discount the future at rate ρ. They are either employed or unemployed. Unemployed individuals sample job offers at the exogenous rate λ. Inasmuch as 95% of French employees are covered by collective agreements, with high binding minimum wages, it is important to account for wage floors. For the sake of simplicity, we first present the model in the case where the wage is exogenous. This restriction is relaxed in Section 3.2 where wages are endogenized through Nash bargaining in the presence of wage floors to account for the fact that some workers, even among the low skilled, earn more than the minimum wages.

It is well known that the standard Mortensen-Pissarides (1994) model, in which productivity dynamics are driven by draws from a stationary distribution with constant Poisson arrival rate, cannot reproduce the relation between seniority and separations described in the previous section since it predicts that jobs are destroyed at a constant rate. By contrast, Shimer (1999) and Prat (2006) have shown that a negative relation between job separation and seniority arises naturally if productivity follows a geometric Brownian motion.[12] In a discrete time set-up, this means that output of a

[11] The benchmark estimation of the model developed below covers this population.

[12] A related finding was already outlined in Jovanovic (1979). Jovanovic's set-up was fundamentally different from the one considered in this paper since he assumed that matches are experience goods with constant average productivity. Then job separation is triggered when the belief about a match productivity falls below some threshold. This mechanism

worker of tenure t, denoted by x_t, starts at value x_0 drawn from an exogenous distribution, and then obeys the following law of motion

$$\log x_{t+1} = \log x_t + \mu + \varepsilon_t, \text{ with } \varepsilon_t \sim \mathcal{N}(0, \sigma^2). \tag{1}$$

The coefficient μ is a deterministic drift which accounts for the accumulation of ***job-specific*** human capital, and ε_t is a normal random variable with zero mean and variance σ^2. Because it simplifies many derivations, we will devise our model in continuous time. Letting the time interval in equation (1) go to zero, while preserving the overall variance of the productivity process, yields a geometric Brownian motion. Then the flow output of a worker with tenure t satisfies the stochastic differential equation

$$\mathrm{d}x_t = (\mu \mathrm{d}t + \sigma \mathrm{d}B_t)\, x_t, \tag{2}$$

where the stochastic component B_t is a standard Brownian motion. One can think of B_t as the sum of uncorrelated normal random variables with zero mean and unit variance per period. Our specification implies that, in the absence of job selection, log-productivity among workers starting from the same x_0 and with tenure t is normally distributed with mean $(\mu - \sigma^2/2)t$ and variance $\sigma^2 t$.

We also introduce a source of uncertainty such that jobs are forced out of business when hit by random shocks which arrive at Poisson rate δ. These separations are the results of discontinuous and strongly negative drops in profitability. Besides those jobs destroyed for reasons exogenous to our model, others will be terminated when their productivity becomes too low to cover the opportunity cost of employment.

We first focus on a simple benchmark case where starting jobs are not covered by job protection and layoff costs are constant from tenure $T > 0$. This stylized case allows us to analytically study the impact of the discontinuity in layoff costs. Then we consider a generalization of this step function where layoff costs, besides being discontinuous at two years, increase with seniority from the start of the job spell. This specification will be used in the empirical part since, as argued in the previous section, it accurately describes the profile of severance payments in France.

3.1 Benchmark case with constant layoff costs for protected jobs

We assume in this subsection that starting jobs can be destroyed at zero cost until they are transformed into protected jobs at tenure T, which corresponds to the two-year threshold presented above. All workers who reach tenure T see their jobs upgraded to protected ones, the destruction of which requires spending the cost $F > 0$. When a job is destroyed and the layoff costs have been paid, the value of the job is equal to zero.

has been embedded into a general equilibrium setting by Pries (2003) and Moscarini (2005).

Protected jobs. We analyze the problem recursively, starting with protected jobs and then reinserting their optimal values to solve for that of non-protected jobs. By definition, the profit flow of a match is equal to output x minus the exogenous wage w, and the firm has to pay the layoff costs F when the job is destroyed. Thus, the profit of a *protected job* with current output x_t, denoted by $J(x_t)$, solves the optimal stopping problem

$$J\left(x_t\right) \equiv \sup_{\tau \geq t} \mathbb{E}^{x_t}\left[\int_t^\tau e^{-(r+\delta)(s-t)}(x_s - w)\mathrm{d}s - e^{-(r+\delta)(\tau-t)}F\right], \tag{3}$$

where $r \equiv \rho + \chi$. The integral is the present value of the profit flow. The term after the integral indicates that profits are expected until some random date τ at which the job becomes unproductive and is destroyed at cost F. The supremum is taken over the set of possible stopping time within $[t, +\infty)$ where the destruction date τ is a stopping time that is measurable with respect to the filtration $\mathcal{F}_t$ generated by B_t. E^{x_t} denotes the expectation operator, conditional on x_t. Given that $J(\cdot)$ is a strictly increasing function of output, separation occurs when x_t breaches the reservation threshold R such that $J(R) = -F$. This is a standard optimal stopping problem, the solution of which can be derived in closed-form. As shown in Appendix A, the values of R and of profits $J(x)$ read

$$J(x) = \frac{x}{r+\delta-\mu} - \frac{w}{r+\delta} - \left(\frac{R}{r+\delta-\mu} - \frac{w-(r+\delta)F}{r+\delta}\right)\left(\frac{x}{R}\right)^{\alpha}, \tag{4}$$

where $\alpha \equiv \frac{1}{2} - \frac{\mu}{\sigma^2} - \sqrt{\left(\frac{1}{2} - \frac{\mu}{\sigma^2}\right)^2 + \frac{2(r+\delta)}{\sigma^2}}$ and

$$R = \frac{\alpha}{\alpha-1}\left(\frac{r+\delta-\mu}{r+\delta}\right)\left[w-(r+\delta)F\right]. \tag{5}$$

The value of a job, $J(x)$, is made of three terms. The first two terms represent the expected value of future profits. The last term is the option value of job destruction. The reservation output R decreases with F as firms procrastinate more before laying-off their workers when firing costs are higher. In the extreme cases where $(r+\delta)F > w$, the reservation output becomes negative and, since output is by definition positive, firing is never optimal. To rule out prohibitive firing costs, we shall henceforth assume that $(r+\delta)F < w$.

Non-protected jobs. Non-protected jobs can be destroyed at zero cost. Although layoff costs do not directly affect the profits of non-protected jobs, firms know that jobs have to be either transformed into protected jobs or destroyed at tenure T. This is why the value of a non-protected job, denoted by $J_n(x,t)$, depends not only on current output but also on tenure t. The terminal condition $J_n(x,T) = J(x)$ ensures that employers anticipate the losses induced by employment protection legislations once they become binding.

Not all jobs reach tenure T either because some are hit by an exogenous destruction shock, or because their output falls below the reservation barrier, denoted by $R(t)$. In the latter case, firm destroy jobs as their value becomes negative, so $J_n(R(t), t) = 0$ for all $t \in [0, T)$. Since the value of jobs varies with tenure, the separation threshold of non-protected jobs cannot be flat. Intuitively, as the transformation date nears, jobs become less valuable and separation more attractive.

Formally, the value of a non-protected job solves the following optimal stopping problem

$$J_n(x_t, t) = \sup_{t \leq \tau \leq T} \mathbb{E}^{(x_t, t)} \left[\int_t^{\tau} e^{-(r+\delta)(s-t)} (x_s - w) \mathrm{d}s + e^{-(r+\delta)(T-t)} J(x_T) \mathbf{1}_{\{\tau = T\}} \right], \tag{6}$$

where the stopping time τ is $\mathcal{F}_t$-measurable and $\mathbf{1}_{\{\cdot\}}$ is the indicator function. The terminal condition implies that the value of non-protected jobs is equal to that of protected jobs if and only if the relationship lasts until tenure T. As explained in Appendix A, the *continuation region* $\mathcal{C}$ is bounded below by the optimal boundary $R(t)$, i.e. $\mathcal{C} = \{(x, t) \in \mathbb{R}^+ \times [0, T) : x > R(t)\}$. The boundary which separates the continuation region from the termination region does not have a closed form solution but can be approximated numerically using the procedure described in Appendix C.

Properties of the separation threshold. Even though we cannot analytically characterize the separation rule for non-protected jobs, we are able to establish its most salient features.

Proposition 1 *The boundary $R(t)$ is strictly increasing in t and in the level of firing costs F, i.e. $R'(t) > 0$ and $R(t; F) < R(t; F + \varepsilon)$ for all $\varepsilon > 0$ and all $t \in [0, T)$.*

Proof. See Appendix A. ■

Proposition 1 states that the continuation region shrinks over time. As seniority increases, the firm anticipates that it will soon have to transform the job into a protected one; this reduces the option value of waiting and so dissuades firms from hoarding labor. This *anticipation effect* implies that layoff costs F have a negative impact on the duration of non-protected jobs. The left-hand panel of Figure 3 illustrates how the reservation threshold increases regularly until the conversion date T, where it drops suddenly as job protection starts to take effect.

Proposition 2 *The separation threshold is discontinuous at T where it exhibits a negative drop (i.e. $\lim_{t \to T^-} R(t) > R$), the size of which is increasing in the firing costs F.*

Proof. See Appendix A. ■

Proposition 2 shows that the magnitude of the drop is increasing in F for two reasons. First, layoff costs raise $R(t)$ before T, because they reduce the value of protected jobs and thus the option value of non-protected jobs. Second, the barrier R prevailing after T shifts down because it becomes more costly to separate, which fosters labor hoarding after T. The impact of an increase in firing costs F is displayed in the right panel of Figure 3.

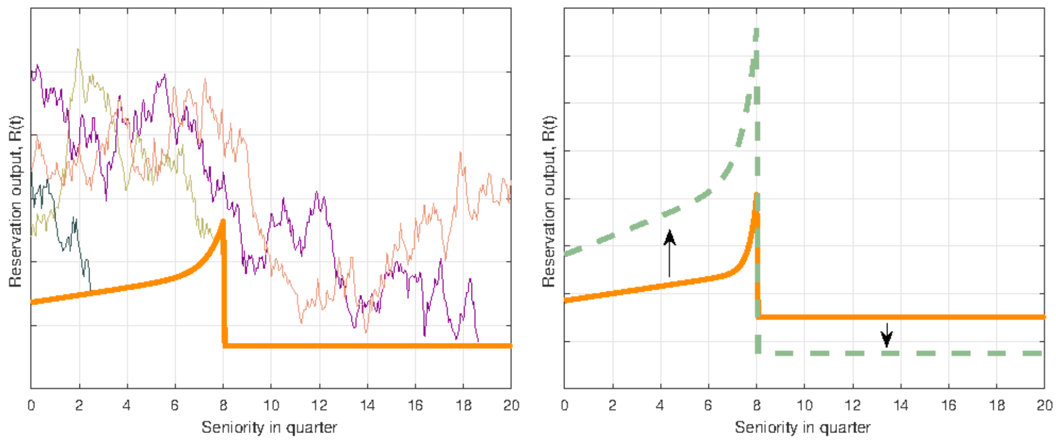

Figure 3: The shape of the reservation productivity (left-hand panel) and the effects of an increase in firing costs (right-hand panel)

Note: In the left-hand panel, the bold line represents the reservation productivity $R(t)$, the thin lines display some output paths; in the right-hand panel, the arrows indicate the effect of firing costs F on the reservation productivity $R(t)$.

To understand why the properties described in Propositions 1 and 2 makes it possible to reproduce the hazard rate displayed in Figure 2, it is useful to recall what happens when the separation threshold is flat. In this case, jobs with unlucky draws are weeded out over time, and the distribution of productivity among surviving jobs becomes more and more skewed to the right. Self-selection mechanically raises average productivity and so exerts downward pressure on the rate of separation, which decreases with seniority.[13]

The natural tendency of the separation rate to decline with seniority is counteracted by the increase in the separation barrier $R(t)$. Balancing the tension between these two countervailing forces enables us to match the empirical hazard rate. Early on, the separation rate decreases because self-selection dominates. Then, as the transformation date nears, the anticipation effect takes over, generating a hike in separations which culminates at the two-year threshold. Finally, when employment protection legislations become binding, the number of separations follows the fall in reservation productivity and drops to a lower plateau.

3.2 Increasing layoff costs and wage bargaining

Figure 1 shows that the average compensation for wrongful termination of labor contracts increases with tenure with an upward jump at the two-year threshold. This suggests that, besides the discrete

[13]For very short job spells, the hazard rate does not always decrease because a job's starting productivity is necessarily higher than its reservation value, so that endogenous separation can only occur after a series of negative shocks. This usually takes some time which explains why the hazard rate is actually hump-shaped in seniority. However, the increasing portion can be made arbitrarily short by setting the initial productivity as close as necessary to the reservation boundary.

hike at two years, layoff costs increase regularly with job seniority. We account for this feature of French legislations by extending our model so as to let layoff costs, now denoted by $F(t)$, increase linearly with tenure t on two segments separated by an upward jump at tenure T:

$$F(t) = \begin{cases} \phi_n t & \text{if } 0 \leq t < T \\ \phi_n T + \phi_0 + \phi_p (t - T) & \text{if } T \leq t \end{cases} \tag{7}$$

where ϕ_n, ϕ_0 and ϕ_p are non-negative parameters.

To account for the fact that there are workers paid above the minimum wage, we extend the model by assuming that wages are bargained over subject to the minimum wage constraint.[14] Wages are continuously negotiated according to the following non-cooperative game. At each point in time, the worker and the employer can make alternating offers. When one of the players offers a wage, the other player either accepts or rejects the offer. If the offer is accepted, then the bargaining ends and the offered contract is implemented. If the offer is rejected, then the game goes on to a next round after a short time delay, denoted by Δ_e if the worker just rejected an offer by the firm or by Δ_f if the firm just rejected an offer by the worker. In the next round, the player who last rejected an offer makes a counteroffer, which again can be either accepted or rejected. The game goes on in this way over an infinite horizon. During the bargaining game, production stops and the worker and the firm bargain according to the value of output at the beginning of the negotiation. The solution to this game, described in appendix B, implies that the worker and the firm agree instantly at each point in time to provide the share $\beta = \Delta_f/(\Delta_e + \Delta_f)$ of the job surplus to the worker. The wage on a job with current output x is

$$w(x) = \max\left\{\beta x + (1-\beta)\, rU + (r+\delta)\left[\beta F(t) + (1-\beta) F_e(t)\right] - \beta \frac{\mathrm{d}F(t)}{\mathrm{d}t} - (1-\beta)\frac{\mathrm{d}F_e(t)}{\mathrm{d}t}, w_{\min}\right\}, \tag{8}$$

where U stands for the value function of an unemployed worker, while $F_e(t) \leq F(t)$ denotes the severance pay which is a component of layoff costs that have different effects on wages than the procedural costs component. In line with the French legislation, we assume that the severance pay increases linearly with tenure t on two segments separated by an upward jump at tenure T:

$$F_e(t) = \begin{cases} \phi_{en} t & \text{if } 0 \leq t < T \\ \phi_{en} T + \phi_{e0} + \phi_{ep}(t - T) & \text{if } T \leq t \end{cases} \tag{9}$$

where ϕ_{en}, ϕ_{e0} and ϕ_{ep} are non-negative parameters.

The negotiated wage increases with the layoff costs because they raise the threat point of the worker in the bargaining process. The wage is negotiated only when current output is above the threshold

$$\bar{x} = \frac{w_{\min} - (1-\beta)\, rU - (r+\delta)\left[\beta F(t) + (1-\beta) F_e(t)\right] + \beta \frac{\mathrm{d}F(t)}{\mathrm{d}t} + (1-\beta)\frac{\mathrm{d}F_e(t)}{\mathrm{d}t}}{\beta}. \tag{10}$$

[14] See Flinn (2006), for a similar approach in a framework without job protection.

Otherwise, when $x \leq \bar{x}$, the minimum wage $w_{\min}$ binds. The job destruction rule is the same as when the wage is exogenous: jobs, the value of which is denoted by $J(x,t)$, are destroyed if output x falls below the reservation productivity $R(t)$ defined by $J(R(t),t) = -F(t)$, with the wage being equal to $w(x)$. Appendix C describes the numerical method used to solve the model with seniority dependent layoff costs and negotiated wages with wage floors.

4 Estimation

4.1 Data and estimation procedure

The parameters of the model are the discount rate r, the death rate χ, the exogenous job separation rate δ, the workers' outside option rU, the bargaining parameter β, the determinist drift μ and the variance σ^2 of the Brownian motion, the coefficients $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$ and $\{\phi_0, \phi_n, \phi_p\}$ parameterizing the effect of tenure on severance payments and on layoff costs, the mean γ and the variance ζ^2 of the distribution of initial productivity which we assume to be lognormal.

The benchmark estimation focuses on the population of individuals aged from 15 to 54, who have completed at most their high school degree and work in the private sector. To avoid overstretching the identification power of the model, we first fix a subset of parameters using direct empirical counterparts. We set the value of the annual discount rate $\rho = r - \chi$ to 4%, while the annual probability of death χ is set to 0.1% , in line with the death rate of the population aged from 19 to 54 years.[15] The exogenous annual job separation rate δ approximates the asymptotic limit of the annual job separation probability, which converges to 2.94% in the Labor Force survey (see Figure 2). The coefficients $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$ parameterizing the effect of tenure on severance payments are chosen to match the average expected severance payments depending on seniority using the Appeal Court rulings data (see Appendix D). The results, displayed in Figure 4, imply that the vector $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$ is equal to $\{0.3368, 0.0401, 0.0422\}$.

We cannot estimate the bargaining power of workers because we would need to observe output per worker, which is not directly available in any data set. We therefore follow the standard practice in the literature and normalize the bargaining parameter β to 0.5. We will assess later the sensitivity of our results with respect to this normalization. A proper identification of the deterministic drift of the Brownian motion μ would also require information about output per worker. Since the returns to tenure are estimated to be very small in France, especially for low skilled workers (Beffy et al., 2006), we use $\mu = 0$ as our benchmark value and report results for different values of μ that are consistent with empirical evidence.

[15]Source: INED, information retrieved on June 5th 2018, at: www.ined.fr/fr/tout-savoir-population/chiffres/france/mortalite-cause-deces/taux-mortalite-sexe-age.

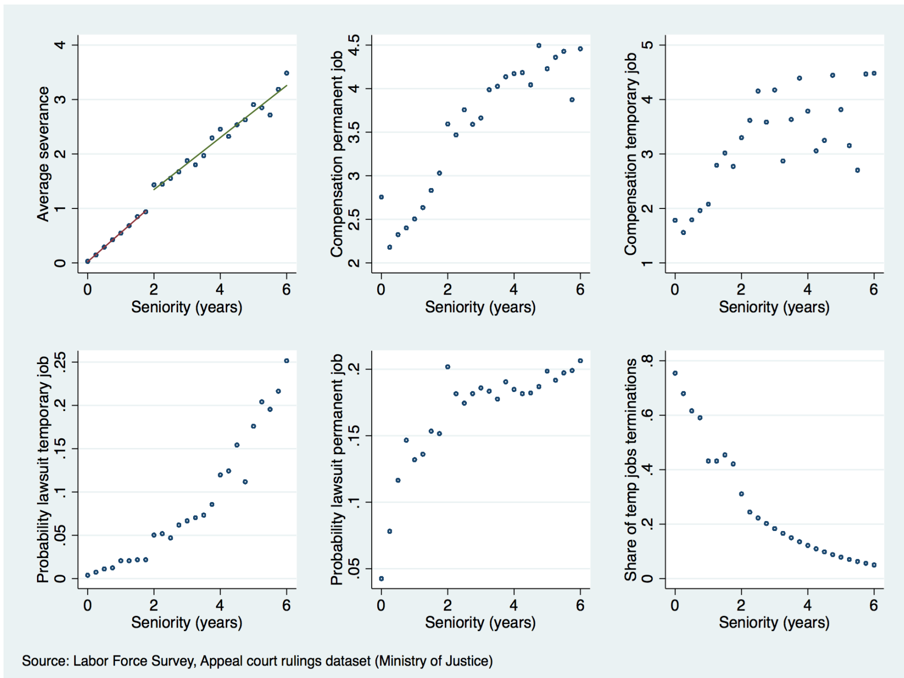

Figure 4: The components of severance payments

Note: The average severance payments are displayed on the top-left graph. They are equal to $F_e(t)$ =stemp(t)[legaltemp(t)+sevtemp(t)] + $[1-$stemp$(t)]$[legalperm(t)+sevperm(t)] where t stands for the years of seniority, stemp(t) is the share of temporary jobs terminations in all job separations (bottom-right graph), legaltemp(t) is the legal severance for termination of temporary contracts, equal to 33% of the monthly wage per year of seniority, sevtemp(t) is the expected compensation in case of wrongful termination of temporary jobs, equal to the probability of prosecution (bottom-left graph) times the average compensation decided by the Appeal courts (top-right graph), legalperm(t) is the legal severance for termination of permanent contracts, equal to 20% of the monthly wage per year of seniority, and sevperm(t) is the expected compensation in case of wrongful termination of permanent jobs, equal to the probability of prosecution (bottom-middle graph) times the average compensation decided by the Appeal courts (top-middle graph). For more details, see Appendix D.

The remaining parameters of the model are estimated by fitting the distribution of quarterly job separation rates according to job seniority and the wage distribution among job entrants. Since it is well known that there is no clear bunching at the minimum wage in France with the Labor Force Survey data (Laroque and Salanié, 2002), we allow for the possibility that hourly wages are not accurately measured and let their observations be contaminated by proportional measurement errors drawn from a lognormal distribution centered around one, with standard deviation denoted by σ_ε.[16] To proceed to the estimation, we split the remaining parameters into two vectors $\Theta \equiv \{rU, \phi_0, \phi_n, \phi_p, \sigma^2\}$ and $\Upsilon \equiv \{\gamma, \zeta^2, \sigma_\varepsilon\}$, and use the following three-stage procedure to estimate their values:

1. For a given value of Θ, we use the numerical algorithm described in Appendix C to simulate the separation barrier $R(t)$.

2. We infer the vector Υ, which contains the mean γ and the variance ζ^2 of initial productivities, by maximum likelihood on the distribution of wages among job entrants.[17] As explained in Appendix E.2, the estimates depend on rU and on the value of the separation barrier for new jobs $R(0)$ inferred in the previous step or, to put it differently, on the vector Θ.

3. We choose the values of Υ and Θ that maximize the likelihood on the distribution of wages among job entrants and that minimize the distance between the predicted and the actual distribution of job separation rates.[18]

Note that our estimators are not 'two-step' estimators because the optimal vector Υ cannot be estimated independently from Θ. Instead Υ depends on Θ and on the vector of observed wages among job entrants. Hence we do not use Υ to minimize the distance between predicted and actual hazard rates, but instead to approximate available information about the distribution of wages. This approach should be favored when one suspects that including Υ in Θ induces a significant bias. Indeed, bypassing the intermediate step of our estimation procedure in favor of full-information estimation would greatly underestimate the amount of wage dispersion observed in the data.

Given that the estimation by maximum likelihood of Υ is rather standard, we relegate its description to Appendix E.2 and focus instead on the last step of the estimation procedure. We describe how the empirical moments are constructed and establish the asymptotic properties of

[16]See Appendix E.2 for the exact specification of measurement errors.

[17]In practice, we include the wages of all workers whose seniority is at most equal to two quarters.

[18]We rely on Monte Carlo methods to infer the job separation rates predicted by the model. We simulate 500.000 paths over a 20 year horizon, with 40 sample points per year. Then we use Brownian bridges to infer the probability that each path crosses the separation barrier. A series of robustness checks indicates that increasing the number of paths or varying the size of the time intervals does not affect the outcome of the simulations.

our estimator. Each observation in our sample provides information about the seniority of the worker and whether the worker has been separated from his job during the quarter. To put it differently, each observation $x_k = \{t_k, d_k\}$ contains two informations: the length of the job spell t_k, and whether the job is destroyed by the end of the quarter

$$d_k = \begin{cases} 1 \text{ if the job is destroyed} \\ 0 \text{ otherwise} \end{cases} .$$

The distribution of job separation rates is estimated from the sample $(x_1, ..., x_n)$. It is assumed that the data come from a statistical model defined up to the unknown vector Θ of parameters. Let us denote by $h_t(\Theta)$ the hazard rate predicted by the model conditional on Θ. Then let us also define, for each tenure t, the indicator function $f(d_i, t)$ which takes value one for every separation occurring at tenure t among all workers who reached tenure t. The moment condition reads

$$\mathbb{E}\left[f(d_i, t) - h_t(\Theta)\right] = 0,$$

where the sample counterpart of $\mathbb{E}\left[f(d_i, t)\right]$ is

$$\hat{h}_t \equiv \frac{\sum_{i=1}^{n} \mathbf{1}_{\{t_i = t\}} d_i}{\sum_{i=1}^{n} \mathbf{1}_{\{t_i = t\}}}. \tag{11}$$

Proposition 3 *Provided that the number of periods T exceeds the size of the parameter vector Θ, an efficient minimum distance estimator of Θ solves*

$$\hat{\Theta} = \arg\min_{\Theta} \left\{\hat{\mathbf{h}} - \mathbf{h}(\Theta)\right\}' \hat{\Omega}^{-1} \left\{\hat{\mathbf{h}} - \mathbf{h}(\Theta)\right\}, \tag{12}$$

where $\hat{\mathbf{h}} = (\hat{h}_1, .., \hat{h}_T)$ is the vector of empirical hazard rates defined in (11), $\mathbf{h}(\Theta) = (h_1(\Theta), ..., h_T(\Theta))$ is its theoretical counterpart, and $\hat{\Omega}$ is a symmetric and positive definite weighting matrix, the expression of which is given in $(E21)$.

Proof. See Appendix E. ■

4.2 Results

The estimates are reported in Table 1. Initial productivities are sampled from a lognormal distribution with mean $\gamma = 0.95$ slightly lower than the hourly minimum wage, the value of which is normalized to one. The workers' outside option rU, which includes the value of searching for a new job, is above the minimum wage. Workers nonetheless find it worthwhile to accept jobs that pay the minimum wage because productivity, and thus future wages, do not remain constant. This *growth option* drives a negative wedge between the reservation productivity and the opportunity cost of employment. The size of the gap is proportional to the volatility parameter σ.

Table 1: Minimum distance estimated parameters

Parameter	Notation	$\mu = 0$	$\mu = 0.005$	$\mu = 0.01$
Workers' outside option	rU	1.2041 (.0179)	1.2370 (.0325)	1.2720 (.0364)
Std. Deviation Brownian motion	σ	.2395 (.0013)	.2360 (.0019)	.2316 (.0040)
Discontinuity of firing costs at 2 years	ϕ_0	.1193 (.0029)	.1161 (.0027)	.1176 (.0017)
Slope firing costs before 2 years	ϕ_n	.2303 (.0028)	.2275 (.0012)	.2328 (.0050)
Slope firing costs after 2 years	ϕ_p	.0653 (.0009)	.0676 (.0004)	.0647 (.0011)
Mean initial productivity	γ	.9460 (.0077)	.8994 (.0208)	.8488 (.0251)
Std. Deviation initial productivity	ζ	.3083 (.0006)	.3042 (.0024)	.2989 (.0037)
Std. Deviation measurement errors	σ_ε	.1200 (.0004)	.1212 (.0011)	.1227 (.0014)

Note: When applicable parameters are for annual frequency. Firing costs parameters are proportional to the annual minimum wage. The model is estimated for different values of the deterministic drift μ of the Brownian motion consistent with empirical evidence in Beffy et al. (2006). μ is set to zero in the benchmark case used in the rest of the paper. Bootstrapped standard errors in parentheses, with 100 resampled data sets and 500,000 paths.

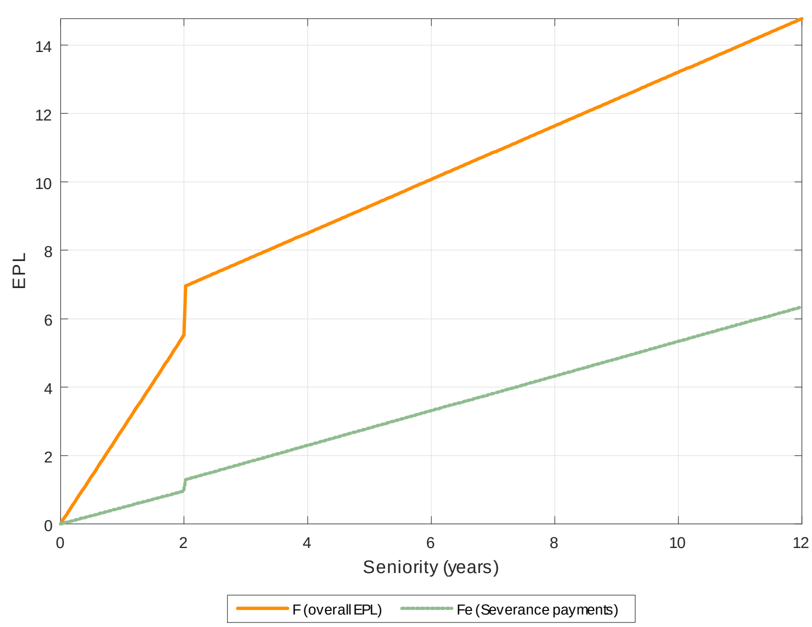

Figure 5: The relation between estimated layoff costs, severance payments and job seniority

Note: Layoff costs F, measured in share of the monthly minimum wage, are estimated from the model. Severance payments F_e are computed from the data as explained in Appendix D.

Total layoff costs, F, which comprise the severance payments and the procedural costs are not observable directly. They are estimated by fitting the job separation rate and the wage distribution of job entrants. Figure 5 shows that total layoff costs increase with seniority with a jump at the two-year threshold where there is a sudden hike equivalent to about two months of the minimum wage. The total layoff costs hike at two years is triggered by the increase in average severance payments displayed in Figures 4 and 5. It is clear that procedural costs are significant, as they are approximately four times larger than the expected severance payments obtained by employees at two years of seniority and 1.5 larger at ten years of seniority. The procedural costs can arise from the time and resources needed to deal with rules and procedures, which are particularly complex in France,[19] from the length of procedures, from the advanced notice and the lawyers fees. It is worth noting that the procedural costs are significant from the beginning of the employment relationship. This finding is consistent with the empirical probability of lawsuit, which is positive for the termination of temporary and permanent jobs, from the very beginning of employment spells, as shown by Figure 4. From the perspective of employers and workers, procedural costs are red tape, which reduce the surplus of jobs, wages, profits and job duration.

Table 1 shows that these results hold for different values of the deterministic drift of the Brownian motion μ consistent with empirical evidence, since the estimated values of ϕ_0, ϕ_n and ϕ_p, which determine the relation between layoff costs and seniority, vary very little with μ. The robustness of the cost estimates underscores the identifying power of our estimation strategy. Its precision stems from the discontinuity in layoff costs and how it interacts with the anticipation channel. While the fall in the separation rate at two years is mostly determined by the size of the hike in layoff costs ϕ_0, the progressive increase in the separation rate before two years results from the interplay between ϕ_0 and the other parameters. Hence fitting how the separation rate evolves with tenure can only be achieved through the identification of a specific combination of parameters.

The accuracy of the model is documented in Figures 6 and 7. Figure 6 displays the simulated hazard rate of job separation and its empirical counterpart. The model does a fairly good job at tracking the data. In particular, the kink at two years of seniority is reproduced by the model. It

[19] Indeed, dismissal procedure is strict and intricate under French law. The employer must give written notice, calling the employee to a preliminary meeting. This notice must mention several pieces of information such as possible assistance for the employee during the preliminary interview. There must be a full 5-working-days delay between the notice and the interview itself. During the preliminary interview, the employer must explain the reasons for the dismissal and has to listen to his employee's remarks. The employer should make no statements as to his decision regarding dismissal. The employer may be assisted by one of his employees but cannot be represented by someone who is not part of the company. Then, following a full 2-working-days delay after the interview, the dismissal letter must be sent by registered mail. The termination letter should mention reasons for dismissal. Content and form of the letter are important. Upon delivery of the dismissal letter, the notice period starts. It generally varies between one and three months depending on the employee's status. More formal rules, requiring state authorization, apply to union representatives and staff representatives. Non compliance with procedural rules is enough to claim damages even when there are proper grounds to dismiss the employee. See Ray (2017).

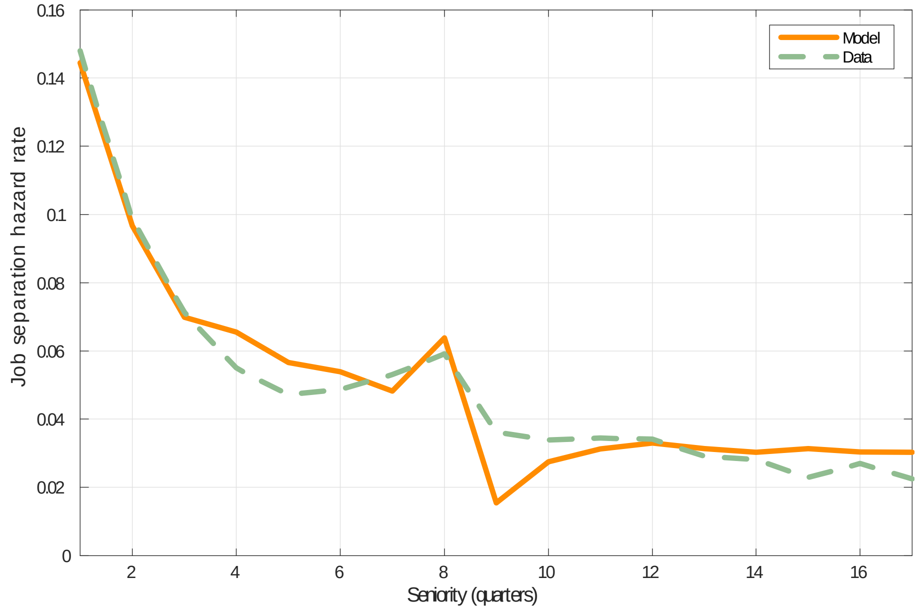

Figure 6: Actual and simulated hazard rates of job separation
Note: Baseline parameters reported in Table 1.

is slightly larger in the simulation than in the data. This could partly be explained by reporting errors around the two-year threshold. We also assess whether our estimates fit the wage distribution among job entrants and among workers with two years of seniority by plotting in Figure 7 the empirical distributions and their estimations by the maximum likelihood procedure described in Appendix E.2. Again, we find that the model accurately replicates the data.

4.3 Counterfactuals

We now use the structural model to simulate the impact of layoff costs on job turnover. Two experiments are conducted: one where the jump at two years of seniority is deleted, i.e. $\phi_0 = 0$; and a more drastic change where all employment protection is eliminated by setting $\phi_0,\ \phi_n$ and ϕ_p equal to zero. Additional scenarios, with more complex reforms, are considered in Appendix G.

As expected, eliminating the upward jump in layoff costs irons out the kink at two years and reduces the job separation rate at the beginning of the employment spell (see Figure 8). This shows that the perverse impact of the upward jump in layoff costs arises because layoff costs *raise* the reservation barrier during the first two years of seniority. The upward jump is counterproductive since it does not protect jobs but instead intensifies their destruction. This is driven by the *anticipation effect* discussed in the previous section: As firms forecast that they will have to pay higher costs in the future in the event of a separation, comparing future with current layoff costs, they find it optimal to anticipate the separation decision. Due to this anticipation effect, the jump at two years of

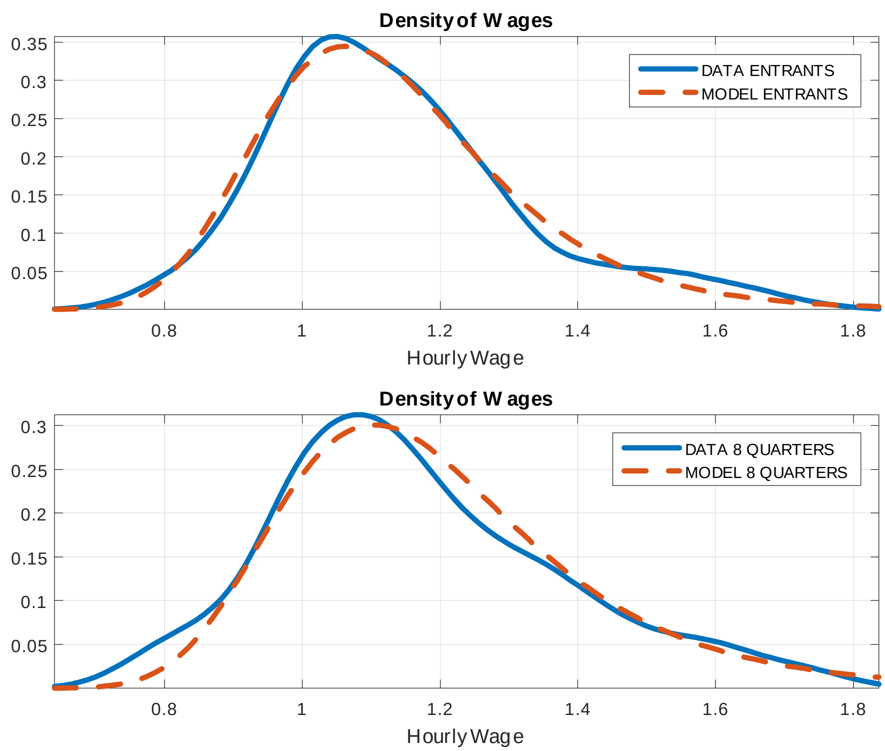

Figure 7: Estimated and empirical densities of hourly wages expressed in units of hourly minimum wages among job entrants

Note: The top graph displays the wage densities of entrant workers with 2 quarters of seniority at most. The bottom graph displays the wage densities of workers with 2 years of seniority. Source: Labor Force Survey (Enquête emploi, INSEE), 2003-2012.

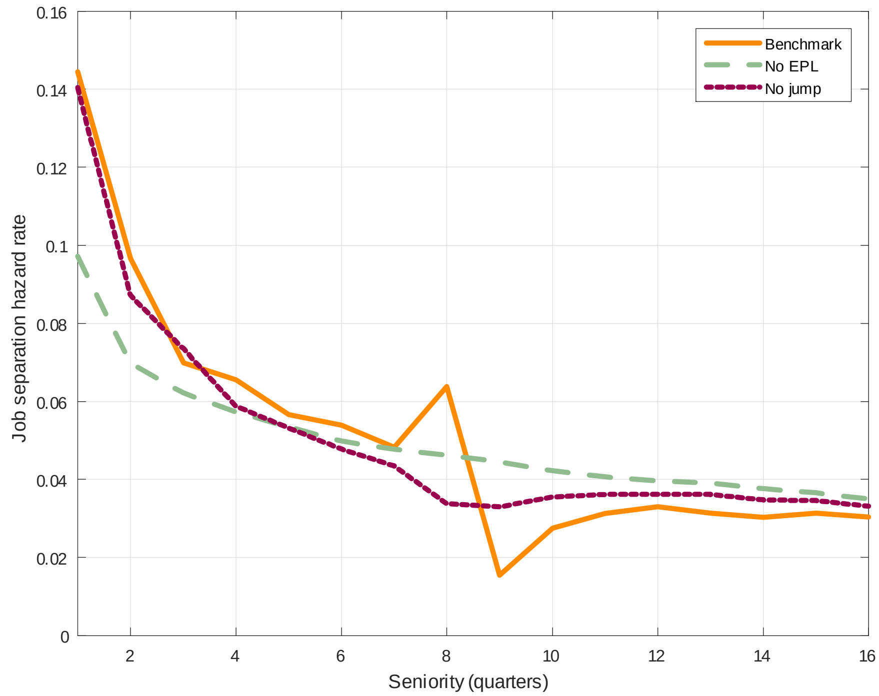

Figure 8: Impact of firing costs on hazard rates Note: Baseline parameters reported in Table 1.

seniority induces an increase in the job separation rate before the two-year threshold. Instead of taking the risk of turning weakly protected jobs with a low productivity into strongly protected jobs, firms prefer to terminate the employment relationship just before the deadline; which results in a sudden hike of the separation barrier and explains the kink in the hazard rate. In our context, this effect is sufficiently strong to *reduce* the average job duration (i.e. the average expected duration from the starting date of jobs) which drops from 5.78 to 5.71 years when the jump in labor costs at two years is introduced (see Table 2, Column (1)). For a given rate of job arrival, this drop in the average job duration is necessarily associated with lower employment.[20] By contrast, job protection increases the average and the median elapsed durations (see Table 2, bottom part of Column (1)). Hence job protection lowers the separation rate for a majority of employed workers, at each point in time. This tension between the impact of job protection on the average expected duration and on the average elapsed duration helps explaining why a majority of workers may support job protection despite its negative effect on employment.

Table 2 indicates that eliminating all layoff costs increases the average expected job duration with respect to its value prevailing under current French regulation. Figure 8 shows that the increase in job duration associated with the elimination of layoff costs is induced by the strong decrease in the job separation rate at the beginning of the employment spell. The effect is quantita-

[20]Let us remind that steady state employment is equal to the job arrival rate times the average expected duration of jobs from their starting date.

Table 2: Average job duration (in year) and layoff costs

	Benchmark		Skilled		Unskilled		Youth	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Mean	Median	Mean	Median	Mean	Median	Mean	Median
Duration (flows)								
Actual EPL	5.7123	1.7627	6.4996	1.9131	5.5292	1.7339	4.6926	1.5558
No EPL	5.9006	2.0456	6.2998	2.0205	5.8300	2.0640	5.1875	1.8943
No jump	5.7846	1.8083	6.6376	2.0080	5.6056	1.7799	4.8242	1.6091
Elapsed duration								
Actual EPL	6.6920	5.2817	6.9509	5.6375	6.6165	5.1816	6.3656	4.8477
No EPL	6.4659	4.9786	6.7440	5.3520	6.3981	4.8898	6.1814	4.5986
No jump	6.6481	5.2148	6.9171	5.5871	6.5731	5.1161	6.3183	4.7688

Note: "Duration (flows)" stands for the expected jobs duration at their starting date. "Actual EPL" stands for the layoff costs estimated in the benchmark case corresponding to the French EPL. "No EPL" is the case without layoff costs, "No jump" corresponds to the actual EPL case with no jump in layoff cost at 2-year seniority threshold, i.e. $\phi_0 = 0$ in Table 2. "Elapsed duration" stands for the average duration of jobs sampled in the stock of jobs. "Benchmark" covers the population of 15-54 years old individuals with at most the high school degree used to estimate the model in the benchmark case. "Skilled" covers the population of 15-54 years old individuals with at least the high school degree. "Unskilled" covers the population of 15-54 years old individuals with less than the high school degree. "Youth" covers the population of individuals with at most the high school degree and who left school at most 10 years ago.

tively important because the mode of the productivity density at the start of the employment spell is close to the threshold value below which jobs are destroyed (see Figure 9). This proximity is instrumental in allowing the model to capture the fragility of jobs with short tenure, which is a well documented stylized fact. Therefore, the strong impact of future layoff costs on the job separation rate at the start of the employment spell relies on a property which is empirically relevant, and that is not captured by standard search and matching models, where productivity changes are driven by a Poisson process which triggers random draws from a stationary productivity distribution. Moreover, this finding is in line with empirical studies relying on natural experiments which consistently find that more stringent job protection increases the turnover of jobs with short tenure, but has either very weak or no significant impact on the turnover of jobs with long tenure (Schivardi and Torrini, 2008; Centeno and Novo, 2012; Hijzen et al., 2017). Overall, in our framework, getting rid of all layoff costs has a significant impact on average expected employment duration, which goes from 5.71 to 5.90 years. Assuming constant job arrival rate, this corresponds to an increase in steady state employment of 3.2%.[21]

Other experiments, in which the profile of layoff costs is changed while their average amount over 20 years of seniority is kept constant are presented in Appendix G. It turns out that removing the jump at two-year seniority still increases expected job duration because job destruction is

[21]Table 5 in Appendix F shows that these results remain qualitatively unchanged when the value of the bargaining parameter β differs from its level of 0.5 in the benchmark estimation.

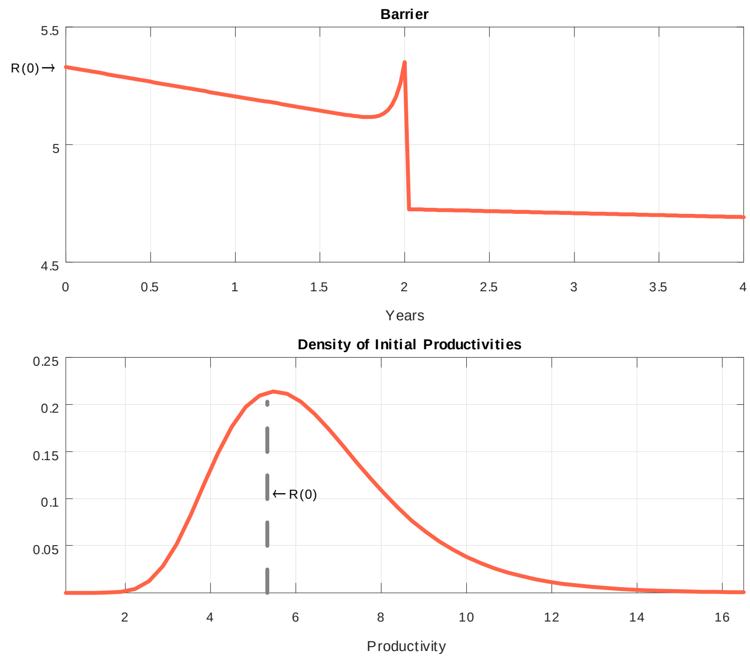

Figure 9: The relation between seniority and the separation barrier (top graph); the density of starting productivities (bottom graph)
Note: $R(0)$ denotes the value of the barrier for workers with zero seniority. The separation barrier is the threshold value of productivity below which jobs are destroyed.

reduced at the beginning of employment spells. Appendix G also shows that the removal of layoff costs before two years of seniority reduces average employment duration, because the negative impact of the anticipation of future layoff costs, after two years of seniority, is no longer counterbalanced by job protection before two years.

It is worth mentioning that the minimum wage amplifies the negative impact of layoff costs on profits since wages cannot adjust downwards when the minimum wage binds. In order to illustrate this phenomenon, we estimate the model for different categories of workers. First, skilled workers, belonging to the same age group, i.e. 15-54 years old, who have at least a high school degree, and whose wage is higher than that of the benchmark population, as shown in Appendix H.[22] We also consider two other categories of workers whose wages are lower: 1) unskilled workers belonging to the same age group, i.e. 15-54 years old, whose level of education is below high school graduation; 2) young workers, who have left school less than 10 years ago, and who have achieved at most high school graduation. Table 2 shows that job protection *increases* the average expected job duration of skilled workers (Column (3)) but has an opposite impact on the average expected job duration of unskilled workers, as in the benchmark case. It is also clear from Table 2, Columns (1), (5), (7),

[22]The population in the benchmark estimation comprises 15-54 years old individuals whose education level is at most equal to high school graduation.

that the negative impact of job protection on the expected job duration is stronger for low wage workers. This amplification is due to the fact that the average surplus of starting jobs is lower for unskilled workers, and that there are less possibilities to compensate the increase in future labor costs induced by job protection with downward wage adjustments when wages are closer to the minimum wage. All in all, these results suggest that job protection may raise the expected duration of skilled employment but is detrimental to the job stability of low wage workers, especially when there is a binding minimum wage. This implies that job protection is likely to be counterproductive for low skilled employment in the presence of wage floors, while its effect on skilled employment is more uncertain. The next section provides further evidence supporting this conclusion.

5 Employment Protection and Unemployment

The previous section identifies the impact of layoff costs on job duration and on employment assuming exogenous job arrival rates. Hence it does not take into account the impact of job protection on job creation. To fix this shortcoming, we now close the model to endogenize the job arrival rate, an extension which enables us to evaluate the impact of employment protection legislations on the unemployment rate. We start by presenting the labor market equilibrium with endogenous job creation. Then, we proceed to the estimation of the parameters of the model before analyzing its predictions about the effects of job protection.

5.1 Labor market equilibrium

In order to account for job creation, it is assumed that there is free entry into the labor market. Firms must post a job vacancy, which costs κ per unit of time, to recruit a worker. Unemployed workers and job vacancies are brought together through a constant returns to scale matching technology which implies that vacant jobs are filled at rate $q(\theta)$, with $q'(\theta) < 0$, where θ denotes the labor market tightness, equal to the ratio of the number of job vacancies over the number of job seekers. Once matches are created, firms draw the output per worker at the start of their employment spell from a sampling distribution, the cumulative distribution function of which is denoted by H. As discussed before, jobs are created only if the output per worker is above the threshold $R(0)$. Thus, the value of a vacant job, denoted by V, satisfies

$$\rho V = -\kappa + q\left(\theta\right) \int_{R(0)}^{\infty} \left[J\left(x,0\right) - V\right] \mathrm{d}H(x), \tag{13}$$

where $J\left(x,0\right)$ stands for the value of a starting job that produces x per unit of time. The free entry condition, $V = 0$, implies that

$$\frac{\kappa}{q\left(\theta\right)} = \int_{R(0)}^{\infty} J\left(x,0\right) \mathrm{d}H(x). \tag{14}$$

The values of $J\left(x,0\right)$ and $R(0)$ have been computed in the benchmark model with exogenous job arrival rates. They depend on one unknown variable only: the discounted expected utility of unemployed workers U. Hence equation (14) has two endogenous variables: the labor market tightness θ and U. The discounted expected utility of unemployed workers satisfies

$$rU = b + \theta q\left(\theta\right) \int_{R(0)}^{\infty} \left[W\left(x,0\right) - U\right] \mathrm{d}H(x), \tag{15}$$

where b is the utility flow of unemployed workers. $W(x,0)$ denotes the discounted expected utility of an employee who starts on a job which produces x per unit of time, and so satisfies the following Bellman equation

$$rW(x,0) = \begin{cases} rU + \frac{\beta}{1-\beta} rJ(x,0) & \text{if } x \geq \bar{x}, \\ w_{\min} + \delta\left[U - W\left(x,0\right)\right] + \frac{\mathbb{E}[\mathrm{d}W(x,0)]}{\mathrm{d}t} & \text{otherwise,} \end{cases} \tag{16}$$

where $\mathbb{E}\left[\mathrm{d}W\left(x,0\right)\right]/\mathrm{d}t$ depends on U, on the parameters of the geometric Brownian motion and on the threshold $\bar{x}$ defined in equation (10). Substituting the definition of $W(x,0)$ into (15) yields a system of two equations, (14) and (15), in two unknowns, θ and U. Equation (14) defines a negative relation between θ and U because $J\left(x,0\right)$, the value of filled jobs, decreases with the reservation utility of workers. Equation (15) defines a positive relation between the two endogenous variables because the expected utility of unemployed workers increases with the labor market tightness θ. Therefore, equations (14) and (15) define a unique solution (θ, U), the existence of which is assumed.

Once the equilibrium values of θ and U are determined, it is possible to get the equilibrium job arrival rate, $\lambda = \theta q(\theta)$, and the equilibrium job separation rate, $h\left(t\right)$. Since the job separation rate depends on job seniority, the unemployment rate depends on the distribution of times elapsed since labor market entry, or, in other words, on the distribution of labor market experience, denoted by e. The relation between labor market experience and unemployment can be computed from the job separation rate $h\left(t\right)$ and from the unemployment exit rate $\lambda\left[1 - H(R(0))\right]$. It is assumed that the unemployment exit rate does not depend on labor market experience. Then the unemployment rate as a function of labor market experience, $u(e)$, obeys a Volterra integro-differential equation

$$\frac{\mathrm{d}u\left(e\right)}{\mathrm{d}e} = \int_{0}^{e} \lambda\left[1 - H(R(0))\right] u\left(s\right) h\left(e-s\right) \mathrm{d}s - \left(\lambda\left[1 - H(R(0))\right] + \chi\right) u\left(e\right), \tag{17}$$

where $(\lambda\left[1 - H(R(0))\right] + \chi) u\left(e\right)$ is the mass of unemployed workers of labor market experience e who find a job or die, while the term $\int_0^e \lambda\left[1 - H(R(0))\right] u\left(s\right) h\left(e-s\right) \mathrm{d}s$ measures the mass of employees of labor market experience e who lose their job. In this integral, $\lambda\left[1 - H(R(0))\right] u\left(s\right)$ is the mass of workers who found a job at labor market experience s, while $h\left(e-s\right)$ is the probability that they lose that job exactly when they reach labor market experience e.

In order to bring the model to the data, it is assumed that the matching function is Cobb-Douglas and homogeneous of degree one, which implies that the number of hires per unit of time is equal to $m\left[1-H(R_0)\right]u^{1-\eta}v^{\eta}$, with $\eta \in (0,1)$, $m > 0$, where u and v stand for the number of unemployed workers and of vacant jobs, respectively.

At this stage, the following parameters of the model remain to be determined: the two parameters of the matching function m and η, the cost of vacancy posting, κ, and the utility flow of unemployed workers, b.

5.2 Estimation of the matching function

The estimation of η, the elasticity of the matching function with respect to job vacancies, relies on data on unemployment and job vacancies for low skilled workers which come from the French employment agency (*Pôle emploi*). Firms can post job vacancies at *Pôle emploi.* This is a free service and *Pôle emploi* estimates that they deal with almost 50% of the total of French vacancies. These data allow us to compute the labor market tightness, as the ratio of the number of job vacancies posted at the employment agency over the number of unemployed workers registered at the employment agency, at the commuting zone level for blue collars and low skilled white collars for each year from 2009 to 2011. There are 348 commuting zones. Data on hires of blue collars and low skilled white collars at the commuting zone level over the same period come from two data sets provided by the French Ministry of labor. The DMMO register (Déclaration Mensuelle de Mouvements de Main d'Oeuvre), which describes establishments job flows (entries, exits, jobs created and lost, etc.) by type of contract, gender, age, occupational category. This is an administrative register which is mandatorily filled by all establishments with more than 50 employees. Information for establishments with fewer than 50 employees relies on the EMMO survey (Enquête sur les Mouvements de Main d'Oeuvre), which is a quarterly survey providing the same information as the DMMO register.

We proceed to the estimation of the parameter η of the matching function taking the logarithm of the job finding rate

$$\log f = \eta \log \theta + \psi,$$

where $\theta = v/u$ and $\psi = \log m\left[1-H(R(0))\right]$. The OLS estimates of this equation are exposed to an endogeneity bias arising from the search behavior of agents on either side of the market. For instance, improvements in the matching technology parameter m can raise the labor market tightness θ and the hiring rate (see Borowczik et al., 2013). This implies a potential correlation between the residuals of the OLS estimation and the labor market tightness which can bias downwards the OLS estimate of the coefficient for the labor market tightness. Thus we need exogenous variations in labor demand to identify the labor market tightness coefficient. We address this issue by using variations across commuting zones over time and we rely on IV estimation following the approach of Bartik (1993). The shift in labor demand in commuting zone j is instrumented by the weighted average of the national rates of growth of the number of entries into employment across industries using commuting zone j industry entries shares as weights.

We measure the tightness (θ_{jt}) and unemployment at the commuting zone level j at date t from the employment agency data and the hires from the establishment data. Using f_{jt} to denote the annual job finding rate, we estimate the following equation

$$\log f_{jt} = a_1 \log \theta_{jt} + \sum_t b_t 1\left[date = t\right] + c_j + \psi_{jt}, \tag{18}$$

where j is one of the 348 commuting zones and the date t varies from 2009 to 2011. The estimation controls for date dummies and commuting zones fixed effects (c_j). Equation (18) is estimated by standard (within) OLS regression, taking first difference to eliminate the commuting zone fixed effect. The shift in labor demand in commuting zone j at date t is instrumented by the variable $z_{jt} = \sum_i \bar{s}_{ij} E_{ijt}$, where $\bar{s}_{ij}$ denotes the average share of entries in industry i in commuting zone j in 2005-2006, and E_{ijt} denotes the growth rate of the number of entries in industry i in year t, in all commuting zones different from commuting zone j. Lagged values of z_{jt} are also used as instruments. These instruments are strongly correlated with the labor market tightness, as shown by Table 3. This table reports the estimates of the coefficient a_1 using OLS in column 1 and IV in column 2. Both estimates are highly significant. However, the OLS estimate is lower than the IV estimate, as expected. Taking the IV estimation as our preferred strategy, the elasticity η of the matching function with respect to the number of vacancies amounts to 0.43. This estimate is in the range of those found in previous studies (see e.g. Petrongolo and Pissarides, 2001; Coles and Petrongolo, 2008 or Borowczik et al., 2013).

Once η, the elasticity of the matching function with respect to the number of job vacancies, has been estimated, it is possible to infer the efficiency parameter m. To do this, we use the one-to-one mapping between the job finding rate $\lambda\left[1 - H(R\left(0\right))\right]$ and the labor market tightness θ defined by the matching function. In order to bring this relationship to the data, we insert the estimate of $1-H(R\left(0\right))$, the share of contacts that lead to actual job creations,[23] into the Volterra equation (17), and then simulate the equation to obtain the relation between the unemployment rate and labor market experience. In line with the data, we assume that individuals with at most a high school degree enter the labor market at age 18, and set the initial level of unemployment equal to its average value of 44.1%. Then we pick the value of λ, assumed to be independent of labor market experi-

[23]The estimates of $R\left(0\right)$ and $1 - H(R\left(0\right))$ are reported in Figure 9 and in Table 4, Column (1), respectively.

Table 3: Estimates of the parameters of the matching function

	(1)	(2)
	OLS	IV
Dep. var.		Labor market tightness (log)
		First stage
Entries		$.63^{***}$
		$(.05)$
Entries (-1)		-1.31^{***}
		$(.17)$
Entries (-2)		-0.65^{***}
		$(.12)$
Dep. var.		Job finding rate (log)
		Second stage
Labor market tightness (log)	$.38^{***}$	$.43^{***}$
	$(.07)$	$(.15)$
Date FE	Yes	Yes
R^2	0,33	
Nb. Observations	879	879

Source : Pôle emploi and EMMO-DMMO. Note: Estimation of the parameter of the job matching function equation (17) on 348 employment pools from 2005 to 2010. Labor market tightness (log) stands for the first difference in the log of the labor market tightness. Job finding rate (log) stands for the first difference in the log of the job finding rate.
(1) Standard OLS ; (2) IV regression. As instruments we include commuting zone fixed effects and we use the Bartik type instrument described in the text. 'Entries' stands for the weighted average of national growth rates of the number of entries into employment across industries using by the commuting zone industry shares averaged on 2005-2006 as weights. For each commuting zone j, the national growth rate of the number of entries in industry i in year t is equal to the growth rate of entries in industry i in year t in all commuting zones different from commuting zone j. 'Entries (-1)' and 'Entries (-2)' are the one year and two year lagged values of 'Entries' respectively. Robust standard errors in parentheses. * significant at 10 percent, ** significant at 5 percent, *** significant at 1 percent.

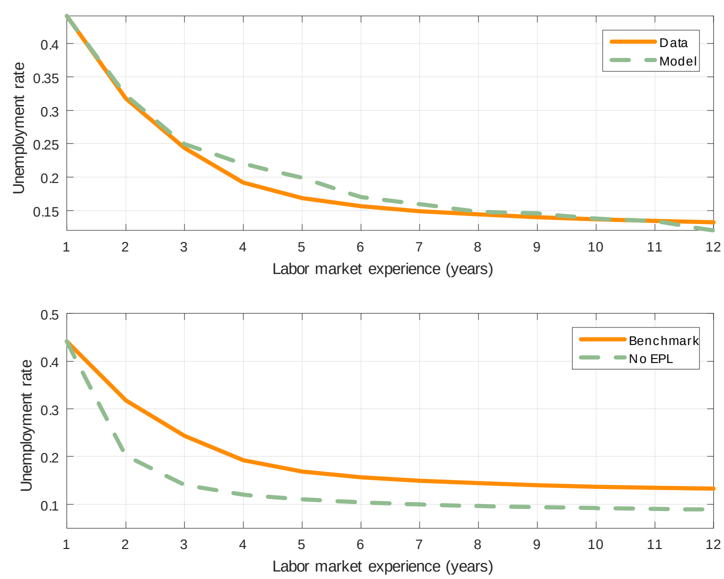

Figure 10: Unemployment rate as a function of labor market experience
Source: Labor Force Suvey (Enquête Emploi, INSEE), 2003-2012. Top panel: Actual and simulated (benchmark) unemployment rate; Bottom panel: Benchmark and EPL free unemployment rate.

ence, which minimizes the distance between the data and the simulation of the Volterra equation for the unemployment series. Our best estimate yields a value of λ equal to $1.06/0.6804 \approx 1.55$ (see Table 4) and is reported in Figure 10. The model matches the data reasonably well, especially considering that we are using only one parameter, namely λ, to fit the whole unemployment series. The equilibrium values of m, the efficiency parameter of the matching function, and of the labor market tightness θ, are then inferred from the value of the job finding rate $\lambda\left[1-H(R\left(0\right))\right]$ and from the average duration for vacant jobs, equal to 0.15 year in our data.[24]

5.3 Calibration and results

At this stage, there remains two parameters to determine: κ, the vacancy creation cost, and b, the utility flow of unemployed workers. Their values are pinned down by the free entry condition (14) and by the estimated value of the opportunity cost of employment (15). Average recruitment and search costs amount to 8 and 5.2 months of the minimum wage, respectively.[25] While the recruitment costs are standard, the search costs are noticeably higher than those commonly used in the search-matching literature. The discrepancy is not specific to our model. As explained in Hornstein et al. (2011), large search costs, or low value of non-market activity, are required to rationalize the amount of wage dispersion observed in the data. The search costs predicted by our model are actually smaller than those resulting from standard search models with constant productivities and exogenous separations for the following two reasons. First, the average lifespan of jobs in our model is proportional to their starting values. This lowers the opportunity cost of accepting a low-paying job because the relationship is not likely to last, so that the worker will soon be free to search for a better job. Second, our estimates are based on the wage distribution among job entrants, whose dispersion is much smaller than that of the wage distribution among all workers.

The features of the economy in the benchmark case, corresponding to individuals aged from 15 to 54 with at most a high school degree, are reported in Table 4, Column (1). The unemployment rate amounts to 14.2% and the annual job finding rate, equal to 1.06, implies that the average unemployment duration is close to one year. Column (2) describes the situation without employment protection, where all the firing costs parameters are set to zero. The results indicate that the regulation has a strong detrimental impact on employment. Removing all employment protection reduces the aggregate rate of unemployment by nearly 5 percentage points, from 14.2% to 9.5%. This strong drop is induced by more job creations and by the lengthening of job durations. The comparison of Tables 2 and 4 shows that the latter effect is qualitatively similar when the arrival of

[24]The job finding rate $m\theta^{\eta}\left[1-H(R\left(0\right))\right]$ and the vacancies duration $\theta^{\eta-1}/m\left[1-H(R\left(0\right))\right]$ define a system in 2 unknown variables, θ and m, insofar as the values of the job finding rate, the vacancies duration, $H(R\left(0\right))$, and η are known.

[25]The average recruitment and average search costs are equal to $\kappa/q\left(\theta\right)$ and $-b/\left[\theta q\left(\theta\right)\right]$.

Table 4: General Equilibrium Impact of EPL

	Benchmark		Skilled		Unskilled	
	(1)	(2)	(3)	(4)	(5)	(6)
	Actual EPL	No EPL	Actual EPL	No EPL	Actual EPL	No EPL
Unemployment rate	14.59%	9.80%	8.53%	6.46%	23.54%	16.01%
Job finding Rate	1.02	1.59	1.65	2.33	0.58	0.91
Successful contacts	68.40%	82.76%	39.69%	47.92%	73.56%	87.56%
Durations						
Mean (flows)	5.71	5.78	6.49	6.19	5.52	5.72
Median (flows)	1.76	2.01	1.91	1.99	1.73	2.02
Mean (elapsed)	6.69	6.42	6.95	6.71	6.61	6.37
Median (elapsed)	5.28	4.92	5.63	5.31	5.18	4.85

Note: "Job Finding Rate" is at annual frequency. "Share Successful Contacts" is the proportion of meetings that induce hires. See Table 2 for the definitions of the other statistics.

job offers is exogenous, but that it is subdued when job creation is endogenized. Quite intuitively, the reaction of job creations implies that the removal of job protection raises the workers' outside option, which leads in turn to more job separations. This feedback effect counteracts the positive impact of deregulation on employment duration identified in the partial equilibrium setting.

The third line of Table 4 indicates that the increase in the job finding rate is partly due to a higher rate of successful contacts, i.e. a greater share of jobs whose initial productivity lays above the reservation productivity $R(0)$. In principle, $R(0)$ could increase or decrease since its adjustment is driven by two opposite forces. On the one hand, the workers' outside option goes up which raises $R(0)$; on the other hand, expected separation costs decrease which lowers $R(0)$. Our simulations show that the second channel largely dominates. Hence the job finding rate is higher because firms intensify their recruitment effort and then hire a greater share of contacted workers.

All in all, accounting for job creation almost doubles the impact of the removal of job protection on unemployment.[26] This result is obtained while the wage elasticity of employment predicted by the model is small relative to empirical evidence. The model predicts that a 1% increase in the minimum wage decreases employment by 0.19 percentage point. This result is consistent with the estimates of the elasticity of labor demand, which belong to the interval $[-0.07, -0.4]$ according to the meta-analysis of Lichter et al. (2015), while empirical evidence points to a higher elasticity for low skilled workers in France.[27] This suggests that the impact of job protection on unemployment is more likely to be underestimated than overestimated by the model.

Table 4, Columns (3) to (6) show that job protection increases unemployment for all categories of workers. Its impact is stronger on unskilled employment. Skilled employment is less sensitive to

[26]From $1-u=L$, one gets $\Delta u = -L\Delta L/L$. Since steady state employment is equal to the job arrival rate times the average expected duration of jobs from their starting date, we can compute $\Delta L/L$ in partial equilibrium from Table 2, Columns (1). We get $\Delta L/L = 0.032$, which yields, with $u = 0.142$, $\Delta u = -0.027$.

[27]See Cahuc, Carcillo and Le Barbanchon (2018) and the references therein.

job protection because the surplus of skilled jobs is on average bigger, and the minimum wage is less often binding. For these workers, job protection increases employment in partial equilibrium since it (slightly) raises job stability, as shown in Table 2, Column (3). But the relation is reversed when job creation is endogenized, even though the elasticity of skilled employment with respect to the minimum wage is equal to 6.5% only.

6 Conclusion

Employment protection legislations are notoriously hard to reform. Many governments have proposed ambitious liberalization agendas, only to water them down after being faced with unrelenting opposition. Given the size of the political costs, one may presume that such reforms are based on sound and ample evidence. Unfortunately the state of research is not as conclusive as expected.

Our paper contributes to filling this gap by proposing an identification strategy which uses a strong hike in the compensation for unfair dismissal in the French labor code. A similar approach could be applied to other countries whose labor regulations feature similar discontinuities. Our results suggest that employment protection does not work as intended by its proponents since it lowers expected job duration for low and medium skilled workers, who are the main targets of the legislations. More protection leads to less security because firms anticipate their separation decisions in order to avoid paying more firing costs in the future. This mechanism implies that job protection is strongly detrimental to employment in the French context because job protection raises job destruction and reduces job creation. Moreover, we do not find that the negative impact of anticipated costs is specifically driven by the hike at two-year seniority. According to our model, introducing single contracts is not likely to noticeably raise aggregate employment as long as dismissal costs remain increasing in job seniority.

The detrimental impact of job protection results from the interaction between the increasing profile of layoff costs and the fragility of short-tenured jobs. In our model, fragility arises because most jobs are born with a productivity close to the reservation threshold. A promising direction for further research would consist in studying models where fragility is an endogenous feature of the environment. In particular, it would be interesting to combine our empirical strategy with a structural model in the spirit of Jovanovic (1979), where firms and workers gradually learn the quality of their matches. It is well known that Jovanovic's model also predicts that jobs become less and less fragile over time. However, when compared to our framework, learning has the additional implication that the variance of the state variable decreases with tenure. Studying the consequences of the anticipation effect in this framework would improve our understanding of the impact of job protection.

Further research could also build on our estimates to characterize the optimal design of employment protection legislations. In our framework, job protection improves the welfare of long tenured workers, while it is detrimental to all other workers, the unemployed included. Hence, the existence of job protection legislations, as analyzed in our paper, can be explained by political economy arguments. We postponed all normative analysis to future work because our framework does not include features that are essential to look at this issue. In this regard, it seems important to endow employees with risk-aversion and a more proactive role. Workers respond to news about the stability of their jobs by adjusting their hidden effort and their search intensity. For instance, empirical evidence suggests that the effort of workers drops when their job is transformed from non-protected to protected.[28] Although such extensions would greatly complicate the analysis by increasing the number of control variables,[29] we believe that they are worth pursuing, if only to assess the normative implications of our findings.

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[28] Ichino and Riphahn (2005), Olsson (2009).

[29] See, for example, the work of Nagypal (2005).

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APPENDIX

A Characterization of the value functions and reservation threshold

In this Appendix, we focus on the stylized model described in Section 3.1 so that wages are taken as exogenous and firing costs are equal to zero if $t < T$ and to $F > 0$ if $t \geqslant T$.

Optimal stopping problem for permanent jobs. The simplest way to approach problem (3) is to analyze it as a free boundary value problem. For all x within the continuation region, the Bellman equation for $J(x)$ reads

$$(r+\delta) J(x) = x - w + \frac{\mathbb{E}^x \left[\mathrm{d}J(x) \right]}{\mathrm{d}t}.$$

According to Ito's lemma, the infinitesimal generator of $J(x)$ is given by $\mathbb{E}^x \left[\mathrm{d}J(x) \right] / \mathrm{d}t = \mu x J'(x) + (\sigma x)^2 J''(x)/2$. Replacing this expression into the previous equation, we obtain

$$(r+\delta) J(x) = x - w + \mu x J'(x) + \frac{\sigma^2}{2} x^2 J''(x).$$

It is well known that the solution of this parabolic differential equation is of the form

$$J(x) = A x^{\alpha_+} + B x^{\alpha_-} + \frac{x}{r+\delta-\mu} - \frac{w}{r+\delta},$$

where α_+ and α_- are the positive and negative roots of the quadratic equation

$$\frac{\sigma^2}{2}\alpha^2 + \left(\mu - \frac{\sigma^2}{2}\right)\alpha - (r+\delta) = 0,$$

while A and B are constants whose values are chosen so as to satisfy the boundary conditions of the problem. First, the option to destroy the job becomes worthless when output diverges to infinity so that

$$\lim_{x\to+\infty} J(x) = \mathbb{E}^x \left[\int_t^{+\infty} e^{-(r+\delta)(\tau-t)} \left(x_\tau - w \right) \mathrm{d}\tau \right] = \frac{x}{r+\delta-\mu} - \frac{w}{r+\delta}.$$

This condition can only be satisfied if $A = 0$. The second boundary condition ensures that the value of the job is indeed equal to the layoff cost at the separation boundary R, i.e. $J(R) = -F$. This *value matching* requirement is satisfied when the constant $B = -\left(\frac{R}{r+\delta-\mu} - \frac{w-rF}{r+\delta}\right) R^{-\alpha_-}$. This implies that

$$J(x) = \frac{x}{r+\delta-\mu} - \frac{w}{r+\delta} - \left(\frac{R}{r+\delta-\mu} - \frac{w-rF}{r+\delta}\right)\left(\frac{x}{R}\right)^{\alpha}, \tag{A1}$$

which corresponds to equation (4). Finally, the value of the barrier is optimal when it maximizes the value of the job, which is true when R is given by (5).

Optimal stopping problem for non-permanent jobs. We wish to solve

$$J_n(x,t) \equiv \sup_{0\le\tau\le T-t} \mathbb{E}^x \left[\int_0^{\tau} e^{-(r+\delta)s} f(x_s) \mathrm{d}s + e^{-(r+\delta)(T-t)} J\left(x_{T-t}\right) \mathbf{1}_{\{\tau=T-t\}} \right],$$

where $x_t(x) = x \exp\left[(\mu - \sigma^2/2)t + \sigma B_t \right]$ and $f(x) \equiv x - w$. The optimal stopping is not standard because the payoff function exhibits a discontinuity at date T. To cast it in a more conventional form, we follow the approach in El Karoui and Karatzas (1991) and introduce the function

$$G(x,t) \equiv \mathbb{E}^x \left[\int_0^{T-t} e^{-(r+\delta)s} f(x_s) \mathrm{d}s + e^{-(r+\delta)(T-t)} J\left(x_{T-t}\right) \right],$$

which is continuous on $\mathbb{R}^{+}\times[0,T]$. Combining the two definitions above, we have

$$V_n(x,t)\equiv G(x,t)-J_n(x,t)=\inf_{0\leq\tau\leq T-t}\mathbb{E}^x\left[\underbrace{\int_{\tau}^{T-t}e^{-(r+\delta)s}f(x_s)\mathrm{d}s+e^{-(r+\delta)(T-t)}J(x_{T-t})\mathbf{1}_{\{\tau<T-t\}}}_{\equiv e^{(r+\delta)t}H(x_\tau,t+\tau)}\right]. \tag{A2}$$

By the strong Markov property of (A2), $V_n(x,t)$ is the minimal loss of the standard optimal stopping problem $V_n(x,t)=e^{(r+\delta)t}\inf_{0\leq\tau\leq T-t}\mathbb{E}^x\left[H(x_\tau,t+\tau)\right]$. As H is continuous on $\mathbb{R}^{+}\times[0,T)$, the value function V_n, and thus J_n, are also continuous on $\mathbb{R}^{+}\times[0,T)$. Furthermore, the convexity of $J(\cdot)$ and the fact that $J(\cdot)\geq 0$, imply that the optimal stopping time is given by

$$\tau^*(x,t)\equiv\inf\{0<\tau<T-t:J_n(x,t)\leq 0\}\wedge(T-t),$$

where $\wedge$ is the minimum operator. The job is therefore destroyed as soon as its value is below or equal to 0. The stopping time can equivalently be defined as an optimal boundary in the output space as[30]

$$\tau^*(x,t)=\inf\{0<\tau<T-t:x_\tau\leq R(t+\tau)\}\wedge(T-t),$$

where

$$R(t)\equiv\inf\left\{x\in\mathbb{R}^{+}:J_n(x,t)=0\right\},\ \text{for } 0\leq t\leq T.$$

In other words, the continuation region is of the form

$$\mathcal{C}_t\equiv\left\{x\in\mathbb{R}^{+}:J_n(x,t)>0\right\}=\left\{x\in\mathbb{R}^{+}:x>R(t)\right\},\ \text{for } 0\leq t\leq T.$$

Properties of the separation boundary. Lemma 4 will be useful to prove Proposition 1.

Lemma 4 *The value of a non protected job $J_n(\cdot,t)$ is decreasing in t.*

Proof. We use a pathwise comparison. Let $\tau^t=\inf\{0\leq s<T-t:X_s(x)\leq R(s)\}\wedge(T-t)$, that is τ^t is the optimal stopping time for (x,t). Consider any $s\in[0,t)$. Since τ^t is feasible for (x,s) we have

$$J_n(x,s)-J_n(x,t)\geq\mathbb{E}^x\left[e^{-(r+\delta)(T-t)}\left(J_n(x_{T-t},T+s-t)-J(x_{T-t})\right)1_{\{\tau^t=T-t\}}\right]\geq 0. \tag{A3}$$

The last inequality holds because: (i) the reward function of unprotected jobs is strictly higher than that of protected ones; (ii) any stopping time τ that is admissible for a protected job is also admissible for a non protected job. Hence, we have $J_n(x,t)\geq J(x)$ for all $t\in[0,T]$ and (A3) follows. ■

Proof. Proposition 1: For any $t\in(0,T)$, $\delta^t\in[0,t]$ and $\varepsilon>0$, we have

$$J_n\left(R(t)+\varepsilon,t-\delta^t\right)\geq J_n(R(t)+\varepsilon,t)>0,$$

where the first inequality follows from Lemma 4, while the second inequality holds because $R(t)+\varepsilon\in\mathcal{C}_t$. The difference above implies in turn that for any $\varepsilon>0$, $R(t)+\varepsilon\in\mathcal{C}_{t-\delta^t}$ or $R\left(t-\delta^t\right)\leq R(t)$.

To determine the impact of F, we use a pathwise comparison similar to that in Lemma 4. Let τ^F denote the optimal stopping time for $(x,t;F)$ with $t<T$. Since firing costs do not affect the feasibility of the stopping times nor the flow surpluses of non protected jobs, we have

$$J_n(x,t;F)-J_n(x,t;F+\varepsilon)\geq\mathbb{E}^x\left[e^{-(r+\delta)(T-t)}\left(J(x_{T-t};F)-J(x_{T-t};F+\varepsilon)\right)1_{\{\tau^{F+\varepsilon}=T-t\}}\right]\geq 0,$$

[30] The shape of the continuation region is a consequence of the fact that $J_n(x,\cdot)$ is increasing in x. A property that can easily be established using an argument similar to the one used for the proof of Proposition 2.1 in Jacka (1991).

for all $\varepsilon > 0$. The last inequality follows from the closed-form expression for $J(\cdot)$ and the fact that $\partial J(x;F)/\partial F < 0$. ■

Proof. Proposition 2: The proposition is a direct consequence of Lemma 4 and Proposition 1. First, we have shown in the proof of Lemma 4 that $J_n(x,0) \geq J(x)$, which implies in turn that $R(0) \geq R$. But we also know from Proposition 1 that $R(t)$ is strictly increasing in t, so that $\lim_{t\to T^-} R(t) > R(0) \geq R$. The size of the drop is increasing in F because the closed-form expression (5) for R is such that $\partial R/\partial F < 0$, while Proposition 1 implies that $\partial R(t)/\partial F > 0$ for all $t \in [0,T)$. ■

B Wage negotiation

In this appendix, we derive the solution to the wage negotiation game given equation (8). Let us remind that the worker and the employer make alternating offers. When one of the players offers a wage, the other player either accepts or rejects the offer. If the offer is accepted, then the bargaining ends and the offered contract is implemented. If the offer is rejected, then the game goes on to a next round after a short time delay, denoted by Δ_e if the worker just rejected an offer by the firm or by Δ_f if the firm just rejected an offer by the worker. In the next round, the player who last rejected an offer makes a counteroffer, which again can be either accepted or rejected. The game goes on in this way over an infinite horizon. During the bargaining game, production stops. The worker and the firm bargain according to the value of output at the starting date of the negotiation. Moreover, the match is destroyed at exogenous rate $\delta + s$, $s > 0$ during the negotiation game (instead of δ when production does not stop because the worker and the firm need time to reach an agreement), because the job is more at risk of being destroyed when production stops. The firm pays the layoff cost F if the job is destroyed during the negotiation game for a reason that is not associated with the event arriving at rate δ (as it is supposed that there is no firing cost for this type of event).

Osborne and Rubinstein (1990) showed that the subgame perfect equilibrium of the negotiation game is a pair of stationary strategies in which the firm (worker) offers the wage w_f (w_e) that makes the other party indifferent between accepting the wage offer instantaneously or waiting his/her turn to make a counteroffer. Let us denote by $W(w,x,t)$, the expected utility of a worker of seniority t employed on a job with current output x and wage w. Let us adopt the similar notation, $J(w,x,t)$ for the value function of the firm. Assuming that Δ_e and Δ_f are small intervals of time, the equilibrium pair of wages (w_f, w_e) solves

$$\begin{aligned} W(w_f,x,t) &= \frac{1}{1+r\Delta_e}\left[\left[1-(s+\delta)\Delta_e\right]W(w_e,x,t) + (s+\delta)\,\Delta_e\,(U+F_e)\right], \\ J(w_e,x,t) &= \frac{1}{1+r\Delta_f}\left[\left(1-(s+\delta)\Delta_f\right)J(w_f,x,t) - s\Delta_f F\right], \end{aligned}$$

where U is the expected utility of an unemployed worker, F_e is the severance payment (the transfer from the employer to the employee) and $F = F_a + F_e$ stands for the total layoff costs, including the administrative costs F_a. These two equations can be written

$$W(w_f,x,t) - W(w_e,x,t) = -\Delta_e\left[(s+\delta)\left[W(w_e,x,t) - (U+F_e)\right] + rW(w_f,x,t)\right], \tag{B4}$$

$$J(w_e,x,t) - J(w_f,x,t) = -\Delta_f\left[(s+\delta)J(w_f,x,t) + sF + rJ(w_e,x,t)\right]. \tag{B5}$$

Both equations imply that $w_f \to w_e$ when $\Delta_e \to 0$ and $\Delta_f \to 0$. Denoting the common limit of w_f and w_e by w, one can write

$$\begin{aligned} \frac{\partial W(w,x,t)}{\partial w} &= \lim_{\Delta_f,\Delta_e\to 0} \frac{W(w_f,x,t) - W(w_e,x,t)}{w_f - w_e}, \\ \frac{\partial J(w,x,t)}{\partial w} &= \lim_{\Delta_f,\Delta_e\to 0} \frac{J(w_f,x,t) - J(w_e,x,t)}{w_f - w_e}. \end{aligned}$$

Using these two equations together with equations (B4) and (B5), we have

$$-\frac{\frac{\partial W(w,x,t)}{\partial w}}{\frac{\partial J(w,x,t)}{\partial w}} = \frac{\Delta_e}{\Delta_f}\frac{(s+\delta)\left[W(w_e,x,t)-(U+F_e)\right]+rW(w_f,x,t)}{(s+\delta)J(w_f,x,t)+sF+rJ(w_e,x,t)}. \tag{B6}$$

Since the wage is a transfer from the firm to the worker, who are both risk neutral, we have $\frac{\partial W(w,x,t)}{\partial w} = -\frac{\partial J(w,x,t)}{\partial w}$. Thus, the wage solution to equation (B6) solves

$$J(w,x,t)+\frac{s}{s+\delta+r}F = \frac{\Delta_e}{\Delta_f}W(w,x,t)-\frac{\Delta_e}{\Delta_f}\frac{s+\delta}{s+\delta+r}\left(U+F_e\right).$$

If s is large with respect to r and δ, this equation can be written

$$(1-\beta)\left[W(w,x,t)-(U+F_e)\right] = \beta\left[J(w,x,t)+F\right], \tag{B7}$$

with $\beta = \frac{\Delta_f}{\Delta_f+\Delta_e}$. As in Cahuc, Postel-Vinay and Robin (2006), the interpretation is that the bargaining power β does not depend on the discount factor and on the destruction rate of operating jobs if these parameters are small enough compared to the job destruction rate during the negotiation process. We assume that this condition is fulfilled.

Let us now derive the expression of the bargained wage of a worker of experience t on a job with current output x. We have

$$rJ(w,x,t) = x-w-\delta J(w,x,t)+\frac{\mathbb{E}\left[\mathrm{d}J(w,x,t)\right]}{\mathrm{d}t}, \tag{B8}$$

$$rW(w,x,t) = w+\delta\left[U-W(w,x,t)\right]+\frac{\mathbb{E}\left[\mathrm{d}W(w,x,t)\right]}{\mathrm{d}t}. \tag{B9}$$

The wage bargaining solution (B7) implies that

$$(1-\beta)\left(\frac{\mathbb{E}\left[\mathrm{d}W(w,x,t)\right]}{\mathrm{d}t}-\frac{\mathrm{d}F_e}{\mathrm{d}t}\right) = \beta\left[\frac{\mathbb{E}\left[\mathrm{d}J(w,x,t)\right]}{\mathrm{d}t}+\frac{\mathrm{d}F}{\mathrm{d}t}\right].$$

Thus, using this equation and substituting (B8) and (B9) into (B7), we get the wage negotiated on a job with current output x

$$w(x) = \beta x+(1-\beta)rU+(r+\delta)\left(\beta F+(1-\beta)F_e\right)-\beta\frac{\mathrm{d}F}{\mathrm{d}t}-(1-\beta)\frac{\mathrm{d}F_e}{\mathrm{d}t}.$$

This wage obtains only if it is higher than the minimum wage $w_{\min}$. Otherwise, the minimum wage applies, as shown by equation (8).

C Solution of the complete model

In this appendix, we detail the methods used to solve the model in which layoff costs increase with tenure and jump at two years as defined by equation (7). The resolution method borrows from the methods used to value American options. The model is solved backwards. It is assumed that the layoff costs reach a plateau after tenure $T_p > T$ that makes the problem stationary from T_p. It is then possible to solve the model analytically at tenure T_p.

Going backwards in tenure from $t = T_p$ to $t = 0$, the model is then solved numerically. In what follows, we analyze the problem recursively, starting with jobs at tenure T_p and then reinserting their optimal values to solve for the values of jobs at tenure $t < T_p$. Tenure from 0 to T_p can be divided into two sub-periods which correspond respectively to the periods during which the jobs are loosely protected ($t < T$) and then heavily protected ($t \geq T$). Let us denote by T the deterministic tenure from which jobs become protected, *i.e.* the tenure from which there is a sharp discontinuity on the extent of employment protection.

From tenure T_p the costs of dismissal are constant. Thus, we get back to the standard infinite horizon problem studied in section 3.1. The value of jobs $J(x)$ and of the reservation output R are given by equations (4) and (5) respectively, where the wage w is defined by equation (8).

Not all jobs reach tenure T_p either because some are hit by an exogenous destruction shock, or because their output falls below the reservation threshold, which we denote by $R(t)$. In the latter case, firms destroy jobs when their value is not sufficient to cover the costs of dismissal, *i.e.* $J_k(R(t),t) \leq -F(t)$, $k = n, p$, for all $t \in [0, T_p)$. Formally, the profit of a job $J_k(x_t, t)$ at tenure $t < T_p$ solves the following optimal stopping problems:

$$J_n(x,t) = \sup_{0\leq t\leq \tau\leq T} \mathbb{E}^{(x_t,t)}\left[\int_t^\tau e^{-(r+\delta)(s-t)}\left[x_s - w(x_s)\right]\mathrm{d}s + e^{-(r+\delta)(T-t)}J_p(x_T,T)\,\mathbf{1}_{\{\tau=T\}}\right], \tag{C10}$$

$$J_p(x,t) = \sup_{T\leq t\leq \tau\leq T_p} \mathbb{E}^{(x_t,t)}\left[\int_t^\tau e^{-(r+\delta)(s-t)}\left[x_s - w(x_s)\right]\mathrm{d}s + e^{-(r+\delta)(T_p-t)}J_p\left(x_{T_p},T_p\right)\mathbf{1}_{\{\tau=T_p\}}\right], \tag{C11}$$

where the stopping time τ is $\mathcal{F}_t$-measurable and $\mathbf{1}_{\{\cdot\}}$ is the indicator function. More generally, for an arbitrary deterministic tenure T, we wish to solve a problem of the form

$$J_k(x,t) \equiv \sup_{0\leq\tau\leq\overline{T}-t} \mathbb{E}^x\left[\int_0^\tau e^{-(r+\delta)s}f(x_s)\mathrm{d}s + e^{-(r+\delta)(T-t)}J_{\cancel{k}}(x_{T-t},T-t)\,1_{\{\tau=T-t\}}\right], \tag{C12}$$

where $x_t = xe^{\left((\mu-\sigma^2/2)t+\sigma B_t\right)}$and $f(x_t) \equiv x_t - w(x_t)$. We have already shown (see Appendix A) that the value $J_k(x,t)$ solves a standard optimal stopping problems. We can now present the numerical resolution method. Let us rewrite (C12) as:

$$\begin{cases} (r+\delta)\,J_k(x,t) = (1-\beta)(x-rU) - \beta(r+\delta)F + \frac{\mathbb{E}^x[\mathrm{d}J_k(x)]}{\mathrm{d}t} & \text{if } x \geq \bar{x}, \\ (r+\delta)\,J_k(x,t) = x - w_{\min} + \frac{\mathbb{E}^x[\mathrm{d}J_k(x)]}{\mathrm{d}t} & \text{otherwise,} \end{cases} \tag{C13}$$

where F is defined by equation (7), $\bar{x}$ by equation (10) and $\frac{\mathbb{E}[\mathrm{d}J_k(x,t)]}{\mathrm{d}t}$ indicates changes in the value of the job over time. Using Ito's lemma, the Hamilton-Jacobi-Bellman equation (C13) reads as:

$$\begin{cases} (r+\delta)\,J_k(x,t) = (1-\beta)(x-rU) - \beta(r+\delta)F + \mu x\frac{\partial J_k(x,t)}{\partial x} + \frac{\sigma^2}{2}x^2\frac{\partial^2 J_k(x,t)}{\partial x^2} + \frac{\partial J_k(x,t)}{\partial t} & \text{if } x \geq \bar{x}, \\ (r+\delta)\,J_k(x,t) = x - w_{\min} + \mu x\frac{\partial J_k(x,t)}{\partial x} + \frac{\sigma^2}{2}x^2\frac{\partial^2 J_k(x,t)}{\partial x^2} + \frac{\partial J_k(x,t)}{\partial t} & \text{otherwise,} \end{cases} \tag{C14}$$

This is a system of two second-order partial differential equations in two independent variables (space and time. These are parabolic equations analogous to e.g. the well-known heat equation in physics. Equation (C14) does not admit a closed form expression and thus has to be approximated numerically. We use finite-difference methods for approximating the solutions to (C14). Our preferred method relies on the Implicit Euler Scheme, which contrary to the explicit method ensures the stability of the approximation.[31] We first define the domain discretization in space and time.

Let us index output by i and tenure by j. Let us denote by δx and δt a small interval of output and tenure respectively. The discretization domain is then given by $\{0, \delta x, 2\delta x, ..., M\delta x\}$ for productivity where $M\delta x = x^{\max}$, and $\{0, \delta t, 2\delta t, ..., N\delta t\}$ for time where $N\delta t = T$. We hence get a grid of $(M+1)\times(N+1)$ points.

Let us define $J_{i,j} \equiv J_k(i\delta x, j\delta t)$, and use finite difference equations to approximate the derivatives of (C14). The model is solved backward, then we use backward approximation to approximate the derivative with respect to time so that

$$\frac{\partial J_k(x,t)}{\partial t} \approx \frac{J_{i,j} - J_{i,j-1}}{\delta t}.$$

[31]Alternatively, we could have used the Crank-Nicolson scheme to approximate (C14). For more details on approximation methods see for example Brandimarte (2006, chapter 5).

We then approximate $\frac{\partial J_n(x,t)}{\partial x}$ and $\frac{\partial^2 J_n(x,t)}{\partial x^2}$ using central approximation and standard approximation respectively, such that

$$\begin{aligned}\frac{\partial J_k\left(x,t\right)}{\partial x} &\approx \frac{J_{i+1,j-1}-J_{i-1,j-1}}{2\delta x},\\ \frac{\partial^2 J_k\left(x,t\right)}{\partial x^2} &\approx \frac{J_{i+1,j-1}-2J_{i,j-1}+J_{i-1,j-1}}{\left(\delta x\right)^2}.\end{aligned}$$

Substituting these expressions in equation (C14), we get, after some algebra

$$J_{i,j} = A_i J_{i-1,j-1} + B_i J_{i,j-1} + C_i J_{i+1,j-1} + D_i, \tag{C15}$$

where

$$\begin{aligned}A_i &= \frac{1}{2}\left(\mu i - \sigma^2 i^2\right)\delta t,\\ B_i &= 1 + \left[\sigma^2 i^2 + (r+\delta)\right]\delta t,\\ C_i &= -\frac{1}{2}\left(\sigma^2 i^2 + 1\right)\delta t,\\ D_i &= -\left[i\delta x - w(i\delta x)\right]\delta t,\end{aligned}$$

where $w(x)$ is defined by (8). We hence link one known value in time layer j to three unknown values in time layer $j-1$. This equation holds for $i = 1, ..., M$ since $J_{0,j}$ and J_{M+1} are not defined. It follows that there are $M-1$ equations for $M+1$ unknowns. The remaining two equations are given by the Dirichlet boundary conditions such that $J_{0,j} = -F\left(j\right)$ and $J_{M+1,j} = J_k\left(x^{\max}, j\right)$. The Implicit Euler Scheme can be written in the following matrix form:

$$\begin{pmatrix} J_{1,j}\\ J_{2,j}\\ J_{3,j}\\ .\\ J_{M-1,j}\\ J_{M,j}\end{pmatrix} = \underbrace{\begin{pmatrix} B_1 & C_1 & 0 & 0 & 0 & 0 & 0\\ A_2 & B_2 & C_2 & 0 & 0 & 0 & 0\\ 0 & A_3 & B_3 & C_3 & 0 & 0 & 0\\ . & . & . & . & . & . & .\\ 0 & 0 & 0 & 0 & A_{M-1} & B_{M-1} & C_{M-1}\\ 0 & 0 & 0 & 0 & 0 & A_M & B_M\end{pmatrix}}_{\mathbf{Q}} \begin{pmatrix} J_{1,j-1}\\ J_{2,j-1}\\ J_{3,j-1}\\ .\\ J_{M-1,j-1}\\ J_{M,j-1}\end{pmatrix} + \begin{pmatrix} D_1 + A_1 J_{0,j-1}\\ D_2\\ D_3\\ .\\ D_{M-1}\\ D_M + C_M J_{M+1,j-1}\end{pmatrix}, \tag{C16}$$

where $\mathbf{Q} \in \mathbb{R}^{M-1,M-1}$ is a positive definite tridiagonal matrix. Knowing the value function at expiration date, we have to solve backwards in time a system of linear equations for each time layer with the following terminal conditions $J_{i,T} = \max\left[J_p(i,T), -F\left(T\right)\right]$ for new jobs and $J_{i,T_p} = \max\left[J_p(i), -F\left(T_p\right)\right]$ for protected jobs. In each case, we are then left with $(M+1-2)\times(N+1-1)$ values to compute, *i.e.* we have to solve for a system of $M-1$ equations with $M-1$ unknowns for N time layers. To solve this system we use a generalization of the Gauss-Seidel iterative method, the Successive-Over-Relaxation (SOR) method[32] whose formulation depends on a relaxation parameter ω. The system (C16) can be rewritten as

$$\underbrace{\begin{pmatrix} B_1 & C_1 & 0 & 0 & 0 & 0 & 0\\ A_2 & B_2 & C_2 & 0 & 0 & 0 & 0\\ 0 & A_3 & B_3 & C_3 & 0 & 0 & 0\\ . & . & . & . & . & . & .\\ 0 & 0 & 0 & 0 & A_{M-1} & B_{M-1} & C_{M-1}\\ 0 & 0 & 0 & 0 & 0 & A_M & B_M\end{pmatrix}}_{\mathbf{Q}} \underbrace{\begin{pmatrix} J_{1,j-1}\\ J_{2,j-1}\\ J_{3,j-1}\\ .\\ J_{M-1,j-1}\\ J_{M,j-1}\end{pmatrix}}_{\mathbf{J}} = \underbrace{\begin{pmatrix} J_{1,j} - (D_1 + A_1 J_{0,j-1})\\ J_{2,j} - D_2\\ J_{3,j} - D_3\\ .\\ J_{M-1,j} - D_{M-1}\\ J_{M,j} - (D_M + C_M J_{M+1,j-1})\end{pmatrix}}_{\mathbf{D}}. \tag{C17}$$

Remarking that the tridiagonal matrix $\mathbf{Q}$ can be written as the sum of diagonal matrix of strictly positive elements ($\mathbf{P}$), a lower triangular matrix ($\mathbf{L}$) and an upper triangular matrix ($\mathbf{U}$) such that $\mathbf{Q} = \mathbf{P} + \mathbf{L} + \mathbf{U}$, the

[32] A detailed description of this method can be found, for example, in Richardson (2009) or Burden and Faires (2011).

system of linear equations $\mathbf{QJ} = \mathbf{D}$ can be rewritten in matrix and vector form for the SOR iterative method as (see e.g. Burden and Faires (2011), chapter 7)[33]

$$(\mathbf{P}+\omega\mathbf{L})\,\mathbf{J}^{(\zeta)} = [(1-\omega)\,\mathbf{P}-\omega\mathbf{U}]\,\mathbf{J}^{(\zeta-1)} + \omega\mathbf{D}, \tag{C18}$$

where ζ is the iteration counter and ω is the relaxation parameter that remains to be determined. It is then possible to rewrite (C18) as

$$\mathbf{J}^{(\zeta)} = (\mathbf{P}+\omega\mathbf{L})^{-1}\left[((1-\omega)\,\mathbf{P}-\omega\mathbf{U})\,\mathbf{J}^{(\zeta-1)} + \omega\mathbf{D}\right]. \tag{C19}$$

The SOR method iteratively solves the left hand side of (C19) using the previous computed value on the right hand side. As the square matrix $(\mathbf{P}+\omega\mathbf{L})^{-1}$ is lower triangular, the elements of $\mathbf{J}^{(\zeta)}$ are easily computed using forward substitution such that

$$J_i^{(\zeta)} = (1-\omega)J_i^{(\zeta-1)} + \frac{\omega}{Q_{i,i}}\left[D_i - \sum_{j=1}^{i-1} Q_{i,j}J_j^{(\zeta)} - \sum_{j=i+1}^{n} Q_{i,j}J_j^{(\zeta-1)}\right].$$

Note that when $\omega = 1$, we get back to the Gauss-Seidel method. It now only remains to calculate the value of the relaxation parameter. Let $\mathbb{T} = \mathbf{P}^{-1}(\mathbf{L}+\mathbf{U})$ and let us denote by $\rho(\mathbb{T})$ the spectral norm[34] of the square matrix $\mathbb{T}$. The optimal relaxation parameter ω that ensures convergence is then defined as (see e.g. Burden and Faires (2011), theorem 7.26 for a more formal statement)

$$\omega = \frac{2}{1+\sqrt{1-[\rho(\mathbb{T})]^2}}.$$

D Severance payments

This appendix starts by presenting the data about severance payments before explaining how the coefficients $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$ parameterizing the effect of tenure on severance payments are computed.

D.1 Data

There are 210 Prud'hommes councils in France and 34 Appeal Courts. Our dataset covers the 2006-2012 period, in which there are 102,779 Appeal Court rulings about dismissals. Appeal Court rulings are extracted from the Dalloz web site which covers all the universe of Appeal Court rulings obtained from the Ministry of Justice. Appeal Court rulings are written in very heterogeneous ways. Usually, they describe the features of the employee and of the employer, the dismissal letter, the claims of the parties, the decision of the Prud'hommes council and the compensation for unfair dismissal.

A first dataset gathered information on 3,092 French Appeal Court rulings from 2003 to 2012 relying on systematic analysis by students in Law in February and March 2016. This information includes the amount of compensations, the seniority, age and gender of workers, the type of contract and a series of other variables. Comparison of rulings comprising different amount of information among the sample of 3,092 rulings indicates that the amount of compensations for unfair dismissal is not correlated with the information available in the rulings. Then, Phyton programming language was used to extract information from all the Appeal Court rulings. This leads to our sample of 27,936 rulings. Comparison of this sample with the initial sample of 3,092 rulings provides consistent results about the relation between seniority and compensation for unfair dismissals.

[33]For a detailed presentation of the SOR method and the pseudo code describing the algorithm necessary for its implementation, see e.g. Brandimarte (2006), Richardson (2009) or Gilli et al. (2011).

[34]Let $\mathbb{T}'$ be the conjugate transpose of the square matrix $\mathbb{T}$. The spectral norm is defined as the square root of the maximum eigenvalue of $\mathbb{T}'\mathbb{T}$, i.e., $\rho(\mathbb{T}) = \max_{1\leq i\leq M+1}\sqrt{|\lambda_i|}$ where λ_i are the eigenvalues of $\mathbb{T}'\mathbb{T}$.

D.2 Computation of $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$

The coefficients $\{\phi_{e0}, \phi_{en}, \phi_{ep}\}$ parameterizing the effect of tenure on severance payments are determined to match the average expected severance payment depending on seniority. To compute these coefficients, we use the dataset of Appeal Court rulings in which we distinguish the separation cost of temporary and permanent contracts.

The share of separations due to temporary and permanent contracts for each quarter of seniority is computed from the Labor Force Survey. These shares are weighted by their associated expected severance payments.

- The expected severance payments for temporary jobs are equal to the sum of the legal severance and of the probability of prosecution times the expected amount of compensation for unfair dismissal in case of prosecution.
  - The legal severance for the termination of temporary contracts amounts to 33% of the monthly wage per year of seniority. It corresponds to the "prime de précarité", equal to 10% of all the remuneration of the temporary contract, which is due by the employer at the termination of temporary contracts which are either not renewed or not transformed into permanent contracts. However, the "prime de précarité" is not due in sectors which make an intensive use of temporary jobs (hotels and restaurants, professional sports, entertainment, audiovisual, film production, education...). See: https://www.service-public.fr/particuliers/vosdroits/F32476. According to the "Fichier National des Allocataires" (UNEDIC), which reports all the labor contracts of all unemployed workers eligible to unemployment benefits, 67% of temporary contracts belong to these sectors. Accounting for these exceptions implies that the legal severance for the termination of temporary contracts amounts to 33% of the monthly wage per year of seniority.
  - To infer the probability of prosecution for the termination of temporary contracts at each quarter of seniority, we compute the share of these terminations in our dataset of Appeal Court rulings. Since Appeal Courts are second instance, our dataset of Appeal Court rulings does not cover all the universe of prosecutions. So, we use the total number of annual prosecutions at Prud'hommes councils provided by Guillonneau and Serverin (2015). Assuming that the probability of appeal of Prud'hommes council's decisions is identical for temporary and permanent contracts and does not depend on seniority, we compute the number of prosecutions at Prud'hommes councils for temporary contracts for every seniority. Then, using the Labor Force Survey, we compute the number of separations for every seniority, which allows us to compute the probability of prosecution for the termination of temporary contracts at each quarter of seniority .
  - The expected amount of compensation for unfair termination of temporary contracts in case of prosecution is computed from the dataset of Appeal Court rulings. This compensation is due if the temporary contract is requalified into a permanent contract. In this case, the severance is equal to the legal severance plus the compensation for wrongful termination. The information about the type of contract (either permanent or temporary) is available in our dataset.
- The expected severance payments for permanent contracts are equal to the sum of the legal severance and of the probability of prosecution multiplied by the average compensation in case of prosecution after terminations of permanent contracts.
  - The legal severance is equal to 20% of the monthly wage per year of seniority.
  - The probability of prosecution at each quarter of seniority is computed as for temporary contracts.
  - The expected amount of compensation for unfair termination of permanent contracts in case of prosecution is computed from the dataset of Appeal Court rulings.

E Minimum Distance Estimator

E.1 Minimum distance estimator

In this Appendix, we establish the consistency of our minimum distance estimator and derive explicitly the efficient weighting matrix. First, let us recall that each observation $x_k = \{t_k, d_k\}$ contains two informations: the length of the job spell t_k, and the dummy variable d_k which takes value 1 solely when the job is destroyed by the end of the quarter. For each of these observations, we can construct the random vector $X^{(k)} = \left(\mathbf{1}_{\{t_k=1\}}d_k, \mathbf{1}_{\{t_k=1\}}, \mathbf{1}_{\{t_k=2\}}d_k, \mathbf{1}_{\{t_k=2\}}, ..., \mathbf{1}_{\{t_k=T\}}d_k, \mathbf{1}_{\{t_k=T\}}\right)$. We shall use $\overline{X}_i \equiv \frac{1}{n}\sum_{k=1}^{n} X_i^{(k)}$, with $i = 1, ..., 2T$, to denote each element of the vector $\overline{X}$ of sample means. The central limit theorem ensures that

$$\sqrt{n}\left(\overline{X} - \mu\right) \longrightarrow \mathcal{N}\left(0, \Sigma\right) \tag{E20}$$

as the number of i.i.d. draws n increases. To ease notation, we shall add a subscript to the elements of the limit vector and use the following decomposition $\mu = \left(\mu_1^1, \mu_1^2, ..., \mu_T^1, \mu_T^2\right)$. The variance-covariance matrix is easily computed once it has been noticed that: (i) $\mathbf{1}_{\{t_k=s\}}d_k$ and $\mathbf{1}_{\{t_k=s\}}$ are binomial variables, (ii) $\mathbf{1}_{\{t_k=s\}}$ and $\mathbf{1}_{\{t_k=\tau\}}$ are mutually exclusive events for all $s \neq \tau$. Thus Σ reads

$$\Sigma = \begin{bmatrix} \sigma_{1,1} & \cdots & \sigma_{1,T} \\ \vdots & \ddots & \vdots \\ \sigma_{T,1} & \cdots & \sigma_{T,T} \end{bmatrix},$$

where[35]

$$\sigma_{i,i} = \begin{bmatrix} \mu_i^1\left(1-\mu_i^1\right) & \mu_i^1\left(1-\mu_i^2\right) \\ \mu_i^1\left(1-\mu_i^2\right) & \mu_i^2\left(1-\mu_i^2\right) \end{bmatrix} \text{ for all } i \in \{1,2,...,T\},$$

and

$$\sigma_{i,j} = \begin{bmatrix} -\mu_i^1\mu_j^1 & -\mu_i^2\mu_j^1 \\ -\mu_i^1\mu_j^2 & -\mu_i^2\mu_j^2 \end{bmatrix} \text{ for all } i,j \in \{1,2,...,T\} \text{ such that } i \neq j.$$

In order to connect $\overline{X}$ to the vector of empirical hazard rates $\hat{\mathbf{h}}$, we now consider the non-linear mapping $g : \mathbb{R}^{2T} \to \mathbb{R}^T$ such that

$$g\left(\mu_1^1, \mu_1^2, ..., \mu_T^1, \mu_T^2\right) = \left(\mu_1^1/\mu_2^1, ..., \mu_T^1/\mu_T^2\right).$$

Since $\mu_2^i > 0$ for all $i \in \{1,2,...,T\}$, g is continuously differentiable in the neighborhood of μ and we can use the *Delta method* to infer the asymptotic property of our estimator. More precisely, taking a first-order approximation of $g\left(\overline{X}_1, ..., \overline{X}_{2T}\right)$ around μ and using $(E20)$, we can apply Slutsky's theorem to conclude that

$$\sqrt{n}\left(g\left(\overline{X}\right) - g(\mu)\right) \longrightarrow \mathcal{N}\left(0, \underbrace{\nabla g(\mu)^T \Sigma \nabla g(\mu)}_{\equiv\Omega}\right),$$

where the gradient of g reads

$$\nabla g\left(\mu\right) = \begin{bmatrix} 1/\mu_1^2 & -\mu_1^1/\left(\mu_1^2\right)^2 & 0 & 0 & ... & 0 & 0 \\ 0 & 0 & 1/\mu_2^2 & -\mu_2^1/\left(\mu_2^2\right)^2 & ... & \vdots & \vdots \\ \vdots & \vdots & \vdots & \vdots & \ddots & 0 & 0 \\ 0 & 0 & 0 & 0 & ... & 1/\mu_T^2 & -\mu_T^1/\left(\mu_T^2\right)^2 \end{bmatrix}.$$

[35] The covariance terms in $\sigma_{i,i}$ are by definition equal to

$$\begin{aligned} Cov\left(\mathbf{1}_{\{t_k=i\}}d_k, \mathbf{1}_{\{t_k=i\}}\right) &= \mathbb{E}\left[\left(\mathbf{1}_{\{t_k=i\}}d_k - \mu_i^1\right)\left(\mathbf{1}_{\{t_k=i\}} - \mu_i^2\right)\right] \\ &= \mathbb{E}\left[\mathbf{1}_{\{t_k=i\}}d_k - \mathbf{1}_{\{t_k=i\}}d_k\mu_i^2 - \mathbf{1}_{\{t_k=i\}}\mu_i^1 + \mu_i^1\mu_i^2\right] = \mu_i^1 - \mu_i^1\mu_i^2. \end{aligned}$$

The second equality follows from $\left(\mathbf{1}_{\{t_k=i\}}\right)^2 d_k = \mathbf{1}_{\{t_k=i\}}d_k$. The expression of the remaining terms in $\sigma_{i,i}$ and $\sigma_{i,j}$ can be obtained using similar computations.

Proposition 3 follows because

$$g\left(\overline{X}\right)=\left(\frac{\sum_{k=1}^{n}\mathbf{1}_{\{t_k=1\}}d_k}{\sum_{k=1}^{n}d_k},...,\frac{\sum_{k=1}^{n}\mathbf{1}_{\{t_k=T\}}d_k}{\sum_{k=1}^{n}d_k}\right)=\left(\hat{h}_1,...,\hat{h}_T\right)=\hat{\mathbf{h}},$$

and $g(\mu)=\mathbf{h}\left(\Theta_0\right)$. The covariance matrix Ω is consistently estimated replacing Σ by $\hat{\Sigma}$ and μ by the vector of sample means $\overline{X}$. Since the inverse of any consistent estimator of Ω is an efficient weighting matrix, replacing

$$\hat{\Omega}\equiv\nabla g\left(\overline{X}\right)^{T}\hat{\Sigma}\nabla g\left(\overline{X}\right) \tag{E21}$$

in (12) yields the efficient minimum distance estimator.

E.2 Maximum likelihood estimation of the productivity distribution

We assume that initial productivities are drawn from the distribution $H(x)$. Furthermore, wages are observed with some noise so that $\hat{w}=w\varepsilon$, where $\hat{w}$ is the observed wage and ε is an i.i.d. noise with CDF $G\left(\varepsilon\right)$. Since wages are set through Nash-bargaining and the minimum wage binds, we have for all new hires

$$w(x)=\begin{cases}\beta x+K & \text{if } x>\bar{x}\\ w_{\min} & \text{if } x\in\left[R_0,\bar{x}\right]\end{cases}, \tag{E22}$$

where $K\equiv(1-\beta)\,rU+(r+\delta)\left[\beta F+(1-\beta)F_e\right]-\beta\mathrm{d}F/\mathrm{d}t-(1-\beta)\mathrm{d}F_e/\mathrm{d}t$, and $\bar{x}=(w_{\min}-K)/\beta$.

1. $w>w_{\min}$: Let us first focus on the observations generated by jobs whose *actual wages* are above the minimum wage. We have by definition

$$\Pr\left(\hat{w}\leq y\middle|\,w>w_{\min}\right)=\int_0^{\frac{y}{w_{\min}}}\Pr\left(w\leq\frac{y}{\varepsilon}\middle|\,w>w_{\min}\right)g\left(\varepsilon\right)\mathrm{d}\varepsilon,$$

where the upper-bound of the integral follows from the fact that $\Pr\left(w\leq y/\varepsilon\middle|\,w>w_{\min}\right)=0$ for all $\varepsilon>y/w_{\min}$. Since $x>\bar{x}$, the solution for wages $(E22)$ implies that

$$\Pr\left(w\leq\frac{y}{\varepsilon}\middle|\,w>w_{\min}\right)=\frac{\int_{w_{\min}}^{\frac{y}{\varepsilon}}h\left(\frac{x-K}{\beta}\right)\mathrm{d}x}{1-H(\bar{x})}=\frac{H\left(\frac{\frac{y}{\varepsilon}-K}{\beta}\right)-H(\bar{x})}{1-H(\bar{x})}.$$

Thus, if we let $\tilde{H}\left(\cdot\right)$ denote the conditional density $\tilde{H}\left(z\right)\equiv\left[H\left(z\right)-H(\bar{x})\right]/\left[1-H(\bar{x})\right]$, the conditional probability above reads

$$\Pr\left(\hat{w}\leq y\middle|\,w>w_{\min}\right)=\int_0^{\frac{y}{w_{\min}}}\tilde{H}\left(\frac{y/\varepsilon-K}{\beta}\right)g\left(\varepsilon\right)\mathrm{d}\varepsilon.$$

As usual, the likelihood of the observation follows by differentiation

$$\begin{aligned}L\left(\hat{w}\middle|\,w>w_{\min}\right) &= \int_0^{\frac{\hat{w}}{w_{\min}}}\frac{\partial\tilde{H}\left(\frac{\hat{w}/\varepsilon-K}{\beta}\right)}{\partial\hat{w}}g\left(\varepsilon\right)\mathrm{d}\varepsilon+\frac{1}{w_{\min}}\tilde{H}\left(\frac{w_{\min}-K}{\beta}\right)g\left(\frac{\widehat{w}}{w_{\min}}\right)\\ &= \frac{1}{1-H(\bar{x})}\int_0^{\frac{\hat{w}}{w_{\min}}}h\left(\frac{\frac{\hat{w}}{\varepsilon}-K}{\beta}\right)\frac{g\left(\varepsilon\right)}{\beta\varepsilon}\mathrm{d}\varepsilon+\frac{1}{w_{\min}}\underbrace{\tilde{H}\left(\bar{x}\right)}_{=0}g\left(\frac{\widehat{w}}{w_{\min}}\right)\\ &= \frac{1}{1-H(\bar{x})}\int_0^{\frac{\hat{w}}{w_{\min}}}h\left(\frac{\frac{\hat{w}}{\varepsilon}-K}{\beta}\right)\frac{g\left(\varepsilon\right)}{\beta\varepsilon}\mathrm{d}\varepsilon.\end{aligned} \tag{E23}$$

2. $w=w_{\min}$: We now turn our attention to workers earning the minimum wage. Due to measurement errors, they do not generate a mass point in the data but instead a non-degenerate distribution

$$\Pr\left(\hat{w}\leq y\middle|\,w=w_{\min}\right)=\Pr\left(\varepsilon<\frac{y}{w_{\min}}\right)=\int_0^{\frac{y}{w_{\min}}}g\left(\varepsilon\right)d\varepsilon=G\left(\frac{y}{w_{\min}}\right),$$

whose associated likelihood reads

$$L\left(\hat{w}\middle|\, w=w_{\min}\right)=\frac{\partial\int_0^{\frac{\hat{w}}{w_{\min}}} g\left(\varepsilon\right)\mathrm{d}\varepsilon}{\partial\hat{w}}=\frac{g\left(\frac{\hat{w}}{w_{\min}}\right)}{w_{\min}}. \tag{E24}$$

The likelihood of an observation is obtained using Bayes' rule to combine the two conditional likelihoods derived above

$$L\left(\hat{w}\right)=L\left(\hat{w}\middle|\, w=w_{\min}\right)\times\Pr\left(w=w_{\min}\right)+L\left(\hat{w}\middle|\, w>w_{\min}\right)\times\Pr\left(w>w_{\min}\right). \tag{E25}$$

The mass of workers earning the minimum wage is given by

$$\Pr\left(w\left(x\right)=w_{\min}\right)=\Pr\left(x\in\left[R_0,\bar{x}\right]\right)=\frac{H(\bar{x})-H\left(R_0\right)}{1-H\left(R_0\right)}, \tag{E26}$$

while the mass of workers with a wage higher than the minimum wage is

$$\Pr\left(w\left(x\right)>w_{\min}\right)=\Pr\left(x>\bar{x}\right)=\frac{1-H(\bar{x})}{1-H\left(R_0\right)}. \tag{E27}$$

Reinserting $(E23)$, $(E24)$, $(E26)$ and $(E27)$ into $(E25)$, we finally obtain

$$L\left(\hat{w}\right)=\frac{1}{1-H\left(R_0\right)}\left[\frac{g\left(\frac{\widehat{w}}{w_{\min}}\right)}{w_{\min}}\left(H(\bar{x})-H\left(R_0\right)\right)+\int_0^{\frac{\widehat{w}}{w_{\min}}} h\left(\frac{\frac{\widehat{w}}{\varepsilon}-K}{\beta}\right)\frac{g\left(\varepsilon\right)}{\beta\varepsilon}\mathrm{d}\varepsilon\right].$$

We assume in our estimation that productivities x and measurement errors ε are lognormally distributed; the distribution of x is Log-$\mathcal{N}\left(\gamma,\zeta^2\right)$ and that of ε is Log-$\mathcal{N}\left(-\sigma_\varepsilon^2/2,\sigma_\varepsilon^2\right)$. Then the log-likelihood of the vector $\hat{\mathbf{w}}$ of wages among job entrants is equal to $\mathcal{L}\left(\hat{\mathbf{w}};\Upsilon\right)\equiv\sum\log\left(L\left(\hat{\mathbf{w}}_i;\Upsilon\right)\right)\omega_i$, where $\Upsilon\equiv\{\gamma,\zeta,\sigma_\varepsilon\}$ is the vector of parameters and ω_i are the weights provided with the survey.

F Robustness of results for different values of the bargaining power β

Table 5: Average job duration (in year) and layoff costs

	$\beta=0.4$		$\beta=0.5$		$\beta=0.6$	
	(1)	(2)	(3)	(4)	(5)	(6)
	Mean	Median	Mean	Median	Mean	Median
Duration (flows)						
Actual EPL	5.8600	1.8356	5.7123	1.7627	5.5386	1.6948
No EPL	6.0022	2.0814	5.9006	2.0456	5.7098	1.9802
No jump	5.9099	1.8730	5.7846	1.8083	5.6395	1.7461
Elapsed duration						
Actual EPL	6.6793	5.2643	6.6920	5.2817	6.6917	5.2816
No EPL	6.4901	5.0110	6.4659	4.9786	6.4198	4.9165
No jump	6.6414	5.2081	6.6481	5.2148	6.6453	5.2060

Note: "Duration (flows)" stands for the expected jobs duration at their starting date. "Actual EPL" stands for the layoff costs estimated in the benchmark case corresponding to the French EPL. "No EPL" is the case without layoff costs, "No jump" corresponds to the actual EPL case with no jump in layoff cost at 2-year seniority threshold, i.e. $\phi_0=0$ in Table 2. "Elapsed duration" stands for the job durations sampled in the stock of jobs. This table reports results for the Benchmark population composed of 15-54 years old individuals with at most the high school degree.

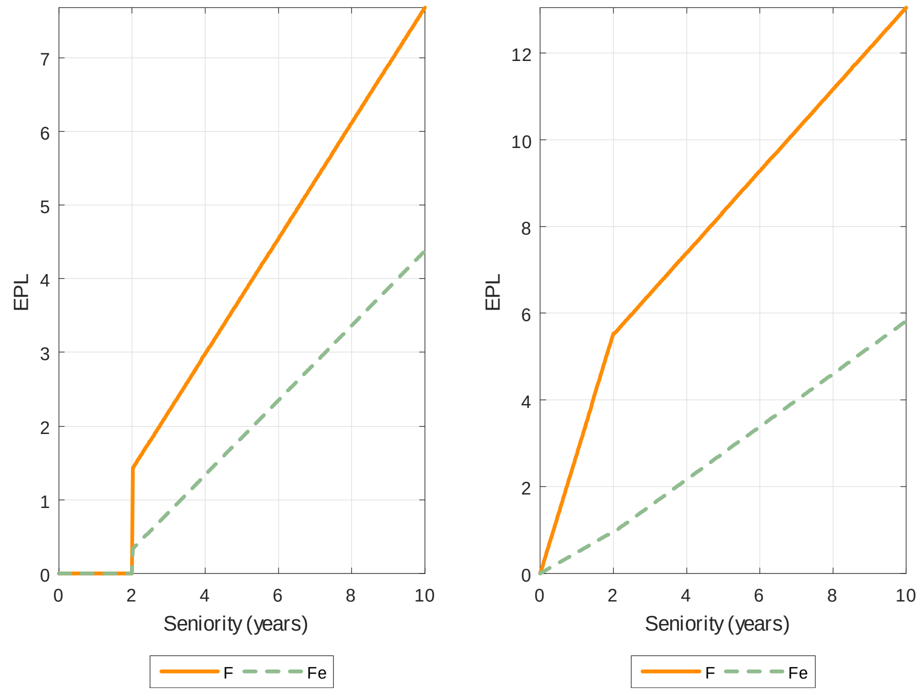

Figure 11: Total layoff costs (F) and severance payments (F_e) in profile 1 (left panel) and profile 2 (right panel).

Note: Profile 1 corresponds to the removal of all layoff costs before two years of seniority, keeping the layoff costs as in the benchmark case depicted in Figure 5 after two years of seniority; Profile 2 corresponds to the removal of the jump at two-year seniority, keeping the average layoff costs as in the benchmark case over twenty years of seniority and assuming that the ratio between the slope of total layoff costs and severance payments (i.e. ϕ_p/ϕ_{ep}) remains unchanged with respect to the benchmark after two years of seniority.

G Changes in layoff costs profile

This appendix presents the effects of changes in the profile of layoffs costs on job durations and on unemployment. Two profiles, displayed in Figure 11, are considered.

1. Removal of all layoff costs before two years of seniority, keeping the layoff costs as in the benchmark case depicted in Figure 5 after two years of seniority.
2. Removal of the jump at two-year seniority, keeping the average layoff costs as in the benchmark case over twenty years of seniority and assuming that the ratio between the slope of total layoff costs and severance payments (i.e. ϕ_p/ϕ_{ep}) remains unchanged with respect to the benchmark after two years of seniority.

The effects of these two layoff costs profiles are reported in Table 5. It turns out that the removal of layoff costs before two years (profile 1) of seniority reduces job durations, which entails a drop in employment at partial equilibrium, when the job arrival rate is exogenous. The decrease in job duration is induced by the fact that the anticipation of future layoff costs is not counterbalanced any more by layoff costs at the beginning of employment spells when there are no layoff costs before two years of seniority. However, the removal of layoff costs at the beginning of employment spells boosts job creation. Hence, despite the drop in job duration, the unemployment rate drops when the layoff costs before two years of seniority are removed. Table 5 shows that the removal of the jump of layoff costs at two-year seniority, keeping the average layoff costs as in the benchmark case over twenty years of seniority, raises job duration and reduces unemployment, as in

Table 6: Effects of different layoff costs profiles

	Benchmark (1)	Profile 1 (2)	Profile 2 (3)
Exogenous job creation			
Durations			
Mean (flows)	5.71	5.41	5.74
Median (flows)	1.76	1.89	1.80
Mean (elapsed)	6.69	6.43	6.64
Median (elapsed)	5.28	4.92	5.194
Endogenous job creation			
Unemployment rate	14.59%	12.51%	14.50%
Job finding Rate	1.02	1.33	1.03
Successful contacts	68.40%	79.12%	68.40%
Durations			
Mean (flows)	5.71	5.25	5.73
Median (flows)	1.76	1.84	1.80
Mean (elapsed)	6.69	6.39	6.63
Median (elapsed)	5.28	4.86	5.19

Note: Profile 1 corresponds to the removal of all layoff costs before two years of senority, keeping the layoff costs as in the benchmark case after two years of seniority; Profile 2 corresponds to the removal of the jump at two-year seniority, keeping the average layoff costs as in the benchmark case over twenty years of seniority and assuming that the ratio between the slope of total layoff costs and severance payments (i.e. f_p/f_{ep}) remains unchanged with respect to the benchmark after two years of seniority. "Job Finding Rate" is the annual job finding rate of unemployed workers. "Share Successful Contacts" is the proportion of matches that induce hiring. See Table 2 for the definitions of the other statistics.

the benchmark counterfactual exercises.

H Model fit for different categories of workers

H.1 Skilled workers

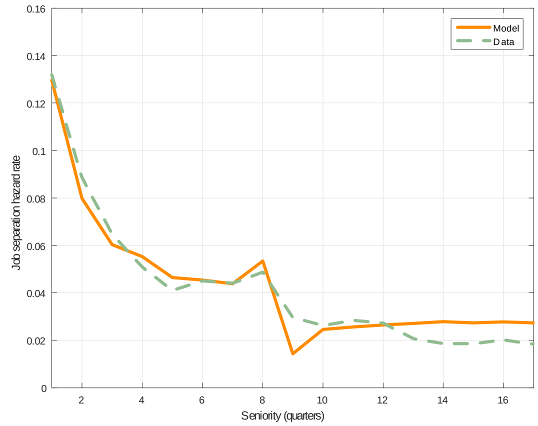

Figure 12: Actual and simulated hazard rate of job separation for 15-54 years old skilled workers, with at least the high school degree.

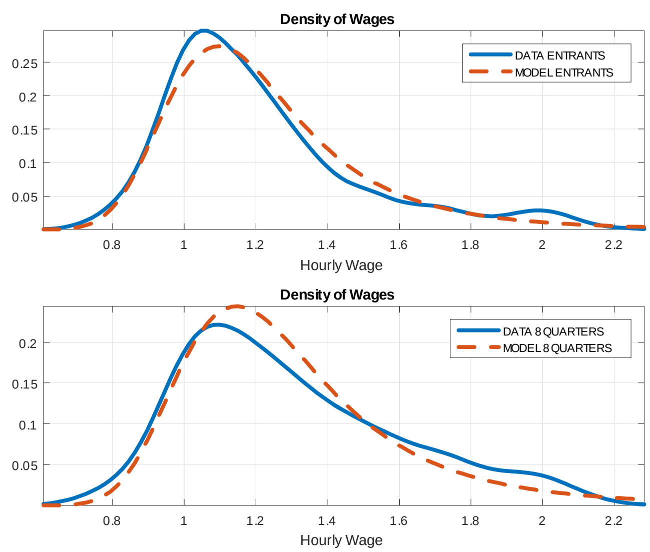

Figure 13: Estimated and empirical densities of hourly wages expressed in units of hourly minimum wages for 15-54 years old skilled workers, with at least the high school degree. Note: Job entrants are workers with 2 quarters of seniority at most. Source: Labor Force Survey (EnquÎte emploi, INSEE), 2003-2012.

H.2 Unskilled workers

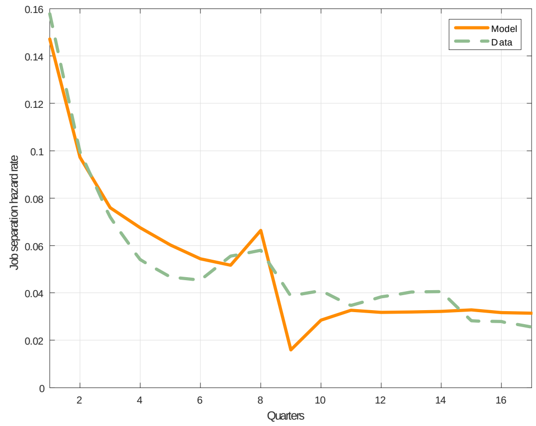

Figure 14: Actual and simulated hazard rate of job separation for 15-54 years old skilled workers, with at less than the high school degree.

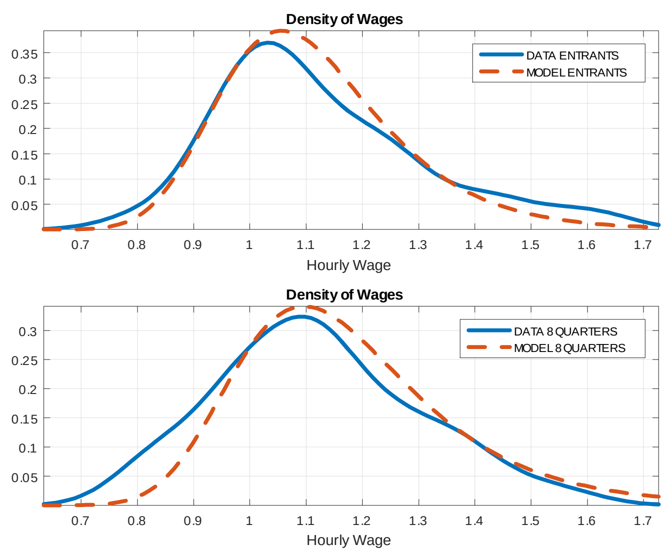

Figure 15: Estimated and empirical densities of hourly wages expressed in units of hourly minimum wages among job entrants for 15-54 years old unskilled workers, with less than the high school degree. Note: Job entrants are workers with 2 quarters of seniority at most. Source: Labor Force Survey (EnquÍte emploi, INSEE), 2003-2012.

H.3 Workers with less than 10 years of labor market experience

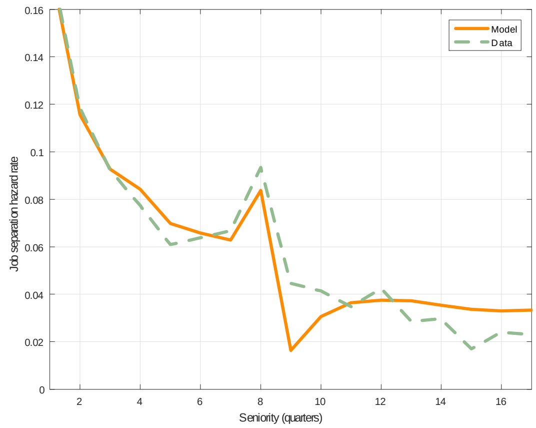

Figure 16: Actual and simulated hazard rate of job separation for workers who left schools less than 10 years ago, with at most the high school degree.

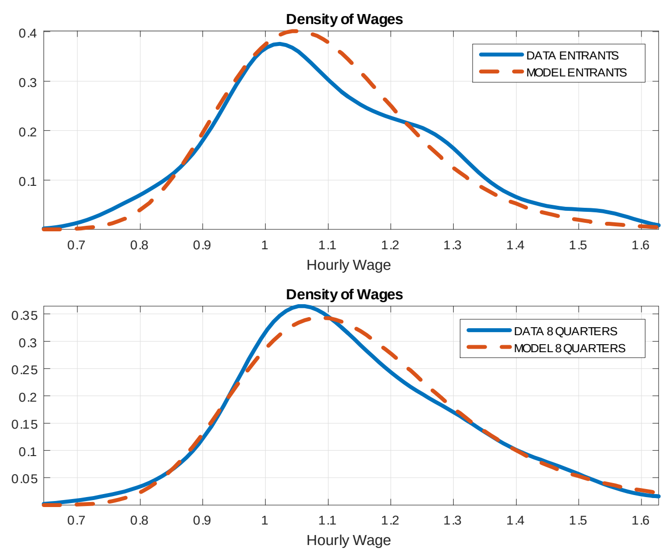

Figure 17: Estimated and empirical densities of hourly wages expressed in units of hourly minimum wages among job entrants for workers who left school less than 10 years ago, with at most the high school degree. Note: Job entrants are workers with 2 quarters of seniority at most. Source: Labor Force Survey (EnquÍte emploi, INSEE), 2003-2012.