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

Thomas Deroyon Thomas Le Barbanchon †

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Abstract

Generous unemployment benefits tend to slow down exits from unemployment and to increase job quality of new matches. In our study, we question 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. In response to Card, Chetty and Weber (AER, 2008), we find spikes at unemployment benefit exhaustion. This new data set also enables us to construct good measures of past employment. In the French system, whether past employment duration is under or over some thresholds, unemployed may receive benefits over a short or longer period. We exploit those discontinuities in the French eligibility system in a regression discontinuity design. We verify that 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 and on subsequent employment duration, but no significant effect on starting wages.

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 any insurance, unemployment insurance is subject to moral hazard issues: too generous unemployment benefits may discourage unemployed to search for jobs and/or to accept reasonable jobs. Thus, understanding the effect of unemployment insurance generosity on jobs exit rate and the quality of jobs is of prime concern. When unemployment insurance is more generous, is the job search activity less intensive? More productive? Does a more generous unemployment insurance lead to better jobs?

According standard job search theory, a more generous unemployment insurance increases the reservation wage of the unemployed. This induces unemployed to be more selective among job offers. They stay longer unemployed and the distribution of jobs accepted should be of better quality. Above 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 to think 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.

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†CREST (Insee), Ecole Polytechnique

Some other mechanisms may lead to a negative link between unemployment insurance generosity and job quality. If the benefit recipient stay unemployed for too long, its 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, the question remains an empirical one.

According to theory, the effect of unemployment insurance on job quality comes through unemployment duration lengthening. 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 Zweimüller (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 along the unemployment spell. Meyer (1990) finds spikes in the unemployment exit rate just before the exhaustion of unemployment benefits. This is evidence that unemployed react to financial incentives. The expected profile of unemployment benefits conditions the search behavior of 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 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 set 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. The data set we built for the study is robust to that critics[1].

Evidence on the effect of unemployment insurance generosity and employment quality is scarce and contrasted (see the review in Addison and Blackburn (2000)). Addison and Blackburn (2000) 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.

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 the impact through a new regression discontinuity design inspired by Card, Chetty, and Weber (2008). We exploit discontinuities in the eligibility rules of the French unemployment insurance system in 2001-2002. Depending on their past employment experience, unemployed can be just below or above some eligibility thresholds which make them entitled to shorter or longer unemployment benefit duration. More precisely, we will compare unemployed who work around 8 months during the year before their unemployment spells start. Working between 6 and 8 months makes workers entitled to 7 months of unemployment benefits, whereas working between 8 and 12 months opens 15 months of benefits.

This 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 spells reported by employers to the French administration for fiscal purpose. 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 the precise

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

past employment duration that makes unemployed eligible to some unemployment benefits durations, information which is not recorded in unemployment registers.

Our paper starts with a complete description of our data and a discussion of how benefit maximum duration can be predicted using the employment data. In a second part, we motivate our regression discontinuity design. In a 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 tends to increase job stability, but do not have any effect on post-unemployment wage. These last results are less robust as selection bias among reemployed workers is not seriously corrected.

2 Data : observing unemployment exits to work, post-unemployment job quality and pre-unemployment job tenure

Our data set is based on the matching of the *Fichier Historique* (FH) of the French Public Employment Agency (ANPE), which records unemployment spells on a daily basis, and the *Déclarations Administratives de Données Sociales* (DADS), which records employment spells for 85% of the 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.

We will focus on the job seekers who enter the Employment Agency between July 2001 and December 2002 and who are entitled to new unemployment insurance benefits . To avoid identification problems caused by the specific policies dedicated to senior job seekers, we drive out of our sample people of 50 years old and more[2]. Finally our sample contains 49 879 job seekers.

Between July 2001 and December 2002, there has been no change in unemployment insurance rules [3]. During that period, recipients are entitled to benefits for a fixed amount of time, the maximal benefit duration, during which the replacement ratio is constant. Recipients can enter into one of the 4 *filières*, which are categories with a specific benefit maximal duration (for more details see annex A). We will focus on 2 specific categories: *filière 2* and *filière 3*. Jobs seekers in *filière 2* are entitled 7 months of unemployment benefits, they will be referred as short benefit duration job seekers. Those in *filière 3* will be referred as extended benefit duration job seekers; they are entitled to 15 month benefits.

In the following subsections, we explain how this new data set enables us to better observe unemployment exits to work. We give descriptive statistics on post-unemployment job quality. Finally, we show how maximal benefit duration can be predicted by observing pre-unemployment job tenure in the DADS.

2.1 Unemployment exits to work

The information in the unemployment registers about the reasons why job seekers leave the Employment Agency is often missing or not accurate enough, as it comes from the job seekers' monthly reports. If a job seeker does not send his monthly report, he is indeed expelled from the Employment Agency, and loses his entitlement to unemployment insurance benefits (such a job seeker is referred to as *absent au contrôle* (AC)). For instance, 31 % of the insured job seekers of our sample left the Employment Agency because they did not send their reports. Their destination is unknown.

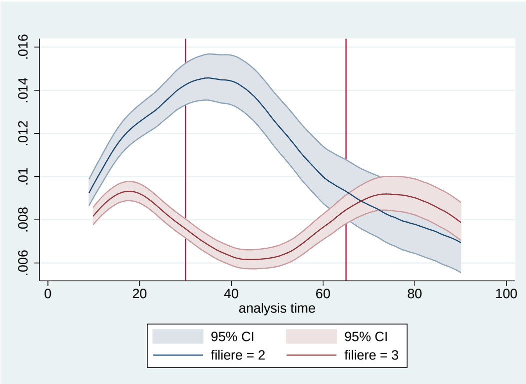
Our data set, matching unemployment and employment spells, enables us to detect and describe the exits to employment, independently of job seekers' reports. Therefore, we define as *unemployment register exits to DADS job* any exits from the Employment Agency with a

[2]We also drop from our main sample job seekers for whom insurance rules are very specific (recurrent temporary workers, artists).

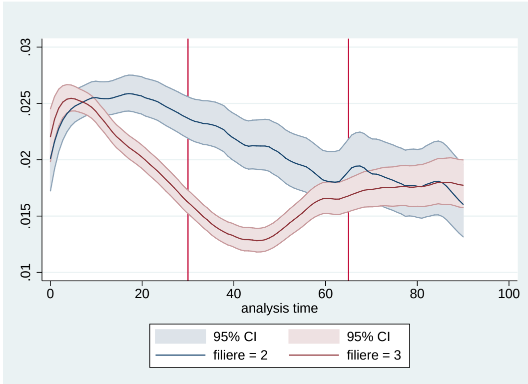
[3]Those changes happen at least every 3 years in France

Table 1: *Weekly unemployment exit rates*

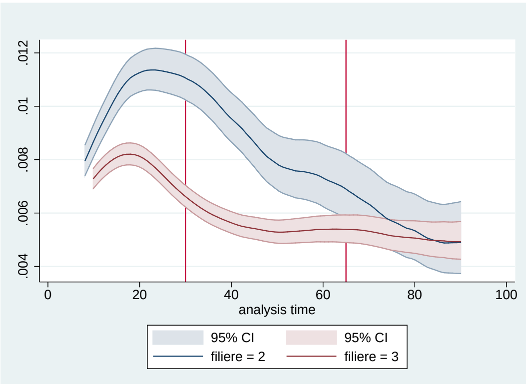
Unemployment register exits
to DADS jobs



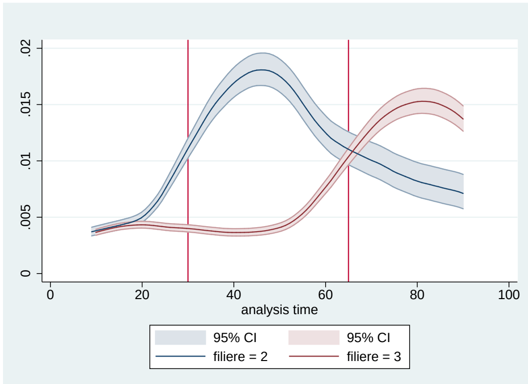
Unemployment exits
to DADS jobs



Unemployment register exits
to FH reported jobs



Unemployment register exits
to FH unknown destination



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

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

Our definition improves the measurement of exits to employment : the part of job seekers who leave the Employment Agency to begin a new job rises from 38 % according to the job seekers' report to 42 % according to the DADS. Yet the exits to work are probably still underestimated : some employment spells are indeed missing in the DADS. Data concerning civil servants and care jobs workers are not collected. Furthermore, some employers do not fill properly their employees identification numbers in the DADS (5 % of them are very likely to be false). This prevents us from observing completely the transitions between employment and unemployment and the lengths spent in each state (see Le Barbanchon and Vicard (2010) for a complete description of each source limitations and the matching process).

Despite this, the information gathered thanks to the FH-DADS sheds a new light on the relationship between unemployment insurance and exits to work. 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 about the existence of spikes at that time. The exit rate to jobs, as declared to the employment agency, indeed rises and declines before the benefit exhaustion. More precisely, the *unemployment register exit to FH reported job* falls by 50 % around the date of exhaustion. In the meantime, the exit rate to FH unknown destination (AC) doubles around the benefit exhaustion (see the lower graphs in panel 1). This certainly points out a change in the reporting behavior of job seekers at benefit exhaustion. Before it, unemployed seem to care more about the reporting of their actual situation, as it conditions their entitlement to remaining benefits[4]. After benefits are exhausted, job seekers have less incentives to report their actual situation, which causes a rise in the exit rates to FH unknown destination.

Nonetheless, the exit rate to DADS job rises before the end of benefit exhaustion, to reach a spike just after it (see the upper left graph in panel 1) : the FH-DADS enables us to detect some exits to work among the spike of exits to FH unknown destination.

We introduce a last measure of transitions to work. Insured job seekers have indeed the right to work in small paid jobs without losing their entire benefits. In this case, job seekers do not leave the Employment Agency. They leave unemployment for they work, but are certainly still searching for a better job. In the FH-DADS, we detect those transitions from unemployment to small paid jobs, which we will then refer to as *unemployment exits to DADS jobs* (as opposed to unemployment register exits). With this new measure, the spike of exits rate at benefit exhaustion disappear, as shown in the upper right graph in panel 1.

2.2 Post-unemployment job quality

The job seekers' reports do not contain any information about the jobs they take. The FH-DADS helps us describe the employment duration of newly employed workers, the wage they get, and the part of their former wages they were able to take back .

In our sample (all *filières* confounded), the job seekers who find a job before December 2004 work 18 months on average in their new firm. The median employment duration is 9 months[5]. The hazard rates out of new jobs show spikes at the usual fixed-term contracts duration : 3, 6, 12 and 24 months (see panel 1). As shown in panel 1, former job seekers with extended benefit durations stay longer in their new jobs than those with short benefit duration : the median of employment duration jumps from 6 to 8 between the two groups.

The average and median daily starting wages are 37.5 euros and 34.5 euros among our full sample[6]. As a comparison point, in 2002, an employee with a full time job paid at the minimum wage could get 36 euros on a daily basis. The left graph of panel 2 shows that starting wages concentrate around the wage of a full time job paid the minimum wage. There is a second mode at the wage of a half-time job paid the minimum wage.

On the right part of panel 2, we compare the distributions of daily starting wages among former short and extended benefit durations job seekers. Former extended durations unemployed seem to have higher starting wages. The difference is however small : the difference in medians is indeed 1 euro.

We are not only interested in the level of starting wages, but on wage loss due to unemployment. So we compute the ratio of starting wage over pre-unemployment wage. We take as the latter the wage computed by the unemployment insurance administration to pay benefits[7].

Half of job seekers lose more than 8 % of their former wages when they get a new job. The wage loss is however lower for the extended benefit durations job seekers (panel 3) : whereas half of workers from short benefit duration category lose more than 11 % of their former wages, the median wage loss rate is less than 6 % among extended durations job seekers.

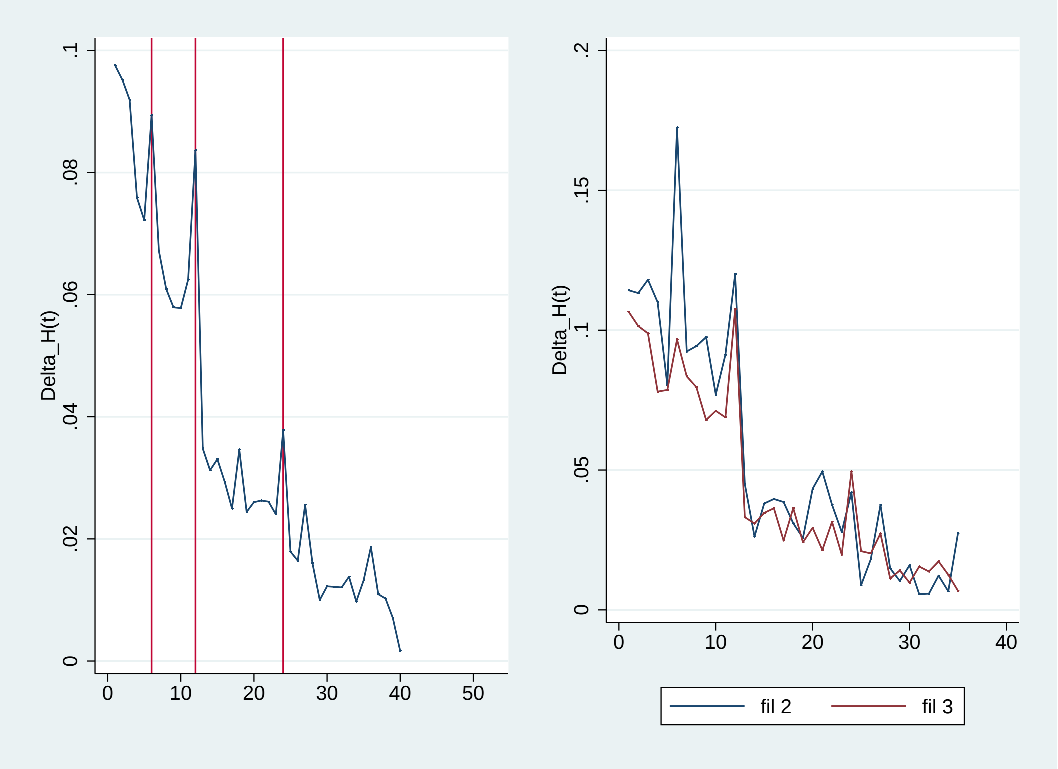
[4]This may seem not so important for those who find a job. But, as soon as they expect to come back later and get remaining benefits, recipients have to report correctly their situation.

[5]Note that 22 % of new jobs spells are censored at the end of the data set (January 2005).

[6]Extreme values (over 100 euros) of the real wage distribution have been dropped.

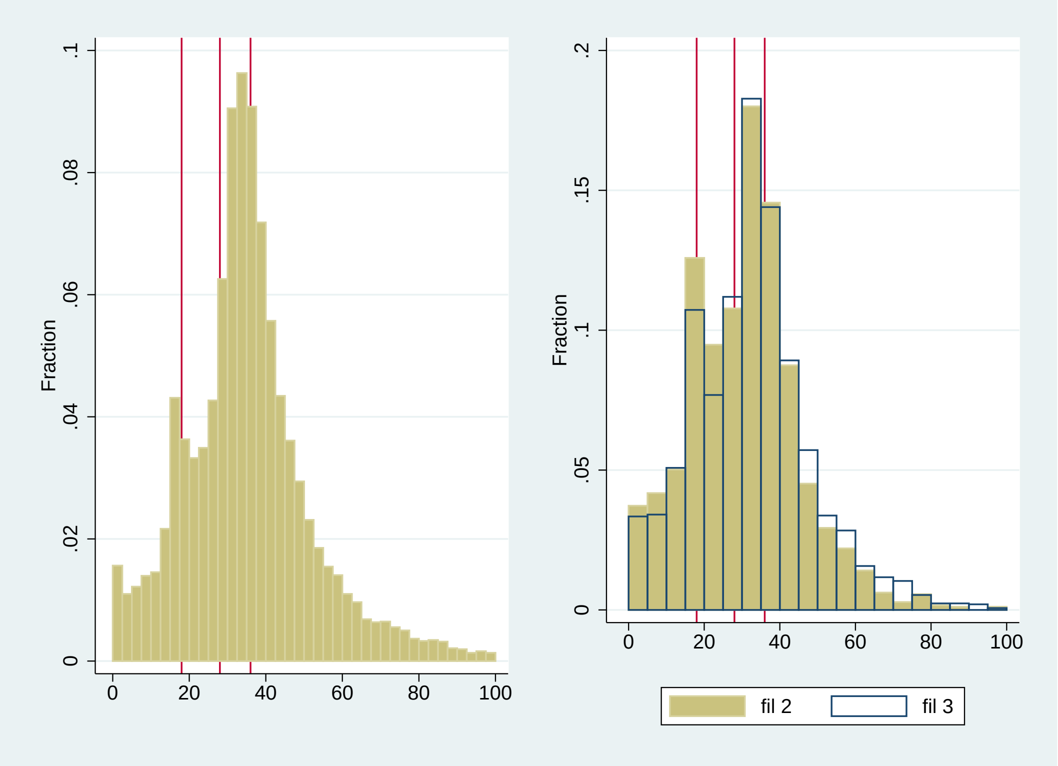
[7]It is the average daily wage the unemployed got in the year before he lost his job

Figure 1: *Monthly job separation rate*



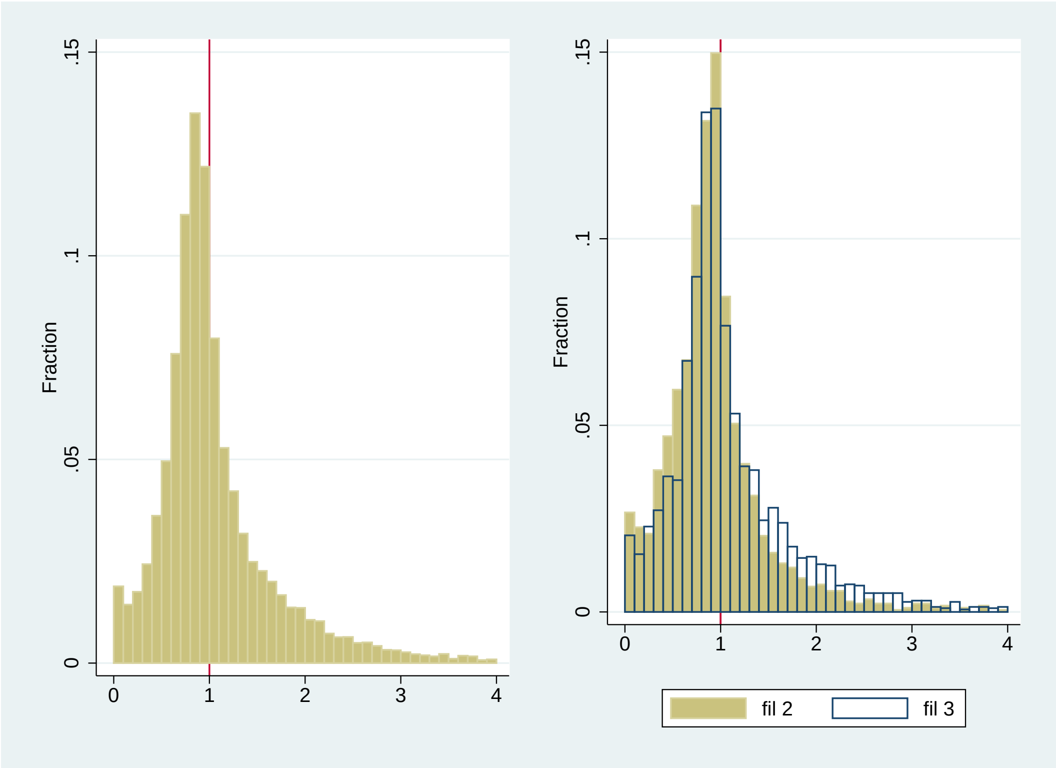
Reading: on the left, job separation is computed for all insured job-seekers. On the right, recipients with benefits duration of 7 months and 15 months are compared. Vertical lines represent typical short term contracts durations (6 months, 1 year, 2 years).

Figure 2: *Density of post unemployment daily wages in euros (base 2000)*



Reading: the histogram on the left is computed for all benefits duration. On the right recipients with benefits duration of 7 months and 15 months are compared. Vertical lines represent the daily earnings of a minimum wage earner in 2002 when he works half time and when he works full time.

Figure 3: *Wage loss density (post / pre unemployment daily wage in euros based 2000)*



Reading: the histogram on the left is computed for all benefits durations. On the right recipients with benefits duration of 7 months and 15 months are compared.

The differences between short and extended benefit durations job seekers cannot be interpreted as direct causal effect of differences in benefit lengths : extended benefit durations job seekers have a longer work experience that enabled them to get entitled to longer benefits. Those structural differences can cause endogeneity bias. To reduce this bias, we select:

- among the short benefit duration unemployed, those who could have been entitled to extended benefit durations had they work only a few weeks more;
- among the extended benefit durations unemployed, those who would have got shorter benefits had they worked a little less.

To implement this regression discontinuity design, we need to know precisely the job spells and their lengths the unemployed used to open their unemployment insurance benefits. This information, missing in the Employment Agency data set (FH), can be computed from the DADS.

2.3 Pre-unemployment job tenure

The way we compute employment duration before job loss is described in appendix A. From this computation, we apply the unemployment insurance eligibility rules and predict the length of benefits a job seeker would be entitled to. We then compare this prediction with the actual maximal benefit duration he gets at the Employment Agency.

Predicted and actual maximal benefit durations are the same for 60 % of job seekers in our sample (table 2). Prediction errors are mainly underestimation of pre-unemployment job tenure : for instance, 24 % of the job seekers would not be entitled to any insurance benefit according to the DADS (radical underestimation). The prediction also overestimates the actual benefit lengths for 7% of the job seekers in our sample.

Misclassification depends on the actual maximal benefit duration : radical underestimation is more important among job seekers with shorter benefit durations : about 30 % for unemployed with benefit durations of 4, 7 or 15 months against 19 % for people with 30 months of benefit duration. Besides, overestimation naturally falls with the actual maximal benefit duration. The part of non radical underestimation (unemployed who are predicted to be entitled to unemployment insurance but whose predicted benefit durations are smaller than the one they actually get) also automatically rises with the length of actual benefits : it jumps from 0 % to 8 % between the shortest and the longest actual benefit durations, whereas the part of overestimation falls from 27 % to 0 %.

Table 2: *Comparison between predicted and actual benefit durations*

Actual maximal benefit duration	Predicted benefit duration					Part in entries
	4 months	7 months	15 months	30 months	None	
4 months	42	9	4	14	31	10
7 months	13	33	11	12	32	9
15 months	6	7	34	21	33	16
30 months	2	2	4	73	19	65
Total	8	6	9	53	24	100

The table 10 in appendix B show the job seekers characteristics according to the comparison between their predicted and actual maximal benefit durations. We thus compare under, over and well classified job seekers. The results clearly indicate that individual characteristics are linked with prediction quality. Well classified job seekers have indeed a stronger relationship to work : they are more often qualified men, with high level of education, higher former wages and longer past tenure lengths, they have been lees registered as unemployed in the past three years. This is no surprise as stable jobs are better reported in the DADS. We also verified that unemployed looking for a job in agriculture or in care jobs are more likely to be misclassified by the DADS. Their former employers were indeed probably not covered by the data set.

In the next parts of this paper, we compare unemployed entitled to benefit durations of 7 and 15 months and control for their estimated past tenure lengths. This is only possible for job seekers whose estimated tenure lengths seem correct, that is whose predicted and actual benefit durations match. That is the reason why we drop from our sample all misclassified job seekers. This entails a problem of external validity for our results : job seekers remaining in our sample have different characteristics than those we drop (see table 10). Only the proportions of low qualified unemployed and job seekers with age between 25 and 34 years old are not significantly different between the two groups. There is therefore a doubt about the extent to which our results can apply to the rest of the population.

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 independent 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 difficult to test. We first explain its credibility in our case. Then we describe how close the covariates of recipients just below and just above the threshold.

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 it 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, post-unemployment employment duration and wages.

We find unlikely that employment duration is manipulated when employers and employees separate. In our sample, the typical employment duration is around 8 months (the forcing variable threshold). Manipulation at separation would mean that fixed term contract are extended, a not very common practice.

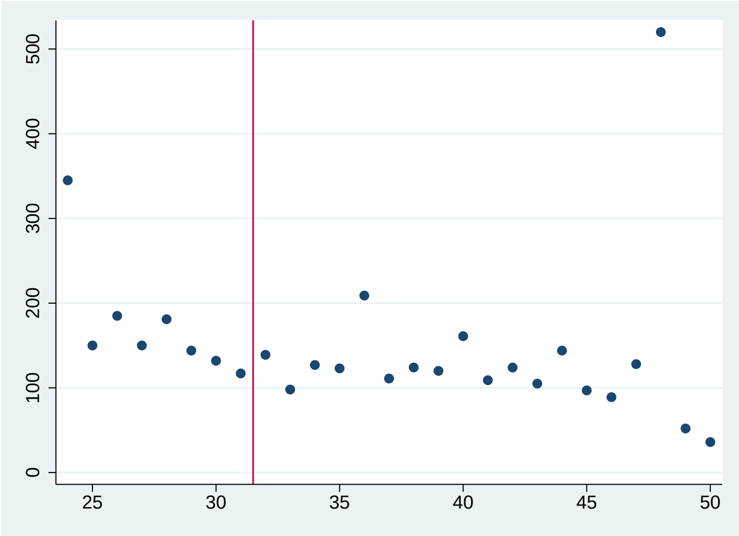
We find unlikely that employment duration is manipulated, before separation, when the match meets. The employment prospects of our sample are structurally small. They are less educated, less qualified than the typical French worker. This limits the ability for a worker of our sample to find a contract that exactly extends his past employment duration in order to meet the eligibility criteria to extended duration. The employment prospects of our sample are all the worse than they enter unemployment during an economic slowdown (2001 and 2002).

Finally, 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. Anyway, fraud at benefit registration is difficult. At benefit registration, job seekers bring employment certificates delivered by their past employers. Collusion between employers and employees are all the less likely that sanctions are severe. To limit fraud bias to our estimate, we select recipients registering in the general unemployment insurance system, as workers under special status (recurrent temporary workers, artists...) are more prone to errors in certificates (see recent reports from the French *Cour des comptes*).

Forcing variable manipulation can be checked by inspecting the population density around the threshold. If employment duration were precisely manipulated, recipients would accumulate just above 8 months (32 "weeks"[8]). We do not see any discontinuity on the graph 4. This is confirmed by a formal discontinuity test (see the annex C). The test has been augmented to account for the fixed term job duration periodicity. Most of the contracts are written in months.

[8]In the DADS, 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 4: *Benefit recipients density along the previous employment duration in weeks (forcing variable)*



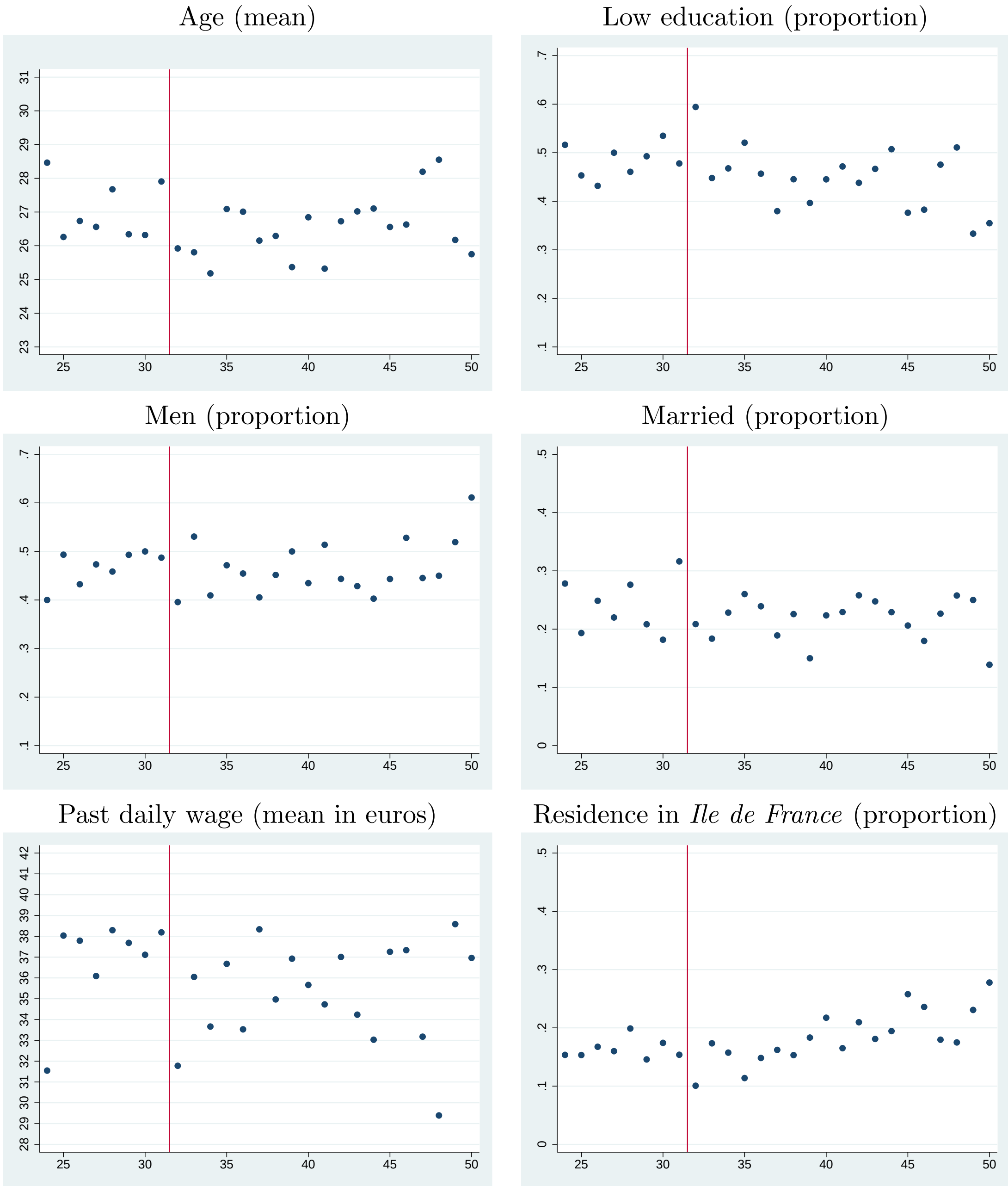
Reading: the vertical line represents the threshold between fil 2 and fil 3. Past employment duration starts from 6 months (24 ”weeks”), this is the minimum employment duration to enter *filière 2*. It goes to 50 ”weeks”, which is over one year in the DADS time scale. This can be the case because workers can cumulate under hours accounting several days for one calendar day worked.

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. To test for discontinuity we run several linear regression discontinuity on different window around the threshold. Results are reported in table 11 in the appendix C. We find a highly significant discontinuity in the pre-unemployment daily wage and moderate discontinuity in age when the test is conducted 2 months around the threshold. For some other covariates, the formal tests show moderate discontinuity. However, for those covariates, those results are not robust when the width of the window around the threshold varies. All in all, age and past wage are the only discontinuous covariates, they represent only 2 of the 11 covariates tested.

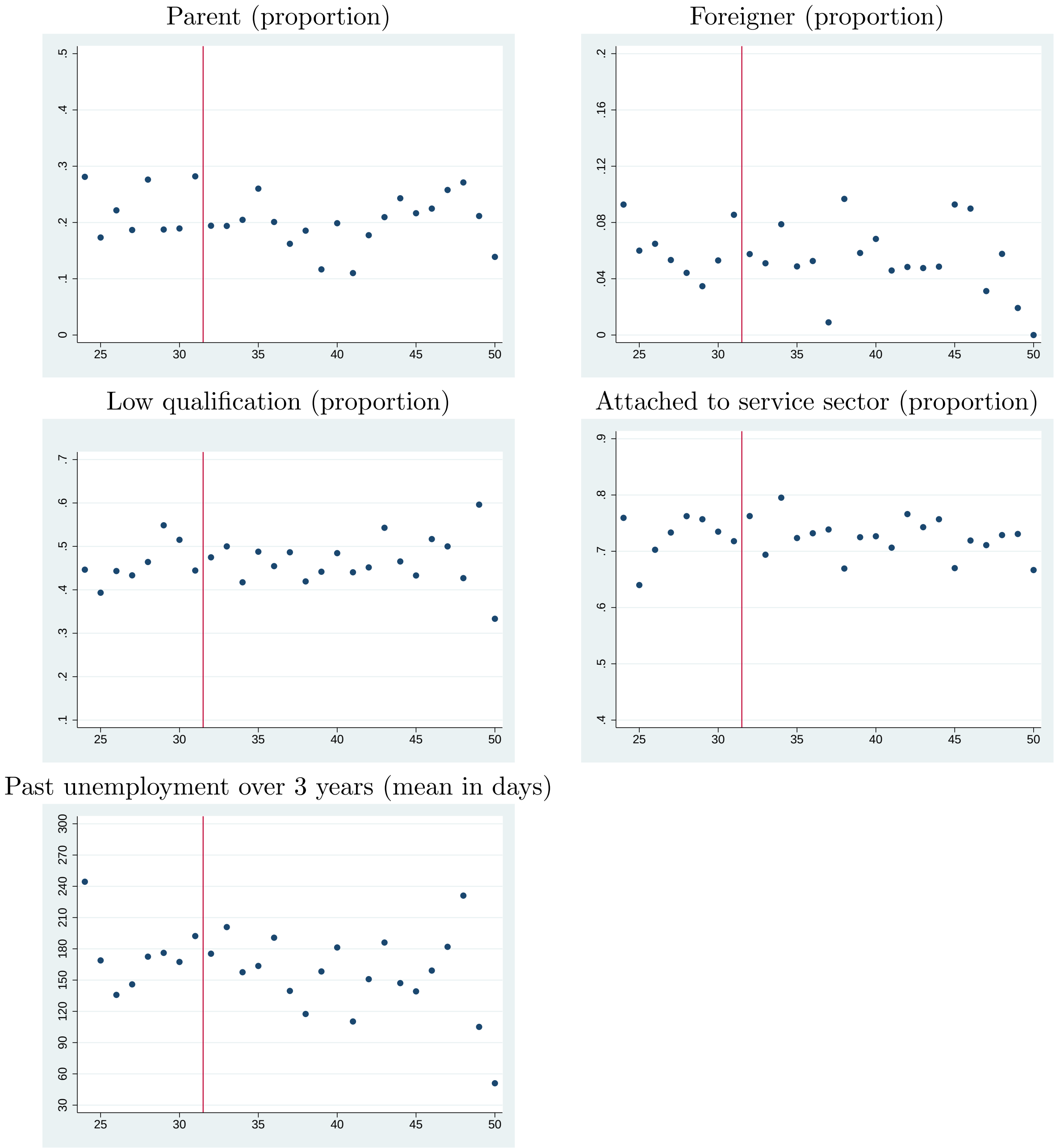
In table 12 in the appendix C, we report a less demanding test, equality of means above and below the threshold for different window. Those comparisons highlight the importance of restricting our estimation around the threshold. Without restriction, extended benefit duration recipients are more qualified and have higher education. They are younger and have lower previous wage. They have spent less time unemployed over the 3 years before their unemployment registration. When the comparison is restricted around the threshold, differences in qualification, education and past unemployment history disappear. However, benefit recipients are still different in age and past wage. This is in line with previous discontinuity tests.

Table 3: *Covariates distribution and past employment duration (I)*



Reading: the vertical line represents the threshold between fil 2 and fil 3.

Table 4: *Covariates distribution and past employment duration (II)*

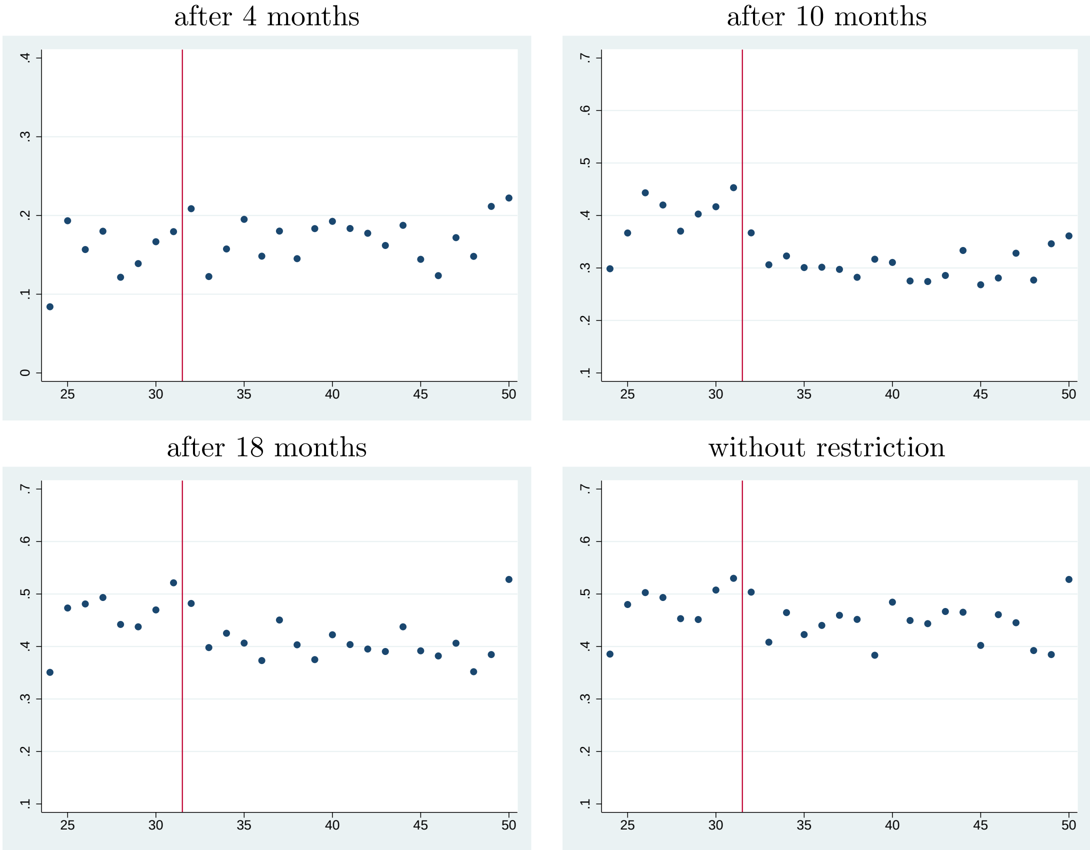


Reading: the vertical line represents the threshold between fil 2 and fil 3.

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 during the first 10 months after registration. In a second step, we model unemployment exits to DADS job *à la* Card, Chetty, and Weber (2008). This model takes into account censoring and spikes at benefit exhaustion.

Table 5: *Unemployment register exits to DADS jobs for different pre-unemployment job tenure in weeks (forcing variable)*



Reading: the vertical line represents the threshold between fil 2 and fil 3.

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 key in the recipient history. At 4 months, unemployed in both categories receive benefits. At 10 months, only unemployed in the extended benefit category receive benefits. At 18 months, all benefits have expired. The standard 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 finds a job during the first m months after registration, d is the past employment duration computed to predict benefit duration, $\bar{d}$ is the eligibility threshold to extended benefits (8 months), X is a set of covariates (gender, nationality, education and age). Our parameter of interest is δ, which is reported in table 6. The effect of duration benefits is estimated, as previously, for different job finding measures:

- unemployment register exits to work reported by employers (DADS jobs)
- unemployment register exits to work recorded at the Employment Agency (FH jobs)
- unemployment exits to DADS jobs (whatever the situation in the Employment Agency registers)

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 less job finding (for all measures). After 10 months the share of unemployed who have deregistered to start a DADS job is 16 points lower in the extended benefit duration category. This is a decrease of 30 % on unemployment exits share.

In the other columns of table 6, we report estimation results on narrower windows. Those restrictions make population above and below the threshold more and more similar. Results are then more robust to misspecification errors (covariates, linear dependance 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).

When the window is narrowed, there is still an effect of extended benefit duration, especially on the measures involving DADS jobs 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 DADS job is 12 points lower in the long benefit duration category. This effect is significant at the 5% level. When the window is 4 months wide, short benefit duration recipients find DADS jobs faster, whatever their unemployment registration behaviors (9 points significant at the 5% level).

The effect of benefit duration on exits to jobs declared to the Employment Agency is less strong than that on other job finding measures. This weakness can be explained by a modification of declaration behavior around the benefits exhaustion (as explained in paragraph 2.1).

4.2 Spikes at benefit exhaustion?

The previous linear regression model does not explain precisely when the effect of benefit duration takes place. 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 unemployment exit rate to job 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:

$$\theta_t = h_t \exp\left(I(d < \bar{d}) \left(\delta + \sum_{k=-3..2} \delta_k I(t \in I_k) \right) + \gamma X \right) \qquad (2)$$

where all notations have already been defined except I_k which represents time after exhaustion. I_0 marks if the observation is in the exhaustion week, I_1 if it is in the 2 following weeks after exhaustion week, I_{-1} in the 2 previous weeks before exhaustion week... One parameter of interest is again δ, but, contrary to the previous estimation, it captures the effect of shortening unemployment benefits duration. All the other δs (δ_0,δ_1...) capture the local effects of shortening benefit duration, the effects around benefit exhaustion.

Table 6: *Effect of extending potential benefit duration.*

	All	Window around the threshold		
		4 months	2 months	1 month
	During the first 4 months			
Unemployment registers exits to DADS job	-.012 (.025)	-.003 (.031)	-.024 (.047)	.047 (.068)
Unemployment registers exits to FH declared job	.0005 (.023)	-.002 (.029)	-.038 (.043)	-.057 (.063)
Unemployment exits to DADS job	.012 (.033)	-.004 (.040)	-.044 (.058)	.022 (.082)
	During the first 10 months			
Unemployment registers exits to DADS job	-.159 (.034)***	-.118 (.041)***	-.124 (.059)**	-.105 (.084)
Unemployment registers exits to FH declared job	-.098 (.032)***	-.063 (.039)	-.084 (.058)	-.134 (.083)
Unemployment exits to DADS job	-.063 (.033)*	-.089 (.040)**	-.077 (.056)	.036 (.079)
	During the first 18 months			
Unemployment registers exits to DADS job	-.107 (.035)***	-.071 (.042)*	-.073 (.061)	-.077 (.085)
Unemployment registers exits to FH declared job	-.084 (.034)**	-.061 (.041)	-.099 (.060)*	-.123 (.085)
Unemployment exits to DADS job	-.061 (.029)**	-.089 (.035)**	-.068 (.049)	.003 (.069)
Observations	4115	2455	1061	486

Standard errors are robust to White heteroscedasticity. Covariates are: gender, nationality, education (lower secondary, BEP-CAP, upper secondary, superior) and age (less than 25 years old, between 25 and 34, between 35 and 49).

Table 7: *Estimation of unemployment exit rate Cox model for job-seekers in a 4-month window around the threshold*

	Unemployment register exits to DADS jobs	Unemployment register exits to declared jobs	Unemployment exits to DADS jobs
Short benefit duration	1.14 (.075)**	1.11 (.081)	1.10 (.052)**
3-4 weeks before exhaustion week	2.12 (.87)*	3.85 (1.89)***	1.01 (.261)
1-2 weeks before exhaustion week	1.88 (.678)*	1.03 (.368)	1.224 (.352)
exhaustion week	5.83 (3.59)***	4.79 (.368)**	1.282 (.514)
1-2 weeks after exhaustion week	1.67 (0.59)	2.34 (1.04)*	1.21 (.360)
3-4 weeks after exhaustion week	1.78 (.686)	2.21 (0.99)*	1.46 (.471)
log-likelihood	-7809.55	-6391.1	-14308.45

Covariates are: gender, nationality, education (lower secondary, BEP-CAP, upper secondary, superior) and age (less than 25 years old, between 25 and 34, between 35 and 49).

Estimation results in table 7 confirm the existence of spikes at exhaustion for unemployment register exits to job. However, for unemployment exits independent of registering behavior, the effect of shortening benefits is not concentrated at exhaustion time; it is spread all over the job search. The effect on unemployment registers exit is very strong during the exhaustion week: the hazard rate is then 5 to 6 times bigger. Exits to declared jobs is also accelerated during the month before exhaustion, whereas faster exits to DADS job are otherwise spread uniformly.

5 Effect of potential benefit duration on job quality

Job quality is by definition observed for job seekers who find a job. From now on, our population of interest is restricted to unemployed who exit unemployment registers to DADS jobs. We compare employment durations and starting wages of job seekers leaving *filière* 2 and 3. This comparison may suffer from a selection bias 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 long benefit duration unemployment spells results in comparing individuals characteristics rather than measuring causal impact of benefit length. Anyway, some evidence show that this bias may not be so dramatic: the fraction of job seekers who finds a job during the maximum observation time after registration is the same among both *filière* (see bottom right corner graphics in panel 5).

5.1 Effects on employment duration

Numerous employment observations are censored at the ending date of our data (22%). So the effect of benefit duration on employment duration is best estimated with a Cox model.

$$\theta_t = h_t \exp\left(\delta I(d >= \bar{d}) + (d - \bar{d})\left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + \gamma X\right) \tag{3}$$

where all notations are as before except θ_t which is now the monthly job separation rate. The parameter δ, as in the standard regression discontinuity model, captures the magnitude of the discontinuity at the threshold. It measures the effect of extended benefit duration.

Table 8: *Cox estimation on employment duration.*

	All	Window around the threshold		
		4 months	2 months	1 month
Effect of extended benefit duration (δ)	.79 (.091)**	.75 (.107)**	.74 (.149)	.79 (.216)
Observations	1665	994	447	213

Covariates are: gender, nationality, education (lower secondary, BEP-CAP, upper secondary, superior) and age (less than 25 years old, between 25 and 34, between 35 and 49).

Job separation rate is 20% to 25% lower when workers are former extended benefit duration recipients. This effect is significant when there is no window restriction or when compared populations have past employment duration in a 4 month window around the threshold (2 months below or 2 months above the threshold). For narrower windows, the effect is not significant any more. This is certainly due to small sample sizes. Workers from extended benefit duration stay 26% to 35% longer in the first firm they enter after unemployment.

5.2 Effects on starting wage

The effect on starting wage is estimated using a standard linear regression discontinuity model[9]. Real starting wage is normalized by past employment wage. Our outcome of interest is thus the logarithm of the ratio between real daily starting wage and real past employment wage. Results are in table 9. In the regression discontinuity context, differences in wages ratio highlighted in graph 3 disappear. This is very surprising when no window restriction is imposed; it is probably due to some model misspecification (linear dependency of distance to the threshold). The effect is positive and quite large when the window is restricted to 4 months, but not significant at conventional level. When the window is narrower, the estimation is very imprecise with standard errors over 10 points. With such imprecision, effects should be 20 points to be detectable at 5% level. Such effects are highly unlikely.

Table 9: *Wage ratio regression discontinuity.*

	All	Window around the threshold		
		4 months	2 months	1 month
Effect of extended benefit duration	.037 (.078)	.115 (.064)	.034 (.123)	.12 (.197)
Observations	1665	994	447	213

Standard errors are robust to White heteroscedasticity. Covariates are: gender, nationality, education (lower secondary, BEP-CAP, upper secondary, superior) and age (less than 25 years old, between 25 and 34, between 35 and 49).

[9]The model writes, with all notations as before,

$$Y = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d}) \left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + \gamma X + u \tag{4}$$

6 Conclusion

In the French case, merging unemployment registers and employment administrative data source enables to improve observation of unemployment exits to jobs, and to a lesser extent job seekers past employment history.

Contrary to Card, Chetty, and Weber (2008), spikes at the unemployment benefit exhaustion date are observed independently of job-seekers declaration behaviors. However there is less evidence of spikes when exits to jobs are measured independently of job-seekers registration behaviors.

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 12 points during the first 10 months of unemployment.

There is also evidence of a causal impact of unemployment benefits on employment duration. Extending benefits duration leads to 25% longer jobs. There is no significant evidence of wage gains due to extended benefits duration. Those 2 last evidences are more fragile, because subject to selection bias into employment.

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A Data selection, Unemployment insurance rules and DADS *filière* classification.

We select from the data unemployed of less than 50 years old entering the Employment Agency between July 2001 and December 2002. Besides, we only keep in our sample job seekers who did not keep any residual rights to unemployment insurance benefits from a former unemployment spell. Therefore, the benefit durations job seekers in our sample get are directly linked with their employment lengths before they enter Employment Agency. Finally, we drop from our sample some kinds of job seekers whose unemployment insurance rules differ from those applied to common workers : temporary workers, artists and technicians working in culture especially.

Unemployment insurance rules in 2001 and 2002 for job seekers of less than 50 years old define four different types of benefit durations, called *filières*: *filières* 1, 2, 3 and 5. Each *filière* is defined by a potential benefit duration and a condition on the cumulated length of employment spells in a certain period of time, called *période de référence* beginning from the date of job firing that caused the entry into Employment Agency. Job seekers eligible to *filière 1* must have worked at least 4 months in the past 18 months, and are entitled to a maximum length of 4 months of benefits. Job seekers eligible to *filière 2* must have worked at least 6 months in the past 12 months and get 7 months of benefits, those eligible to *filière 3* must have worked 8 months in the past 12 months and get 15 months of benefits, and those eligible to *filière 5* must have worked 14 months in the past 24 months, and get a maximum benefit length of 30 months. A job seeker naturally enters the *filière*whose benefit duration is the highest among those he is eligible to.

For each entry into Employment Agency, we select in the DADS the job firing date closest to the date of entry and which probably caused it. We calculate for each *filière* the beginning of the *période de référence* and collect all employment spells with some days in common with it. For employment spells not included in the *période de référence*, but whose intersection with it is not empty, we only collect the period of time they have in common. We then cumulate the length of those truncated employment spells : the estimated job tenure for each *filière*. Job seekers are classified in the best *filière* they are estimated to be eligible to and their job tenure is then the cumulated employment spells length calculated for the *filière* they are classified in.

B Comparison of under, well and over classified job seekers' characteristics

Table 10: *Comparison of under, well and over classified job seekers*

	under classified	well classified	over classified
Men	42^{***}	52^{***}	49^{***}
No qualification	33^{***}	20^{***}	29^{***}
Low qualification	47^{***}	48^{***}	45^{***}
Intermediate profession	6^{***}	11^{***}	8^{***}
Management	4^{***}	10^{***}	6^{***}
Lower secondary	23^{***}	10^{***}	13^{***}
BEP-CAP	38^{***}	36^{***}	33^{***}
Upper secondary	18^{***}	20^{***}	23^{***}
Superior	18^{***}	30^{***}	26^{***}
Age under 25 years old	34^{***}	35^{***}	56^{***}
Age between 25 and 34	33^{***}	35^{***}	27^{***}
Age between 35 and 49	33^{***}	30^{***}	16^{***}
Foreigner	12^{***}	5^{***}	7^{***}
Past daily wage under 20 euros	31^{***}	12^{***}	14^{***}
– between 20 and 35 euros	24^{**}	20^{**}	23^{***}
– between 35 and 45 euros	26^{***}	28^{***}	38^{***}
– more than 45 euros	19^{***}	39^{***}	25^{***}
Past unemployment over last 3 years			
No	43^{***}	57^{***}	43^{***}
Less than one year	38^{***}	36^{***}	43^{***}
Between 1 and 2 year	13^{***}	5^{***}	10^{***}
More than 2 years	6^{***}	2^{***}	5^{***}

Note : values marked by three stars (resp. two, one) are significantly different at the 1% level (resp. 5%, 10%).

C Discontinuity tests on population density and covariates

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

$$N_d = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d}) \left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + \beta I_{debut} + v \tag{5}$$

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 and I_{debut} indicates that d corresponds to entire months. We test whether there is a discontinuity, whether δ is equal to 0. We estimate $\hat{\delta} = 38.7$ with standard error 70.9. The test is accepted. Note that the entire month dummy is highly significant ($\hat{\beta} = 103.4$ with standard error 33.3).

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

$$Y = \alpha + \delta I(d >= \bar{d}) + (d - \bar{d}) \left(\delta_{-1} I(d < \bar{d}) + \delta_1 I(d >= \bar{d})\right) + v \tag{6}$$

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*

	All	Window around the threshold		
	(	4 months	2 months	1 month
Man	-.085	.054	-.089	-.100
	(.035)**	(.039)	(.061)	(.085)
Foreigner	.020	.054	-.014	-.016
	(.017)	(.039)	(.030)	(.043)
Age (log)	-.023	.054	-.055	-.078
	(.018)	(.039)	(.032)*	(.045)*
Lower secondary	.040	.054	.028	.102
	(.025)	(.039)	(.046)	(.065)
CAP-BEP	-.028	.054	.004	.087
	(.033)	(.039)	(.058)	(.078)
Upper secondary	.0003	.054	-.041	-.095
	(.030)	(.039)	(.051)	(.073)
Superior	.0002	.054	.041	-.049
	(.031)	(.039)	(.053)	(.075)
Parent	-.029	.054	-.054	-.142
	(.030)	(.039)	(.052)	(.074)*
Married	-.018	.054	-.055	-.128
	(.030)	(.039)	(.053)	(.075)*
Residence in IdF	-.035	.054	-.019	-.058
	(.026)	(.039)	(.042)	(.056)
No qualification	.024	.054	-.005	-.051
	(.032)	(.039)	(.056)	(.079)
Low qualification	-.025	.054	.005	.093
	(.035)	(.039)	(.061)	(.085)
Intermediate profession	-.025	.054	-.032	-.058
	(.022)	(.039)	(.038)	(.054)
Management	-.005	.054	.024	.007
	(.015)	(.039)	(.024)	(.033)
Previous wage (log)	-.196	.054	-.199	-.228
	(.030)***	(.039)	(.055)***	(.079)***
Days unemployed during last 3 years	19.529	.054	6.554	-9.089
	(17.793)	(.039)	(32.107)	(45.510)
Attached to service sector	.007	.054	.035	.013
	(.031)	(.039)	(.054)	(.075)
Obs.	4115	2455	1061	486

Table 12: *Covariates comparison: fil 2 vs fil 3 on different windows around the threshold*

	All		Window around the threshold					
			4 months		2 months		1 months	
	fil 2	fil 3	fil 2	fil 3	fil 2	fil 3	fil 2	fil 3
Men	46	46	46	45	48	44	49	44
No qualification	30^{*}	27^{*}	30	29	28^{*}	33^{*}	30	30
Low qualification	46	46	46	46	49	45	48	49
Intermediate	10	11	10	11	10	8	11	8
Manager	5	6	5	5	5	4	4	3
Lower secondary	15	14	15	14	13^{*}	17^{*}	16	17
CAP BEP	32	31	32	31	34	34	34	34
Upper secondary	24	21	24	24	24	23	24	23
Superior	26^{***}	31^{***}	26	28	26	23	24	24
Age less than 25	52	54	52^{***}	58^{***}	54^{*}	59^{*}	53^{***}	65
Age 25 to 34	28	28	28	26	28	25	29^{**}	20^{**}
Age 35 to 49	20^{**}	17^{**}	20^{***}	16^{***}	18	16	18	15
Foreigner	6	6	6	6	5	6	7	5
Past daily wage under 20 euros	11^{***}	21^{***}	11^{***}	18^{***}	5^{***}	19^{***}	6^{***}	18^{***}
— between 20 and 35 euros	28	26	28	26	28	28	26	28
— between 26 and 45 euros	40^{***}	32^{***}	40^{***}	35^{***}	42^{**}	36^{**}	45	39
— greater than 45 euros	20	20	20	21	25^{***}	17^{***}	24^{***}	15^{***}
Past unemployment over last 3 years								
No	34^{**}	37^{**}	34	34	34	33	29	30
Less than 1 year	47^{**}	43^{**}	47	47	48	46	52	50
Between 1 and 2 years	12^{***}	15^{***}	12	13	12	14	11	13
Between 2 and 3 years	8^{***}	5^{***}	8^{*}	6^{*}	6	6	7	7
Observations	4115		2455		1061		486	

Note : values marked by three stars (resp. two, one) are significantly different at the 1% level (resp. 5%, 10%).