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The effect of potential unemployment benefits duration on unemployment exits to work and on job quality

Effet de la durée maximale d'indemnisation du chômage sur le retour à l'emploi et sur sa qualité

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Avertissement : cette version annule et remplace la première version diffusée en mai 2011, suite à détection d'une erreur dans la construction de la variable « salaire précédant l'entrée au chômage ». Par rapport à la version initiale, le principal changement concerne l'effet de la durée maximale d'indemnisation sur la décôte salariale qui n'apparaît plus significatif. Le texte a été amendé en conséquence, s'agissant notamment de la présentation des statistiques descriptives sur les salaires (p. 5) et les covariates (p. 8), et des résultats d'impact sur la décôte salariale (p. 15).

Effet de la durée maximale d’indemnisation du chômage sur le retour à l’emploi et sur sa qualité

Thomas DEROYON et Thomas LE BARBANCHON

Résumé :

Selon les modèles standards de recherche d'emploi, des allocations chômage plus généreuses ralentissent les sorties du chômage et accroissent la qualité de l'emploi retrouvé. Cette étude teste empiriquement l’impact d’un allongement de la durée maximale d’indemnisation à l’assurance chômage en mobilisant une nouvelle source de données où sont appariées à un niveau individuel les listes de demandeurs d'emploi inscrits à Pôle emploi et les déclarations d'emploi des entreprises, source dite FH-DADS. Cette nouvelle source permet de caractériser le temps passé en emploi avant l'épisode de chômage, de mieux repérer les reprises d'emploi parmi les sorties des listes de Pôle emploi et de mesurer la qualité de l’emploi retrouvé (stabilité et salaire versé).

L’effet de la durée maximale d’indemnisation est estimé en comparant le devenir de nouveaux allocataires des filières 2 et 3 de l'assurance chômage entre 2000 et 2002. Sur cette période, lorsqu'un nouvel allocataire avait travaillé entre 6 et 8 mois pendant l'année précédant son entrée au chômage, il était indemnisé pendant une durée maximale de 7 mois (filière 2). Lorsqu'il avait travaillé au-delà de 8 mois sur l'année précédente, il pouvait être indemnisé pendant 15 mois (filière 3). Les allocataires des filières 2 et 3 représentent 28% des nouveaux allocataires sur la période 2000-2002. Ils sont plus jeunes, moins qualifiés et occupaient, avant leur inscription sur les listes, des emplois moins stables et moins rémunérateurs que la moyenne des nouveaux allocataires. L’estimation menée dans cette publication porte donc sur une population particulière des allocataires de l’assurance-chômage.

L'effet causal d'un doublement de la durée maximale d'indemnisation est estimé en comparant le devenir des allocataires de la filière 2 qui auraient été en filière 3 s'ils avaient travaillé un mois de plus (i.e. allocataires ayant travaillé entre 7 et 8 mois) et le devenir des allocataires de la filière 3 qui auraient été en filière 2 s'ils avaient travaillé un mois de moins (i.e. allocataires ayant travaillé entre 8 et 9 mois). Cette procédure d'estimation, connue sous le nom de « régression discontinue », suppose pour être valide que la répartition des allocataires autour du seuil de 8 mois puisse être considérée comme aléatoire. Elle suppose donc que les allocataires qui avaient travaillé entre 7 et 8 mois et entre 8 et 9 mois aient des caractéristiques observables proches (en termes de sexe, âge, niveau de formation, emplois antérieurs…), ce qui est vérifié ici. Elle suppose aussi que les allocataires de la filière 3 n’ont pas cherché à travailler davantage pour atteindre exactement le seuil d'éligibilité de la filière la plus généreuse. Si ce type de comportement était fréquent, le nombre d’allocataires ayant travaillé un peu plus de 8 mois serait sensiblement supérieur au nombre d’allocataires ayant travaillé un peu moins de 8 mois, ce qui n’est pas le cas.

Les résultats de l’estimation montrent que lorsque la durée d'indemnisation est portée de 7 à 15 mois:

- la part d'allocataires qui ont repris un emploi pendant les 10 premiers mois de chômage diminue de l'ordre de 6 points,
- la stabilité de l'emploi retrouvé et le salaire versé ne sont pas significativement améliorés.

L’impact de l’allongement de la durée maximale d’indemnisation sur le retour à l'emploi se décompose en un effet constant sur toute la période et un effet concentré avant l'expiration des droits.

The effect of potential unemployment benefits duration on unemployment exits to work and on job quality.

Thomas Deroyon Thomas Le Barbanchon

July 26, 2012

revised version - it replaces the "May 9, 2011" version

Abstract

According to standard job search theory, generous unemployment benefits tend to slow down exits from unemployment and to increase job quality of new matches. In our study, we query those effects in a new French data set, merging both unemployment and employment registers. This new data set enables us to construct good measures of exit to employment and job quality and observe past employment duration. In the French system, if past employment duration is under or over certain thresholds, the unemployed may receive benefits over a shorter or longer period. We exploit those discontinuities in the French eligibility system in a regression discontinuity design. We verify that the unemployed do not seem to precisely manipulate this forcing variable and we find a causal effect of unemployment benefits duration on the exit rate to employment, but no significant effects on subsequent employment duration and starting wage.

Keywords : unemployment benefits, job match quality, regression discontinuity
JEL codes : J64 J65 C41

1 Introduction

Unemployment benefits have a double objective: to insure workers against the loss of revenue due to job separation and to give them adequate financial means to look for another job. Nevertheless, as with any insurance, unemployment insurance is subject to moral hazard issues: too generous unemployment benefits may discourage the unemployed from searching for jobs and/or from accepting reasonable jobs. Thus, understanding the effect of unemployment insurance generosity on job finding rates and on the quality of jobs is of prime concern. When unemployment insurance is more generous, is job search activity less intensive? More productive? Does a more generous unemployment insurance lead to better jobs?

According to standard job search theory, a more generous unemployment insurance increases the reservation wage of the unemployed. This induces the unemployed to be more selective among job offers. They stay unemployed longer and the distribution of jobs accepted should be of better quality. Apart from this effect on labor supply, generous unemployment insurance may give the unemployed the opportunity to take advantage of increasing returns in job search. At the beginning of an unemployment spell, it certainly takes time for the unemployed person to think through his job project and to start searching in the right employer pool. Of course, we also expect job search activity to feature decreasing returns when enough time has been devoted to searching.

However other mechanisms, when they are not fully internalized by the unemployed, may lead to a negative link between unemployment insurance generosity and job quality. If the benefit recipient stays unemployed for too long, his or her human capital may depreciate. This depreciation induces a negative link between unemployment duration and job quality. The time spent unemployed may also be a screening device for the employers. This also induces a negative link between unemployment duration and job quality, and thus a negative link between unemployment insurance generosity and job quality. As theory is ambiguous about the effect of unemployment insurance on job quality, we need to query this effect from an empirical point of view.

We expect the effect of unemployment insurance generosity on job quality to be linked to the lengthening of unemployment duration. According to numerous empirical studies (see van Ours and Vodopivec (2006a) for a survey), there is a positive link between unemployment insurance generosity and unemployment duration. In more recent studies, authors focus on identifying a causal relationship through difference in difference methods (van Ours and Vodopivec (2006a) ; Lalive, Ours, and Zweimuller (2006)) or through regression discontinuities (Lalive (2008); Caliendo, Tatsiramos, and Uhlendorff (2009)).

Moreover, the effect of unemployment benefits on the job finding rate is not homogenous over the unemployment spell. Meyer (1990) finds spikes in the unemployment exit rate just before the exhaustion of unemployment benefits. This is evidence that the unemployed react to financial incentives. The expected profile of unemployment benefits conditions the search behavior of the unemployed. So changes in this profile are informative about the link between unemployment benefits and unemployment duration. Building on this idea, Dormont, Fougere, and Prieto (2001) also verify the existence of spikes at the exhaustion of unemployment benefits in France. Spikes are especially large when estimated on unemployed persons who were well paid before their unemployment spell. Evidence on spikes at exhaustion has recently been criticized by Card, Chetty, and Weber (2008): it is usually based on data in which destinations of unemployment exits are only correctly observed before the exhaustion. Using a richer Austrian data set, they show that spikes disappear. Our data set is robust to that criticism[1].

Evidence on the effect of unemployment insurance generosity and employment quality is scarce and contrasting (see the review in Addison and Blackburn (2000)). Addison and Blackburn (2000) and Card, Chetty, and Weber (2007) find limited effects on post unemployment earnings, whereas Ehrenberg and Oaxaca (1976) find positive effects. Tatsiramos (2006), and Caliendo, Tatsiramos, and Uhlendorff (2009) find positive effects on job stability, whereas van Ours and Vodopivec (2006b) and Belzil (2001) find limited effects. More precisely, Caliendo, Tatsiramos, and Uhlendorff (2009) find that unemployed persons who find jobs just before their unemployment benefits run out accept less stable jobs than comparable unemployed persons who benefit from longer entitlement.

In this paper, we estimate the causal impact of potential benefit duration on unemployment exits to work and on subsequent employment duration and wage. Our main contribution is to identify this impact through a new regression discontinuity design inspired by Card, Chetty, and Weber (2007) and Card, Chetty, and Weber (2008). We exploit discontinuities in the eligibility rules of the French unemployment insurance system in 2000-2002. Depending on their past employment experience, the unemployed fall into different benefit categories, each category defined by a specific maximal benefit duration. Thus, when job-seekers' past employment duration crosses eligibility thresholds, they are entitled to longer unemployment benefit duration. More precisely, we will focus on unemployed persons who work around 8 months during the year before their unemployment spell starts. Working between 6 and 8 months makes workers entitled to 7 months of unemployment benefits, whereas working between 8 and 12 months opens up 15 months of benefits. Crossing the 8 month threshold doubles the generosity of unemployment insurance.

Our empirical objective relies on an original French data set which merges, at the individual level, unemployment spells recorded at the French Employment Agency, and employment

[1] See Boone and van Ours (2009) for another example

spells reported by employers to the French administration for fiscal purposes. This new administrative data set gives unique information on the destination of unemployment exits and on the quality of exit jobs. It also gives unique information on past employment experience, which is crucial to our identification strategy. In this new data set, we observe past employment durations, information which is not precisely recorded in unemployment registers.

Our paper starts with a complete description of our sample selection. In the second part, we motivate our regression discontinuity design. In the third part, we show that extended benefit duration tends to slow down unemployment exits. We also present evidence of spikes at benefits exhaustion. In a fourth part, we show that extended benefit duration does not have any significant effects on job stability, nor on post-unemployment wage. These last results are less robust as selection bias among reemployed workers is not seriously corrected.

2 Sample selection and measurement issues

Our sample is selected from a new French matched unemployment-employment registers data set (a complete description can be found in appendix A). These data give unique information on the prior and posterior employment spells of benefit recipients. This information is crucial to implementing our regression discontinuity design and to inspecting post unemployment job quality.

We select a flow of new unemployment benefit recipients who enter the Employment Agency from 2000 to 2002[2]. To avoid identification problems caused by the specific policies aimed at senior job seekers, we exclude from our sample people 50 years old and more[3].

Between January 2000 and December 2002, new unemployment benefit recipients might enter into one out of 4 categories, called *filières*. Each *filière* has a specific maximal benefit duration which depends on past employment duration over a reference period. Because *filière 2* and *filière 3* share the same reference period, they can easily be compared in a regression discontinuity design. We will thus focus on those 2 categories. Jobs seekers in *filière 2* are entitled to 7 months of unemployment benefits; they will be referred to as short benefit duration job seekers. Those in *filière 3* will be referred to as extended benefit duration job seekers; they are entitled to 15 month benefits. In both categories, job seekers benefit from the same replacement rate. The replacement rate rule changed in July 2001. Between January 2000 and June 2001, the replacement rate was smoothly decreasing with time elapsed in unemployment.[4] After July 2001, it was constant during the whole benefit duration.

Short and extended benefit duration job seekers represent 28% of all new unemployment benefit recipients. The majority (63%) of recipients are entitled to 30 month benefits (*filière 5*). Those recipients were employed during at least 14 months before registering as unemployed, i.e. a longer period than recipients in our sample of interest. Thus we will identify the impact of maximal benefit duration on recipients with relatively low employability. Table 10 in appendix A shows the job seekers' characteristics for different *filières*. Job-seekers entering in *filière 2* and *filière 3* (first column) are younger and have lower education and qualification than those entering *filière 5* (second column). The proportions of women and foreigners are higher. Their previous job position was less stable and less rewarding : only

[2]We only retain new unemployment benefit recipients, meaning they do not have any residual benefits left from a former unemployment spell. Therefore, their potential benefit durations are directly linked with their employment spells since their last unemployment spell. When benefit recipients have residual benefits, a complicated rule extends their residual benefits according to their last employment spells.

[3]We also drop from our sample job seekers entitled to very specific insurance rules (recurrent temporary workers, artists, technicians working in the culture sector...).

[4]After 4 months, recipients were to lose around 15% of their benefits; after 10 months, there was a further 15 % decrease. Decreasing replacement rates make the difference in generosity between categories less important before than after July 2001. This difference turns out to be minor and does not induce great change in estimated effects.

14% had a permanent contract before their job separation, and their daily wage was 25% lower than in the broader group. Finally, they had spent almost a year unemployed during the last 3 years.

We further restrict our sample to recipients with "consistent" employment records in the employment registers. More precisely, we drop recipients whose past job spells, as recorded in the employment registers, are not consistent with their maximal benefit duration recorded in the unemployment registers (43%)[5]. This restriction is crucial for our regression discontinuity design : we need to observe the pre-unemployment job spells[6]. As displayed in table 10 in appendix A, "consistent" job seekers have a stronger relationship to work than the unrestricted sample : they are more often qualified men, with high levels of education, higher former wages and longer past tenure lengths, and they have been less often registered as unemployed in the past three years. This is no surprise as stable jobs are better reported in the employment registers. We also verified that unemployed persons looking for a job in the agriculture or in the care sectors are more likely to have inconsistent employment records. Their former employers, probably in the same sector, are not covered by the employment registers.

Despite those inconsistencies, adding employment registers information to unemployment registers clearly increases the quality of measurement of unemployment exits to work. In our sample, 35% of the unemployed leave the Employment Agency reporting they have found a job. However, 29% of the unemployed leave the Employment Agency without reporting their new situation to their caseworkers, and the Employment Agency drops 9% of the unemployed for administrative reasons (not showing up for counselling...). Those benefit recipients may also have found a job. Indeed, 41% of the unemployed leaving the Employment Agency start a job recorded in the employment registers around their exit date[7].

Measuring jobs in employment registers not only increases the levels of exits to jobs, it also alters their timing. The lack of information due to missing job seekers' reports usually blurs the variations of exit rates to employment at benefit exhaustion and casts doubts on the existence of spikes at that time. The exit rate to jobs, as reported to the Employment Agency, does indeed rise and decline before benefit exhaustion (see the second graph in panel 1). However the exit rate to jobs, as recorded in the employment registers, rises before the end of benefit exhaustion and reaches a spike just after it (see the first graph in panel 1). This certainly highlights a change in the reporting behavior of job seekers at benefit exhaustion.

The unemployment registers do not contain any information about the exit jobs of benefit recipients. The employment registers help us to describe the employment duration of newly employed workers, their wages, and thus the part of their former wages they were able to recover.

In our sample, the median employment duration is 6 months[8]. The monthly job separation rate shows spikes at the usual temporary contracts durations : 6, 12 and 24 months (see the first graph in panel 2). Former job seekers with extended benefit duration stay longer in their new jobs than those with short benefit duration : the median of employment duration increases by 1 month between the two groups.

Half of job seekers gain more than 2 % of their former real hourly wages when they start a new job[9]. The wage gain is higher for the extended benefit duration job seekers (see the second graph in panel 2) : whereas more than one half of workers from the short benefit

[5]29% have no employment spell recorded in employment registers, before unemployment.

[6]The past employment duration recorded in unemployment registers is not precise. When the unemployed claim their benefits, caseworkers manually enter past employment duration in the information system of the French Employment Agency. They usually enter the minimal past employment duration that makes the unemployed person eligible in his *filière*. There is no requirement to record the exact past employment duration, nor automatic control.

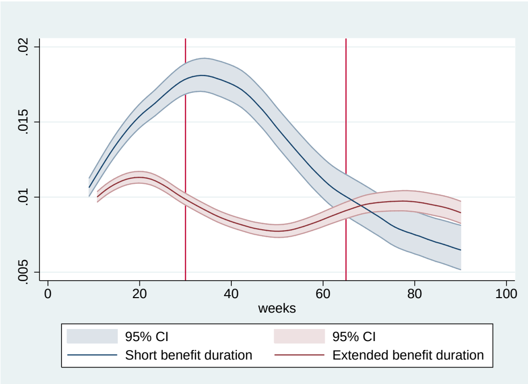
[7]The corresponding employment spell should begin at most sixty days before or after the actual exit date and it should not end before it.

[8] Note that 14 % of new jobs spells are censored at the end of the data set (December 2004).

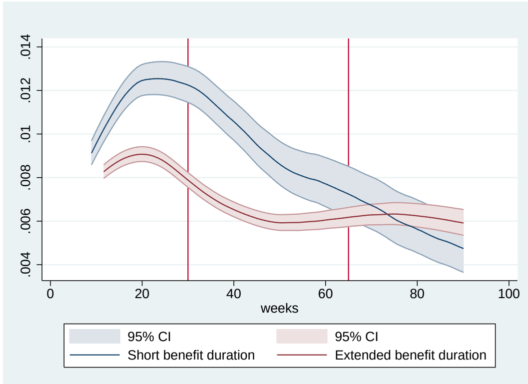
[9]Wage loss is computed as the ratio of starting wage over pre-unemployment wage as computed in the employment registers.

Table 1: *Weekly unemployment register exit rates*

to jobs recorded in employment registers



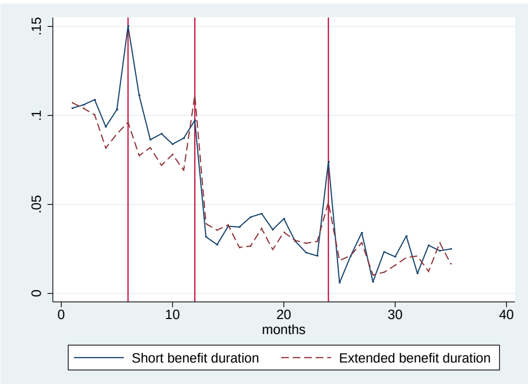
to jobs reported in unemployment registers



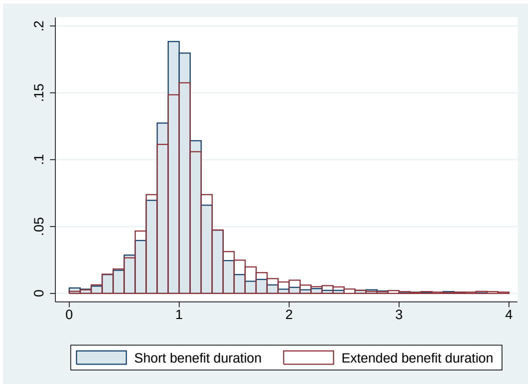
Reading: vertical lines represent dates of benefit exhaustion for short benefit duration (*filière 2*) and extended benefit duration (*filière 3*).

Table 2: *Subsequent job quality*

Monthly job separation rate



Wage loss density



Reading: on the left graphics, vertical lines represent typical temporary contracts durations (6 months, 1 year, 2 years).

duration category do not recover their previous wage, the median wage gain is more than 3 % among extended duration job seekers.

The previous descriptive statistics show that job-seekers entitled to longer benefits take more time to find a new job (see panel 1). Their new jobs last longer and are more rewarding. Those differences shed some light on the link between unemployment generosity and return to work and job quality. However they could reflect the fact that recipients with extended benefit duration have worked during a longer period before becoming unemployed. They could have both observable and unobservable characteristics which make them less effective in job search, but more productive in their new jobs. In the following, our comparisons will be robust to this endogeneity bias : we use a regression discontinuity design.

3 Regression discontinuity

Comparing individuals who have been randomly assigned extended potential benefit duration is the ideal design to estimate its causal effect. In a regression discontinuity framework (see Imbens and Lemieux (2008)), assignment to the extended benefit duration is locally random around the threshold of one forcing variable, here past employment duration. Then any difference in outcomes between recipients who are just below and just above the threshold can be attributed to the effect of extended potential benefit duration. The randomness assumption is impossible to test. However we first explain its credibility in our case. Then we test whether recipients just below and just above the threshold have similar observable characteristics.

3.1 Is employment duration precisely manipulated?

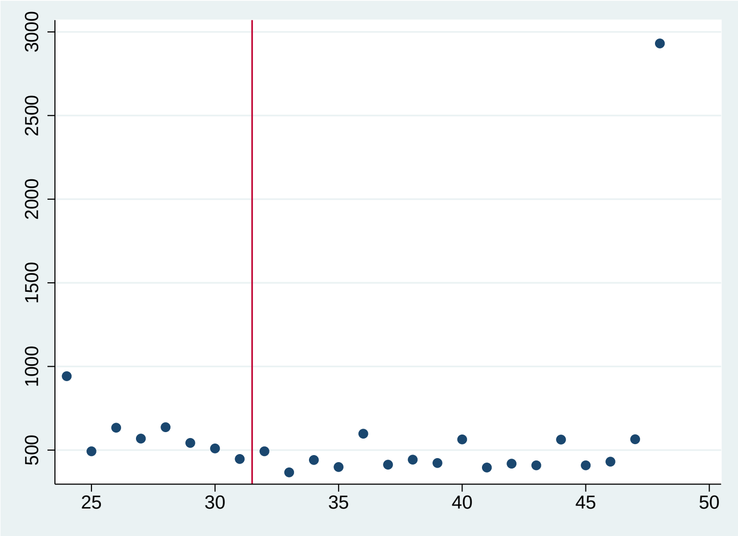
Local randomness of the forcing variable is not verified if some benefit recipients are able to precisely manipulate their employment duration. If that were the case, those individuals who manipulate employment duration would be just above the threshold, and the comparison of benefits recipients just below and just above the threshold would be biased. Actually, individuals who manipulate their employment duration are likely to have special characteristics highly correlated with unemployment exit rates, subsequent employment duration and wages.

Manipulation could occur at different stages : at benefit registration, when employer and employee separate, or when they first meet. Our measure of past employment is robust to fraud at benefit registration. We observe past employment from an external source, not from administrative recordings at benefit registration, and we drop observations with inconsistent past employment history. By the way, our sample excludes workers whose past employment certificates shown at benefit registration are more often erroneous, namely recurrent temporary workers and technicians working in the culture sector (see recent reports from the French *Cour des comptes*). Because most of the job seekers in our sample separate from temporary contracts, we believe that manipulation at job separation is less a concern than in the general case. The use of temporary contracts, and their extensions, is highly regulated in France. However we cannot exclude that employer and employee collude when they first meet, and set the contract duration so that it exactly extends the worker's past employment duration to meet the eligibility criteria to extended duration. The only argument which could limit the prevalence of collusion is that the employment prospects of our sample are structurally small. They are less educated and less qualified than the typical French worker. This should limit their ability to bargain such a trade off.

Turning to statistical argument, forcing variable manipulation can be checked by inspecting the population density around the eligibility threshold. If employment duration were precisely manipulated, recipients would accumulate just above 8 months (32 "weeks"[10]). We do not see any discontinuity on the graph in figure 1. This is confirmed by a formal discontinuity test (see the appendix B). The test has been augmented to account for the fixed term job duration periodicity. Most of the contracts are actually written in months.

[10] In the employment registers, time is scaled such that each month has 30 days, each year has 360 days. Hence, we define an employment "week", as one fourth of a month.

Figure 1: *Benefit recipients density along the past employment duration in weeks (forcing variable)*



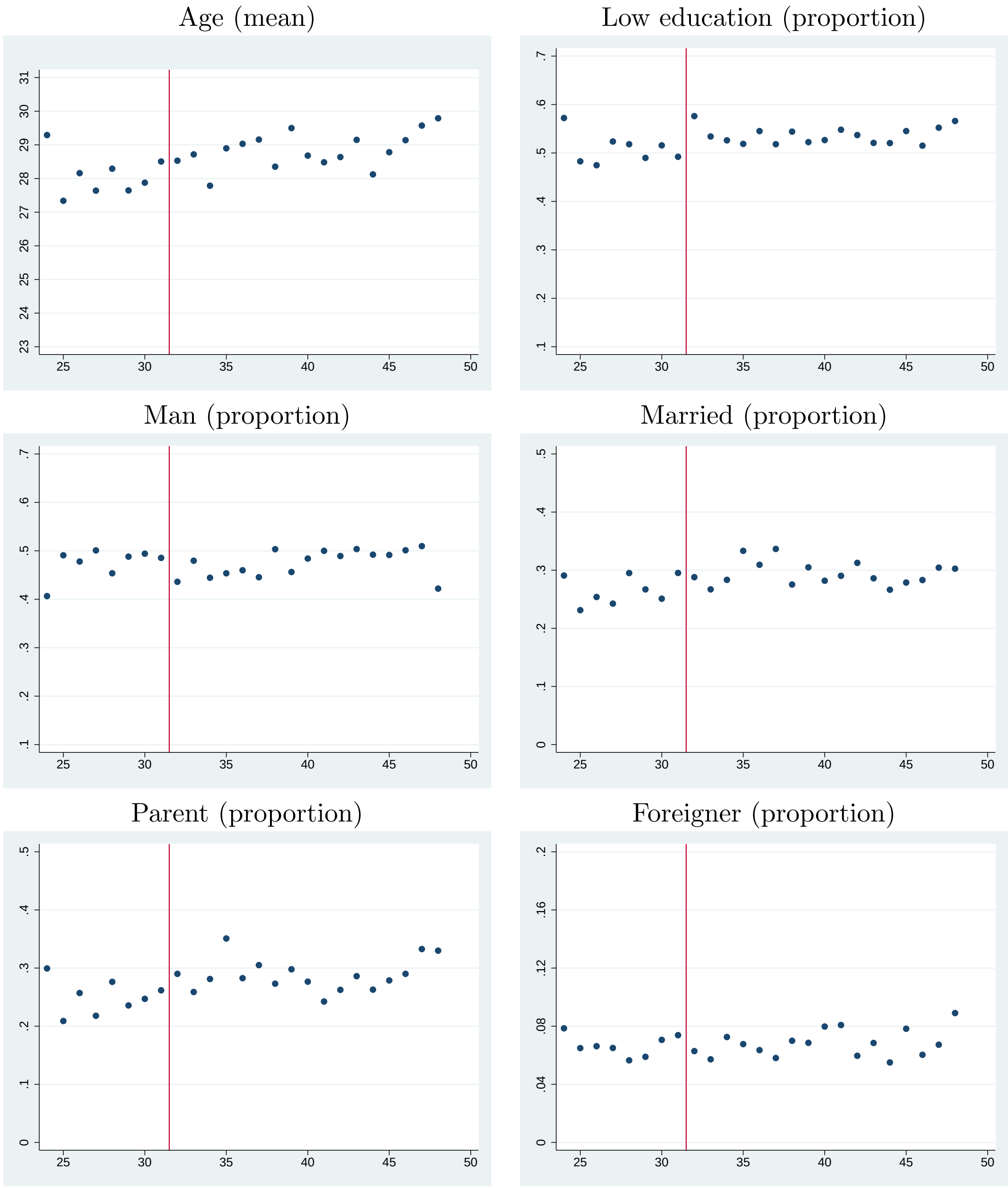
Reading: on the x axis, we report past employment duration in "weeks"; it starts from 6 months (24 "weeks"), this is the minimum employment duration to enter *filière 2*. The vertical line represents the threshold between short and extended benefit duration. Mass points are found at typical contract duration (6 and 12 months).

3.2 Covariates around the threshold

Further evidence of the forcing variable exogeneity can be found by inspecting recipients' characteristics around the threshold. There should be no discontinuities in the proportion of men, low qualified workers... This is illustrated by the graphics in tables 3 and 4. We can see moderate discontinuities in education (upper right graphics in table 3) and the type of contract before unemployment (lower right graphics in table 4).

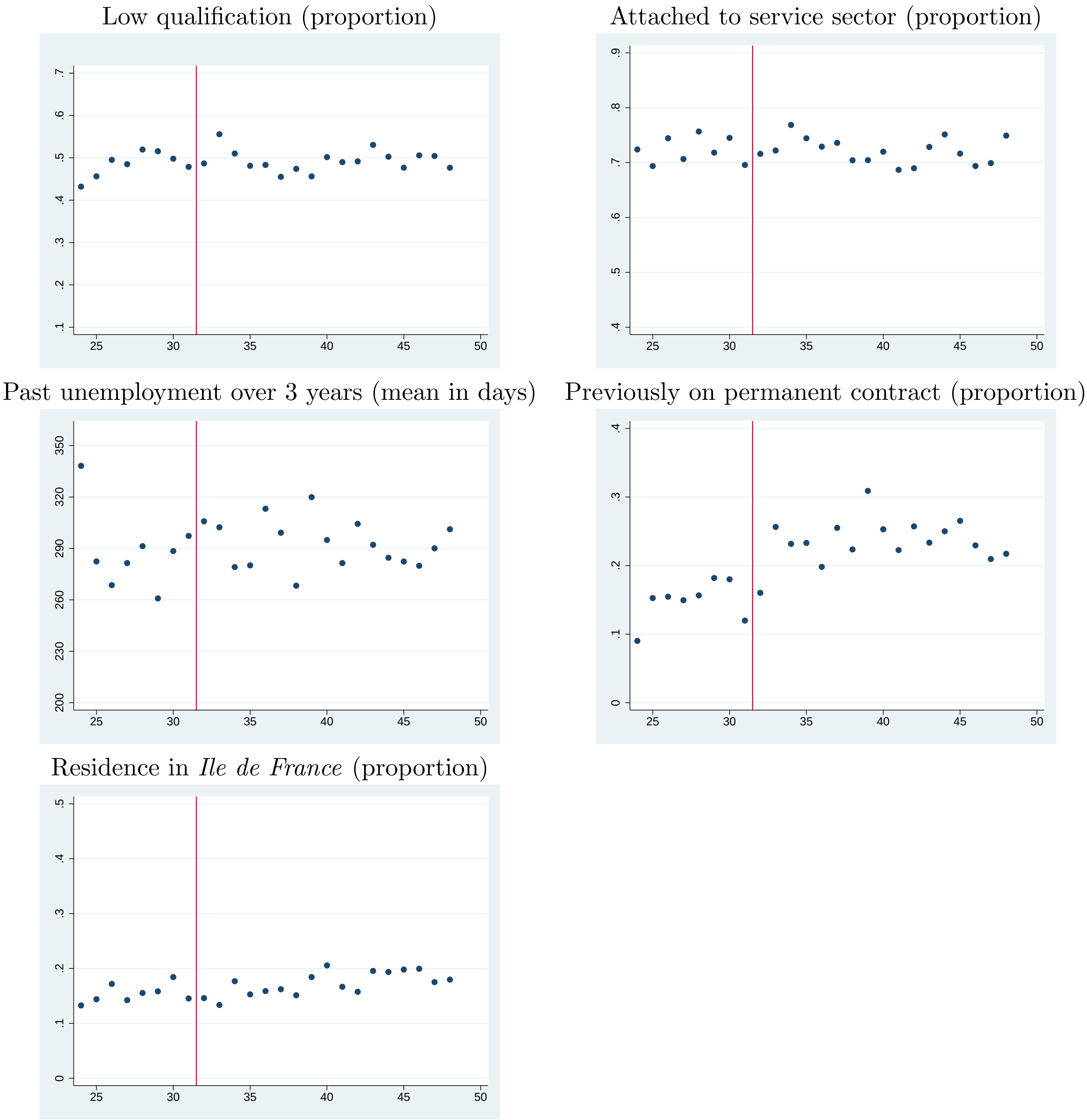
To test for discontinuity we run several linear regression discontinuity estimations on different windows around the threshold. Results are reported in table 11 in appendix B. We confirm significant discontinuities in the education level; the result of the test is robust to varying the window around the threshold down to 1 month. We also confirm moderate discontinuities in the type of the previous contract. We find moderate discontinuities in gender and in the separation date seasonality (no graphics shown). For all the other covariates, the formal tests show no discontinuities when the width of the window around the threshold is less than 2 months. A conservative conclusion would be that 4 out of 13 dimensions tested show discontinuities. We can thus be confident with our "no manipulation" assumption, but we will verify that our regression discontinuities results are robust to covariates controls.

Table 3: *Covariates distribution and past employment duration in weeks (forcing variable) (I)*



Reading: on the x axis, we report past employment duration in ”weeks”; it starts from 6 months (24 ”weeks”), this is the minimum employment duration to enter *filière 2*. The vertical line represents the threshold between short and extended benefit duration.

Table 4: *Covariates distribution and past employment duration in weeks (forcing variable) (II)*



Reading: on the x axis, we report past employment duration in "weeks"; it starts from 6 months (24 "weeks"), this is the minimum employment duration to enter *filière 2*. The vertical line represents the threshold between short and extended benefit duration.

4 Effect of potential benefit duration on exits to work

The effect of potential benefit duration on exits to work is first estimated using standard regression discontinuity linear models. Estimation confirms what can be inferred from graphics in table 5: there is a strong effect of extended benefit duration on the job finding rate around the short duration benefit exhaustion date (7 months). Before exhaustion, for example 4 months since registration, the share of benefit recipients who have found a job seems to be continuous while crossing the threshold of the forcing variable. After exhaustion, for example 10 months since registration, it jumps while crossing the threshold. In a second step, we model job finding rate *à la* Card, Chetty, and Weber (2008). This model takes into account censoring and spikes at benefit exhaustion.

Table 5: *Unemployment register exits to jobs and past employment duration in weeks (forcing variable)*

Reading: on the x axis, we report past employment duration in "weeks"'; it starts from 6 months (24 "weeks"), this is the minimum employment duration to enter *filière 2*. The vertical line represents the threshold between short and extended benefit duration.

4.1 Exits to work during the first 4, 10 and 18 months after registration

In the following we consider job finding during periods of 4, 10 and 18 months after registration. Those dates are representative of different periods in the recipient history. At 4 months, unemployed in both categories receive benefits. At 10 months, only unemployed in the extended duration category receive benefits. At 18 months, all benefits have expired. The linear regression discontinuity model we estimate is the following:

$$S_m = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d})\left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + \gamma X + u \qquad (1)$$

where S_m is equal to 1 if the unemployed person finds a job during the first m months after registration[11], d is the past employment duration, $\bar{d}$ is the eligibility threshold to extended benefits (8 months), X is a set of covariates. $I(d >= \bar{d})$ indicates treatment. When past employment (d) crosses the eligibility threshold ($\bar{d}$), job seekers are entitled to extended benefit duration. Thus our parameter of interest is δ, whose estimates are reported in table 6. δ_{-1} and δ_1 capture any linear dependencies of the outcome to past employment duration, which can be different before and after the threshold.

In the first column of table 6, there is no window restriction to select recipients. We find no effects of extended benefit duration on unemployment exits during the first 4 months. However, as soon as the outcome is observed after the short benefit duration (7 months), receiving benefits for an extended duration is associated with reduced job finding. After 10 months the share of unemployed who have deregistered to start a new job is 10 points lower in the extended benefit duration category. This is a substantial decrease of 25 % on unemployment exits share.

In the other columns of table 12, we report estimation results on narrower windows. For example, in column 2, the estimation sample is restricted to job seekers in a 4 month window around the 8 month threshold : their past employment duration is thus between 6 and 10 months. Those restrictions make populations above and below the threshold more and more similar. Results are then more robust to misspecification errors (covariates, linear dependency of the distance to the threshold). This advantage comes at the price of precision loss. Indeed, no effect is significant when the window is one month wide (15 days below and 15 days above the threshold, column 4).

When the window is gradually narrowed from 4 to 2 months, there is still an effect of extended benefit duration, especially when destination is observed after 10 months. When the window is 2 months wide, the share of unemployed who have deregistered during the 10 first months to start a new job is 6 points lower in the extended duration category. This effect is significant at the 5% level. When the window is 4 months wide, short benefit duration recipients find jobs faster (7 points significant at the 5% level). In appendix C, estimations are conducted without controls; results are robust.[12]

4.2 Spikes at benefit exhaustion?

The previous linear regression model does not explain precisely when the effect of benefit duration is felt. Neither does it model unemployment spells censoring. The following Cox model solves those shortcomings; it is inspired from Card, Chetty, and Weber (2008). Namely, this Cox model tests for spikes at benefit exhaustion. Formally, the job finding rate at time t after registration (θ_t) depends on a baseline hazard rate (h_t), on benefit duration category, on time to benefit exhaustion and on covariates as follows:

$$\theta_t = h_t \exp\left(I(d < \bar{d}) \left(\delta + \sum_{k=-2..2} \delta_k I(t \in I_k) \right) + \gamma X \right) \qquad (2)$$

where all notations have already been defined except I_k which represents time before or after exhaustion (7 months after registration for short duration benefits). I_0 marks if the observation is in the exhaustion week, I_1 if it is in the 2 following weeks after exhaustion

[11]Unemployed persons find a job if they quit the unemployment registers and start a new job reported in social security data at almost the same date. The effects have also been estimated on the competing definition of job finding: jobs reported by the unemployed to the Unemployment Agency. As the descriptive statistics suggest, the timing of the effect is different.

[12]Estimations with higher polynomial degrees ($(d - \bar{d})^2$ and $(d - \bar{d})^3$) have also been conducted. Results are robust when the polynoms are cubic. They are less important when the polynoms are quadratic.

Table 6: *Effect of extending potential benefit duration with covariate controls.*

Exits to work	Window around the threshold			
	All	4 months	2 months	1 month
After 4 months	-.005	.008	-.015	.015
	(.013)	(.016)	(.023)	(.032)
After 10 months	-.102***	-.074***	-.059**	-.031
	(.017)	(.021)	(.030)	(.041)
After 18 months	-.085***	-.058***	-.045	-.028
	(.018)	(.022)	(.031)	(.043)
Without restriction	-.044**	-.025	-.019	.014
	(.018)	(.022)	(.031)	(.043)

Standard errors are robust to White heteroscedasticity, as in all the following regressions. All covariates tested in the previous section are included : gender, nationality, age, education, parent, married, residence in Parisian region, qualification, past wage, past employment history, preferred sector, seasonality and time dummy.

week, I_{-1} in the 2 previous weeks before exhaustion week... One parameter of interest is again δ. Note that contrary to the previous estimation, it captures the effect of shortening unemployment benefits duration ($I(d < \bar{d})$). δ correspounds to a global effect, constant all over the unemployment spell. All the other δs (δ_0, δ_1...) capture the local effects of shortening benefit duration concentrated within 2 months around benefit exhaustion.

Estimation results in table 7 confirm the existence of local effects at short duration benefit exhaustion (spikes at benefit exhaustion). As above, each column reports the estimation results on a specific subsample of job seekers whose past employment duration is concentrated around the 8 month eligibility threshold. Above a constant effect all over the unemployment spell (δ), the unemployment exit rate 2 months around the exhaustion week is between 50 % and 130 % higher for short duration benefit recipients than for extended duration benefit recipients who have been working between 8 and 10 months (column 2). The spikes are still large when the comparison is restricted to a 2 month window around the eligibility threshold (column 3). However, spikes just after potential benefit exhaustion are no longer significant. As found in the previous section, effects diminish in the 1 month window estimation; whatever the time around the exhaustion week, spikes are not signicant (in column 4). The broad picture is the same should we exclude covariates from the regression (see table 7 in appendix C).

Spikes are quite widespread around the potential exhaustion week. There are two reasons for this. Firstly, unemployment deregistration is not "continuous", as the Employment Agency controls job search on a monthly basis. Then, some unemployed may find a job exactly during the exhaustion week, but they may be deregistered a few weeks later. Secondly, 7 month benefit duration is not calendar tested. French unemployment insurance rules feature schemes giving claimants incentives to work : "reduced activity". When benefit recipients find low-paid part-time jobs, they can cumulate their wages and benefits. Such schemes delay the exhaustion week. For example, consider a claimant who works 15 days with the same wage rate as in his former job. His exhaustion week will be 2 weeks later than 7 months since registration. "Reduced activity" concerns more than one third of claimants in our sample. It is difficult to control for, because benefit duration may also have an impact on the "reduced activity" take-up. In the previous estimation, we consider potential exhaustion week, i.e. 7 months since registration. This should imply conservative spikes estimates.

Spikes at benefit exhaustion do not capture the entire effect of shortening benefit duration. In table 7, we estimate a large global effect (δ in line 1). Thoses estimates are more conservative than local effects : shortening benefit duration increases job finding rate by 15% all along the unemployment spell (column 3, taken to be representative of line 1). Effects are even significant when the comparison is restricted to benefit recipients who worked between 7 months and a half and 8 months and a half (one month window, column 4).

Table 7: *Effect of shortening potential benefit duration on unemployement exit rate with covariate controls.*

		Window around the threshold		
	All	4 months	2 months	1 month
Short duration benefit (δ)	1.137***	1.113***	1.155***	1.255***
	(.031)	(.039)	(.059)	(.094)
3-4 weeks before exhaustion week (δ_{-2})	1.559***	1.924***	1.898**	1.663
	(.232)	(.427)	(.601)	(.830)
1-2 weeks before exhaustion week (δ_{-1})	1.469***	1.648**	1.618*	.986
	(.205)	(.328)	(.467)	(.389)
short duration benefit exhaustion week (δ_0)	1.704***	1.836**	1.589	1.583
	(.352)	(.543)	(.732)	(1.126)
1-2 weeks after exhaustion week (δ_1)	1.787***	1.496**	1.573	.598
	(.276)	(.304)	(.472)	(.267)
3-4 weeks after exhaustion week (δ_2)	2.023***	2.271***	1.760*	1.441
	(.320)	(.530)	(.538)	(.647)
Subjects	16692	8352	3837	1817
Log-likelihood	-67126.58	-31824.18	-13687.92	-5654.558

All covariates tested in the previous section are included : gender, nationality, age, education, parent, married, residence in Parisian region, qualification, past wage, past employment history, preferred sector, seasonality and time dummy.

5 Effect of potential benefit duration on job quality

Job quality is by definition observed for the job seekers who find a job. From now on, our population of interest is restricted to unemployed persons who exit unemployment registers to start a new job. As before, we compare employment durations and starting wages of job seekers in short and extended duration categories. Because the comparison is now on job finders, it may suffer from a selection bias into employment which we do not take into account in this paper. Indeed the job seekers induced to exit unemployment because of shorter benefits duration may be a very special population with intrinsic characteristics that make them work in different jobs. Then comparing characteristics of jobs found after short and extended benefit duration unemployment spells results in comparing individual characteristics rather than measuring the causal impact of benefit length. Anyway, some evidence shows that this bias may not be so dramatic: the fraction of job seekers who find a job during the maximum observation time after registration is the same across benefits duration categories (see bottom right corner graphics in panel 5).

We divide employment spells after unemployment exits into two broad categories : those lasting strictly less than 8 months and those lasting more than 8 months. We thus separate jobs into typical short temporary contracts and more stable employment relations (panel 2). 8 months is an interesting threshold : it is indeed the extended benefit duration eligibility threshold. Then former job seekers who find a job lasting more than 8 months are entitled to extended benefit duration. Were they already in this category, 8 months can be understood as a renewal threshold. The effect on this binary variable is estimated using a linear regression discontinuity model[13] (equivalent to model 1).

When we compare job seekers who worked between 6 and 10 months (column 2 in table 8), extended benefit duration seems to have a positive effect on employment duration. Extending benefit duration increases the proportion of jobs lasting more than 8 months by 8 points.

[13]We can abstract from censoring issues : there are virtually no employment spells censored before 8 months.

However, when the window is less than 2 months, the effect of extended benefit duration on employment duration is lower and not significant (columns 3 and 4). As a consequence, there is no evidence of a causal impact of extended benefit duration on employment duration. This conclusion is robust, when covariates are excluded (see appendix C).

Table 8: *Effect of extending potential benefit duration on employment duration with covariate controls.*

		Window around the threshold		
	All	4 months	2 months	1 month
Extended benefit duration benefit	.039	.076**	.020	.017
	(.026)	(.033)	(.045)	(.063)
Obs.	7537	3872	1835	845

All covariates tested in the previous section are included : gender, nationality, age, education, parent, married, residence in Parisian region, qualification, past wage, past employment history, preferred sector, seasonality and time dummy.

The effect on starting wage is also estimated using a linear regression discontinuity model (equivalent to model 1). Real hourly starting wage is normalized by past employment wage. Our outcome of interest is thus the logarithm of the ratio between real hourly starting wage and real past employment wage. Differences highlighted by the descriptive statistics in graph are not confirmed in the regression discontinuity estimation (table 9). The estimated effect is small and not significant, whatever the restriction imposed on past employment duration and whether controls are added or not (see appendix C).

Table 9: *Effect of extending potential benefit duration on wage ratio with covariate controls.*

		Window around the threshold		
	All	4 months	2 months	1 month
Extended benefit duration benefit	.0009	-.009	-.030	.010
	(.024)	(.028)	(.040)	(.056)
Obs.	7537	3872	1835	845

All covariates tested in the previous section are included : gender, nationality, age, education, parent, married, residence in Parisian region, qualification, past wage, past employment history, preferred sector, seasonality and time dummy.

6 Conclusion

In the French case, merging unemployment registers and employment administrative data sources certainly improves observation of unemployment exits to jobs and job seekers' past employment history. Contrary to Card, Chetty, and Weber (2008), spikes at the unemployment benefit exhaustion date are still observed when destinations of unemployment exits are observed independently from unemployment registers.

In a regression discontinuity design inspired by Card, Chetty, and Weber (2008), we find that potential unemployment benefit duration has a significant large impact on unemployment exits to work. When job-seekers are entitled to 15 months of benefits instead of 7 months, only because they cross the 8 months past employment threshold, their exits to jobs are slowed down by 6 points during the first 10 months of unemployment.

There is however no significant evidence of longer employment spells or higher starting wages due to extended benefits duration. Those 2 last results are more fragile, because subject to selection bias into employment.

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A Sample selection.

Our data set is based on the matching of the *Fichier historique* of the French Public Employment Agency (ANPE), which records unemployment spells on a daily basis, and the employment registers or *Déclarations Administratives de Données Sociales* (DADS), which record employment spells for 85% of French workers. It is a $1/24^{th}$ sample of unemployed who registered at the Employment Agency between 1999 and 2004. Spells before 1999 are included, but they are all censored in December 2004.

Table 10: *Effects of sample selection on covariates*

	Short or extended benefit duration *filières* 2 and 3	Long benefit duration *filière* 5	Final sample restricted *filières* 2 and 3
Man	0.46	0.49	0.48
Foreigner	0.09	0.06	0.07
Age	29.58	32.28	28.75
Lower secondary education	0.21	0.14	0.15
Professionnal degree	0.38	0.42	0.37
Upper secondary education	0.19	0.18	0.21
Superior education	0.19	0.24	0.24
Parent	0.34	0.43	0.28
Married	0.33	0.46	0.29
Residence in Parisian region	0.16	0.20	0.17
No qualification	0.31	0.20	0.27
Low qualification	0.47	0.50	0.49
Intermediate profession	0.07	0.10	0.09
Management	0.04	0.08	0.05
Previous real hourly wage	6.42	8.88	7.95
Days unemployed during last 3 years	311.67	106.19	286.40
Attached to the service sector	0.71	0.72	0.72
Previously on permanent contract	0.14	0.40	0.15
Job separation during first quarter	0.24	0.23	0.24
Job separation during second quarter	0.21	0.20	0.20
Job separation during third quarter	0.28	0.31	0.28
Job separation during fourth quarter	0.27	0.26	0.29
Job separation before July 2001	0.49	0.55	0.51
Observations	31945.00	71184.00	16692.00

B Discontinuity tests on population density and covariates

To test for the discontinuity in the population density, we estimate the following model :

$$N_d = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d}) \left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + v \qquad (3)$$

where d is pre-unemployment employment duration (in "weeks"), N_d the population size of recipients with pre-unemployment employment duration d , $\bar{d}$ the threshold. We test whether there is a discontinuity, i.e. δ is equal to 0. We estimate $\hat{\delta} = 34$ with standard error 83. The test is accepted. In this baseline estimation, the last point in graphic 1 corresponding to jobs seekers who have worked one year before job separation has been discarded. The result of the test is robust to its inclusion. It is also robust to controlling for any "entire month" effect and for polynoms of past employment duration with higher degree.

To test for the discontinuity in the covariates distribution around the threshold, we estimate linear regression discontinuity models :

$$Y = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d}) \left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + v \qquad (4)$$

where Y is our covariate of interest and all other notations defined as above. In table 11, the estimate of δ is reported for different populations around the threshold.

Table 11: *Covariates discontinuity test on different windows around the threshold*

		Window around the threshold		
	All	4 months	2 months	1 month
Man	-.074*** (.018)	-.062*** (.022)	-.054* (.032)	-.042 (.044)
Foreigner	.001 (.009)	-.00006 (.011)	-.019 (.016)	-.013 (.022)
Age (log)	.031*** (.009)	.020* (.012)	.009 (.017)	-.0008 (.023)
Lower secondary education	.065*** (.013)	.057*** (.016)	.024 (.022)	.039 (.030)
Professionnal degree	.005 (.018)	.016 (.022)	.055* (.031)	.080* (.043)
Upper secondary education	-.022 (.015)	-.014 (.019)	-.016 (.026)	-.014 (.036)
Superior education	-.042*** (.015)	-.058*** (.019)	-.055** (.027)	-.088** (.037)
Parent	.055*** (.016)	.050** (.020)	.016 (.029)	.021 (.039)
Married	.028* (.016)	.015 (.020)	-.001 (.029)	-.021 (.040)
Residence in Parisian region	-.022* (.013)	-.025 (.016)	-.020 (.023)	.016 (.030)
No qualification	.050*** (.016)	.043** (.020)	.015 (.029)	.033 (.039)
Low qualification	-.032* (.018)	-.004 (.023)	.041 (.032)	.018 (.044)
Intermediate profession	-.024** (.010)	-.027** (.013)	-.017 (.018)	-.026 (.026)
Management	-.012 (.008)	-.015* (.009)	-.016 (.013)	-.016 (.018)
Days unemployed during last 3 years	48.513*** (11.166)	30.094** (13.817)	19.879 (19.748)	12.293 (27.572)
Attached to the service sector	.003 (.016)	.013 (.020)	.011 (.029)	.012 (.040)
Previously on permanent contract	.025** (.012)	.010 (.015)	.031 (.020)	.048* (.025)
Job separation during first quarter	-.015 (.015)	-.033* (.019)	-.020 (.027)	-.098*** (.037)
Job separation during second quarter	-.024 (.015)	-.006 (.018)	-.048* (.026)	.003 (.035)
Job separation during third quarter	.062*** (.016)	.082*** (.020)	.091*** (.028)	.073* (.038)
Job separation during fourth quarter	-.023 (.017)	-.043** (.021)	-.024 (.029)	.022 (.040)
Job separation before July 2001	.067*** (.018)	.054** (.023)	.037 (.032)	-.002 (.044)
Number of observations	16 692	8352	3 837	1817

C Robustness : regression discontinuity without covariates controls

Table 12: *Effect of extending potential benefit duration.*

Exits to work		Window around the threshold		
	All	4 months	2 months	1 month
After 4 months	-.029**	-.010	-.028	-.0003
	(.013)	(.017)	(.024)	(.032)
After 10 months	-.139***	-.102***	-.073**	-.044
	(.017)	(.021)	(.031)	(.042)
After 18 months	-.120***	-.084***	-.056*	-.040
	(.018)	(.022)	(.032)	(.043)
Without restriction	-.078***	-.049**	-.028	.005
	(.018)	(.023)	(.032)	(.044)
Observations	16 692	8352	3 837	1817

Table 13: *Effect of shortening potential benefit duration on unemployement exit rate.*

		Window around the threshold		
	All	4 months	2 months	1 month
Short duration benefit (δ)	1.185***	1.161***	1.226***	1.338***
	(.032)	(.040)	(.062)	(.099)
3-4 weeks before exhaustion week (δ_{-2})	1.596***	1.963***	1.923**	1.663
	(.237)	(.436)	(.609)	(.830)
1-2 weeks before exhaustion week (δ_{-1})	1.502***	1.682***	1.628*	.987
	(.210)	(.334)	(.469)	(.389)
short duration benefit exhaustion week (δ_0)	1.746***	1.868**	1.594	1.583
	(.361)	(.553)	(.734)	(1.125)
1-2 weeks after exhaustion week (δ_1)	1.829***	1.523**	1.577	.600
	(.282)	(.309)	(.473)	(.268)
3-4 weeks after exhaustion week (δ_2)	2.074***	2.307***	1.759*	1.442
	(.328)	(.538)	(.538)	(.648)
Subjects	16692	8352	3837	1817
Log-likelihood	-68152.81	-32256.6	-13883.82	-5741.443

Table 14: *Effect of extending potential benefit duration on employment duration.*

	All	Window around the threshold 4 months	2 months	1 month
Extended benefit duration benefit	.044*	.075**	.026	.017
	(.026)	(.033)	(.046)	(.063)
Obs.	7537	3872	1835	845

Table 15: *Effect of extending potential benefit duration on wage ratio.*

	All	Window around the threshold 4 months	2 months	1 month
Extended benefit duration benefit	-.011	-.012	-.031	.028
	(.023)	(.028)	(.040)	(.056)
Obs.	7537	3872	1835	845