Partial Unemployment Insurance and Hour Decisions

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Abstract

How do financial incentives embedded in unemployment programs affect job uptake? Partial Unemployment Insurance (PUI) programs allow jobseekers to keep their benefits when working, if the job abides by eligibility conditions. PUI programs operate as an in-work benefit scheme that aims towards labor market reintegration: allowing benefits as a top-up on earnings increases the value of employment relative to unemployment's. We exploit a reform of the French PUI scheme introduced in 2006: the benefit eligibility hour threshold was decreased by 20%, offering a quasi-experimental setting. This paper studies labor supply responses when benefit availability is restricted. Using unique administrative data on unemployment spells and employment episodes, we estimate competing risks models with correlated risks to determine the propensity to exit towards PUI job intensities depending on whether they allow for benefits. We show the reform significantly increased the conditional probability to take up a PUI job below the new hour threshold. Hence, narrowed benefit availability contributes to a substantial decline in worked hours for PUI claimants. We use our parameter estimates to compute the hour elasticity to PUI earnings, which equals .142. Its small magnitude stems from labor market rigidities; yet, it is consistent with the bunching literature.

JEL codes: C41, D04, J64, J68

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1 Introduction

The disincentive effect of unemployment insurance (UI) generosity on employment has long been studied in the economic literature (see Tatsiramos and van Ours (2014), for a modern survey). A less studied aspect is the fact that policies may encourage returns to work by increasing the value of employment, namely with financial incentives such as complementary benefits. Partial Unemployment Insurance (PUI) schemes have been implemented for a while in many European countries and North America. Under specific eligibility conditions, these schemes enable UI claimants to keep a share of their benefits while working in typically part-time, temporary jobs. The primary aim is to foster reintegration into the labor market or, at least temporarily, to maintain some attachment to the labor market in order to prevent the decay of jobseekers' social and professional skills. While PUI jobs can thus be seen as stepping stone towards more stable employment, they also increase incentives to work and provide a substantial top-up on earnings, acting in this way as an alternative form of in-work benefit.

These schemes have expanded gradually in many countries. While 12% of UI claimants in OECD countries work while on claim, the proportion of PUI claimants can be as high as 33% in Sweden (Kyyrä, 2010) or 20% in some US states (McCall, 1996). In France, the proportion of people on PUI schemes out of all UI claimants has increased from 15% in the mid-1990s to 28% more recently, amounting to about 600,000 persons. Despite this relative success, the economic literature on PUI remains rather limited. Most of the research has focused on professional outcomes stemming from PUI schemes, and essentially addresses the question of whether PUI improves the probability to find a permanent job (Caliendo et al., 2012, Fremigacci and Terracol, 2013, Godoy et al., 2014, Gurgand, 2002, Kyyrä, 2010, Kyyrä et al., 2013).[1] To our knowledge, only two studies have investigated behavioral responses to the nature itself of PUI schemes, *i.e.* the propensity to take PUI jobs depending of the benefit availability conditions contained in this program. These studies exclusively focus on behavioral responses in relation to the PUI earnings eligibility condition in the US: this condition makes that for every dollar earned above the earnings disregard, current benefits are reduced at a 100% withdrawal rate. McCall (1996) examines responses at the extensive margin by studying an increase in the earnings threshold. He finds a positive effect of the reform on claimants' likelihood to enter both part-time and overall employment. Le Barbanchon (2016) considers movements at the intensive margin by studying the potential bunching around the kink created by the earnings threshold. Using state variation in threshold levels, he recovers behavioral elasticities and calibrates a dynamic model in order to assess the rate of PUI withdrawal that

[1] This strand of literature points to an overall positive effect of PUI on job opportunities. Yet, jobseekers are first locked in the PUI activities before their stepping stone effect operates.

would be welfare-improving.

In this paper, we exploit a reform of the French PUI scheme to study responses at the intensive margin. It relates to an eligibility condition on hours worked enforced in the French system. In France, PUI benefits decrease with hours worked but are also fully withdrawn if earnings are larger than a given threshold (creating a notch rather than a kink) and if worked hours are above a given level (136 hours per month before 2006). The latter condition was reformed in 2006 mainly for budgetary reasons but also to somehow align both earnings and hour eligibility conditions for a person around the minimum wage. Subsequently, the eligibility hour threshold has been decreased by 20% and went down to 110 hours per month, offering a quasi-experimental setting to study labor supply responses.

Using administrative data (*Fichier Historique* and *Fichier National des Allocataires*) for years surrounding the reform, we evaluate its causal effect with competing risks models with correlated risks. This unique dataset matches monthly information on unemployment spells, benefit/compensation characteristics and the record of PUI employment episodes. Considering potential workers for whom the earnings condition binds above the old and new hour thresholds, we test whether UI claimants are incentivized to change the hour levels at which they would take up a PUI activity. Identification relies on a recent approach developed by Van den Berg et al. (2014), which combines regression discontinuity and duration analysis. To identify the treatment effect, we use a rich dataset containing information on PUI spells around the reform implementation; we then exploit variation at the reform implementation date.

We find that the reform significantly increased the conditional probability of entering the PUI scheme below the new hour threshold, and decreased the probability to take up a PUI above the new threshold. This suggests that UI claimants are less induced to join PUI activities in the hour interval where benefits are fully taxed away (the new hour notch). We also find the reform contributes to a substantial decline in total working time among PUI claimants. Results are stable to alternative specifications of duration dependence and to additional controls for local unemployment rate and remaining time before benefit exhaustion. We translate our estimates of the relative change in conditional probabilities of entering each PUI-type in hour equivalents. The ratio of PUI intensity variation over the change in implicit taxation provides an estimate of the hour elasticity to net-of-tax wages which stands around .142.

This paper contributes to the literature by adding novel evidence on behavioral responses to the implicit taxation of PUI income. To our knowledge, it is the first paper of this kind for Europe and, more generally, one of the few attempts to study non-marginal adjustments to reforms of the PUI eligibility conditions. In this sense, it is closely related to McCall (1996) and complementary to the literature on bunching at kinks or notches. Studying a policy change and non-marginal responses actually seems appropriate to the French context with a relatively rigid labor market. It consistently leads to estimated elasticities that are larger than the intensive margin response found in the US using bunching at kinks.[2] Moreover, our elasticity is driven by responses from differentiated entry rates into PUI across types of contracts, hence less subject to frictions related to hour constraints. Nevertheless, our elasticity is consistent with bunching-based estimates under a no-friction scenario (see Gelber et al. (2016a)).

The rest of this paper is organized as follows. Section 2 explains the institutional background, the reform and provides theoretical insight. Section 3 presents the empirical strategy, data and the selection process. Results are discussed in section 4 while section 5 concludes.

2 Partial Unemployment Insurance System

We start with a simple description of the French PUI system. Denote Y the monthly level of gross earnings from PUI, Y^r the monthly reference income (the pre-unemployment earnings) and B the maximum amount of monthly unemployment benefits. Then, monthly disposable income from PUI, i.e. from cumulating earnings and some benefits, is computed as follows:

$$R = Y + B\left(1 - \frac{Y}{Y^r}\right), \tag{1}$$

i.e., benefits are withdrawn at rate Y/Y^r. This withdrawal rate gets larger with PUI job earnings. Aside from the progressive benefit taxation, two conditions lead to a complete withdrawal of PUI benefits. Firstly, gross PUI earnings should not exceed 70% of pre-unemployment earnings Y^r. Secondly, hours worked under PUI, denoted H hereafter, should be lower than a threshold set at 136 hours per month before 2006. The reform under study consists in a 20% decrease of this threshold down to 110 monthly hours for

[2] Changes in kinks or notches have been successfully used to estimate labor supply elasticities, for instance kinks generated by tax credits in the US Saez (2010), Chetty and Saez (2013) or notches in tax systems Kleven and Mazhar (2013)), cf. Kleven (2016) for a recent survey. Closest to us, Le Barbanchon (2016) finds a substantial degree of bunching at the kink created by the earnings condition of the PUI on a sample of former wage-earners (i.e. an eligibility condition for claiming in the US). In contrast, the condition on earnings in the French PUI system generates very little bunching, as shown by Gonthier and LeBarbanchon (2016).

any new PUI jobs started after January 1, 2006. Conditions for benefit eligibility under PUI temporary contracts are summarized as:

$$B = 0 \text{ if : } \begin{cases} Y \geq 0.7Y^r \\ \begin{cases} H \geq 136 \text{ before } 2006 \\ H \geq 110 \text{ after } 2006 \end{cases} \end{cases} \tag{2}$$

If at least one of the two conditions is binding, benefits are completely withdrawn and unpaid benefits are postponed to the next uncompensated unemployment episode. For ongoing PUI spells that started before the reform, UI claimants can continue to cumulate earnings and benefits under the old rules until the contract reaches its term.

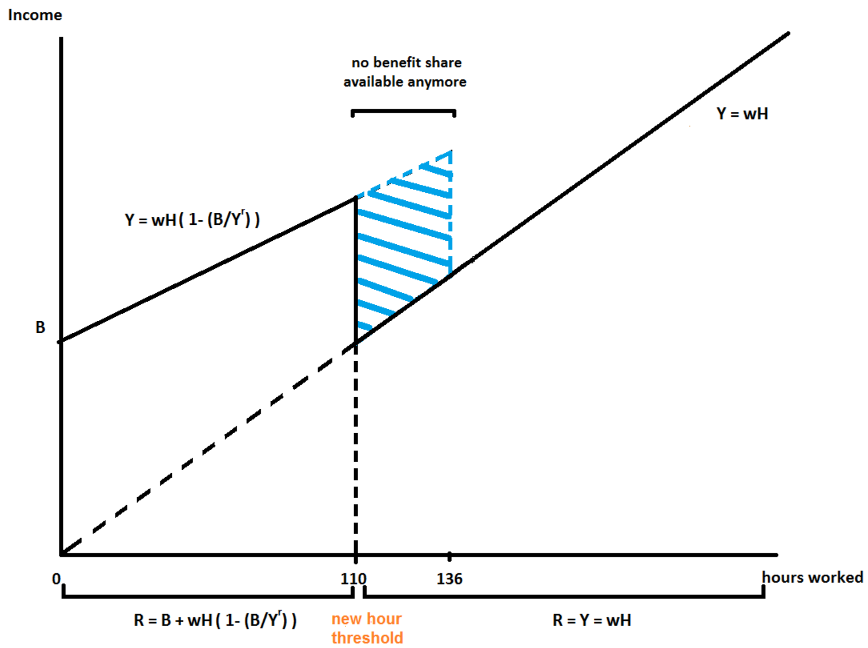


Figure 1: Budget Constraints Before and After the Reform

To illustrate the system and the reform, denote w the wage rate, so that current earnings are $Y = wH$, and rewrite equation 2 as:

$$R = B + wH\left(1 - \frac{B}{Y^r}\right). \tag{3}$$

This relationship between disposable income and worked hours is depicted in Figure 1. Budget constraints are shown before and after the reform. The slope remains unchanged after the reform and corresponds to the wage net of the replacement rate B/Y^r. The graph presents the notch generated by the hour constraint. It shows how this notch moved to the left with the hour threshold shift from 136 to 110 hours per month. Thus, compared to claimants entering PUI in 2005, their 2006 counterparts lost a surplus encompassed by the blue area. In other words, the reform reduced opportunities to complement earnings with benefits for contracts between 110 and 136 hours per month.

Note that even if the region just to the right of the 136-hour threshold was strictly dominated before the reform, a certain density of temporary workers are nonetheless observed there. Indeed, being above this threshold guarantees full-time activities that may be similar to what UI claimants are actually seeking as permanent jobs (136 hours, *i.e.* 32 hours per week, correspond to some types of full-time contracts in France while the main legal work duration is 35 hours per week). Similarly, the reform creates a new notch in a region that coincides with partial activities (from 25 to 32 hours per week) that may still attract some UI claimants despite the fact that cumulating benefits is no longer possible. This is corroborated by the hour distribution presented on Figure 2 (data and selection leading to this graph are described in the next section). We observe a substantial mass both above 136 before the reform and above 110 after the reform.

Moreover, we observe declarative data and do not seize empirically any substantial bunching behavior. Gonthier and LeBarbanchon (2016) test the bunching behavior of the unemployed towards the PUI earnings constraint: they find no bunching when considering the declared earnings, but find bunching when considering PUI earnings that are proved to the Unemployment Agency. Hence, there is no clear bunching in working decisions (because of the labor market rigidities among other aspects) but the optimization behavior is based on self-reported justified work experiences (Kleven et al., 2011). We only have declarative data: we can assume that picking precise amounts of hours worked remains rather limited, so the real choice relies on picking a PUI contract that allows for benefits or not. Then, maybe individuals bunch at the notch when reporting proofs of work episodes to the Unemployment Agency, but so far no data containing that information has been made available.

This conveys that studying bunching at the hour notch(es) is not the most adapted framework to our study, all the more so as the French labor market is relatively rigid (the unemployed cannot precisely choose the amount of hours worked in PUI jobs, cf. Gonthier and LeBarbanchon (2016)). Thus, analyzing behavioral responses at the intensive margin around the hour threshold(s) would not provide sensible results. More appealing is the reform itself, which may induce UI claimants to take up PUI jobs at lower worked hours after January 1, 2006. Indeed, Figure 2 shows a move to the left with the reform, i.e. the frequency of part-time jobs (notably the peak at 80 hours) increases relative to full-time activities in the 110-136 monthly hours range.[3] The rest of the paper will attempt to precisely measure this effect using transitions throughout the reform in a competing risk model with unobserved heterogeneity.[4]

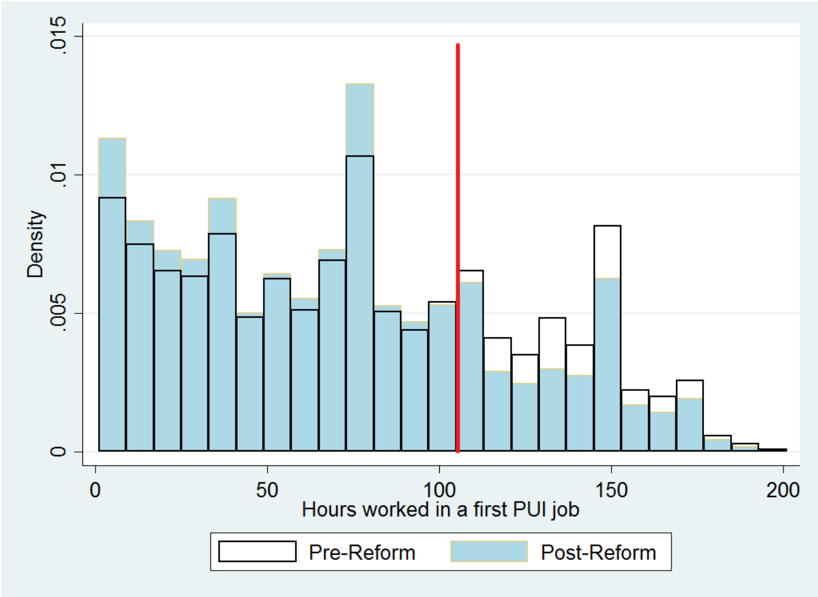


Figure 2: Distribution of Hours Worked in PUI Before and After the Reform

Finally remark that the reform was motivated by at least two considerations. First, narrowing benefit accumulation possibilities was aimed at reducing the UI deficit, as explained in the Collective Agreement of December 2005 (CFDT et al. (2005)). It is also likely that the reform was designed to align the two eligibility conditions together.[5] Hence, this does not constitute a policy initially targeted at the unemployed.

Moreover, an overview of the French labor market situation provides insight on employment incentives over the period 2004-2006. Table 4 (in Appendix) shows the economy remained relatively stable with a moderate but steady and positive GDP growth. The unemployment rate had a similar path with a non-substantial decrease over the period

[3]Note that the data selected for this graph corresponds to individuals for whom the hour constraint binds before the earnings constraint, both before and after the reform, so that the hour responses are not affected here by any restriction due to the earnings constraint. Note also that the decline in hours concerns contracts that are also above 136 hours (and not only in the 110-136 hours range). It is possible that the reform has contributed to better inform about (non)cumulating possibilities, i.e. a change in frictions due to information as investigated in Gelber et al. (2016b).

[4]Note that we expect two effects of the reform: an overall decrease in the propensity to join PUI jobs and a substitution towards PUI jobs at low working hours. The former, overall change in unemployment exits towards PUI activities remains unindentified since many other factors (macroeconomic conditions, other policies) may affect the probability to take up a PUI job.

[5]Let us consider the following UI claimant's profile: a person working at the official full-time work duration (35 hours per week, i.e. 151.6 hours per month) before unemployment and who obtain the same hourly wage in a PUI job as she had before unemployment. This profile can reasonably be thought as the reference one used by policy makers. Thus, the earnings constraint is reached at the same level as the new hour constraint ($\frac{wH}{w^r H^r} = \frac{H}{H^r} = \frac{110}{151,6} \approx 0.7$).

considered. The constant employment rate is paired with a small increase of part-time jobs: this indicates the 2006 reform was implemented in a context of stable labor outcomes and opportunities. There is no evidence of large movements of jobseekers towards stable employment after the reform. The stability of the French labor market also argues in favor of a stable selection process over the reform time span.

3 Empirical Identification Strategy

We evaluate causal effects of the treatment (i.e., exposure to the 2006 reform) on duration before one takes up a PUI activity. Using a competing risks model we test, before and after the reform, which event happens first between two alternative PUI intensities, i.e. below or above 110h/month. We expect jobseekers to be relatively less likely to resort to highly intensive PUI activities when benefits cannot be cumulated anymore. In addition, we expect an increased recourse to PUI activities that allow for additional benefits. The evaluation of these effects is addressed in a dynamic treatment evaluation framework.

3.1 Dynamic Treatment Evaluation

We follow the main notations and general framework from Abbring and Van den Berg (2003) and Van den Berg et al. (2014) with a time-continuous approach. Let τ^* define the implementation date of the reform. The dynamically assigned treatment can be defined by a sequence of mutually exclusive treatments characterized by the index $s \in \{0, \infty\}$ where s represents the time elapsed in unemployment before τ^*. Moreover, $s = \infty$ stands for the non-treatment situation. To each treatment s corresponds the random variable $T(s) \geq 0$, that is the potential duration outcome if affected to treatment in s. To isolate treatment effects, we can contrast by simple difference of hazard functions, the distribution of potential duration outcomes $T(\infty)$ and $T(s')$ when the reform occurs before s'.

To seize the relative propensity to join each PUI intensity, we decompose potential duration outcomes in two subcases: below and above 110 hours/month. Observed durations around this threshold brings information on the relative proportion to take up each PUI intensity with respect to the benefit availability rules. $T_1(s)$ and $T_2(s)$ are defined as the elapsed durations before starting a PUI, respectively below or above 110 hours/month linked to treatment s. The absolute distribution of duration cannot be observed: instead, we observe the minimum of both durations attached to a type of PUI job, $(T(s), J)$:

$$T(s) = min(T_1(s), T_2(s)) \qquad \begin{cases} J = 1 \text{ if } T(s) = T_1(s) \\ J = 2 \text{ if } T(s) = T_2(s) \end{cases} \tag{4}$$

The policy effect's identification strategy borrows from the regression discontinuity approach, the difference being that we estimate hazards instead of densities (as in Van den Berg et al. (2014)). The causal effect of the reform is identified with only two assumptions of non-anticipation and exogeneity of the treatment assignment.

The first assumption (henceforth A.1) defines as exogeneous the policy change conditional on unobserved and observed individual characteristics. The treatment assignment S is assumed to be random conditional on unobserved and observed variables: $S \perp \{T(s)\}|(X, \nu)$. This assumption is consistent with this reform which was exogeneously enforced for all current and future claimants in 2006. The treatment assignment is also independent from unobserved factors given observables: $S \perp \nu | X$. This implies that the treatment date is independent from the outcome conditionally on X and ν. The second assumption (A.2) states that the reform is unanticipated by UI claimants. Unemployed individuals do not have private information on the reform implementation date, or do not act on such information if they do. Given that this change in benefit rule is technical-oriented, it is extremely likely that this reform has been confidential for UI claimants who had not yet taken up a PUI activity. Non-anticipation is implied by the fact that the policy and its enrollment rules are only known to the unemployed from its implementation date onwards. [6]

Under these two assumptions, randomization should be verified at the beginning of each spell, *i.e.* $\nu \perp S | X$. This means that the date at which one is treated is independent from the outcome of the treatment conditionally on X and ν. However, the dynamic selection of survivors at each period could impair the independence thereafter, and the conditional distribution of unobserved characteristics ν, could be deformed. Van den Berg et al. (2014) establish the (non-parametric) identification proof of the instantaneous treatment effect of each treatment s by using spells before, after and ongoing spells throughout the reform.

The basic insight is that the policy change is an exogenous time-varying binary explanatory variable whose discontinuity point varies independently across spells that started before τ^* and lasted up to τ^*. When the policy reform takes place for a specific cohort at an elapsed duration t_0, we compare in t_0 survivors from this cohort and survivors from

[6] We check if the announcement of the reform has modified the behavior of claimants at any period prior exposure to the reform as a robustness check in Section 5.2.

earlier cohorts, who were treated at higher elapsed durations. Assuming that the reform is unanticipated by UI claimants (A.2), the impact of the future treatment does not matter on the before-treatment duration (dynamic identification). All jobseekers belong to the same group, so they have the same selection structure towards PUI over time (A.1).

Let us consider a cohort crossing the reform date at elapsed duration t_0, and another cohort, that entered unemployment previously: hence, the sub-population of survivors from both cohorts has the same composition at duration t_0. It is crucial that the sub-populations come from populations that are identical to each other when they enter into the state of interest. Hence, a cross-cohort comparison of outcomes conditional on survival up to t_0 identifies the average causal effect and is not contaminated by selection effects.

Figure 3 illustrates the identification process. This figure presents two particular spells among all ongoing spells when reform is implemented (τ^*). In cohort 0, individuals experience changes in benefit rules at the elapsed duration $S = t_0$. Cohort 1 individuals entered their unemployment spell earlier so they are not subject to treatment at the same elapsed duration. The dynamic selection process is the same for both cohorts over time during the first t_0 periods spent in unemployment. Therefore, the distribution of unobserved characteristics is the same on the two sub-populations at the elapsed duration t_0. Thus, any change in hazard rates at t_0 between the two cohorts can be attributed to reform effects.

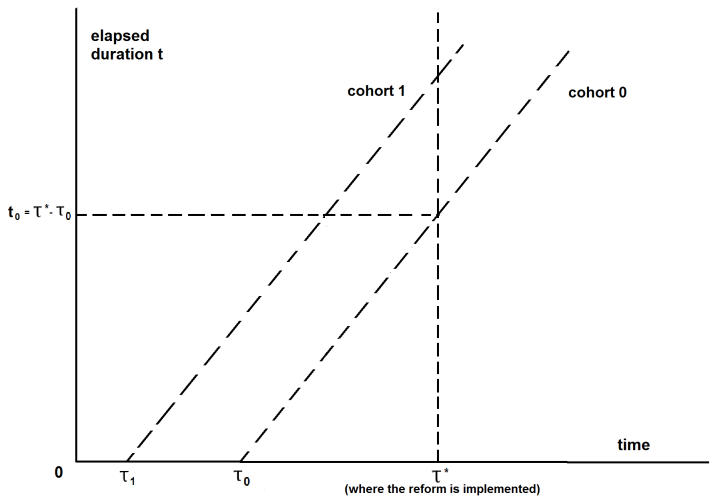


Figure 3: Illustrating Identification with Two Cohorts

This methodology provides a counterfactual for cohorts post-reform: we can compare individuals exiting for PUI jobs after the same elapsed duration in unemployment. We obtain both the treated and the untreated (pre-reform) counterpart cleared from time effects related to the selection process towards PUI jobs for the group in general. This methodology borrows from the regression discontinuity approach: the evolution of all controlled variables is smooth around the reform period; the unobserved heterogeneity evolution is specified parametrically. Hence, the only registered jump in hazard rates should contain behavioral responses to the reform. Thus, this methodology isolates treatment effects with a before-after comparison, while taking care of time-related selection issues. The estimated treatment effect of changes induced by the reform is an average treatment effect on survivors (ATS) at a given elapsed duration. That is, ATS(t_0|X) measures the difference in participation rates between treated and untreated survivors at the elapsed duration t_0. The ATS is identified by the distribution of unobserved heterogeneity. The selection process until elapsed duration t_0 in unemployment is identified by the Mixed Proportional Hazard (MPH) specification presented below; the unobserved heterogeneity distribution is also identified with a Heckman-Singer specification. Indeed, the selection process remains rather stable over the period studied. Following the regression discontinuity approach, the smoothness in calendar time effects around the policy implementation date τ^* implies that the only registered change in durations after the reform stem from policy effects.

3.2 Exit Hazards

In this paper, we retain a competing risks framework. Our model allows duration variables to be dependent by way of unobserved determinants, with each single risk having its own Mixed Proportional Hazard (MPH) model specification. The MPH parametric structure identifies the competing risks model as in Abbring and Van den Berg (2003), provided that there is enough variation within the regressors and enough exits towards PUI jobs: our dataset validates the two requirements above[7].

Hazard equations specified by the MPH model are defined as follows:

$$\begin{cases} \theta_1(t|X, S = s, \nu_1) = \lambda_{01}(t)exp(X\beta_1 + I\{t \geq s\}\alpha_1)\nu_1 \\ \theta_2(t|X, S = s, \nu_2) = \lambda_{02}(t)exp(X\beta_2 + I\{t \geq s\})\alpha_2)\nu_2 \end{cases} \tag{5}$$

where λ_0 is the baseline hazard function, specified in a flexible way using a piecewise constant exponential function. Empirically chosen intervals (in months) are $[0; 3],]3; 6],]6; 24]$. This interval partition is consistent with jobseekers taking up PUI activities in the first 6 months of their unemployment spell (cf. Figure 5, Appendix A.2)[8] . X denotes the vector of explanatory variables, which essentially consists in pre-treatment individual characteristics. We represent the treatment variable with a time-varying in-

[7]Abbring and Van den Berg (2003) identifies this competing risks model for cases of multiple spells - present in our data.

[8]We shall relax this restriction by allowing the treatment effect to be time-dependent.

dex, assuming a unique treatment effect that is independent from the elapsed duration in unemployment when treated. $I\{t \geq s\}$ is an indicator function taking the value 1 if the condition – the instant t is post-reform s – is verified, and 0 otherwise. Consequently, α is the treatment effect. Finally, (ν_1, ν_2) represents unobserved individual heterogeneity for both risks.

The two PUI activity types materialize different job intensities but both constitute PUI jobs, which is where the correlation lies: unobserved heterogeneity characteristics (e.g. motivation, or ability) influence the risk to take up a PUI job regardless of its intensity level. We account for the correlation between risks, thanks to an unobserved heterogeneity specification *à la* Heckman-Singer (Heckman and Singer, 1984). The distribution of unobserved heterogeneity follows a univariate distribution defined by two mass points taking values (ν_1, ν_2). We estimate p the probability of belonging to be linked to a given profile of unobserved heterogeneity.

In this study, we isolate treatment effects that do not encompass earnings adjustment frictions or inertia (*e. g.* Gelber et al. (2016b), Kleven and Mazhar (2013)). Indeed, we solely consider first entries in PUI jobs before and after the reform, so there are no adjustment costs to the reform.

3.3 Data and Selection

Administrative Data Estimations are conducted on a 5% sample of all UI claimants in France. This sample is drawn from two administrative files made available by the French Unemployment Agency, namely the *Fichier Historique* (FH) and the *Fichier National des Allocataires* (FNA). The merged data is unique in the sense that it combines the unemployment history of individuals in the FH data with the benefit claims (and compensation path) of the unemployed in the FNA, as well as the required information regarding pre-unemployment (such as the past reference wage). Both datasets are matched on common information regarding the episodes of PUI activity. The data contains the monthly record of ongoing unemployment spells and includes new monthly entrants, providing us with both a calendar of compensated periods and a calendar of PUI episodes for each unemployment spell. It allows for multiple spells per individual.

Selection and Censorship Our initial dataset records monthly unemployment spells for workers who are entitled to receive benefits at the start of their unemployment spell (non-claimants are not affected by the reform and, hence, of no interest). From there, we select jobseekers who are in the "general regime", since other particular unemployment regimes are not concerned by the reform. We also exclude those unemployed for over two years. The data structure corroborates the treatment identification, since it contains information on monthly cohorts of new entrants over years 2005 and 2006: data provides enough information for valid counterfactuals before and after the reform.

Up to this stage, our sample is a very broad representation of UI claimants in France. A necessary additional selection step consists in keeping only those for whom the earnings eligibility constraint does not bind before the hour constraint. The 2006 reform would be neutral if the earnings constraint binds at monthly working hours below 110h, and possibly also if it is binding between 110 and 136 hours. In those cases, benefits would be withdrawn due to the earnings constraint below 110h/month anyway, hence we cannot expect any behavioral response motivated by a change in the hour threshold. A conservative selection thus consists in discarding all the claimants whose earnings contraint binds at monthly hours below 136h [9]. We keep claimants for whom:

$$\begin{aligned} & 0.7Y^r = wH \text{ at } H \geq 136 \\ \Leftrightarrow\ & 0.7Y^r \geq w \times 136 \end{aligned} \tag{6}$$

Since in the vast majority of cases, PUI jobs are paid at the minimum wage (i.e. $w \geq w^{\min} \approx 8$ euros per hour), we keep claimants for whom $Y_r \geq w^{\min} \times 136/0.7 \approx 1554$ euros per month. This selection step reduces our sample and, potentially, the external validity of the analysis. Yet, it has the merit to rely on a simple rule based on the pre-unemployment earnings.

Observations are right-censored. For the sake of our duration model with two competing risks, we make the following choice: censorship takes place if an individual leaves the unemployment spell without starting a PUI employment episode. In this category, we pool individuals who have never exited unemployment for a PUI job, whatever the reason (e.g. leaving UI to take a stable job, losing track of the unemployment spell when the jobseeker fails to send back to form, no matter the reason). Note that the question of censoring of exits towards regular employment may be asked. Precisely, the reform may increase jobseekers' interest for regular job search because the range of PUI activities does no longer provide complementary benefits. We argue that this issue does not undermine our identification strategy. Indeed, we focus on the relative attraction of PUI activities at

[9]The earnings threshold binds before the hour threshold if jobseekers' (pre-unemployment) reference wage is sufficiently low compared to PUI employment wage. Precisely, if $Y \geq 0.7Y^r$ for $H \leq 110$, the earnings constraint binds below the hour constraint both before and after the reform, which is expected not to alter incentives. If $Y \geq 0.7Y^r$ for $H \in \{110; 136\}$, the earnings constraint binds first before 2006 (with the 136 hours threshold); after 2006, jobseekers have to pick an hour contract below 110 hours to be able to cumulate part of their benefits, and may be affected by the reform. If $Y \geq 0.7Y^r$ for $H \geq 136$, the hour constraint always binds before the earnings constraint. These claimants will constitute the baseline selection for our analysis.

different hours, not on the absolute incentive to exit towards PUI relative to other types of exit. Moreover, the move of the notch to the left makes that PUI below 110h also dominates B (pure job seach) for those who would have chosen 136h in case of possible cumulation (cf. Figure 1).

Our raw sample is composed of 601,508 total observations over 2005-2006 while the selected sample comprises 59,211 monthly unemployment spells over the period. Hence, our selection still manages to keep 34% of the initial sample of UI claimants. Tables in Appendix A.1 show the evolution of entry rates in PUI employment below and above 110 hours around the reform, both for the raw data on UI claimants (Table 4) and for our selection (Table 5). In both cases, we observe an overall increase in PUI employment. It is likely to be context-related since temporary job offers broadly soared over that period (Magnier and al. (2008)). Most interestingly, the upward trend is mainly driven by PUI contracts under 110 hours, underlining a difference in the extent to which individuals resort to PUI jobs with respect to benefit availability. These average changes reflect the moves in hour density already observed in Figure 2. Descriptive statistics are reported in Table 6.

4 Main Results

4.1 Competing Risks Estimates

Estimates of the competing risks model with dependent risks are presented in Table 1. In the baseline model (model 1), we control for duration dependence, and for unobserved heterogeneity among the two risks. Duration dependence coefficients indicate that those who tend to be longer-term unemployed also exit for PUI employment more rapidly. This may simply stem from a mechanical lock-in effect due to time spent in PUI activities, and even more in PUI jobs below 110 hours. The policy reform, *i.e.* treatment effect indicates both a significant increase in the probability to enter PUI below 110 hours, and a decrease in the probability to enter in PUI above 110 hours.

In the next specification (model 2), we additionally control for socio-demographic characteristics (gender, marital status) and the replacement rate. These variables seem to have a relatively balanced effect over the two risks, yet jobseekers with higher replacement rates tend to take up PUI jobs above 110 hours. The larger the replacement rate, the larger the benefit share taken away in PUI (Gurgand, 2002). These jobseekers are then less likely to take up any PUI activity, to moderate the expected loss. We also control for regional quarterly unemployment rates (model 3). When local unemployment increases, UI claimants have less side opportunities, which motivates them to pick up PUI jobs earlier. The PUI scheme thus operates as a means to adjust to time and context-dependent difficulties. Finally, we add the months left with entitlement to benefits (model 4). Indeed, a large strand of the economic literature addresses jobseekers' propensity to go back to employment when their benefits are soon to be expired (Meyer, 1990). The take up of a PUI job seems to accelerate when the benefits are close to exhaution, yet this effect is not identified given the way duration dependence varies in this last model. In any case, the policy effect remains extremely stable among various specifications. In particular, adding local unemployment levels only slightly increases the relative probability of exiting under 110 hours.

Estimates in Table 9 (in Appendix) present additional results whereby the duration dependence specification is a little more refined. Precisely, we now use a piece-wise constant exponential form with the following month intervals: $[0;3],]3;6],]6;9],]9;12],]12;18]$ and $]18;24]$. The results are similar to the previous set of estimates and, once again, very stable across model specifications.

	Model (1)		Model (2)		Model (3)		Model (4)	
Variables [10]	**< 110h**	**> 110h**	**< 110h**	**> 110h**	**< 110h**	**> 110h**	**< 110h**	**> 110h**
betw. 3 and 6 months in unemp	0.986***	1.009***	0.986***	1.009***	0.999***	1.019***	0.368***	0.345***
	(0.0199)	(0.0185)	(0.0199)	(0.0185)	(0.0199)	(0.0185)	(0.0165)	(0.0151)
above 6 months in unemp	2.051***	2.021***	2.050***	2.022***	2.067***	2.030***	1.231***	1.132***
	(0.0238)	(0.0217)	(0.0237)	(0.0217)	(0.0238)	(0.0217)	(0.0220)	(0.0197)
mean gross replacement rate			-1.359***	-0.782***	-1.476***	-0.889***	-2.459***	-1.957***
			(0.175)	(0.161)	(0.171)	(0.158)	(0.178)	(0.163)
married (yes=1)			-0.0677***	-0.0796***	-0.0633***	-0.0760***	0.00202	-0.00144
			(0.0128)	(0.0112)	(0.0127)	(0.0112)	(0.0137)	(0.0120)
man (yes=1)			0.0382***	0.0356***	0.00841	0.0110	0.0239*	0.0260**
			(0.0132)	(0.0116)	(0.0131)	(0.0116)	(0.0141)	(0.0124)
regional quarterly unemp rate					0.155***	0.129***	0.172***	0.144***
					(0.00430)	(0.00382)	(0.00470)	(0.00414)
UB remaining months							-0.0361***	-0.0396***
							(0.000926)	(0.000816)
post-2006 reform (yes=1)	0.0959***	-0.0531***	0.0961***	-0.0530***	0.167***	0.00820	0.152***	-0.00844
	(0.0160)	(0.0137)	(0.0160)	(0.0137)	(0.0160)	(0.0137)	(0.0168)	(0.0143)
Constant	-4.962***	-4.534***	-4.219***	-4.092***	-5.425***	-5.090***	-1.546***	-1.281***
	(0.0264)	(0.0236)	(0.0988)	(0.0909)	(0.103)	(0.0943)	(0.108)	(0.0980)
p_1	0,744		0,745		0,755		0,3276	
ν_1	2.379***		2.376***		2.347***		-2.307***	
	(0.0194)		(0.0194)		(0.0192)		(0.0184)	
ν_2	2.145***		2.141***		2.114***		-2.062***	
	(0.0178)		(0.0178)		(0.0176)		(0.0168)	
Observations	601,508		601,508		601,508		495,422	
Log-likelihood	-285930.99		-285865.22		-284912.01		-265363.77	
Δ Exit Risk Post-Reform	10,06 %	-5,17 %	10,09 %	-5,16 %	18,18 %	0,82 %	16,42 %	-0,84 %

Standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 1: Competing Risks with Unobs. Heterogeneity: Baseline Estimates

4.2 Magnitude and Elasticities

So far, we have discussed hazard estimates. However, to interpret the magnitude of the policy effect, it is necessary to compute hazard ratios. From the competing risks estimates in the vector β, we compute: $([exp(\beta) - 1] * 100)$. Hazard rates corresponding to post-reform exit rates variations are presented at the bottom line of Table 1. Using the parcimonious model 4 of Table 1, the estimates .152 and $-.0844$ represent a 16.4% increase and a 0.8% decrease post-reform in the propensities of new entrants to resort to PUI jobs respectively below and above 110 hours.

This behavioral response at the broad intensive margin, i.e. between PUI employment

[10] For duration dependence, the reference interval is between 0 and 3 months in unemployment.

below and above 110 hours per month, is driven by a change in the budget constraint, or in other words, by a change in implicit taxation at different worked hours. To compute an elasticity of hours worked with respect to net-of-tax wages, we first translate our policy effect in terms of hour variations. We use the average working time in contracts below and above 110 hours (respectively 50 hours and 142 hours per month, with little variation across the 2005-2006 period) and weight them by the proportion of exits in these two options (24.5%/28.5% and 6.3%/6.3%) to calculate the mean working time before reform. As indicated in the first column of Table 2, we find a mean of monthly hours worked of around 70. Then, using the +16% and −8% estimates on the respective probabilities of exit, we compute the new proportions altered by the post-policy responses. We then find the mean post-policy hours worked is 67.3. The decline in mean hours (−2.9%) is then divided by the mean increase in implicit taxation at 110-136 hours (−20.4%), which gives an hour elasticity of .142. Given the very precise policy effect estimates, we find a narrow confidence interval of $[.137, .146]$ for this elasticity.

	Exit rates	
	Before 2006	Predicted response (2006)
Below 110h	24.5%	28.5%
Above 110h	6.3%	6.3%
Mean hours worked (all)	69.32	67.3

Table 2: Elasticity of Hours Worked with Respect to Earnings in PUI Activity

This elasticity seems consistent with the bunching literature estimates, e.g. the notch approach by Kleven and Mazhar (2013) and Gelber et al. (2016a) and the parametric kink approaches by Chetty et al. (2010), Chetty (2012) and Gelber et al. (2016b) (see Kleven (2016) for a recent survey). Bunching estimates are typically low, since they are affected by frictions such as hour constraints and adjustment and information costs. In our context, elasticities are driven by infra-marginal adjustments to contracts in two broad hour categories, i.e. above and below 110 hours per month. The bunching literature usually computes elasticities for within-job adjustments: in those cases, inertia is caused by the limited possibility to change hours worked within the same job and/or by the seach costs of switching jobs. However, this paper studies broad intensive margin responses for first entries in PUI jobs so we do not have to account for these adjustment costs. Yet, our low elasticity can be attributed to labor market rigidities, namely the difficulty to obtain much hour options within the same job. There is ample evidence that much of the labor supply adjustments goes through job changes rather than contract changes within a job (for instance, Blundell et al. (2008)); indeed, when it comes to job proposals, hours and wages are tied together (cf. Blundell and MaCurdy (1999)). Moreover, we measure hour supply elasticities rather than income elasticities as in Le Barbanchon (2016) or McCall

(1996). The literature on the elasticity of taxable income (ETI), either based on panel data estimation with tax reforms or on bunching at kinks or notches, generally point to lower elasticities than the traditional labor supply elasticities of worked hours. This is illustrated in direct comparison of the two approaches in studies like Thoresen and Vattø (2015), who find an ETI around .05 for Norway and structural labor supply elasticities in the range .10–.29.

In fact, the quasi-experimental literature on bunching points to wide variation in estimates due to very different intensities of frictions. While early studies point to little bunching among wage earners and elasticities below .02, more recent studies find estimates of a closer magnitude to ours. For instance, Le Barbanchon (2016) finds intensive margin elasticities around .10-.20 using bunching around the kink of state-specific earnings threshold of the PUI scheme in the US. Using bunching among US old-age wage-earners at the kink of the Social Security Annual Earnings Test (AET), Gelber et al. (2016b) find an estimate around .35. Overall, Kleven (2016) argues that structural elasticities may be larger than observed bunching-based elasticities by an order of magnitude. For example, using the mass of individuals observed in dominated regions above notches, Kleven and Mazhar (2013) find elasticities around .01-.04 for salary workers. Yet, they indicate that about 90% of workers do not adjust labor supply due to some form of optimization frictions: if not for frictions, bunching at notches would be 10 times larger than observed. Note that Gelber et al. (2016b) observe individuals that continue to bunch at the AET kink even when they are no longer subject to the AET. They derive an elasticity under the assumption of zero adjustment costs as high as .58. Finally, it is relevant to compare elasticities obtained using bunching versus policy reforms. Using the same Danish register data, Chetty et al. (2011) obtain a kink-based elasticity of .01 while Kleven and Schultz (2014) find a reform-based elasticity of .20. The latter study relies on a difference-in-difference approach to identify the elasticity from variation across tax brackets over time, which is arguably less sensitive to the types of adjustment costs that may affect bunching.

5 Additional Checks

5.1 Heterogenous Effects

We use a competing risks model where interaction terms test for heterogenous reform effects with respect to time spent in unemployment. We use the duration dependence partition depicted in Table 1. Competing risks estimates are presented in Table 10 in Appendix.

Pure time dependence coefficients are larger in this regression compared to previous estimates, both for PUI below or above 110h/month. These values are robust to the implementation of various additional controls.

We notice a significant and differentiated reform effect with respect to time spent in unemployment. The decrease in PUI job uptake is mainly caused by the behavior of those unemployed for less than 6 months. And more precisely, the treatment effect magnitude is larger for those seeking a job for less than 3 months compared to those between 3 and 6 months in unemployment. For those over 6 months in unemployment, the treatment effect is an increased PUI uptake: this effect is rooted in the increased necessity to find an activity over time in order to maintain one's income level when benefits' expiration date is either already reached or soon to be reached. Hence, even though we seize a differentiated effect per PUI intensity, it does not contradicts the previous results found in our paper. For PUI contracts above 110h/month, negative effects are larger and positive effects are smaller compared to less intensive PUI contracts, which is in line with estimates obtained in Table 1.

Similarly, elasticities of hours worked in PUI can be computed with respect to time spent unemployed. Estimates of probabilities to join PUI per unemployment duration are available in Table 10. Corresponding elasticities and confidence intervals are presented in Table 3 below.

	Heterogeneity: Unemployment duration		
	Less than 3 months	For 3 to 6 months	Over 6 months
Elasticity of hours worked	0.044	0.061	0.355
Confidence Interval	[0.039 ; -0.049]	[0.056 ; 0.067]	[0.345 ; 0.365]

Table 3: Elasticities of Hours Worked by Unemployment Duration

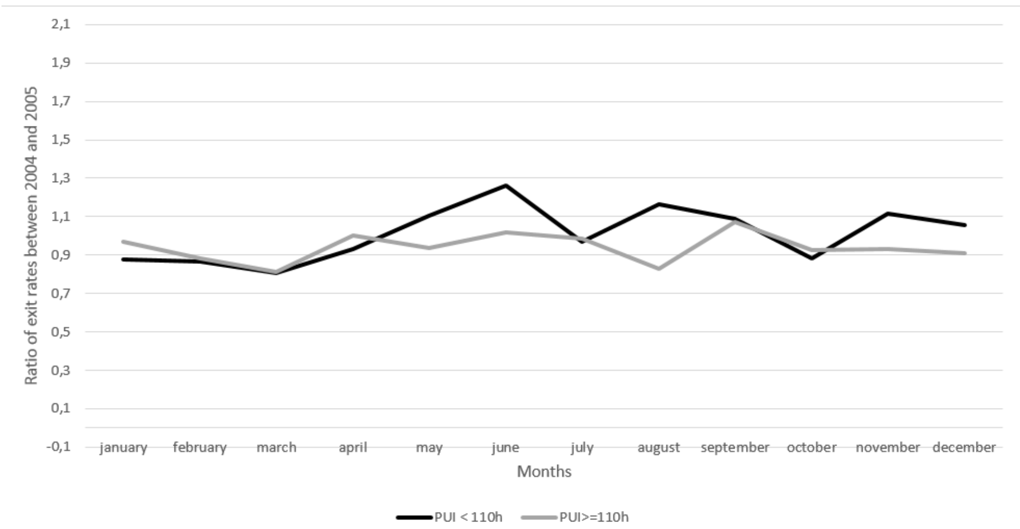
The elasticity of hours worked is rather similar for both groups of unemployed for 0 to 3 months, or for 3 to 6 months. Their magnitude is rather small and, due to precise estimates of the effect of the reform, elasticities are contained within a tight confidence interval. The table suggest that the more time spent unemployed, the larger the elasticity of hours worked with respect to PUI earnings. However, cases of long-term unemployed with ongoing benefits are more rare over time, and encompass individuals who worked long enough, had a large enough wage before being unemployed: these individuals have a PUI behavior which highly depends on the compensation capacity of the PUI system.

5.2 Checking for non-anticipation

An important hypothesis in this model is the non-anticipation of the treatment. The context provides the 2006 reform has been announced in late 2005. Only two references signalled the reform before its implementation in January 1, 2006: the unions' agreement of December 22, 2005 relative to unemployment compensation [11], the earliest reference of the reform being a report from the French Office of unemployment insurance management (UNEDIC) dated October 6, 2005 [12]. The reform was officially declared in the Official Bulletin issued in January 18, 2006. Its late official notification gives us insight that the reform was not likely to be anticipated by the unemployed, without having to suggest a myopic behavior of the unemployed.

Moreover, the observation of monthly exits towards PUI in both 2004 and 2005 present no anticipation behavior of the unemployed. Exit rates towards PUI below or above the threshold remain steady over the last months of 2004 and 2005, as presented in Figure 4. Data corroborate the stability of PUI exit rates, arguing in favor of a non-anticipating behavior towards it.

Figure 4: Monthly PUI Exit Rates Between 2004 And 2005 (Ratios)



Another assumption on which relies the identification strategy is that the selection process remains constant over the period studied: it implies a constant composition of monthly cohorts of entrants in unemployment. The composition of cohorts is also stable with respect to several variables : gender, experience, age, for which tables are presented below (Figure 6, in Appendix).

[11]Accords du 22 décembre 2005 relatif à l'aide au retour à l'emploi et à l'indemnisation du chômage

[12]Délibération du bureau de l'unédic du 6 octobre 2005 pour un suivi et accompagnement des demandeurs d'emploi

6 Conclusion

This paper addresses the issue of how financial incentives embedded in unemployment programs affect job uptake. Partial Unemployment Insurance covers a large number of claimants in France, allowing them to cumulate earnings and a share of their unemployment benefits. This analysis assesses the 2006 reform which narrowed the hour threshold allowing benefit cumulation from 136 to 110 hours per month. We analyze the direct effect of the reform on labor supply at a broad intensive margin, *i.e.* working above or below 110 hours per month. Estimates from competing risks models point to a substantial response to the reform in the form of a 16.4% increase of the relative probability of exiting towards part-time PUI employment. Put against the implicit change in net-of-tax wage induced by the reform, this yields an hour elasticity to PUI earnings of 0.142.

Results are stable to various controls and different duration dependence specifications. Yet, we seize a differentiated treatment effect with time spent in unemployment: essentially, the higher the elapsed duration in unemployment, the larger the reform-induced effect. This is rooted in the necessity to maintain one's income level and generate new benefit rights when current benefits are soon to be expired. The paper provides one of the first applications of the time discontinuity approach suggested in Van den Berg et al. (2014), which borrows from the regression discontinuity framework and requires few assumptions.

This study shows that jobseekers tend to pick jobs that allow for a benefit share. From a policy point of view, the 2006 reform contributed to increase part-time activity, which is unlikely to be a policy objective. For instance, policy measures in France regarding in-work benefits have explicitly put a bonus on full-time activity. UI claimants who would have taken up very intensive PUI contracts if the reform was not implemented would have been more likely to benefit from the stepping stone effect of the PUI program: hence, narrowing the threshold diverts the program from its original purpose. This reform was not initially targeted at the unemployed and was motivated by budget restrictions, so these incentives were probably not accounted for when eligibility conditions were designed. Yet, our results indicate voluntary adjustments to benefit availability.

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A Data Description

A.1 The French Labor Market Context

Table 4: French Labor Market (2004-2006)

Year	Real GDP growth	Employment rate	Unemployment rate	Share of part-time jobs
2004	2.8	51.53	8.9	17.03
2005	1.6	51.48	8.88	17.2
2006	2.4	51.4	8.83	17.25

Sources: World Bank, French Public Employment Service, INSEE.

A.2 PUI Entries and Descriptive Statistics

Table 5: PUI Entries (volumes and rates)

		Pre-Selection		**Post-Selection**	
		2005	2006	2005	2006
Total yearly UI claimants		85,164	131,642	27,116	45,072
New PUI Entries	PUI <110h	21,719	48,361	6,646	14,906
	PUI >= 110h	4,656	10,528	1,715	3,743
PUI Entry rates	PUI <110h	25.5%	36.74%	24.51%	33.07%
	PUI >= 110h	5.47%	8%	6.32%	8.3%

Entry rates are the share of new entrants in PUI on the total yearly UI claimants available (who did not exit for PUI or regular employment).

Table 6: PUI Variations Over the Reform (Post-Selection Data)

Δ Entry rates in PUI	PUI <110h	8.56%
	PUI >= 110h	1.98%
	Total PUI variation	6.58%

Table 7: Detailed PUI Entries (volumes and rates)

		Post-Selection	
		2005	2006
Total yearly UI claimants		27,116	45,072
New PUI Entries	PUI <110h	6,646	14,906
	PUI $\in$ [110;136]	794	1,514
	PUI>136	921	2,229
PUI Entry rates	PUI <110h	24,5%	33,1%
	PUI $\in$ [110;136]	2,9%	3,4%
	PUI>136	3,4%	4,9%

Table 8: Descriptive Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
Male (Yes=1)	59,211	0.6	0.5	0	1
Married (Yes=1)	59,211	0.5	0.5	0	1
Nb. Of Children	59,211	0.9	1.1	0	9.0
Age	59,211	36.3	9.1	18.0	56.0
Unskilled Employee (Yes=1)	59,211	0.1	0.3	0	1
Skilled Employee (Yes=1)	59,211	0.3	0.5	0	1
Experience (in years)	59,211	7.7	8.1	0	41.0
Unemp. Duration	59,211	10.2	6.2	1	24.0
Mean gross replacement rate	59,211	0.5	0	0	0.7
Daily ref. wage	59,211	76.4	35.8	51.8	620.5
Monthly ref. wage	59,211	2291.1	1075.3	1554.3	18615.9

A.3 Duration Dependence

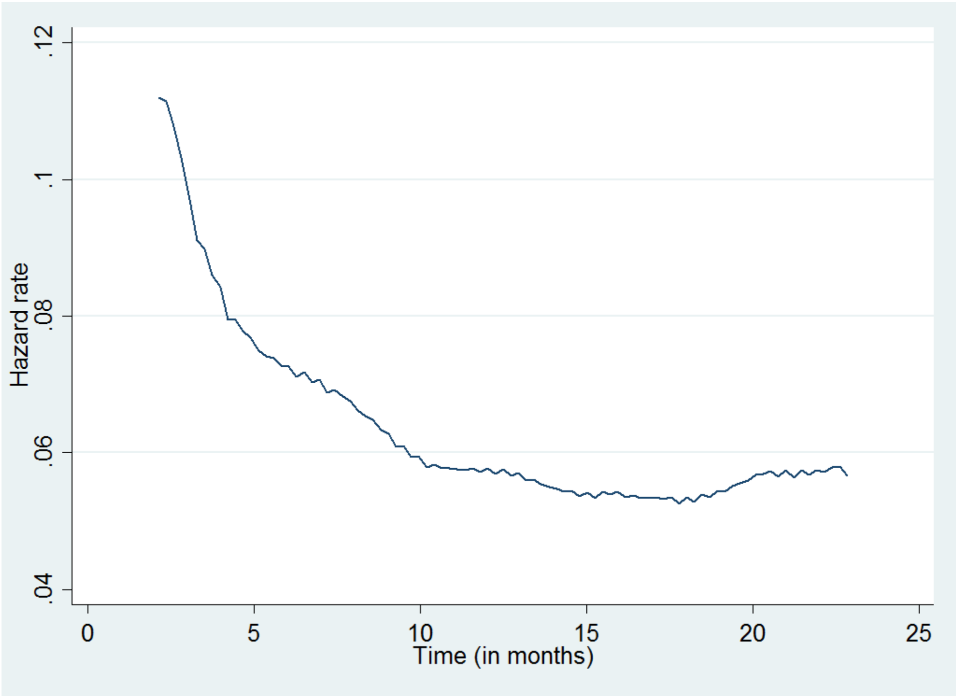

Figure 5: Empirical Hazard Before a First PUI Episode

B Competing Risks Model (detailed)

This appendix derives the likelihood function for the competing risks model defined in this paper. We index by $\{1; 2\}$ the risks to start a PUI activity respectively below and above 110 hours/month. This model accounts for multiple-spell data. The vector $\Omega = \{\lambda_{11}, ...\lambda_{J1}, \lambda_{12}, ...\lambda_{J2}, \beta_1, \beta_2\}$ contains baseline hazards and coefficients related to covariates for both risks.

B.1 Baseline Hazards' Definition

We use a Mixed Proportional Hazard model where the specification of the duration dependence follows a piece-wise constant exponential form. We compute baseline hazards after partitioning months in J intervals. For each interval $j \in J$ we obtain two baseline hazards λ_{j1} and λ_{j2}.

Let τ_j an interval bound with $j \in J$. UI claimants in their t-th unemployment month belong to interval j if $t \in \{\tau_j; \tau_{j+1}\}$.

Let $t_j \in \{\tau_j; \tau_{j+1}\}$. For a given risk i, the baseline hazard λ_{ji} is assumed to have the following form:

$$\lambda_{ji}(t) = \lim_{\Delta \to 0} Pr(t_j \leq t \leq t_j + \Delta | t \geq t_j) \tag{7}$$

For identifiability purposes, we assume that effects are constant within each time interval, which is not overly restrictive since time intervals are relatively short.

B.2 Likelihood Function

The likelihood function for competing risks model with dependent risks is given by:

$$L(t, \Omega | X, \nu) = (\theta_1(t|X, \nu_1))^{d_1} S_1(t|X, \nu_1)(\theta_2(t|X, \nu_2))^{d_2} S_2(t|X, \nu_2) \tag{8}$$

where d_1 and d_2 are dummy variables equal to 1 if the individual exits unemployment for PUI respectively below or above 110 hours, and 0 otherwise. Hazard functions follow a MPH specification defined in equ.(5). We include them:

$$L(t, \Omega | X, \nu) = (\lambda_{j1}(t) exp(X\beta_1)\nu_1)^{d_1} S_1(t|X, \nu_1)(\lambda_{j2}(t) exp(X\beta_2)\nu_2)^{d_2} S_2(t|X, \nu_2) \tag{9}$$

Recall that only the duration before a first PUI: $T = min(T_1, T_2)$ and the type of PUI intensity are observed. Hazard functions θ_1 and θ_2 are dependent.

We account for unobserved heterogeneity with a specification *à la* Heckman-Singer (1984). Let G(ν) the distribution of unobserved heterogeneity: this distribution has a discrete support with 2 mass points.
Class probabilities are determined following a multinomial logit (MNL) parametrization, with $p_i = \frac{exp(\alpha_i)}{1 + \sum exp(\alpha_i)}$, where α_i denotes the probability of having a given combination of unobserved heterogeneities. Class probabilities belong to $\{0; 1\}$ and $\sum_i p_i = 1$.

We integrate unobserved heterogeneity terms in the likelihood function:

$$\Leftrightarrow L(t, \Omega | X, \nu) = \sum_{i \in \{1,2\}} p_i (\lambda_{j1}(t) exp(X\beta_1)\nu_1)^{d_1} S_1(t|X, \nu)(\lambda_{j2}(t) exp(X\beta_2)\nu_2)^{d_2} S_2(t|X, \nu)$$

	Model (1)		Model (2)		Model (3)		Model (4)	
Variables [13]	**< 110h**	**> 110h**	**< 110h**	**> 110h**	**< 110h**	**> 110h**	**< 110h**	**> 110h**
between 3 and 6 months in unemp.	0.326***	0.345***	0.326***	0.345***	0.327***	0.347***	-0.00289	0.000599
	(0.0139)	(0.0128)	(0.0139)	(0.0128)	(0.0139)	(0.0128)	(0.0132)	(0.0120)
between 6 and 9 months in unemp.	0.757***	0.758***	0.756***	0.758***	0.758***	0.757***	0.506***	0.489***
	(0.0143)	(0.0131)	(0.0143)	(0.0131)	(0.0144)	(0.0133)	(0.0151)	(0.0138)
between 9 and 12 months in unemp.	1.462***	1.434***	1.457***	1.431***	1.426***	1.394***	0.889***	0.812***
	(0.0223)	(0.0204)	(0.0224)	(0.0205)	(0.0235)	(0.0215)	(0.0275)	(0.0249)
between 12 and 18 months in unemp.	2.460***	2.321***	2.445***	2.312***	2.275***	2.171***	1.703***	1.534***
	(0.0303)	(0.0260)	(0.0303)	(0.0261)	(0.0313)	(0.0271)	(0.0378)	(0.0319)
above 18 months in unemp.	3.281***	3.223***	3.274***	3.221***	3.211***	3.169***	2.635***	2.491***
	(0.0307)	(0.0262)	(0.0306)	(0.0261)	(0.0298)	(0.0256)	(0.0371)	(0.0314)
mean gross replacement rate			-0.848***	-0.277*	-0.976***	-0.394**	-2.026***	-1.469***
			(0.181)	(0.167)	(0.185)	(0.171)	(0.196)	(0.178)
married (yes=1)			-0.0691***	-0.0823***	-0.0729***	-0.0860***	-0.0392***	-0.0423***
			(0.0131)	(0.0116)	(0.0135)	(0.0119)	(0.0139)	(0.0122)
man (yes=1)			0.0369***	0.0330***	0.0109	0.0112	0.0282**	0.0272**
			(0.0136)	(0.0120)	(0.0140)	(0.0124)	(0.0143)	(0.0126)
quarterly reg. unemp. rate					0.155***	0.133***	0.163***	0.138***
					(0.00528)	(0.00467)	(0.00562)	(0.00495)
remaining months with UB							-0.0165***	-0.0201***
							(0.000892)	(0.000803)
post-2006 reform (yes=1)	-0.101***	-0.219***	-0.102***	-0.220***	-0.0563***	-0.180***	-0.0811***	-0.207***
	(0.0157)	(0.0136)	(0.0158)	(0.0136)	(0.0163)	(0.0141)	(0.0171)	(0.0146)
Constant	-5.367***	-4.809***	-4.889***	-4.632***	-5.953***	-5.546***	-2.151***	-1.946***
	(0.0279)	(0.0242)	(0.103)	(0.0946)	(0.111)	(0.102)	(0.118)	(0.107)
p_1	0,492		0,492		0,491		0,560	
ν_1	2.775***		2.762***		2.581***		-2.548***	
	(0.0230)		(0.0230)		(0.0223)		(0.0244)	
ν_2	2.436***		2.425***		2.267***		-2.181***	
	(0.0199)		(0.0199)		(0.0198)		(0.0211)	
Observations	601,508		601,508		601,508		495,422	
Log-likelihood	-282052.74		-282006.75		-281370.44		-263440.34	
Δ Exit Risk Post-Reform	-9,61%	-19,67%	-9,70%	-19,75%	-5,47%	-16,47%	-7,79%	-18,70%

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 9: Competing Risks Estimates: Detailed Duration Dependence Specification

[13] Standard errors in parentheses. For duration dependence, the reference interval is between 0 and 3 months in unemployment.

B.3 Additional Checks : Heterogenous Effects and Non-Anticipation

	Model (1)		Model (2)		Model (3)	
Variables [14]	**< 110h**	**> 110h**	**< 110h**	**> 110h**	**< 110h**	**> 110h**
betw. 3 and 6 months in unemp.	0.791***	0.848***	0.792***	0.849***	0.808***	0.862***
	(0.0344)	(0.0317)	(0.0345)	(0.0317)	(0.0345)	(0.0317)
above 6 months in unemp.	1.132***	1.372***	1.132***	1.371***	1.119***	1.355***
	(0.0393)	(0.0345)	(0.0393)	(0.0345)	(0.0394)	(0.0345)
mean gross replacement rate			-1.251***	-0.699***	-1.395***	-0.830***
			(0.177)	(0.162)	(0.172)	(0.159)
married (yes=1)			-0.0723***	-0.0834***	-0.0668***	-0.0789***
			(0.0128)	(0.0113)	(0.0127)	(0.0112)
man (yes=1)			0.0429***	0.0396***	0.0131	0.0151
			(0.0132)	(0.0116)	(0.0131)	(0.0116)
regional quarterly unemp. rate					0.160***	0.133***
					(0.00429)	(0.00381)
$\leq$ 3 months and post-reform (yes=1)	-0.579***	-0.594***	-0.579***	-0.595***	-0.546***	-0.566***
	(0.0367)	(0.0343)	(0.0368)	(0.0343)	(0.0369)	(0.0344)
for 3 to 6 months and post-reform (yes=1)	-0.271***	-0.339***	-0.272***	-0.341***	-0.239***	-0.312***
	(0.0245)	(0.0221)	(0.0245)	(0.0221)	(0.0248)	(0.0223)
$\geq$ 6 months and post-reform (yes=1)	0.574***	0.243***	0.573***	0.243***	0.651***	0.309***
	(0.0249)	(0.0194)	(0.0249)	(0.0194)	(0.0247)	(0.0192)
Constant	-4.465***	-4.137***	-1.435***	-1.631***	-4.985***	-4.735***
	(0.0335)	(0.0307)	(0.101)	(0.0931)	(0.105)	(0.0964)
p_1	0,744		0,745		0,756	
ν_1	2.349***		-2.345***		2.322***	
	(0.0192)		(0.0192)		(0.0191)	
ν_2	2.113***		-2.109***		2.089***	
	(0.0176)		(0.0175)		(0.0174)	
Observations	601,508		601,508		601,508	
Log-likelihood	-285113.49		-285047.88		-284030.55	

Standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 10: Competing Risks Estimates: Heterogenous Effects

[14] Standard errors in parentheses. For duration dependence, the reference interval is between 0 and 3 months in unemployment.

Figure 6: Distribution of Entries in Unemployment by Individual Characteristics