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ABSTRACT

Unemployment Duration and the Take-up of Unemployment Insurance*

A large fraction of the eligible unemployed workers does not claim for unemployment insurance (UI) and, among claimants, many do not register immediately upon layoff. This paper argues that, to understand this intriguing phenomenon, one needs to model jointly job search and take-up efforts and to allow for heterogeneity in both dimensions. Estimating such a model using French administrative data, we find substantial heterogeneity in both search and claiming frictions. If half of the sample faces high claiming frictions, many have good employment prospects and exit unemployment quickly. The burden of the claiming difficulties is concentrated on 10% of the sample that suffers both from claiming and job search difficulties. For that reason, the alleviation of the complexity of the claiming process is likely to have very heterogenous effects but little effect on aggregate unemployment duration. Additionally we show that the link between claiming and job search efforts has important implications when measuring how UI parameters impact unemployment duration.

JEL Classification: J64, J65, C41
Keywords: unemployment insurance, take-up, job search

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1 Introduction

Similar to other social benefits (e.g. Currie, 2006 or Hernanz et al., 2004), the take-up of unemployment benefits is far from complete, ranging between 30% and 70%, depending on the country and the nature of the data[1]. Although these numbers are striking, most of the empirical studies devoted to unemployment insurance put aside this issue and concentrate on the sole claimants (e.g. Chetty, 2008 or Card et al., 2015 among many others), mostly because of data constraints. In a similar manner, most of the theoretical papers assume that workers collect benefits - Kroft (2008) or, more recently, Auray et al. (2019) being notable exceptions.

Our main contribution is to investigate the link between job search and the take-up of the unemployment insurance at the individual level. By means of a structural model and administrative data for France, we show that both are interrelated. First, when take-up requires costly efforts, good employment prospects reduce the take-up intensity. This explains a large share of the observed non take-up. Second, using data for France, we nevertheless find that a quarter of the eligible face both job search and claiming difficulties. They need time to complete the claiming process and they can't compensate by intensifying their search intensity given high search costs. We show that these findings have two implications. The first is that claiming frictions can't be considered as filtering the entry into unemployment insurance, selecting those most in need of unemployment benefits. They mostly hurt job seekers that have low unemployment prospects. The second is more methodological. When measuring the effects of benefits, especially the elasticity of unemployment duration, accounting for endogenous take-up and distinguishing between claimed and unclaimed unemployment durations is important. This is due both to the selection into unemployment compensation and to the fact that all eligible workers react to the change even if only a fraction actually claim.

In the article, take-up is the result of a utility-maximizing decision that balances unemployment benefits and take-up costs. As we focus our empirical analysis on a particular

[1]The fact that this range is so wide is due both to differences between countries and, more importantly, to differences between administrative and labor force survey data. Administrative data give much lower estimates because they allow to observe short unemployment spells where the take-up rate is very low. For the US see Blank and Card (1991), McCall (1995) or Anderson and Meyer (1997), for Canada see Storer and van Audenrode (1995), for Austria Fontaine and Kettemann (2019) and for the UK the British Department for Work and Pensions publishes yearly estimates of take-up rates.

population of eligibles with no information problem about their eligibility[2], these take-up costs, possibly heterogeneous among agents, are thought as capturing the practical difficulties to make a claim. There is now a large literature showing that the take-up of social programs is strongly affected by such costs (see Currie, 2006, Kopczuk and Pop-Eleches, 2007, Kleven and Kopczuk, 2011 or Gupta, 2017 among many others). In the context of unemployment insurance, these costs are transaction costs induced, for example, by all the paper work needed to receive unemployment benefits - the documents to collect, the form to fill etc. If our model borrows from the vast literature on the take-up of social benefits that started with Moffitt (1983)'s seminal contribution, there is one important departure: we model take-up costs as *frictions* in an analogy with search frictions and not as a fixed cost. This is key to explain *temporary* non take-up, that is the fact that many claimants do not register immediately[3]. In our framework, it takes *time* and *effort* to collect benefits and it may be easier for some workers than others. Some eligibles are not observed as receiving unemployment benefits because they leave unemployment quickly and made few efforts to collect. Others will be observed as claimants but with a lag between job loss and benefit collection. Facing high degree of frictions and having low employment prospects, they need time to go through the claiming process and more time to find a job.

Considering the interaction between search and claiming, and looking at durations in claimed *and* in unclaimed unemployment is new compared to the existing literature. Most of the existing papers on take-up use static choice models (Blank and Card, 1991, McCall, 1995 or Anderson and Meyer, 1997), where the individual is considered as choosing between claiming and not claiming when she enters unemployment, and the duration until registration is of no concerns. In terms of welfare though, what matters is not only whether the individual eventually claims or not, but also how long she stays without receiving unemployment insurance. Our framework allows for temporary non take-up and we study the distribution of duration with and without benefits. Moreover, although optimistic job expectations are usually reported as one of the main determinants for non claiming (see descriptive evidence in Vroman, 2009), no study has focused yet on this specific determinant. Notable exceptions are provided by Anderson and Meyer (1994)

[2]The fact that some individuals are not aware of their eligibility may be important in other contexts (e.g. the examples in Currie, 2006).

[3]If we were to assume fixed take-up costs, we would not be able to replicate the observed distribution in the duration between the entry into unemployment and the receipt of UI benefits.

and Storer and van Audenrode (1995) who discuss the selection problem, but do not derive or measure its implications for the unemployment insurance.

In terms of data, we exploit a unique administrative dataset which records the entry into unemployment, the take-up of unemployment insurance and the durations in claimed and unclaimed unemployment. Most of the existing studies rely on survey data. The only exceptions, to the best of our knowledge are Anderson and Meyer (1997), Ebenstein and Stange (2010) and Fontaine and Kettemann (2019). When concerned with the analysis of take-up behaviors, administrative data present two main advantages: they are usually more reliable to determine eligibility and they include short unemployment spells. This is important since temporary non take-up is a key feature. We provide a first pass on the data using a multivariate mixed proportional hazard model where we jointly model the exit rate towards employment and the duration to the receipt of benefits. We use the results to understand the importance of unobserved heterogeneity and as guidelines in the estimation of the structural model. We then estimate our structural model by maximum likelihood and run a number of counterfactual exercices.

As predicted by the theoretical model, we find that changes in the parameters of UI affect the composition of claimants. Lower claiming costs significantly moves the composition of claimants towards those facing claiming and job search difficulties. This means that the actual complexity of the claiming process hurts those the most in need of unemployment benefits. Besides, this complexity is not an effective tool to speed up exits from unemployment. In our simulations, the rise in take-up has a limited effect on durations: a one-percent decrease in the claiming costs increases the total duration by only 0.07%.

Regarding unemployment benefits, we find that a one-percent benefit rise increases take-up by 1.3% and that higher benefits improves the pool of claimants. In terms of unemployment duration, it has quantitatively very different effects whether we consider the claimants or the whole population, the duration in claimed unemployment or the total duration. We show that this is due both to the selection into unemployment compensation and to the fact that all eligible workers react to the change even if only a fraction actually claim. Accounting for endogenous take-up and distinguishing between claimed and unclaimed unemployment durations is important to estimate the effects of a change in the unemployment insurance parameters.

The model is presented along with stylized facts in section 2. In section 3, we discuss our dataset and we provide some first results using a multivariate mixed proportional hazard model. Section 4 discusses the estimation, analyses the results and provides counterfactual exercises.

2 A job search model with endogenous unemployment insurance take-up

2.1 The French UI system and some stylized facts

We consider the French UI system between July 2001 and December 2002 and the model is designed to mimic its main features. Details about the system and the data are presented in Appendix A and B and we only provide here a general picture. First, eligibility for UI depends on past employment duration. For example, a worker who has 14 months over the past 24 months gets 30 months. Second, the benefit profile is flat and its level hinges on previous wages. Third, the claiming process is complicated and time consuming. An unemployed worker has first to contact her local unemployment agency, she has to fill a form, describing precisely her situation and she has to provide different documents to prove her entitlement rights. Eventually, she has to show up at her local agency within the first week following her claim. Notice that if a worker completes the process for example four months after her entry into unemployment, she will not be retroactively compensated for these four months. Last, over the period of time we consider, there is in practice no search intensity requirement.

We use the data from the FH-DADS, a match between the records of the French national employment agency (FH) and the French administrative linked employer-employee data (DADS). The analysis sample is presented in section 3 but it is useful to provide some basic empirical evidence before going into the model presentation. This motivates how we model the take-up and gives some insights about potential economic mechanisms. The estimation sample includes male between 30 and 50 years of age and this selection, discussed more carefully later, is made to limit measurement errors of eligibility. Besides, individuals eligible to long duration are most likely to be aware of their entitlement given that they easily pass the minimum eligibility criteria[4]. This helps to separate the mechanism we want to study (the link between job search and claiming) from considerations about eligibility awareness.

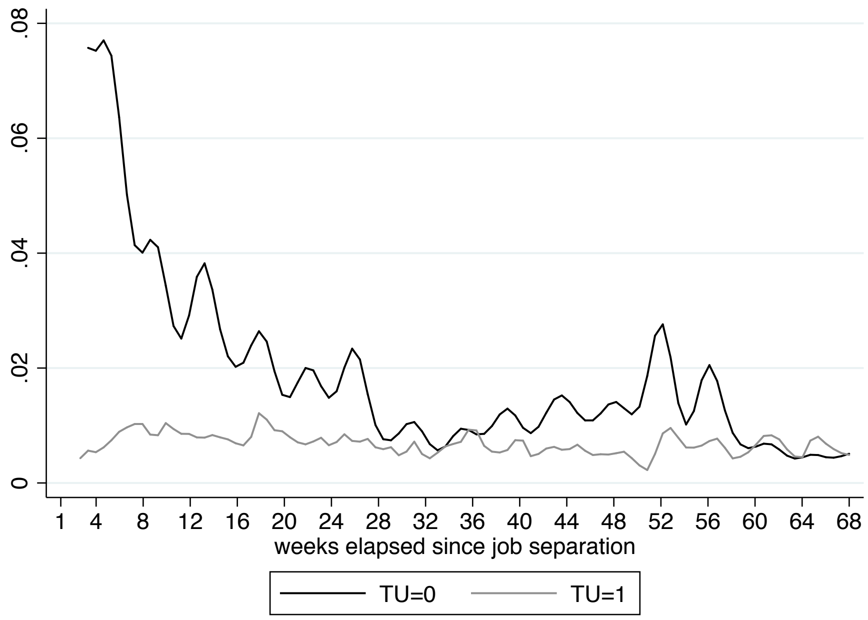
We consider as claimant someone who has gone through all the steps of the administrative process and who is recognized as being eligible by the unemployment insurance agency. We study both the total unemployment duration and the *claimed* unemployment duration (the terminology is chosen in reference to Landais, 2015[5]), that is the duration of unemployment starting when the individual is tagged as eligible. In our sample, the take-up rate, that is the share of eligibles who collects benefits during the non-employment spell, is 30.3%. It is 31.7% on all eligibles. It is important to note that we only exclude unemployment spells shorter than one week and thus keep short spells. This choice is motivated by the fact that even short spell are relevant to measure claiming frictions. In an environment where there are no such frictions, even the individuals who expect very short unemployment spell would collect benefits[6]. Beyond the fact that they are entitled to compensation irrespective of that duration, uncertainty about the next job makes immediate claim relevant. Of course, the welfare loss from non take-up is very limited in this case and our structural framework will take that into account. A second motivation for that choice to retain short spells is that we want to link take-up decisions and job search. For some individuals, high search intensity/short unemployment spells could be explained by their difficulties to collect benefits that make unemployment very costly.

[4]Entitlement to the longest benefit duration requires 14 months of work activity in the past two years, while only 4 months in the past year are required for the shortest benefit duration.

[5]Landais (2015) distinguishes between the total duration, the *claimed* unemployment duration and the *paid* unemployment duration. The difference between claimed and paid durations is that there is sometimes a waiting period before actually receiving benefits. We discuss this issue when presenting the data.

[6]In some cases, the worker still needs to wait before the actual benefits collection. This is for example the case if the worker was fired with severance payments *above* the legal limit. Otherwise, the waiting period is only one week. In practice, it would have been impossible to account explicitly for the waiting period: it is only observed for the claimants, it is impossible to predict precisely with our data for the non-claimants, and its inclusion in a structural model would have been very complicated without clear benefit for what we are eventually looking at.

Figure 1: Unemployment exit rates by take-up status (in weeks)



TU=0, non-claimants. TU=1, claimants. Sample: 30-50 year-old males who experience a job loss between 07/2001 and 12/2002, are eligible for UI at job separation and are entitled to the longest benefit duration. Kernel-smoothed hazard estimates for the total unemployment duration. Source: FH-DADS.

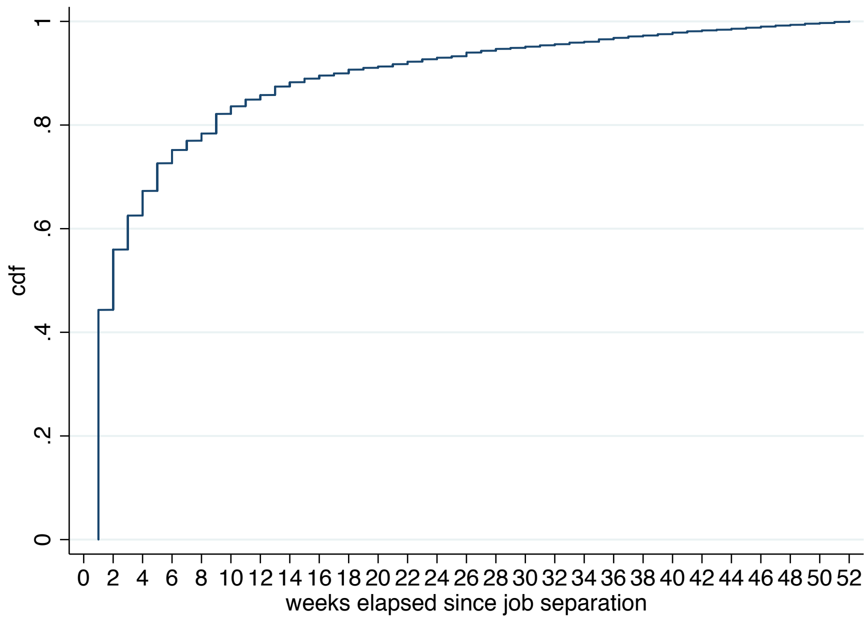
Figure 1 displays the unemployment hazard rates[7]. We make a distinction between workers who are observed as receiving unemployment benefits at some point of their unemployment spell and those who are not. Interestingly, for the latter, the exit rate from unemployment is much higher in the weeks just following separation. There is then is a huge drop in the exit rate during the first four months. This comes from the fact that most of the workers who do not claim for benefits find a job quickly. As documented by Vroman (1991), expectation of a short unemployment spell is declared by the individuals as the most important reason for not claiming[8]. It is likely that this induces a form of selection. Workers who exit unemployment quickly have the most efficient search technology. This both increases mechanically the probability to exit before completing the claiming process and decreases the incentive to make an effort to collect benefits. For the other workers, those who are observed as taking up benefits, the hazard rate is overall flat.

[7]Notice that the spikes in the hazards correspond to the fact that most of the jobs end at the end of a month and start at the beginning of a month or a week.

[8]Using March supplementary CPS data, he finds that 37% of eligible job seekers did not claim for benefits because they "expected to get another job soon/be recalled". Claiming difficulties and stigma account together for 18% of the answers.

Among claimants, the transition to the state where they collect benefits is not instantaneous. Figure 2 displays the distribution of durations in non-claimed unemployment among those who receive benefits at some point during their unemployment spell. While take-up is completed within a week for half of the sample claimants, there are still 20% of claimants that stay more than 12 weeks without unemployment benefits. Among the sole claimants, average duration until the completion of a claim is of 6.7 weeks. For some workers, claiming takes time. Remark that, if the reason for the gap between the entry into unemployment and claiming was that some workers had the wrong beliefs about their job opportunities, we would see an increase in the claiming probability over time as they learn about their true opportunities. We do not observe such an increase in the data.

Figure 2: Distribution of durations (cdf) in non-claimed unemployment among takers (in weeks)



Sample: 30-50 year-old males who experience a job loss between 07/2001 and 12/2002, are eligible for UI at job separation and for the longest benefit duration, and who claim benefits during their unemployment spell. Source: FH-DADS.

Some implications for the model of UI take-up can be derived from this overview of the UI system and description of claiming behaviors. First, a model of UI take-up has to allow for claiming costs given that UI receipt is neither automatic, nor simple. Second, we need to model claiming as a process that takes time and where individuals make costly efforts. In an analogy to job search, we thus assume that there are claiming *frictions*. Last, it is necessary to model jointly job search and claiming to model selection into take-up based on expected and effective unemployment durations, and to account for the impact of frictions on the exit rate.

2.2 A search model with endogenous take-up

The model. Our model is designed to account for the interaction between job search and claiming behaviors, providing a natural way to model selection into insured unemployment. As in our estimation sample, all workers are eligible. Time is continuous and the labor market is assumed to be at the steady state.

When a worker enters unemployment, she may claim for unemployment benefits and, at the same time, search for a job. This state is noted N. If she completes the claiming process before finding a job, she starts receiving unemployment benefits. She continues to search, but may be changing her search intensity or reservation wage. This state is noted P and named claimed unemployment. We denote u_j the utility flows in state j (with $j = \{N, P\}$). Differences between u_N and u_P come from the receipt of unemployment benefits in P, but also from possible costs of being registered at the public employment agency (due to stigma or to compulsary meetings with caseworkers). The worker chooses the job arrival rate, noted λ_j, at a cost $c_E(\lambda_j)$ (with $c_E(.) > 0$, $c_E(0) = 0$, $c'_E(.) \geq 0$ and $c''_E(.) > 0$). For the sake of simplicity, the wage offer distribution $F(.)$ is not state dependent. The exit rate from the unemployment state j to employment thus reads $\lambda_j (1 - F(R_j))$, with R_j the reservation wage in state j.

We model the take-up decision as the result of an effort to deal with the complexity of the administrative process. The claiming process is costly and takes time, as the worker has to understand the administrative requirements, collect the documents needed and fill a claim. Claiming frictions are modeled in a similar way as search frictions. A worker switches from state N to state P, where she receives benefits, at a rate γ and at a cost $c_P(\gamma)$ (again a function with the usual properties). This cost function, to be estimated, is an index of frictions in the claiming process. As it will be more apparent when we discuss estimation, it can be worker specific, like all the other primitives of the model (the job arrival rates, the utility flows, the parameters of the cost functions etc.). Such a model of take-up costs as claiming frictions permits to have a dynamic process and different durations until registration between job seekers. Assuming a fixed cost of claiming would indeed only separate the eligible population between those who claim immediately and those who never claim.

When the worker receives benefits, we assume that the insurance ends at a rate μ. Assuming that the end of the insurance is stochastic induces stationary search behaviors in state P and may be considered as an important simplification. Remember though that the focus of the model is on the claiming of unemployment insurance, that is the transition between state N and state P, and on the unemployment duration, rather than on the exact time profile of the exit rates from claimed unemployment.

We now introduce the value functions. We denote ρ the discount rate. The intertemporal values of unemployment in states N and P, respectively V_N and V_P, read:

$$\rho V_N = \max_{\lambda_N,\gamma}\left\{u_N - c_E(\lambda_N) - c_P(\gamma) + \lambda_N \int_{R_N} (V_E(x) - V_N)dF(x) + \gamma \max\{V_P - V_N, 0\}\right\}$$

$$\rho V_P = \max_{\lambda_P}\left\{u_P - c_E(\lambda_P) + \lambda_P \int_{R_P} (V_E(x) - V_P)dF(x) + \mu(V_{N'} - V_P)\right\}$$

with $V_E(x)$ the value of a new job with a wage x and $V_{N'}$ the intertemporal value of unemployment after benefit exhaustion.

The function value of unemployment after exhaustion is assumed to be similar to the one of uninsured unemployment, except for the fact that claiming is not allowed, as the individual is no longer eligible. It then reads:

$$\rho V_{N'} = \max_{\lambda_{N'}}\left\{u_{N'} - c_E(\lambda_{N'}) + \lambda_{N'} \int_{R_{N'}} (V_E(x) - V_{N'})dF(x)\right\}$$

To define the value of a job, we allow for job-to-job transitions (with probability λ_E and a search cost $c_E(\lambda_E)$) and job destructions[9]. With probability q the worker loses her job and becomes unemployed, in which case she gets V_j. To simplify the model (and its estimation), both probabilities μ and q are assumed to be exogenous. Moreover, we make

[9]Notice that the model would be much simpler if we did not allow for job destruction and job-to-job transitions, in which case we would simply have $V_E(x) = u_E/(1-\rho)$. However, this would be an important simplification.

the simplifying assumption that the individual expects to return to the unemployment state j where she was before being employed. Hence, if a worker finds a job before collecting benefits (that is in state N), she has to redo the claiming process from start in case of job loss. If she was receiving benefits (state P), she assumes she will be collecting benefits again. This is arguably a simplifying assumption but it speeds up dramatically the computation of the optimal solutions and thus the estimation in a framework where we allow for observed and unobserved heterogeneity in the benefit levels, search costs and claiming costs[10]. It is also not such a bad proxy of the actual rules[11]. The value of a job paid x and found when the individual is in the unemployment state $j \in \{N, P, N'\}$ then reads:

$$\rho V_E(x) \quad = \quad \max_{\lambda_E} \left\{ u(x) - c_E(\lambda_E) + \lambda_E \int_x (V_E(x') - V_E(x)) dF(x') + q(V_j - V_E(x)) \right\}$$

The optimal behaviors. We focus now on job seeker's optimal behaviors in state N and P. Remember that job search and claiming activities are simultaneous. Workers choose their search efforts, reservation wages and claiming intensities to maximize their intertemporal utility. The reservation wages are such that workers are indifferent, if paid at that wages, between staying unemployment and accepting the job. Formally, the reservation wages R_N and R_P are such that $V_E(R_N) = V_N$ and $V_E(R_P) = V_P$. Not surprisingly, the higher the value of unemployment, the higher the reservation wage. Interestingly, it means that changes in the parameters of unemployment insurance affect the reservation wages in both claimed and unclaimed unemployment states: a more generous unemployment compensation means higher reservation wages even for workers who are not currently collecting benefits if they have a positive probably to succeed in claiming before finding a job.

The search efforts are functions of the relative returns to job search and the optimal claiming rate is the result of a comparison between the value of unemployment as recipient and as non-recipient. The first order conditions are

[10]Under this assumption, the reservation wage in P simply satisfies $u(R_P) = u_P$.

[11]If an individual on benefits finds a job but then goes back to unemployment, she can collect the benefits for the remaining insurance duration (potentially increased if she qualified for new rights).

$$c'_E(\lambda^*_N) = \int_{R_N} (V_E(x) - V_N)\, dF(x) \tag{1}$$

$$c'_E(\lambda^*_P) = \int_{R_P} (V_E(x) - V_P)\, dF(x) \tag{2}$$

$$c'_P(\gamma^*) = Max\{V_P - V_N, 0\} \tag{3}$$

In some cases, the worker has few incentives to claim unemployment benefits and γ^* is close to zero. This is especially true if unemployment benefits are small with respect to claiming costs. Besides, the value of claimed unemployment (P) impacts job search behaviors of all non-employed workers. A rise in the benefits increases V_P and, at the same time, V_N. For that reason, if the compensated workers are better off, workers who still have to claim also reduce their search efforts and increase their reservation wages. This induces a fall in the exit rates from unemployment in both N and P. Notice that the impacts of UI parameters on search behaviors in N are higher when claiming costs are low because the individual is then more likely to collect benefits.

Endogenous take-up and duration elasticities to unemployment benefits All these elements point to the fact that the elasticity of unemployment duration with respect to UI benefits is affected by take-up. The existing literature focuses on the elasticity for UI recipients. As noted by Anderson and Meyer (1994) or Krueger and Meyer (2002), the true elasticity accounts for the effect on the take-up rate, together with the effect on durations, both for recipients and non recipients. The elasticity of the total unemployment duration ε^D equals:

$$\varepsilon^D = \varepsilon^N s_N + \left(\varepsilon^{TU} + \varepsilon^P\right)(1 - s_N) \tag{4}$$

with ε^N the elasticity of the duration in state N, ε^{TU} the elasticity of the take-up probability, ε^P the elasticity of the claimed unemployment duration and s_N the share of the duration in state N in the total unemployment duration. ε^P and ε^{TU} are both positive. Indeed, a more generous system pushes workers to decrease their exit rate from unemployment but it also increases their claiming rate. As for ε^N, it can be either positive or negative depending on the relative effect on job search behaviors *vis-à-vis* claiming behaviors.

Hence, when measuring the effect of a change in the benefit level, the account for take-up raises a number of issues. First, we need to define which part of the unemployment spell one considers. There is a difference between the overall unemployment spell (states N *and* P) and the *claimed* unemployment spell (state P). Job search behaviors are affected by the level of unemployment benefits even before the individuals collect benefits. The existing papers on the effect of unemployment benefits on unemployment duration have made different choices, mostly due to data constraints, and usually without explicitly accounting for this selection problem. Recent papers consider eligible UI takers (the claimants). Some focus on the period where individuals receive benefits (e.g. Card et al., 2015), some make the distinction between claimed or paid unemployment and total non-employment (Landais, 2015), others consider non-employment duration, including the take-up period and the claimed unemployment period, but under the condition that workers collect within a month (e.g. Card et al., 2007 or Chetty, 2008).

Second, one needs to account for selection. A change in the benefit level modifies the pool of workers who claim. The direction of this change depends on the link between the claiming cost heterogeneity and the job search cost heterogeneity which is usually not modeled explicitly. The control of the endogenous selection is thus complicated. For example, strategies using variations across states and time (as in Chetty, 2008) to measure duration elasticity to benefits can't control for the fact that the composition of claimants changes with the benefit levels. When applied on the total unemployment duration of all eligible job seekers, this approach captures the effect of benefit changes on take-up, duration in N and duration in P, without disentangling the specific effects on each of these three components. When applied on claimed unemployment for takers only, it assumes that the change does not affect the unobserved characteristics of the takers, which is unlikely in presence of endogenous take-up. Strategies based on discontinuities or kinks (see Card et al., 2007 or Card et al., 2015), assume that the heterogeneity evolves smoothly around the discontinuity/kink. If it is true, the correct elasticity is identified locally[12] even if the benefit level does change the pool of claimants. However, as it will be clear in our counterfactual exercises, a treatment implying a substantial change is likely

[12]More precisely, what is identified is a weighted average of the effect of the benefit generosity on duration.

to violate the smoothness assumption[13].

In order to see how the different measures of the non-employment duration, the take-up rate and the claimant composition pool respond to a change in unemployment insurance parameters, we will measure in the last section how workers would respond to a change in the replacement rate. Before, we estimate our model on administrative data.

3 Data and Descriptive Analysis

We begin this section by presenting the data and the selected sample. Both are described in details in Appendix B. We then offer a first analysis of our data using a multivariate mixed proportional hazard model. Results obtained will help to understand their main features and will serve as guidelines for the specification and the parametrization of the model.

3.1 The data

Our data come from a merge between two longitudinal administrative datasets: the records of the French employment agency, that include all compensated unemployment spells since 1999, and the matched employer-employee dataset which contains, for the private and semi-public sector, all employment periods since 1976. The original datasets are 1/24 nationally representative samples and the merge contains any individual who appears in one of these two records between January 1st, 1999 and December 31st, 2004.

These data present a number of features which make them well-suited for the study of the UI take-up. First, in comparison to survey data (for example used by Blank and Card, 1991), administrative data provide information which is less likely to be plagued with measurement errors and misreporting. Second, we can sample on the flows out of employment, which is crucial here, as we can observe all unemployment spells, even those of short duration. Survey data usually sample from the stock of unemployed, which leads to an over-representation of longer unemployment spells. This leads to overestimate the take-up rates because, as it is clear in the model, unemployment durations and take-up probability are positively correlated. Third, our panel has enough retrospective

[13]For that reason, regression kink designs as in Card et al. (2015) are likely to provide a better measure of the true local elasticity of paid unemployment duration than designs using a big jump in one of the UI parameters as in Card et al. (2007).

information to predict eligibility which depends on the worker's past employment spells. Finally, we have *daily* information about job separation, reemployment, registration and compensation. In comparison, Anderson and Meyer (1997), who also use administrative data, only have quarterly information about job separations.

3.2 Eligibility and sample selection

The analysis sample includes eligible unemployment spells starting between July 2001 and December 2002, a period where the UI rules have been stable. Even with administrative data, the determination of the eligibility status can be an issue and has to be handled carefully. Eligibility depends on past employment duration (see Table A.1 in Appendix A) and, since our data trace all past jobs, we observe whether a worker satisfies this criterion[14]. However, we cannot distinguish between genuine unemployment and inactivity due to schooling, retirement or child care. For that reason, we restrict the sample population to male workers between 30 and 50 years of age[15]. In the same way, as it is the case in most studies on the UI take-up, we have no information about the causes of job separation. This may be a problem since voluntary quits are not eligible. However this problem is likely to be limited: using the Labor Force Survey (*Enquête Emploi*) over the same period, we compute that only 1.8% of 30-50 years-old males declare that the job ended because they quit. Moreover, a worker who voluntarily quits can become eligible after four months of unemployment and we do not see a spike at four months in the hazard to registration. Nevertheless, because a worker who voluntarily quits is likely to quit his current job to another (more attractive) job, we exclude very short unemployment spells (less than a week).

We obtain a sample of 18,034 eligible non-employment spells. We finally perform the analysis on the subsample of the 15,453 spells eligible for the longest UI duration. Including workers entitled to shorter benefit durations put us at risk of false positives and thus downward-biasing the take-up rate estimates (see in Appendix B[16]). By focusing

[14]Note that we use the strict definition of eligibility to avoid keeping workers that are not eligible. Since some workers can benefit from additional eligibility rules (given that they do not meet the standard criteria), we are likely to exclude from the analysis some eligible workers.

[15]Other types of inactivity such as maternity leave or sickness leave do not generate an exit from employment in the DADS dataset and do not affect the eligibility status, and are hence not problematic to us.

[16]Potential benefit duration is recorded in the public employment agency records, allowing us to test the quality of our eligibility predictions, but for takers only.

on workers eligible to thirty months of benefits, we render negligible the probability of tagging a worker as eligible while she is not, though we perfectly recognize that this choice induces selection towards workers with more stable labor market trajectories. Table 1 shows that, in comparison with the overall population, our sample is indeed a bit older, slightly more skilled and less likely to have worked in services or as a plant worker.

We follow the individuals from the entry into unemployment (state N) to the exit to employment (state E), if any, and observations are censored 72 weeks after job separation. Entry in claimed unemployment (state P) occurs when we observe the completion of a registration at the public employment agency. Notice that it doesn't mean that unemployment benefits are received immediately. For some workers, those who receive severance pay above the legal minimum or who had been paid for outstanding vacation days, there is a waiting period that cannot exceed 75 days. This is the case for 23% for our sample. However, it is unlikely that the waiting period explains the duration in unclaimed unemployment (that is in N). First, there is no reason to wait before completing the administrative process: if somebody has not finished that process in time, there is no retrospective payments, payments start when registration is completed. Thus, if anything, by waiting the individual risks of being paid with delay. Second, for those who claim, our data include the information about the waiting period and we can check if the waiting time is correlated with the unclaimed duration. This is not the case, if anything the correlation is slightly negative (≈ -0.05) and remains negative after controlling for observables like location, age, replacement rate or industry. Hence, the waiting period doesn't explain the time between the entry into unemployment and the entry in claimed unemployment.

3.3 Descriptive analysis

Table 1 shows the take-up rates in our analysis sample, by individual characteristics, entitlement level and occupation. The observed take-up rates are around 32%, comparable to the one obtained by Anderson and Meyer (1997)[17]. There are not much differences in take-up rates across individual characteristics. Older workers have lower take-up rates but there are no clear patterns emerging in terms of skills or sector of activity. On the

[17] Depending on the sample selection, the authors found a take-up rate between 24 and 54%. After cleaning for spurious separations or job-to-job quits, they obtained 39%. The highest number is based on a sample considering only mass layoffs.

contrary, workers entitled to shorter compensation periods have higher take-up rates. One of the possible reasons is that these workers usually have little or no wealth, and little labor market prospects, and thus the cost of bypassing unemployment benefits is higher for them than for other workers.

Table 1: Sample composition and take-up rates

Eligibles	All		Maximum duration only	
	Composition	TU rate	Composition	TU rate
N	18034		15453	
Overall take-up rate (%)		31.7		30.4
Age category in 2001 (%)				
31-35	40.9	33.2	40	31.2
36-40	23.2	30.5	23.5	29.9
41-45	22.8	31.2	23	30.4
46-50	13.1	29.6	13.5	28.6
Skilled workers (%)	36.2	32.3	39.2	31.9
Previous job occupation(%)				
Managers and prof.	17.1	31.5	18.8	31.7
Tech. and associate prof.	19.1	33.1	20.4	32.1
Employees	10.8	36.7	9.9	36
Plant workers	53	30.2	50.8	28
Previous employer sector(%)				
Agriculture	4.5	27.9	3.8	28.1
Industry	19	33.4	21	32.7
Construction	14.5	27.3	15.1	24.8
Retail	24.2	33.1	25.6	32.2
Service	36.5	32	33.1	30.2
Social	1.3	30.7	1.3	29.7
Benefit duration (%)				
4 months (type 1)	4.7	43.8		
7 months (type 2)	3.3	44.3		
15 months (type 3)	6.3	33.8		
30 months (type 4)	85.7	30.4		

Sample: 30-50 year-old males eligible for UI who experience a job loss between 07/2001 and 12/2002. Source: FH-DADS.

Before going to the estimation of our structural model, we estimate a multivariate mixed proportional hazard model (e.g. Van den Berg, 2001; details of the model are presented in Appendix C) to serve as guidelines for our empirical specification. Such a model is very similar to the job search model presented above. It indeed jointly models the distribution of the duration to registration and the unemployment duration. Therefore, it accounts for the endogenous censoring of the duration to registration that happens when a worker exits unemployment quickly.

We allow for common determinants between job search and benefit claiming. For example, the value of insurance is a determinant of take-up, so that take-up is related to past wages, but past wages are also an indicator of employability and thus impact the exit rate from unemployment. Moreover, unobserved heterogeneity, reading ability for instance, might affect both the search efficiency and the ability to go successfully through the claiming process. We assume a discrete distribution for the unobserved heterogeneity with a factor-loading specification. The unobserved heterogeneity parameters, denoted ω_E (for the unemployment-to-job hazard rate) and ω_P (for the registration rate), can take two values, one being zero, the other to be estimated. We hence end up with four types and we estimate the probability to belong to each of these types.

Table 2 reports the results of two specifications, one without unobserved heterogeneity and one with correlated unobserved heterogeneity in the two hazards. In the first specification, individual characteristics significantly affect the hazard rate to registration. But once we control for unobserved heterogeneity, age and occupation level have much lower or no significant effect on collecting UI benefits (exit to state P). This could mean that variations in take-up rates, if any, come from unobserved heterogeneity or variations in the exit rates to employment, which are related to worker observed heterogeneity. Indeed, worker characteristics do affect the exit rate from unemployment in the two specifications: as expected, younger individuals or individuals with higher past wages find jobs more quickly. On the contrary, other variables being kept constant, more skilled workers (higher occupation) need more time to find a job.

The distribution of unobserved heterogeneity reveals that about 40% of the sample has high search abilities and low claim abilities. Very few individuals are of high types in both claiming and searching for a job (less than 1% of the sample)[18]. Still, 75% of the sample have low claiming abilities ($\omega_P = 0$) and around one third faces both high claiming frictions and low job finding prospects ($\omega_P = 0$ and $\omega_E < 0$): they are probably the ones for which unemployment has the highest welfare cost.

As emphasized by our model, exit rates are functions of abilities (captured by the heterogeneity in the search and claiming cost functions) and efforts, the latter being also impacted by the value of employment *versus* the value of unemployment insurance. This makes a definitive interpretation of the results from Table 2 complicated. For example, past wages are found to have no significant effect on the exit to claimed unemployment (P). This does not mean that past wages have no effect on the costs. We could indeed

[18]In our specification, being a high type means to have $\omega_P > 0$ for claiming and $\omega_E = 0$ for job finding since the estimation gives $\omega_E^b < 0$.

imagine that better paid workers claim more easily but, because they have better job prospects, they put more efforts in job search and lower their claiming efforts. This would induce the coefficients related to past wage in the claiming hazard to be not significant from zero. This is one of the difficulties of such a reduced-form approach.

Hence, an important takeaway is that it is crucial to account for unobserved heterogeneity and to consider jointly job search and claiming behaviors if one wants to model and understand the UI take-up[19]. The reduced-form estimation brings evidence about the determinants of take-up but it can be usefully complemented by a structural approach to reveal the underlying mechanisms at play.

[19] We do not allow for duration dependence since claiming generally happens in the very first weeks/months of unemployment - where human capital depreciation is probably very limited - and because it is always challenging to identify separately unobserved heterogeneity from true duration dependence.

Table 2: Multivariate mixed proportional hazard model

	Independent durations		Dependent durations	
	to P	to E	to P	to E
log(past weekly wage)	0.313	4.013***	0.407	5.938***
	(0.441)	(0.364)	(0.718)	(0.623)
log(past weekly wage) squared	-0.008	-0.329***	-0.022	-0.474***
	(0.037)	(0.031)	(0.061)	(0.053)
Age	-0.012***	-0.008***	-0.008*	-0.021***
	(0.003)	(0.002)	(0.004)	(0.003)
Skilled	-0.179***	-0.106***	-0.052	-0.208***
	(0.049)	(0.034)	(0.083)	(0.053)
Previous job duration (in months)	-0.002***		-0.001	
	(0.000)		(0.001)	
Intercept	-5.224***	-15.831***	-7.182***	-19.869***
	(1.310)	(1.077)	(2.155)	(1.845)
ω_k^b			4.803***	-2.671***
			(0.042)	(0.026)
$P(\omega_P = \omega_P^b, \omega_E = 0)$			0.81%	
			(0.003)	
$P(\omega_P = 0, \omega_E = 0)$			41.13%	
			(0.007)	
$P(\omega_P = \omega_P^b, \omega_E = \omega_E^b)$			24.36%	
			(0.005)	
$P(\omega_P = 0, \omega_E = \omega_E^b)$			33.71%	
			(0.007)	
No. of Obs.	15453		15453	
Log-likelihood	-75032.766		-68603.365	

Sample: 30-50 year-old males eligible for the longest UI compensation who experience a job loss between 07/2001 and 12/2002. Standard errors for the probabilities are calculated using the delta method. Source: FH-DADS.

4 Estimation and results

We estimate our model on the sample presented in the previous section using maximum likelihood. We first present our parametrization choices, then the estimation results and finally the results of our counterfactual exercise.

4.1 Estimation and parametrization

Our parametric assumptions are driven by data constraints or are necessary for identification. For example, we estimate a version of the model where the benefit exhaustion rate μ is set to 0 mainly because we do not observe any transition towards social assistance in our sample, rendering impossible the identification of parameters for that labor market state. Moreover, given the nature of our data, which do not provide information about wealth or consumption, we do not estimate the utility function but assume a log utility function[20]. It is specified as being a function of the worker's past wage with $u_j(w_{\text{past}}) = \ln(a_j + b_j w_{\text{past}})$ for $j = \{N, P\}$, where w_{past} is the average weekly wage in the last employment spell. In state P, when the worker receives unemployment benefits, we calibrate b_P to 0.75 which corresponds to the average *net* replacement rate over the period (OECD, 2014). We estimate the three other parameters (a_N, a_P and b_N). This specification accounts for the fact that, while not receiving benefits, the worker's utility can still be a function of past wages, for example because of savings correlated with past earnings. It also allows for the existence of stigma or for costs related to additional administrative requirements for registered individuals. Finally, the weekly discount factor ρ is set to 0.001 which amounts to a five percent annual discount.

Consistent with previous literature, we assume a quadratic search cost function and we make the same assumption for the claiming cost function. However, we allow for heterogeneity since, for each cost function, there is a scale factor which varies according to observed and unobserved heterogeneity. We note $(\omega_{\gamma i}, \omega_{\lambda i})$ the individual vector of unobserved heterogeneity affecting claiming and job search respectively, and X_i the vector of observable characteristics. The cost functions read

$$c_\lambda(\lambda) = \eta_\lambda \lambda^2, \text{ with } \eta_\lambda = e^{X_i\beta_\lambda + \omega_{\lambda i}}$$
$$c_\gamma(\gamma) = \eta_\gamma \gamma^2, \text{ with } \eta_\gamma = e^{X_i\beta_\gamma + \omega_{\gamma i}}$$

As for the observed heterogeneity, we follow the duration model presented in subsection 3.3 : we account for the worker's past occupation level, her age and the duration of her previous job. Remember that, even if the cost is not in itself a function of past wage, the search and claiming intensities are functions of past wage, since it impacts the utilities and thus the incentives to search and collect benefits. Regarding unobserved heterogeneity, like in the reduced-form model, we assume a factor-loading specification. $\omega_{\lambda i}$ and $\omega_{\gamma i}$ can each take two values with the normalization that the first one equals zero

[20]It is in principle feasible but the identification would heavily rely on parametric assumption.

while the second one, noted respectively $\bar{\omega}_\gamma$ and $\bar{\omega}_\lambda$, is to be estimated. The associated probabilities are

$$p_1 = Pr(\omega_\gamma = \bar{\omega}_\gamma, \omega_\lambda = \bar{\omega}_\lambda) \qquad p_2 = Pr(\omega_\gamma = 0, \omega_\lambda = \bar{\omega}_\lambda)$$
$$p_3 = Pr(\omega_\gamma = \bar{\omega}_\gamma, \omega_\lambda = 0) \qquad p_4 = Pr(\omega_\gamma = 0, \omega_\lambda = 0)$$

As for the wage offer distribution, we assume that wage job offers are drawn from a truncated Weibull distribution with cdf

$$F(x) = 1 - e^{-\left(\frac{\ln(x)-\ln(w_{\min})}{m_1}\right)^{m_2}}$$

where x is the wage in level, $\{m_1, m_2\}$ are shape and scale parameters to be estimated, and $w_{\min}$ is chosen such that it truncates the distribution at ten percent below any wage observed in the data[21]. Remark that the wages we observed are not hourly wages but weekly salaries. Some jobs correspond to part-time jobs while others are fully time. Our wage distribution is hence a mix of full-time wages, part-time wages, low hourly wages and high hourly wages. Since workers' decisions are assumed in the model to be made according to the salary rather than the hourly wage, the weekly salary is the relevant variable[22]. Finally, because we would have to follow the worker history after reemployment, it is difficult to estimate the job destruction rate q. We thus calibrate it to match the average employment duration in our data ($q = 0.0019$).

The model is estimated by maximum likelihood. Given parameter values, we solve numerically for the optimal behaviors (the search intensities, the reservation wages and the claiming effort) and then compute the individual contributions to the likelihood. The likelihood is pretty straightforward to derive and we postpone its derivation to Appendix D. However, it is important to note that we allow for measurement errors in wage by assuming that observed log-wages equal actual log-wages plus an error term. This error term is iid across individuals and distributed log-normally with a standard deviation σ_ν.

[21]This minimum wage in level is about 85 euros in term of weekly wage. This cutoff, which has to be below the minimum observed wage, would be poorly identified if we tried to estimate it. In practice, this number doesn't affect the results if it is chosen within a reasonable range.

[22]Additionally, we average the wage along the re-employment spell (up to one year) to limit the noise that could come from workers who vary their working hours over the spell.

4.2 Estimates and fit

Estimated parameter values are reported in Table 3. The utility in unclaimed unemployment (state N) is, as expected, positively affected by past wages. The constants of the utility functions are positive but only the one associated to unclaimed unemployment is significant and higher than the one in claimed unemployment (P). One of the possible explanations for this result is the presence of stigma or administrative constraints that would shift down the utility when the individual enters claimed unemployment. Consider now the cost functions. Search efficiency is positively affected (negative values mean lower costs) by worker's age and skill level while seniority in the previous job has no significant effect. Age, skill level or seniority have no significant effect on the ability to claim. This is consistent with the results of the duration model where most of the parameters affecting the registration hazard turn out to be insignificant when we control for unobserved heterogeneity. This confirms that the difficulties to claim, if any, are related to unobserved heterogeneity rather than to the observed characteristics. This doesn't mean though that the take-up decision is unrelated to the observed characteristics of the worker: both the utilities and the search costs are affected by past wage, age and skill level, and they eventually affect the incentives to claim. The claiming efforts, and thus the claiming rate γ, do depend on worker characteristics, but indirectly via the fact that they impact the value of unemployment. This is why it is important to control for job search when trying to explain take-up behaviors and how it varies among workers.

Table 3: Parameter Estimates.

	Est.	s.d.
Parameters of the utility function		
a_N	86.25	(0.00)
a_P	56.41	(35.44)
b_N	0.34	(0.00)
Unobserved heterogeneity		
$\bar{\omega}_\lambda$	−4.32	(1.14)
$\bar{\omega}_\gamma$	−8.94	(2.95)
$p_1 = Pr(\omega_\gamma = \bar{\omega}_\gamma, \omega_\lambda = \bar{\omega}_\lambda)$	0.11	(0.17)
$p_2 = Pr(\omega_\gamma = 0, \omega_\lambda = \bar{\omega}_\lambda)$	0.40	(0.07)
$p_3 = Pr(\omega_\gamma = \bar{\omega}_\gamma, \omega_\lambda = 0)$	0.18	(0.08)
Search cost : observed heterogeneity (η_λ)		
Age (log)	−0.73	(0.15)
Skilled	−1.65	(0.49)
Previous job duration (in months)	−0.03	(0.03)
Intercept	9.39	(0.95)
Claiming cost : observed heterogeneity (η_γ)		
Age (log)	1.42	(1.59)
Skilled	0.15	(1.51)
Previous job duration (in months)	−0.04	(0.04)
Intercept	5.48	(1.81)
Wage distribution and error term		
μ_F	0.93	(0.00)
σ_F	1.55	(0.00)
σ_ν	0.29	(0.17)

Sample: 30-50 year-old males eligible for the longest UI compensation who experience a job loss between 07/2001 and 12/2002. Note: standard errors are calculated using the delta method. Source: FH-DADS.

Remember that in the reduced-form model the unobserved terms are applied to the hazard rates, while in the structural model, they are attached to the costs functions. Here, finding a job at low cost corresponds to $\omega_\lambda < 0$ and this implies high exit rates. In the same way, facing low claiming frictions and thus high take-up rates equates to having $\omega_\gamma < 0$. In terms of unobserved heterogeneity, workers' abilities vary substantially. Some job seekers collect benefits at low costs but have no advantages in terms of job search, while others have no particular ability in claiming but benefit from low search costs.

Consider the first group, that is workers with low search costs *and* low claiming costs[23].

[23] These individuals have $\omega_\gamma = \overline{\omega}_\gamma < 0$ and $\omega_\lambda = \overline{\omega}_\lambda < 0$.

As with the reduced-form estimation, we find that the share of workers belonging to this group is not significantly different from zero. Still, this group serves as an interesting point of comparison. *A priori*, the total effect on the take-up rate of this combination of high search and claiming abilities can go both ways. Our model allows us to compute the average take-up rate for each group and, in the case of the first group, the expected claiming rate is 74.4%, much higher than the population average, which means that the lower claiming costs offset the high ability to search for a job. The second group has lower search costs ($\overline{\omega}_\lambda < 0$) but no particular advantage in claiming ($\omega_\gamma = 0$). Therefore, they tend to find a job quickly while needing time to register. Their take-up rate is estimated to be very low on average, at 1%, as they find a job quickly before completing the claiming process. This group is the biggest one, accounting for forty percent of the sample. This is consistent with our data, where a lot of workers exit unemployment quickly without claiming, and with the reduced-form estimates. Individuals belonging to the third group (18% of the sample) have the opposite pattern of heterogeneity: they collect benefits at a relatively low cost ($\overline{\omega}_\gamma < 0$) but they have no particular ability to search for a job ($\omega_\lambda = 0$). At the end, their claiming intensity and thus take-up rate are likely to be very high, and we indeed get a take-up rate of 98.8%. Finally, the group with both high search costs and high claiming costs ($\omega_\gamma = \omega_\lambda = 0$ for 31% of our population) has a take-up rate of 11.4%. Although they need time to register, their exit rate to employment is low enough such that in some case we observe take-up.

As before, unobserved heterogeneity plays a very important role and, again, estimates show that it is the sole statistically significant source of heterogeneity for the claiming cost. For the job search cost, a variance decomposition of log-η_λ indicates that 91% of its variance comes from unobserved heterogeneity. This is not surprising if one considers that we already control for past-wage in the utility function and if one remembers that the value of unemployment in N and P impacts the search effort and the claiming intensity, and thus the probability to be a claimant and the exit rate to employment. The past wage is probably already capturing a good share of the individual observed heterogeneity.

4.3 Fit

In the following subsections, we provide a number of counterfactual exercises. Before, we check how we fit some important features of the data. For that purpose, we simulate a large number of individuals entering unemployment. For each of the 15,453 workers in the sample, we draw 100 times in the unobserved heterogeneity distribution and thus simulate 100 trajectories with the same observed characteristics but different $\{\omega_\gamma, \omega_\lambda\}$. We thus eventually simulate $15,453 \times 100$ individual labor market histories with a similar observation window as in the data (72 weeks). We compare in Table 4 the data moments and the simulated moments.

Table 4: Fit of the data moments.

	Real data	Simulated data
Duration moments (in weeks, T_j duration in state j)		
$\mathbb{E}(T_N)$	22.2 (0.22)	23.4
$\mathbb{E}(T_P)$	50.3 (0.35)	49.3
$\mathbb{E}(T_N)$ given $N \to P$	6.7 (0.16)	7.8
Transition moments (in %)		
Take-up rate	30.4 (0.37)	29.7%
Conditional on collecting		
Share of $N \to P \leq 2$ weeks	55.5 (0.72)	40.3
Share of $N \to P \leq 4$ weeks	66.7 (0.69)	61.7
Share of $N \to P \leq 8$ weeks	77.7 (0.61)	80.1
Share of $N \to P \leq 12$ weeks	85.0 (0.52)	86.6
Share of censoring in P	62.1 (0.37)	66.6
Among all workers		
Share of $N \to E$ transitions	52.7 (0.40)	51.2
Share of $N \to E \leq 2$ weeks	7.7 (0.21)	9.2
Share of $N \to E \leq 4$ weeks	15.5 (0.29)	16.2
Share of $N \to E \leq 12$ weeks	32.1 (0.70)	32.2
Wage moments (log-wage)		
$\mathbb{E}(w^{new})$	5.83 (0.005)	5.85
$\mathbb{V}(w^{new})$	0.23 (0.004)	0.21
$\mathbb{E}(w^{new} \mid N \to E)$	5.84 (0.005)	5.82
$\mathbb{V}(w^{new} \mid N \to E)$	0.23 (0.004)	0.20

Note: We simulate given the estimated parameter values and a similar observation window as in the data. For each of the 15,453 workers in the sample, we draw 100 times from the unobserved heterogeneity distribution. We end up with 1,545,300 simulated labor market histories. The standard errors of the empirical moments are obtained by bootstrap.

The overall fit of the data is good. We match the take-up rate (we are within the confidence interval of the empirical rate) and the model delivers a very good fit of the duration moments (durations in N and P), only overestimating by a week the expected duration without benefits. Our data have two important features: a good share of workers exit directly towards employment in the very first weeks ($N \rightarrow E$ transitions in the table) and, for those who receive the benefits, claiming usually happens within the first unemployment weeks ($N \rightarrow P$ transitions). The model is able to replicate these patterns. Conditional on collecting, the overall match of the durations is good: the model underestimates the share of individuals who claim within the first two weeks but then offers a good fit of the distribution of durations. It also matches well the distribution of duration in N for those who find a job directly, only slightly over-predicting the share of workers finding a job the first two weeks. Finally, the model generates a little bit too much censoring when the workers are collecting the benefits (67% vs 62%).

To conclude, the estimated model is doing a good job in matching the most distinctive features of our data and the key moments. In the next two subsections, we use it to provide a number of counterfactual exercises. First, we look at the effect of claiming frictions on unemployment durations. We test whether a decrease in claiming frictions reduces or increases the average unemployment duration and we identify which workers are the most affected by such a change. Second, we look at the effect of unemployment benefits on durations when we account for the endogeneity of the take-up decision. For each exercise, we simulate the model for $1,545,300$ individuals, that is, for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but with a new draw in the unobserved heterogeneity distribution.

4.4 How bad are claiming frictions for the unemployment duration?

One could argue that claiming frictions may not be such an issue. They could be viewed as a supplementary incentive to exit unemployment quickly, operating a selection between workers who find a job easily and those who really need insurance because they face longer unemployment spells[24]. However, for this latter worker type, claiming frictions could have the sole effect of slowing down benefit collection without having any significant

[24] Kleven and Kopczuk (2011) theoretically analyze how the complexity in social programs can be used as a selection device, to screen applicants, creating incomplete take-up. Their analysis is related to the difficulty for the principal to monitor the eligibility of applicants.

effect on the exit rate from unemployment. In that case, it would eventually increase unemployment duration. In this section, we investigate this issue, considering a change in claiming frictions and looking at its effect on durations.

We simulate a change in the claiming process that *decreases* the claiming costs by ten percent. For that we first simulate at the estimated parameters' values and then resimulate for the same individuals but with 10% lower claiming costs[25]. Each time we compute the take-up rate, the durations in each of the unemployment states, as well as the distribution of heterogeneity. Table 5 reports the results translated in terms of elasticities.

Table 5: Elasticities of durations with respect to claiming costs (aggregate level)

Moments	Elasticities
Take-up rate	0.14
Total u. duration ($N + P$)	0.03
Duration in N	-0.03
Duration in P	-0.01
$\Delta(\eta_\lambda \| new)$	-0.88

Note: Results are translated in elasticities, that is the effect of a *one* percent decrease in the claiming costs. We simulate the model for 1, 545, 300 individuals and for four years of weekly labor market histories. That is, for each of the 15, 453 individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution. This is done twice: one time at the estimated parameter values, another time with a 10% decrease in the claiming costs. $\Delta(\eta_\lambda|new)$ denotes the difference in η_λ between the *new* claimants (workers simulated as not collecting before the change in claiming costs and that now collect benefits) and those who were getting benefits before.

The elasticity of the take-up rate is limited (0.14): a ten percent decrease in the costs translates to an aggregate 1.4% increase in the take-up rate[26]. Since the returns to claiming efforts are improved by the decrease in claiming frictions, workers claim more intensively and lower their search efforts in unclaimed unemployment, easing transitions from N to P. Finally, despite lower search intensities, the *average* duration in N falls slightly: a decrease in the cost by 10% lower durations in N by 0.3%.

[25] Given our parametric assumption, for any level of γ, a reduction of 10% percent of $c_\gamma(\gamma)$ corresponds to a decrease in η_γ by 10% percent.

[26] The existing literature shows that the take-up of social benefits reacts to transaction costs (for a quick survey see Currie, 2006 and the references in Kleven and Kopczuk, 2011) but there is little work on the unemployment insurance. A notable exception is Ebenstein and Stange (2010). Using aggregate American data, they find that the development of remote UI claiming (by phone or the Internet) has no statistically significant effects on take-up rates nor the characteristics of claimants. However, the authors recognize that they cannot rule out moderate effects.

Looking now at what happens in claimed unemployment (state P), the *average* duration also decreases but only by 0.1% if claiming costs decrease by 10%. As claiming costs do not affect the incentives to search while recipients, this change comes from a shift in the composition of claimants. Indeed, after the cost reduction, new workers[27] enter in P and they are of different types : the average search cost parameter (η_λ) is indeed estimated to be 8.8% lower after the change. The pool of job seekers collecting benefits thus improves, as predicted, but this does not affect much the *average* duration in P since the increase in the take-up is rather limited. To reconcile the effect on durations in each unemployment state to the total duration ($N+P$), it is worth remembering that there are now more workers in state P and, although they have better job search prospects, they reduce their search efforts when they receive unemployment benefits (compared to when they do not), so that their total duration increase, leading to longer total unemployment on average in the counterfactual scenario. However, overall the effects are limited since a decrease in the claiming costs by 10% only increases the average unemployment duration by 0.3%.

To get a better understanding of these results, it is useful to dig further into the heterogeneity of individual responses and to account properly for selection. For that purpose we analyse transitions out of unemployment by estimating proportional hazards models on our simulated data. Here duration is the addition of the times spent in claimed and unclaimed unemployment and the data are the ones used in Table 5. We control for observed and unobserved heterogeneity (with dummies for the heterogeneity groups) and include a dummy for the treatment (10% reduction in claiming costs). We also estimate the model using all individuals and then re-estimate the model separately for each heterogeneity group to get a sense of how heterogeneous the treatment effect is. We also estimate probit models on the same data for the take-up responses. Table 6 reports the effects of the cost reduction[28], detailed results being provided in Appendix E. Note that the results for the take-up probability are reported with covariates set to their mean values, while the results from the proportional hazard models are by construction independent of the covariates.

[27] When we talk about *new* workers, we are considering workers who, at the initial level of the claiming costs, were not observed as claimants and who are now observed as such in our simulations.

[28] Remember that the expected duration is the inverse of the hazard rate so if β is the effect on the hazard rate $-\beta$ is the effect on expected durations.

Table 6: Effects of a 10% *decrease* in claiming costs on take-up and unemployment durations.

Sample/Semi-elasticities of	TU probability	Non-emp. duration
All obs.	0.055	0.007
	(0.003)	(0.001)
Group 1	0.012	0.004
	(0.002)	(0.004)
Group 2	0.109	0.005
	(0.019)	(0.001)
Group 3	0.000	−0.000
	(0.000)	(0.006)
Group 4	0.078	0.019
	(0.006)	(0.004)

Note: Results from the estimation of probit models and proportional hazard models. We report semi-elasticities that is rate of change of the take-up probability (reported at the means of covariates) or of durations as the claiming costs decrease by 10%. The dataset is the merge of two simulated datasets, one where claiming costs are at their estimated values, another with a 10% decrease in the claiming cost. Each time, we have $1,545,300$ individual observations, that is for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution. Group 1: workers with low claiming and job search frictions; Group 2: workers with high claiming frictions but low job search frictions; Group 3: workers with high search frictions but low claiming frictions; Group 4 : workers with high claiming and search frictions.

First, the four categories of job seekers (in terms of unobserved heterogeneity) react very differently to a change in claiming costs. The 10% decrease rises the take-up probability by 5% for a job seeker with the average values of the covariates and increases the duration (for all workers) only slightly by 0.7% (we found 1.4% and 0.3% in terms of macro-elasticities[29]). These averages mask a lot of heterogeneity. Individuals from groups 2 and 4 are those whose claiming probabilities increase the most, by 10.9% and 7.8% respectively. Group 1 reactions are very limited (around 1%) and members of group 3 do not react at all. Remember that the groups 2 and 4 are the ones where claiming costs are the highest. Any alleviation of the claiming frictions has strong effect for them. That also means that any additional complexity in the claiming process hurt the most those who already have difficulties to collect benefits. For all the groups, the effects on durations are very limited. They are stronger for the last group, group for which both

[29]The numbers are different for a number of reasons. First the results of the probit are reported for an individual at the mean value of the covariates. Second, both rely on particular functional assumptions about the link between the dependent variables and the heterogeneity. Finally, the duration models account explicitly for censoring.

the search and claiming frictions are the highest, but even here they are very limited since a reduction of the claiming costs by 10%, a rather large decrease, only increases their unemployment duration by 1.9%. A the end, complexity in the claiming process do not appear as an efficient filtering instrument. More frictions in the claiming process would decrease substantially the take-up for those who already have hard time accessing unemployment insurance, lowering durations only by a very limited amount.

4.5 The elasticity of unemployment benefits on unemployment spells

We now consider an increase in the unemployment benefits by simulating the model for two different levels of the replacement rate b_P: one at its calibrated value and another one one-percent higher. Again, we simulate each time our model for 1545300 individuals, that is, for each of the 15453 individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw in the unobserved heterogeneity distribution. We then estimate a number of proportional hazard models on simulated durations controlling for observed heterogeneity (past wage, skill level, age, past employment duration), unobserved heterogeneity (dummy variables for the heterogeneity groups that we do observe in simulations) and including a dummy for the treatment (a higher level of benefits). In line with most of the existing empirical papers (e.g. Chetty, 2008, Landais, 2015 or Card et al., 2015), we first consider the claimants only[30]. We look at their exit from claimed unemployment (P) and total unemployment ($N + P$). Then we look at all eligible workers, even if they don't claim, and look at their exit from unemployment. We aim at showing that, when estimating the effects of a change in the insurance parameters, the imperfect take-up and the selection into claimed unemployment matter. The detailed results of the estimations are reported in Appendix E (Tables E.3 and E.4). Table 7 presents the main results, translated in terms of duration elasticities.

[30]We thus include all unemployment spells with positive duration in claimed unemployment. The elasticities are captured by the coefficient of the dummy variable.

Table 7: Impact of the unemployment benefits on durations (elasticities)

Moments	Elasticities
Claimants	
Duration in P	1.81***
Total duration $(N + P)$	0.62***
All individuals	
Take-up rate	1.30***
Total duration $(N + P)$ (control for ω)	2.30***
Total duration $(N + P)$ (no control for ω)	1.13***
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$	

Note: We simulate the model for $1,545,300$ individuals and for four years of weekly labor market histories. That is, for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution. We display two sets of results: one for the individuals observed as claiming, one for all workers including those who exit without collecting benefits.

We have previously shown that, at the individual level, the elasticity of the total unemployment duration can be decomposed into the take-up elasticity, the unclaimed-duration elasticity and the claimed-duration elasticity (see equation 4). First, when the benefits are rising, the individuals lower their efforts to find a job, even in unclaimed unemployment, but they exert more efforts to collect benefits (see equations (1) and (3)). As a consequence, the take-up probability increases. Here, for a one-percent increase in benefits we find that the take-up rate increases by 1.3%.

Second, the elasticity of the claimed unemployment duration is positive, since higher benefits decrease the incentives to search for a job. Controlling for unobserved heterogeneity we find an elasticity of 1.81. It is at the high end of the range of the usual estimated values (see Card et al., 2015), but we consider a particular sample here (male workers with relatively high experience on the labor market). Moreover, looking now at the effect on the total duration for the claimants, we find a much lower elasticity of 0.62, more in line with the median estimates from the literature. For the claimants, the increase in the duration spent on benefits is thus balanced by shorter durations in unpaid unemployment. This demonstrates the importance of looking at total unemployment durations, even when one considers only those who collect benefits.

Finally, if one looks now at both claimants *and* non-claimants, the elasticity of the *total* unemployment duration reaches 2.30. The number is very different because we consider now the whole population and not a selected sample of claimants. This shows how both populations react to the change in the insurance generosity and how looking at the sole claimants (especially if only in claimed unemployment) can produce an underestimation of the impact of that change. Notice that if we don't control for unobserved heterogeneity we get a much lower elasticity at 1.13. The difference between the two estimates is informative. Indeed, when we don't control for unobserved heterogeneity, the estimate of the treatment effect is biased by the selection into *claimed* unemployment[31]. Here, when the replacement rate increases, the pool of workers receiving benefits changes: in our simulated data, job seekers in claimed unemployment are now more efficient, both in terms of job search and claiming, as ($\mathbb{E}(\eta_\lambda|P)$ and $\mathbb{E}(\eta_\gamma|P)$ decrease by 1.45 and 4.9 percent respectively). This simple exercise shows that accounting for the endogeneity of the take-up and distinguishing between claimed and unclaimed unemployment durations is crucial to estimate the effect of a change in the unemployment insurance parameters[32].

5 Conclusion

The imperfect take-up of the unemployment insurance is an intriguing phenomenon still understudied in the economic literature. This paper argues that, to get a better understanding of the non take-up, one needs to consider jointly the claiming efforts and the job search activity and to model claiming in a way that allows for temporary non-takeup. In that case, it becomes apparent that, when claiming takes time, part of the imperfect take-up is driven by fast exit from unemployment.

Moreover, when individuals are heterogeneous in terms of job search and claiming abilities, the cost of claiming frictions is concentrated on those with both high search and claiming frictions. Our estimations on French administrative data show that 10% of the individuals in our sample are in this category. For that reason, reducing claiming frictions would increase take-up with very little effects on the aggregate unemployment durations. It would relax the constraint for job seekers who have anyway low exit rates from unemployment.

This paper ends with a more methodological implication. When measuring the elas-

[31] If there wasn't such a selection, that is if the take-up was immediate for all, the estimated treatment effect even without control for unobserved heterogeneity would be unbiased, conditional on the fact that the observed and unobserved characteristics are orthogonal, which is the usual assumption.

[32] As such, this result is reminiscent of Heckman et al. (1999) or Meyer and Mok (2007) who discuss how estimates can be biased when the endogeneity of take-up is not accounted for.

ticity of unemployment duration to the unemployment benefits, accounting for take-up and distinguishing between claimed and unclaimed unemployment durations is important. Indeed, a change in the UI parameters impacts all the eligible job seekers even if only a fraction collects benefits eventually and the elasticity substantially differs whether one considers the claimants only or the whole population of the eligible, the duration in claimed unemployment or the total unemployment duration.

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A Eligibility rules

A.1 Eligibility criteria

To benefit from the unemployment insurance, an individual must satisfy certain requirements and fill a claim within the year following job separation. They are two sets of eligibility requirements. First, the worker needs to be previously employed in the private sector[33]. Second, the individual must have worked a minimum number of days or hours in a base period (in the 12, 24 or 36 months before the start of the spell of unemployment). Notice that individuals who are fired for cause are disqualified for the receipt of benefits. Job quitters can receive benefits, but only after an extended waiting period of four months and conditional on an administrative decision that considers their job search activities in the past four months. There are in principle job search criteria but in practice, during that period of time, they were not enforced and the sanction rate was almost null.

A.2 Entitlement

Unemployment benefits are paid monthly for a limited period of time. Age and previous work experience determine the potential benefit duration. Table A.1 displays the UI duration as a function of the number of months worked in the past months.

Table A.1: Potential benefit duration (January 2001 - December 2002).

	Work experience	**Age**	**Benefit duration**
1	4 months in the past 18 months	-	4 months
2	6 months in the past 12 months	-	7 months
3	8 months in the past 12 months	$<$ 50 years old	15 months
	$\geq$ 50 years old	21 months	
4	14 months in the past 24 months	$<$ 50 years old	30 months
	$\geq$ 50 years old	45 months	
5	27 months in the past 36 months	between 50 and 54 years old	45 months
	$\geq$ 54 years old	60 months	

Source: Unedic.

The amount of compensation depends on past wages. It is calculated on the basis of a daily benefit and a "reference wage", which is the average of monthly wages received during the last 12 months. In case of part-time work, the amount of benefits is propor-

[33]Corporate executives and independent workers do not benefit from the unemployment insurance.

tional to part-time hours relative to full-time hours. There exists a non linear function linking the reference wage to benefits paid. The amount of benefits is bounded, its daily amount cannot go below 25.01 euro or above max($10.25euro + 0.4 \times$ reference wage; $0.75 \times$ reference wage). Generally, the net replacement rate is 75% (OECD, 2014). Once the unemployed worker has exhausted her benefits eligibility, she enters the welfare system.

A.3 Claiming process and timing in benefit collection

Even when the worker is eligible, the receipt of benefits is not automatic. From 2001 to 2009, the unemployed worker must complete several administrative steps with various intermediaries at the local and national levels. At that time, the UI system had two main entities: one for the job search support (ANPE) and the other one for the paiement of benefits (Unedic-Assedic). To claim the support to which he is entitled, the individual first had to contact his local Assedic agency or go to the city hall to complete a 8-pages form, wherein he reported precisely his situation[34]. The Assedic office examined the claim and calculated the benefit entitlement. Eventually, the claimant had to show up at the local employment agency (ANPE) within the first month after filing his claim. Hence, to make successfully a claim, a worker had to understand and follow different administrative steps. The coexistence of two distinct institutions, one providing benefits, the other in charge of counseling, added to the complexity of the claiming process.

For all workers, there is a minimum waiting period of 7 days. For some workers, those who receive severance pay above the legal minimum or who had been paid for outstanding vacation days, the waiting period can be longer but always below 75 days. Note that the 7-days waiting period is imputed automatically starting from the completion of the claiming process. The other waiting days, if any, start from the date of job loss. Consider an individual who has a waiting period of 60 days. If he registers the day of job separation or within 60 days, then the benefit is first paid on the 67th day after the job loss. If he successfully registered after 70 days of unemployment, he is paid on the 77th day if still unemployed. The timing of the registration has no impact on the potential entitlement duration and there is no retrospective payment: if the worker of our example finds a job after 73 days, he does not collect any benefits.

[34]The form had to be returned completed and signed, with different supporting documents necessary to determine entitlement. The registration became effective the day the Assedic received all the needed documents.

B The data

The FH-DADS data combine two administrative datasets. The first one, from the national employment agency, records the registered unemployment spells of individuals (*Fichier Historique Statistique*, FH hereafter). The second one is an employer-employee administrative panel which describes the employment spells any worker had in the private and semi-public sectors[35] (*Déclarations Annuelles de Données Sociales*, DADS hereafter). The original datasets are 1/24 nationally representative samples[36]. The merge of these datasets includes any individual who appears in one or another of these records between January 1st, 1999 and December 31st, 2004. Both data are longitudinal and give daily information on workers back to January 1st, 1976, on registered unemployment history back to January 1st, 1993 and on insured unemployment back to January 1st, 1999.

For any job, we observe the starting and ending dates, the wages, the number of working days and the employer's sector. We can thus predict the worker's entitlement to the unemployment insurance (see below). Besides, for any registration at the public employment agency, we know the dates and reasons of registration and deregistration. For workers who receive unemployment benefits, we observe the amounts and dates of first and last payments and we know the potential benefit duration. Putting these elements together, we observe for all the non-employment spells, the unemployment-to-employment transitions, the take-up decisions, and the compensated and uncompensated unemployment durations.

B.1 Sample selection

We address the question of eligibility in the next subsections. For a moment, we put the problem aside and we describe the sample selection. We sample, from the outflow from employment, all *eligible* non-employment spells which start between July, 1st 2001 and December, 30th 2002. During that period of time, the parameters of the unemployment insurance system were stable. We exclude all the non-employment spells shorter than a week. Maternity leaves and sickness absences do not involve an exit from employment in the dataset, such that these non working periods cannot be wrongfully counted as

[35]These sectors cover the firms affiliated to the unemployment insurance system are in the DADS sample.

[36]Workers born on October of an even-numbered year are sampled.

uninsured non-employment. However, we cannot distinguish between unemployment and inactivity due to schooling, retirement or family care. For these reasons, we restrict the sample to males aged between 30 and 50 on July 2001. Additionally, we exclude workers who exit a job from the semi-public sector because they have complex and very specific eligibility rules. They account for 6.3% of the outflows from employment in 2001 and 2002.

The unemployment insurance allows individuals to have a paid work activity for a short period of time while remaining registered[37]. We follow the rule and consider these short employment periods as part of the unemployment spell and more precisely, as part of the P state.

We obtain a sample of 18,034 spells. As entitlement prediction has proved to be better for individuals entitled to the longest compensation duration (type 4), we perform the analysis on the subsample of non-employment spell giving right to this type of compensation. This subsample contains 15,453 spells.

B.2 Definition of the unemployment states

We assume that the worker enters in state N (called unpaid unemployment) as soon as the non-employment spell starts, that it at job separation. The worker leaves state N to employment (state E) or to claimed unemployment (state P).

Our measure of take-up (a switch from N to P) is the opening of a so-called compensation period. This does not mean that the individual receives unemployment benefits right away. This means that he completed the registration process and that he is flagged has eligible[38].

Entry into employment occurs when there is a reemployment associated with the closing of the compensation period. We want to consider "sustainable" reemployment and do not consider as entry into E a reemployment that do not lead to a deregistration. These are very short employment spells coming from the right to receive benefits and working a few hours per week.

[37]This is the so-called"reduced activity" scheme. Periods of reduced activity are declared as such in the records of the employment agency. Because of possible under-reporting of reduced activity periods by individuals to the unemployment insurance administration, we also consider as reduced activity periods employment spells observed in the DADS panel that start and end during a registration period.

[38]In some rare cases, the process is completed before the actual end of the employment period. Our model cannot accommodate directly for these observations. In that case, we consider that the individual only spends one week in state N and then switches to state P.

The individual is right-censored in a given state if he occupies this state 72 weeks after job separation. For each non-employment period, we calculate the weekly average of the net wages received the year before the spell starts and the year after it ends (if any job observed). Wage information is expressed in constant euros (base 1998). We trim the wage distributions and delete the first and last percentiles for the preceding and reemployment wages.

B.3 Eligibility status and potential entitlement

Both eligibility and the maximum length of compensation depend on the number of days the worker has worked in the covered sectors in the past months[39]. The exact criterion for our period of interest are displayed in Table A.1. Using the employer-employee matched data, which trace the sequence of jobs since the first entry into the labor market, we sum the number of working days in the covered sectors from the beginning of the reference period to the date of job separation. As a single employment period cannot be used twice for compensation, the reference period is truncated at the closing date of the previous compensation period, if any.

Our data have a number of limitations. First, some individuals may have remaining rights from previous registration periods which we cannot account for. In case they have not worked enough days to renew their eligibility, we would count them as ineligible. Because they are some extra rules that can be applied to become eligible if the usual criterion are not satisfied, some eligibles are dropped from the analysis and we tend to underestimate the maximum length of the benefit entitlement. For the workers who claim and receive unemployment benefits we can correct our predictions if necessary.

Second, as already mentioned, our dataset does not cover all the employed workers. For that reason, if a worker enters in unemployment from a covered sector but exit to, for example, the public sector before receiving benefits, he will be tagged as not claiming and being still unemployed although he has left unemployment and is no longer eligible. However, we are confident that this is not quantitatively such an issue for our analysis. First, the DADS data cover about 82% of total employment (source: LFS, 2002). Most importantly for our analysis, using the French Labor Force Survey, we can measure the

[39]Entitlement criteria applying to the claimants are those in place at the time of job separation, and not at the time of registration.

occurrence of such a problem. Consider the individuals who declare at the first interview to be unemployed after having worked in the sectors covered by the DADS data, and those who declare to be working in these covered sectors and who experience a job-to-job or an unemployment-to-job transition in the following year. Only 2.9% of them are observed taking up a job which is outside the range of the DADS panel.

Finally, we do not observe the reasons for separation. This may be a problem for workers who voluntary quit. However, notice that even a worker who resigns can pretend to benefits after four months of unemployment. Moreover, using the 2002 LFS survey, we can see that only 1.8% of 30-50 years-old males (our sample of analysis) declare that the job ended because they quit.

We use past employment durations to determine eligibility following the rules displayed on Table A.1. As mentioned before, we may underestimate eligibility because some individuals have remaining rights from previous registration periods. It is possible to see this if we consider those who claim in our data. For that subsample, we do observe the eligibility and Table B.1 displays the fit of our initial entitlement predictions to the actual entitlement.

Misclassification is mostly due to residuals rights from previous compensation periods[40] and 16.2% individuals are wrongly estimated as ineligible. However, the fit is especially good for workers entitled to the longest benefit duration: 90.83% of the workers who are considered as entitled to the longest benefit duration (30 months) are indeed entitled to this duration.

Moreover, for those who register, we correct the predicted entitlement status based on past employment duration if needed. However, for those who are not observed as claiming, we rely on our prediction. Our final estimation sample only includes workers eligible (after correction) to the longest duration.

[40]The rules used to compute residuals rights are too complicated to be used given the information in our dataset.

Table B.1: Predicted versus observed entitlement for claimants (in %).

	Actual entitlement				Total
	4 months	7 months	15 months	30 months	
Predicted entitlement					
Not eligible	23.09	12.10	12.26	52.55	100
4 months	50.66	23.58	10.92	14.85	100
7 months	18.82	44.09	19.35	17.74	100
15 months	9.45	10.42	46.91	33.22	100
30 months	2.02	2.73	4.42	90.83	100
Total	7.23	6.36	8.32	78.10	100

Sample: 30-50 year-old males who experience a job loss between 07/2001 and 12/2002 and are compensated during their unemployment spell.
Note: 90.83% of compensated workers who are *predicted* as entitled to 30 months of benefits are indeed eligible to 30 months
Source: FH-DADS.

C Reduced-form analysis

Non parametric analysis. To describe the distribution of the duration in unregistered unemployment (state N), we consider a competing risk framework. Unregistered unemployment can end for two reasons: successful claim (entry in P) or reemployment (entry in E). The occurrence of reemployment precludes the occurrence of registration. The two competing events, entry in P and entry in E, cannot be considered as independent, as they are determined by common factors. In such a setting, the appropriate characterization of the duration distribution is given by the cumulative incidence functions CIF_k of cause $k = P, E$[41], defined as

$$CIF_k(t) = P(T \leq t, D = k) = \int_0^t h_k(s) \exp\Big(- \int_0^s h_P(u)du - \int_0^s h_E(u)du \Big) ds$$

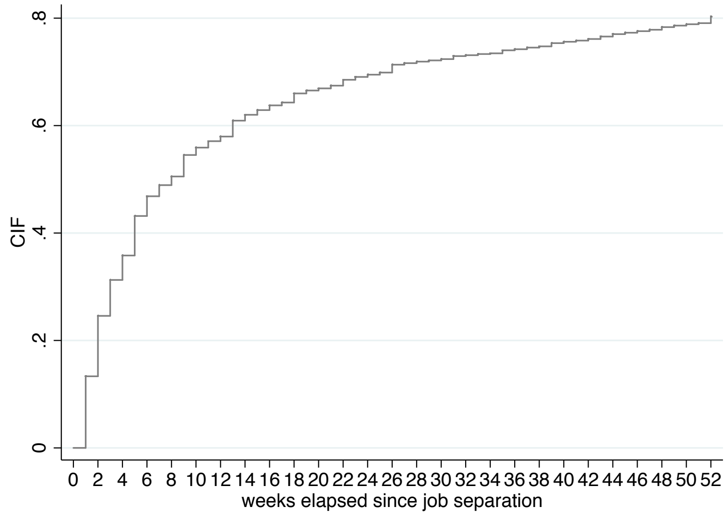
where $h_k(.)$ is the k-specific hazard (the instantaneous probability of exiting to state k given that exit to k has yet to happen).

The overall cumulated hazard in N plotted in Figure C.1 shows that the probability of exiting N within a month, for reemployment or completion of the claiming process, is about 35%. Figure C.2 plots the risk-specific cumulative incidence functions $CI_P(t)$

[41]The Kaplan-Meier estimate indeed assumes independence of the competing events and overestimates the true failure probability. Basically, workers who found a job before getting registered, will never register. The Kaplan-Meier estimate considers the observations of these individuals as censored and treat them as if they could fail. As a result, it overestimates the probability of failure and the corresponding cumulative hazard. The cumulative incidence function explicitly accounts for the competing risks and allows each risk-specific duration distribution to depend on the duration distribution of competing risks.

and $CI_E(t)$. It shows that within 6 weeks the probability of exiting to P is greater than the probability of exiting to E. For instance, the probability of completing the claiming process by 4 weeks is 20% and the probability of finding a job within the first 4 weeks after separation is 16%. After 6 weeks, the risk of reemployment becomes greater than the risk of registration.

Figure C.1: Overall cumulative incidence function of the duration in N (in weeks)

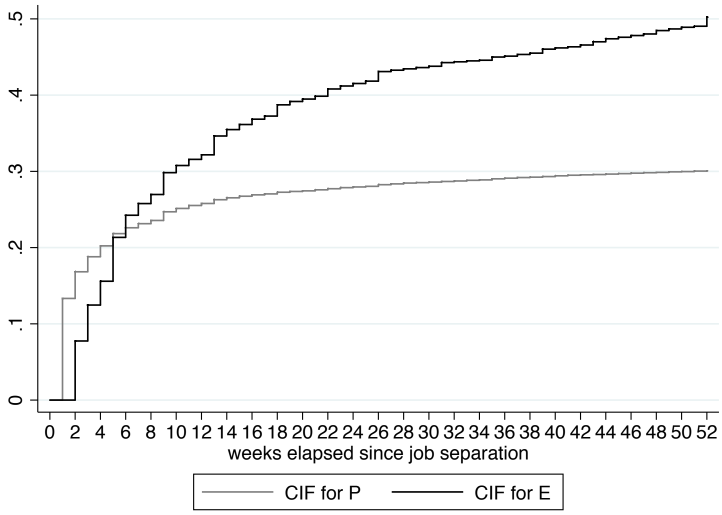


Sample: 30-50 year-old males who experience a job loss between 07/2001 and 12/2002, are eligible for UI at job separation and are entitled to the longest benefit duration.
Lecture: the curve CI is the complement of the Kaplan-Meier estimate of the duration in N, considering both reemployment and registration as exits. Source: FH-DADS.

A multivariate duration model. We estimate a bivariate duration model with unobserved heterogeneity to analyze the determinants of the duration to registration and those of the unemployment duration. This multivariate mixed proportional hazard model (Lancaster (1999), van den Berg (2001)) permits to account for dependence of the two duration processes: there may be common unobserved component jointly explaining both processes. For instance, recent working periods affect entitlement, that are incentives to register, as well as employability, *i.e.* the outcome of job search.

Note T_N the duration in unclaimed unemployment, that is the time between job separation and the entry into state P. It is censored if the worker is still in state N at the end of the observation period (after 72 weeks) or if he gets a job before the completion of the claiming process. The censoring for reemployment is explicitly taken as endogenous

Figure C.2: Risk-specific cumulative incidence functions of the duration in N(in weeks)



Sample: 30-50 year-old males who experience a job loss between 07/2001 and 12/2002, are eligible for UI at job separation and are entitled to the longest benefit duration.
Lecture: the curve CIF for P is the cumulative incidence function for registration considering reemployment as a competing risk. The curve CIF function for E is the cumulative incidence function for reemployment from N, considering registration as a competing risk.
Note: This corresponds to our estimation sample, but the picture is similar for all the eligibles.
Source: FH-DADS.

thanks to the joint model of the two durations.

We note T_U the total duration in unemployment (claimed or unclaimed). The spell ends when the individual gets a job or is censored if the individual is still unemployed (either in state N or in state P) after 72 weeks of unemployment. Completion of the claiming process does not censor the duration to reemployment. As the structural model assumes that the search technology is the same whichever the type of unemployment, we assume here as well that the realization of T_P does not affect the shape of the hazard of T_U.

We consider the joint distribution of T_P and T_U. Dependence is introduced via the unobserved components ω_P, ω_E. Conditional on some observed characteristics X and the two specific unobserved heterogeneity components, T_P and T_U are assumed independent:

$$T_P \mid X, \omega_P \perp T_U \mid X, \omega_E$$

The individual likelihood function writes

$$L(t_P, t_U \mid X, \omega_P, \omega_E) = \sum_{m}^{M} P_m f(t_P \mid X, \omega_P^m) f(t_U \mid X, \omega_E^m)$$

with

$$\begin{aligned} f(t_P \mid X, \omega_P^m) &= h_P(t_P \mid X, \omega_P^m)^{c_P} S(t_P \mid X, \omega_P^m) \\ f(t_U \mid X, \omega_E^m) &= h_U(t_U \mid X, \omega_E^m)^{c_E} S(t_U \mid X, \omega_E^m) \end{aligned}$$

with $c_P = 1$ if the individual gets registered and 0 otherwise, and $c_E = 1$ if he gets a job and 0 otherwise.

We consider here constant hazards. We take a discrete distribution for the unobserved heterogeneity with $M = 4$ types:

$$\begin{array}{ll} P_1 = Pr(\omega_P = \omega_P^1 = \omega_P^b, \omega_E = \omega_E^1 = \omega_E^b) & p_P = Pr(\omega_P = \omega_P^2 = 0, \omega_E = \omega_E^2 = \omega_E^b) \\ P_3 = Pr(\omega_P = \omega_P^3 = \omega_P^b, \omega_E = \omega_E^3 = 0) & p_P = Pr(\omega_P = \omega_P^4 = 0, \omega_E = \omega_E^4 = 0) \end{array}$$

with $0 < P_m < 1$ and $\sum_m P_m = 1$. P_1, P_2, P_3, ω_P^b and ω_E^b are estimated.

D The likelihood function of the structural model

Consider a worker i, with $i \in \{1, \ldots, I\}$, of the unobserved type m, with $m \in \{1, \ldots, 4\}$, and with the observed attributes X_i. Let t_{Ni} and t_{Pi} denote the time spent by the worker in state N and state P respectively. $t_{Pi} = 0$ if she does not enter state P.

Let us first look at the events possibly occurring in state N and remark that our model implies constant hazard rates. The individual receives offers at rate λ^*_{Nim} and a job offer W is accepted if it is above reservation wage R^*_{Nim}. λ^*_{Nim} and R^*_{Nim} correspond to the individual's optimal choices and are functions of X_i and m. We denote θ_{NEi} her exit rate from state N to employment E, $\theta_{NEi} = \lambda^*_{Nim}\bar{F}(R^*_{Nim})$. Similarly, she completes the claiming process and enters state P at an endogenous rate γ^*_{im}, so that the hazard to claimed unemployment is $\theta_{NPi} = \gamma^*_{im}$. The worker survives in state N for t_{Ni} weeks if she does not get an acceptable job offer and does not complete the claiming process within her first t_{Ni} weeks of unemployment. The survival in N for t_{Ni} periods is then given by

$$S_N\left(t_{Ni}|X_i, \omega_{Pm}, \omega_{Em}\right) = e^{-(\theta_{NEi}+\theta_{NPi})t_{Ni}} = e^{-\left(\lambda^*_{Nim}\bar{F}(R^*_{Nim})+\gamma^*_{im}\right)t_{Ni}}$$

Now let us consider the time-events in state P. In P, an individual is only facing a risk of employment and exits claimed unemployment at an endogenous rate $\theta_{PEi} = \lambda^*_{Pim}\bar{F}(R^*_{Pim})$. Hence, the survival probability in P for t_{Pi} periods is given by

$$S_P(t_{Pi}|X_i, \omega_{Pm}, \omega_{Em}) = e^{-\theta_{PEi}t_{Pi}} = e^{-\lambda^*_{Pim}\bar{F}(R^*_{Pim})t_{Pi}}$$

A job offer W is a draw from the wage distribution $F(.)$. In the absence of measurement error, the probability of getting a wage W is simply

$$f(W|W \geq R^*_{jim}) = \frac{f(W)}{\bar{F}(R^*_{jim})} \text{ with } j \in \{N, P\}$$

We introduce measurement errors on wages so that we only observe $W^o = W \times \varepsilon$, with $\varepsilon \sim_{i.i.d.} \log\mathcal{N}(0, \sigma_\nu)$. We note $h(.)$ the probability density function of ε. The measurement error can be written as a function of the observed wage and the true underlying wage:

$\varepsilon = \frac{W^o}{W}$ with W ranging from R^*_{jim} to W_{max}. Hence, the probability of observing the reemployment wage W^o is

$$\frac{1}{\bar{F}(R^*_{jim})} \int_0^{W^o/R^*_{jim}} f\left(\frac{W^o}{\varepsilon}\right) h(\varepsilon) d\varepsilon$$

Using these elements, we can derive the individual contributions to the likelihood for the four types of time-events we may observe in the data. First, the individual contribution to the likelihood for a worker i of type m who is censored in N after t_{Ni} weeks of unemployment reads:

$$\ell_{im1}(t_{Ni}, t_{Pi}, W^o | X_i, \omega_{Pm}, \omega_{Em}) = S_N(t_{Ni} \mid X_i, \omega_{Pm}, \omega_{Em}) = e^{-\left(\lambda^*_{Nim}\bar{F}(R^*_{Nim})+\gamma^*_{im}\right)t_{Ni}}$$

Second, the individual contribution to the likelihood for a worker who gets back to employment directly from state N after t_{Ni} weeks of unemployment and is observed getting a job paid W^o reads:

$$\begin{aligned}\ell_{im2}(t_{Ni}, t_{Pi}, W^o \mid X_i, \omega_{Pm}, \omega_{Em}) &= \theta_{NEi} S_N(t_{Ni}|X_i, \omega_{Pm}, \omega_{Em}) \frac{1}{\bar{F}(R^*_{Nim})} \int_0^{W^o/R^*_{Nim}} h(\varepsilon) f(\frac{W^o}{\varepsilon}) d\varepsilon \\ &= \lambda^*_{Nim} e^{-\left(\lambda^*_{Nim}\bar{F}(R^*_{Nim})+\gamma^*\right)t_{Ni}} \int_0^{W^o/R^*_{Nim}} h(\varepsilon) f(\frac{W^o}{\varepsilon}) d\varepsilon\end{aligned}$$

Third, a worker who enters state P after t_{Ni} weeks and who gets censored after having spent t_{Pi} weeks in state P contributes to the likelihood with:

$$\begin{aligned}\ell_{im3}(t_{Ni}, t_{Pi}, W^o \mid X_i, \omega_{Pm}, \omega_{Em}) &= \theta_{NPi} S_N(t_{Ni}|X_i, \omega_{Pm}, \omega_{Em}) S_P(t_{Pi}|X_i, \omega_{Pm}, \omega_{Em}) \\ &= \gamma^*_{im} e^{-\left(\lambda^*_{Nim}\bar{F}(R^*_{Nim})+\gamma^*\right)t_{Ni}} e^{-\lambda^*_{Pim}\bar{F}(R^*_{Pim})t_{Pi}}\end{aligned}$$

Last, consider a worker who enters state P after t_{Ni} weeks and who gets a job paid with the observed wage W^o after having spent t_{Pi} weeks in state P. Her contribution to the likelihood reads

$$\ell_{im4}(t_{Ni}, t_{Pi}, W^o | X_i, \omega_{Pm}, \omega_{Em}) = \theta_{NPi} S_N(t_{Ni} | X_i, \omega_{Pm}, \omega_{Em}) \times$$

$$\theta_{PEi} S_P(t_{Pi} \mid X_i, \omega_{Pm}, \omega_{Em}) \frac{1}{\bar{F}(R^*_{Pim})} \int_0^{W^o/R^*_{Pim}} h(\varepsilon) f(\frac{W^o}{\varepsilon}) d\varepsilon$$

$$= \gamma^*_{im} e^{-(\lambda^*_{Nim} \bar{F}(R^*_{Nim}) + \gamma^*) t_{Ni}} \lambda^*_{Pim} e^{-\lambda^*_{Pim} \bar{F}(R^*_{Pim}) t_{Pi}} \int_0^{W^o/R^*_{Pim}} h(\varepsilon) f(\frac{W^o}{\varepsilon}) d\varepsilon$$

The likelihood function is the product of the individual contributions. When we sum over the distribution of the unobserved heterogeneity, we obtain the following likelihood function:

$$\mathcal{L} = \prod_{i=1}^{I} \sum_{m=1}^{4} p_m \times \ell_{im1}^{(1-\delta_{NEi})(1-\delta_{P_i})} \times \ell_{im2}^{\delta_{NEi}} \times \ell_{im3}^{\delta_{Pi}(1-\delta_{PEi})} \times \ell_{im4}^{\delta_{Pi}\delta_{PEi}}$$

where p_m is the probability of being of the unobserved type m, δ_{NEi}, δ_{Pi} and δ_{PEi} take value 1 if the worker exits to employment when in state N, if the worker exits to claimed unemployment and if the worker exits to employment from state P, respectively.

E Additional Tables

E.1 Claiming cost reduction

Table E.1: Effects of a claiming cost decrease by 10% on the take-up probability (probit).

	(1) All	(2) Group 1	(3) Group 2	(4) Group 3	(5) Group 4
Lower costs	0.0395***	0.0315***	0.0399***	0.0200*	0.0462***
	(0.00246)	(0.00491)	(0.00684)	(0.0103)	(0.00334)
log(past wage)	-2.424***	-11.77***	-5.880***	-5.738***	1.680***
	(0.0388)	(0.123)	(0.0881)	(0.297)	(0.0585)
log(past wage) sq.	0.252***	1.064***	0.552***	0.602***	-0.102***
	(0.00325)	(0.0107)	(0.00725)	(0.0272)	(0.00487)
Age	-0.838***	-1.115***	-0.709***	-0.623***	-0.772***
	(0.00879)	(0.0172)	(0.0251)	(0.0355)	(0.0120)
Skilled	-0.419***	-0.954***	-0.697***	-0.643***	-0.118***
	(0.00425)	(0.00880)	(0.0127)	(0.0166)	(0.00562)
Previous Job Duration (in months)	0.00170	-0.0158***	-0.00155	0.00535	0.0113***
	(0.00111)	(0.00218)	(0.00316)	(0.00452)	(0.00151)
Heterogeneity group=1	1.908***				
	(0.00297)				
Heterogeneity group=2	-1.149***				
	(0.00376)				
Heterogeneity group=3	3.579***				
	(0.00533)				
[1em] Constant	7.392***	37.25***	15.69***	17.80***	-4.722***
	(0.121)	(0.362)	(0.287)	(0.826)	(0.182)
Observations	3090600	327274	1231526	553776	978024

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note : The dataset is the merge of two simulated datasets, one where claiming costs are at their estimated values, another with a 10% decrease in the claiming cost. Each time, we have $1,545,300$ individual observations, that is for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution. We report here the estimates, they are translated into treatment effects at covariates' mean values in Table 6.

Table E.2: Proportional Hazard Models : Effect of a claiming cost decrease.

	(1) All	(2) Group 1	(3) Group 2	(4) Group 3	(5) Group 4
Lower costs	-0.00708***	-0.00440	-0.00510***	0.000175	-0.0192***
	(0.00146)	(0.00377)	(0.00182)	(0.00635)	(0.00364)
log(past wage)	5.448***	8.537***	6.614***	7.152***	9.717***
	(0.0282)	(0.0912)	(0.0350)	(0.164)	(0.0908)
log(past wage) sq.	-0.593***	-0.902***	-0.665***	-0.874***	-1.028***
	(0.00245)	(0.00799)	(0.00300)	(0.0151)	(0.00814)
Age	0.563***	0.580***	0.482***	0.547***	0.808***
	(0.00510)	(0.0132)	(0.00642)	(0.0220)	(0.0126)
Skilled	1.073***	1.114***	1.023***	1.215***	1.390***
	(0.00256)	(0.00685)	(0.00325)	(0.0107)	(0.00589)
Previous Job Duration (in months)	0.0148***	0.0181***	0.0125***	0.0246***	0.0152***
	(0.000647)	(0.00167)	(0.000811)	(0.00282)	(0.00161)
Heterogeneity group=1	2.104***				
	(0.00273)				
Heterogeneity group=2	2.881***				
	(0.00237)				
Heterogeneity group=3	-0.648***				
	(0.00366)				
Observations	3090600	327274	1231526	553776	978024

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note : The dataset is the merge of two simulated datasets, one where claiming costs are at their estimated values, another with a 10% decrease in the claiming cost. Each time, we have $1,545,300$ individual observations, that is for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution.

E.2 Unemployment benefits increase

Table E.3: Proportional Hazard Models : exit rates from non-employment ($N + P$ to E) - Claimants.

	(1) Duration in P	(2) Total duration (claimants)
Higher benefits	-0.0181***	-0.00620*
	(0.00358)	(0.00358)
log(past wage)	4.901***	4.656***
	(0.0793)	(0.0792)
log(past wage) sq.	-0.638***	-0.621***
	(0.00708)	(0.00707)
Age	0.489***	0.415***
	(0.0126)	(0.0126)
Skilled	1.078***	1.104***
	(0.00699)	(0.00701)
Previous Job Duration (in months)	0.0202***	0.0218***
	(0.00159)	(0.00159)
Heterogeneity group=1	2.896***	3.517***
	(0.0112)	(0.0113)
Heterogeneity group=2	2.898***	3.213***
	(0.0162)	(0.0163)
Heterogeneity group=3	0.0202*	0.612***
	(0.0115)	(0.0115)
Observations	923172	923172

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note : The dataset is the merge of two simulated datasets, one for $b_P = 0.75$, another with $b_N = 1.01 \times 0.75$. Each time, we have $1,545,300$ individual observations, that is for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution.

Table E.4: Proportional Hazard Models : exit rate from non-employment ($N + P$ to E) - All.

	(1) Total duration (all) no unobserved heterog.	(2) Total duration (all) control for unobserved heterog.
Higher benefits	-0.0114***	-0.0229***
	(0.00146)	(0.00146)
log(past wage)	2.909***	5.418***
	(0.0280)	(0.0282)
log(past wage) sq.	-0.324***	-0.590***
	(0.00242)	(0.00245)
Age	0.297***	0.553***
	(0.00510)	(0.00511)
Skilled	0.579***	1.068***
	(0.00254)	(0.00257)
Previous Job Duration (in months)	0.00648***	0.0143***
	(0.000647)	(0.000647)
Heterogeneity group=1		2.083***
		(0.00273)
Heterogeneity group=2		2.880***
		(0.00237)
Heterogeneity group=3		-0.659***
		(0.00367)
Observations	3090600	3090600

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note : The dataset is the merge of two simulated datasets, one for $b_P = 0.75$, another with $b_N = 1.01 \times 0.75$. Each time, we have $1,545,300$ individual observations, that is for each of the $15,453$ individuals in the data, we simulate 100 individuals with the same observable characteristics but each time with a new draw from the unobserved heterogeneity distribution.