Are the Spanish Long-Term Unemployed Unemployable?*

Samuel Bentolila†
CEMFI

J. Ignacio García-Pérez
Universidad Pablo de Olavide and Fedea

Marcel Jansen
Universidad Autónoma de Madrid and Fedea

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Abstract

Long-term unemployment currently comprises around 57% of the unemployed in Spain. In this paper we describe its enormous buildup in the wake of the Great Recession and analyze empirically the individual and aggregate determinants of entering and leaving this state, which is concentrated among mature and low-educated individuals. We also find a large effect of the receipt of unemployment benefits and strong duration dependence. Declared reservation wages are found to respond to the cycle but little to unemployment duration. In view of these findings, we derive some prescriptions on the nature and scope of active labor market policies that are needed to fight long-term unemployment.

Keywords: Long-term unemployment, great recession, duration model, survival probability, Spain

JEL Classification: J63, J64, J65, C41

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†Corresponding author. CEMFI. Casado del Alisal, 5. 28014 Madrid. Spain. bentolila@cemfi.es.

1 Introduction

During the Great Recession, unemployment in Spain shot up, going from 8% in 2007Q3 to a staggering peak of 27% in 2013Q1, and it is currently still at 19% (2016Q3). As a result, long-term unemployment (LTU) reached a peak of 16% of the labor force and, as a share of long-term unemployed, of 64%. Although employment has grown by 2.5% per year over the past three years, LTU still comprises 57% of the unemployed. Such a modest progress in reducing LTU suggests that economic growth alone cannot solve this problem.

The existence of a large pool of long-term unemployed workers poses tremendous challenges for society. Due to various factors, the probability of leaving unemployment falls with duration, so that reenfranchising these workers becomes progressively harder, causing an increasing risk of social exclusion. As a result, policymakers need to confront this issue head on. This motivates our work, which provides a detailed analysis of the factors that contributed the rise of LTU in Spain in the aftermath of the Great Recession, a topic which has attracted very little research so far.

To a large extent, the factors that account for Spain's high unemployment rate also underlie its high LTU. Spain suffered a longer and deeper Great Recession than most European Union (EU) countries. From 2008 to 2013, GDP fell by 9% and employment by 16%. Spain was hit by the 2008 international financial shock and then by the Eurozone crisis, i.e. the so-called double-dip recession. These external shocks were compounded by the bursting of a housing bubble and a strong credit crunch, which had a very large economic impact due to a large buildup in private indebtedness during the boom (Jimeno and Santos 2014, Bentolila *et al.* 2016). The collapse of the construction sector alone added 1.7 million, mostly low-educated males, to the ranks of the unemployed. These external shocks interacted with institutional factors, namely the dual nature of the Spanish labor market –with temporary employees representing around one-quarter of all employees– and an insider-outsider collective bargaining system. These institutional features resemble those in other European countries like France or Italy, but they adopt a stronger form.[1] The result has been very high worker-turnover and real-wage rigidity.[2] In this paper we take the shocks to the Spanish economy as given and pursue an analysis of the mechanisms generating unemployment persistence leading to an enormous buildup of LTU.

Our paper adds to a long literature on long-term unemployment in Europe. Good overviews of work in this area as of the 1990s appear in Layard *et al.* (1991) and Machin and Manning (1999). Thereafter the incidence of LTU started to fall and so did academic work on the issue. However, the rise in LTU in the US during and in the aftermath of the Great Recession sparked an intense debate about the driving forces behind the increase in the incidence of LTU and the low job finding probabilities of the

[1] See Bentolila and Jimeno (2006) for a description of basic features and Bentolila *et al.* (2012a) for a comparison between France and Spain.

[2] From 2010 to 2012, three labor market reforms modified collective bargaining but not duality, see Bentolila *et al.* (2012b) and García-Pérez and Jansen (2015).

long-term unemployed. In spite of their varying approaches, a common finding from these studies is that the low unemployment exit rates of the long-term unemployed are due to negative duration dependence. On the contrary, changes in the pool of the unemployed due to dynamic selection play at best a modest role. Let us briefly review those who are closer to our goals.

In a seminal study, Kroft *et al.* (2016) construct an augmented matching model to assess the relative role of composition effects, duration dependence, and transitions into nonparticipation in the buildup of LTU and the outward shift of the Beveridge curve. These authors back out duration dependence from observed transition rates and show that this is one of the dominant forces at work, together with cyclical changes in nonparticipation.

In a related study Krueger *et al.* (2014) analyze the role of unemployment duration on wages, reemployment possibilities, and labor force withdrawal. The structure of their paper is similar to ours, but the authors use logistic models to analyse how transition rates to employment and nonparticipation vary with duration. Unlike our case, the authors use data from different sources that allow a distinction between unemployment and nonparticipation, but the logistic models that they use do not contain controls for unobserved heterogeneity, which could significantly bias the estimates of duration dependence.

Our paper is closest in spirit to Abraham *et al.* (2016), who use matched firm-worker data to estimate the impact of duration on subsequent employment outcomes. The estimation technique compares the difference in pre- and post-employment shares for unemployed workers in seven different duration categories. This procedure allows the authors to control for time-invariant unobserved heterogeneity and for differences in prior employment outcomes. Nonetheless, there are relevant differences with our work. While we observe the entire working history of the individuals in our sample, Abraham *et al.* (2016) only consider the employment outcomes eight quarters before and after the measurement of unemployment. Moreover, their estimation technique is essentially a difference-in-differences estimator, while we use a duration model.

In our analysis we adapt duration estimation methods to address some of the problems that have been found in the previous literature, by dealing with unobserved heterogeneity, state dependence, and self-selection. Recent work on LTU in the US does not treat the problem of selection into unemployment. By contrast, we do it by estimating a duration model for the probability of entering unemployment from employment, which depends on worker observed and unobserved characteristics. Since these characteristics also determine the probability of leaving unemployment, we estimate both duration models jointly.

Why do we need to use duration models? For two reasons. First, we have censored spells, for both employment and unemployment. Censoring cannot be dealt with using linear probability models, but ignoring censoring can lead to bias, since then the error term does not satisfy the standard properties in a regression model. Second, we wish to control well for explanatory variables that change with duration, unemployment benefits in particular. To do this we need a duration model that takes into account the

individual situation each month.In this vein, our paper also belongs in the literature on the estimation of the determinants of the probability of leaving employment and unemployment.[3]

We also pay close attention to two key institutional factors: the dual structure of the labor market and the unemployment benefit system. Studying LTU in a dual labor market raises some interesting issues that, to the best of our knowledge, have not been addressed yet in the European literature. In particular, is it reasonable to set the unemployment duration counter to zero for a long-term unemployed worker who is hired for a very short temporary job? Are temporary contracts a useful work sharing arrangement during a crisis? We look at these issues by separately analyzing exits from unemployment to temporary and permanent contracts, by studying what happens when exits to very short jobs are censored, and by differntiating betwee exits from these two types of contracts to unemployment.

This paper is structured as follows. In Section 2 we describe the main facts of the buildup of long-term unemployed in Spain using the Labor Force Statistics. We show that LTU especially affects older and less educated workers. In Section 3 we estimate a two-state competing-risks duration model to analyze the factors that explain the LTU inflow and outflow. The estimation is performed using the Continuous Sample of Working Lives. Our empirical results indicate that in Spain the conditional probability of entering LTU is very large and it is significantly raised by receipt of unemployment benefits, mature age, low experience, and, to a lesser extent, low education and low skill. Duration dependence and not dynamic selection is the primary source of the low job finding rates of the long-term unemployed. Temporary contracts help to reduce the risk of LTU conditional on unemployment, but they also cause huge inflows into unemployment.

Finally, in Section 4 we explore an analysis of reservation wages during the Great Recession. We find that declared reservation wages do adjust with the business cycle and also, though quite slowly, with unemployment duration. In Section 5 we summarize our findings and argue that higher aggregate demand alone will not solve the LTU problem. Expanding and, especially, improving active labor market policies and linking them to the receipt of unemployment benefits would help. The Appendix includes further empirical results.

2 The legacy of the Great Recession

In this section we provide several pieces of descriptive evidence on the nature of LTU in Spain. We begin by showing the unprecedented increase in LTU as a result of the Great Recession. Next, we discuss the incidence of LTU for different population groups and

[3]For Spain, see García-Pérez (1997), Bover *et al.* (2002), Bover and Gómez (2004), García-Perez and Muñoz-Bullón (2004), Arranz *et al.* (2010), Arranz and García-Serrano (2014), Carrasco and García-Pérez (2014), Nagore and van Soest (2016b) and many others. This paper especially builds on Rebollo and García-Pérez (2015).

Figure 1: Duration-specific unemployment rates in Spain, 1976Q3-2016Q3 (% of labor force)

Note: Recessions are shaded. Sources: INE, Encuesta de Población Activa and, for recession dates, Spanish Economic Association, Spanish Business Cycle Dating Committee.

we end with a review of the evolution of the main labor market flows during the crisis. Our discussion of the flows includes preliminary evidence on the existence of duration dependence and highlights the thin line between unemployment and non-participation.

2.1 The unprecedented rise in LTU

Figure 1 plots unemployment rates by duration, expressed as a share of the labor force, for the period between 1976 and 2016. The three lines represent, respectively, short-term unemployment (STU, up to one year), long-term unemployment between up to two years, and very long-term unemployment (VLTU), which refers to spells of two or more years. The shaded bars indicate recessions. Their purpose is to highlight the cumulative effect of the two recent recessionary periods, the first one unleashed by the financial crisis in the US after the fall of Lehman Brothers and the second one caused by the Eurozone crisis.

Inspection of Figure 1 shows that all three duration-specific unemployment rates reached record levels at distinct moments during the crisis. While STU peaked at the onset of the first recession, LTU continued to grow until the end of the second recession and VLTU peaked even later. At its peak, 10.6% of the Spanish labor force was unemployed for more than two years. Since that time, the incidence of VLTU has

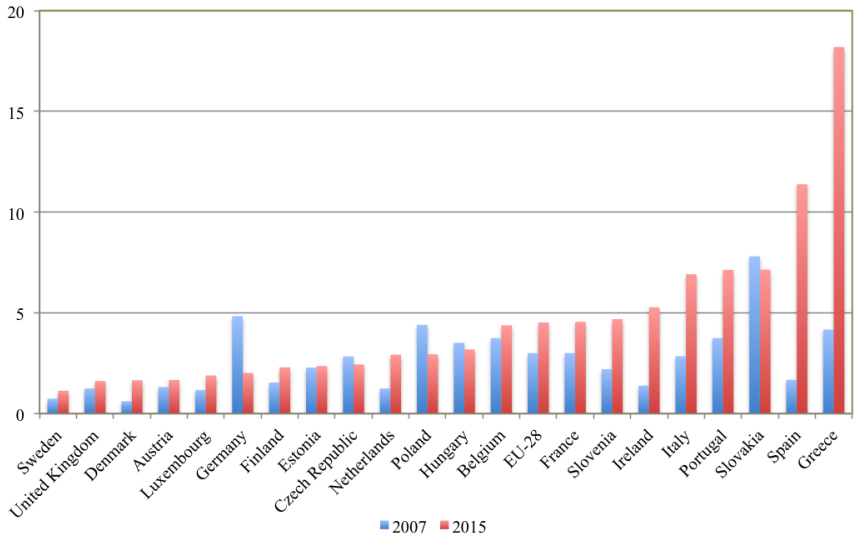


Figure 2: Long-term unemployment before and after the Great Recession in selected EU countries, 2007 and 2015 (% of labor force)

Source: OECD.Stat, Annual Labour Force Statistics.

dropped to a level of 7.9% in 2016Q3, but almost two-thirds of the people remaining in this group (63.4%) have been unemployed for at least four years.

The levels of LTU are not only striking from a historical perspective. Spain also stands out internationally, as is shown in Figure 2. In 2015, its LTU rate (11.4%) was the second-highest in the EU, exceeding the EU average by almost 7 percentage points.[4] Interestingly, four out of the six EU member states with the highest LTU rates participated in an EU-sponsored bailout program. This evidence points at the importance of the deep and prolonged contraction of aggregate demand in these countries as one of the main drivers of the rise in LTU. Moreover, Greece, Italy, Portugal, and Spain are countries with dual labor markets that entered the crisis with a rigid system of collective bargaining. These institutional features and the resulting delay in wage adjustments also seem to have played an essential role in the buildup of LTU in these countries (Bentolila and Jansen 2016).

Next, returning to the case of Spain, we observe a considerable degree of cross-sectional variation in the incidence of LTU. In most cases there exists a strong correlation between the rise in STU and LTU by demographic group. However, changes in non-participation and shifts in the distribution of labor demand also play a role. By way of illustration, Table 1 provides a breakdown of the changes in labor force participa-

[4]Notice that Eurostat reports lower figures for the incidence of LTU than the Spanish Statistical Institute (INE), as it uses a different definition. Eurostat measures the length of unemployment spells as the period since the end of a person's last job. INE, on the contrary, exploits information on search duration.

tion, employment, and unemployment for different worker characteristics. Besides the standard unemployment rate, the table also reports cohort-specific LTU shares. The latter are defined as the proportion of long-term unemployed among the unemployed in each group. The LTU shares provide a proxy for the conditional probability that an unemployed worker in each of the groups ends up in LTU.

The reported data confirm a number of well-known facts. The drop in employment rates affected men more than women, and was strongly concentrated among youth, low-educated workers with at most mandatory education, and immigrants. In 2016 the same cohorts also present the highest LTU rates as a share of the labor force (not shown),[5] but in the cases of youth and immigrants these high LTU rates are mostly driven by the relatively high inflow into unemployment. Indeed, both groups of workers have below-average LTU shares. While 59.5% of the unemployed in our reference period were long-term unemployed, this figure is reduced to 40.2% for youth and to 55.9% for non-nationals. On the contrary, for the remaining demographic groups the LTU shares are all close to 60%. This uniformity is a striking finding that highlights the widespread nature of LTU in Spain. College education provides some protection against the risk of LTU, but the difference between the LTU share of college-educated workers and those with at most primary education is equal to only 8 percentage points (pp). Similarly, even for prime-age workers we obtain an LTU share of 59%.

Finally, older unemployed workers clearly seem to be the most vulnerable group. They face a relatively low risk of unemployment, but more than two-thirds of the unemployed in the age bracket from 45 to 54 years old are long-term unemployed, and for those who are 55 and older this share is a staggering 77.6%. Given these comparatively high LTU shares, one might have expected a rise in non-participation for this group, but the reverse is true. Between 2007 and 2016 the labor force participation rate of the oldest cohort increased by 10.3 pp from 45 to 55.3%. Some part of the increase in the LTU rate of the oldest workers may therefore be driven by a drop in the transition rate to non-participation. Two other cohorts for which we observe marked changes in participation rates during the crisis are youth and women. In our econometric analysis below we avoid the potentially confounding role of these changes in participation rates by restricting the analysis to males aged 25 to 54 years old.

2.2 The role of mismatch and composition effects

Besides the drop in aggregate labor demand, the economic crisis has also produced profound shifts in the relative demand for skills and the distribution of employment by industry. The two major shifts are the drop in the demand for low-educated workers and the collapse of the construction sector. In 2007 the construction sector accounted for 13% of total employment and nine years later this share has fallen to 6%. In absolute terms, the collapse of the construction sector caused a maximum loss of 1.74 million,

[5] Notice that cohort-specific LTU rates can be computed as the product of cohort-specific unemployment rates and LTU shares.

Table 1: Participation, employment, and unemployment rates, and LTU share before and after the Great Recession in Spain, 2007Q2-2016Q2 (%)

	Participation rate		Employment rate		Unemployment rate		LTU share	
	2007Q2	2016Q2	2007Q2	2016Q2	2007Q2	2016Q2	2007Q2	2016Q2
Total	72.1	74.6	66.4	59.6	8.0	20.0	27.8	59.8
Gender								
Male	82.0	79.6	76.9	64.9	6.2	18.5	24.1	58.0
Female	62.0	69.5	55.5	54.3	10.4	21.9	30.8	61.5
Age								
16-24 years old	52.4	37.3	42.9	20.0	18.1	46.5	16.0	40.2
25-34 years old	86.6	87.7	80.1	68.0	7.5	22.4	24.9	56.1
35-49 years old	84.5	90.5	78.8	75.4	6.7	16.8	27.2	58.8
45-54 years old	76.8	84.5	72.1	69.9	6.1	17.3	39.1	67.9
55-64 years old	45.0	55.3	42.4	46.0	5.7	16.9	55.6	77.6
Education								
Primary	53.5	54.3	47.8	35.2	10.6	35.2	33.0	65.2
Secondary, 1st stage	70.9	69.6	64.1	51.0	9.6	26.8	26.5	62.0
Secondary, 2nd stage	74.3	72.6	68.3	58.1	8.0	20.0	26.3	55.7
College	85.3	88.0	80.8	77.7	5.3	11.7	26.3	56.7
Nationality								
Native	57.1	57.4	52.9	47.1	7.3	17.9	31.3	60.8
Foreign	76.2	73.2	67.2	55.1	11.9	24.8	15.2	55.9

Note: Data are shares of source population. Source: INE, Encuesta de Población Activa.

mostly low-skilled jobs during the crisis, but this is not the whole picture.[6] From peak to through Spain lost three million jobs that were previously occupied by workers with no more than compulsory education. Moreover, the sharp drop in the employment of low-educated workers contrasts with a slight increase in the employment of university graduates.

A straightforward procedure to illustrate the degree of mismatch between the demand and supply of labor and to gauge its importance for the buildup of LTU is to compare the characteristics of the pool of employed and unemployed workers. The former provides an approximation to labor demand while the composition of the pool of unemployed workers provides a proxy for the characteristics of labor supply. Once again we distinguish between STU, LTU, and VLTU. Moreover, in addition to the personal characteristics listed before, in Table 2 we also provide a breakdown by the industry of the unemployed workers' previous occupation. This decomposition, unavailable from the Labor Force Survey (LFS), is computed from the administrative records of the Muestra Continua de Vidas Laborales (see Section 3.2), while the rest of the data are computed using data from the Spanish LFS.

Our main interest concerns the groups of workers that are more prominent among the unemployed than among the employed. Careful inspection of the data shows that this is the case for youth, immigrants, low-educated individuals –i.e., those without completed upper-secondary education– and workers from the construction sector. For the latter two groups the degree of mismatch is increasing roughly monotonically with unemployment duration, while the opposite is true for youth and immigrants. For example, the construction sector currently accounts for 6.1% of employment and the workers from this sector represent 11.3% of the short-term unemployed and 19.3% of the very long-term unemployed, while the corresponding figures for individuals with primary education are, respectively, 6.6%, 12.4%, and 16.4%. The individuals who belong to these two groups therefore face both an above-average probability of entering unemployment and a comparatively high conditional probability of ending up in VLTU.

By contrast, youth and immigrants enter unemployment more frequently than the average worker, but they also exit unemployment more quickly. The latter explains why their share in the pool of unemployed is negatively related to duration. Finally, workers above 45 years of age make up a comparatively small share of the short-run unemployed and a comparatively large share of the long-term unemployed, and this share is increasing with duration. In other words, relatively few older workers become unemployed, but those who do so face a high conditional probability of ending up in VLTU.

The relatively strong concentration of disadvantaged workers among the unemployed with spells longer than two years poses a challenge to policymakers. Our analysis below indicates that these workers have very low job finding rates, which have barely improved in recent times. Nonetheless, the contribution of composition effects to the buildup of

[6] This calculation is based on data from the Spanish Labor Force Survey and takes 2008Q1 as the reference period. The breakdown of employment by industry is not consistent before this date.

Table 2: Employment and unemployment breakdown by duration and population group in Spain, 2016Q2 (%)

		Unemployed		
	Employed	Short-term	Long-term	Very long-term
Gender				
Male	54.5	51.3	45.8	48.5
Female	45.5	48.7	54.2	51.5
Age				
16-24 years old	4.4	22.5	18.4	7.0
25-34 years old	20.7	26.0	27.4	20.8
35-44 years old	31.9	26.1	23.7	25.8
45-54 years old	27.5	18.3	19.9	28.2
55-64 years old	15.5	7.0	10.6	18.2
Education				
Primary	6.6	12.4	12.5	16.4
Secondary, 1st stage	27.4	37.6	37.2	42.6
Secondary, 2nd stage	23.9	26.1	23.3	21.5
College	42.1	23.8	26.9	19.5
Industry of previous job				
Primary	4.1	0.1	3.6	3.0
Manufacturing	14.0	8.6	9.7	13.5
Construction	6.1	11.3	9.2	19.3
Wholesale and retail trade	16.6	16.3	16.4	14.5
Finance and real state	3.1	1.5	1.9	1.7
Professional and business serv.	20.5	24.7	19.0	20.5
Education	6.7	3.5	4.1	3.0
Health care	8.1	5.5	5.4	4.6
Leisure and hospitality	7.8	16.5	13.1	4.4
Scientists, artists and other	13.0	12.0	17.6	15.5

Note: The breakdown by industry corresponds to 2015. Columns add up to 100 by characteristic. Sources: INE, Encuesta de Población Activa and for industry breakdown Ministerio de Empleo y Seguridad Social, Encuesta Continua de Vidas Laborales. Short-term denotes unemploment for up to one year, long-term between one and two years, and very-long term for more than two years.

the current stock of long-term unemployed seems to have been modest compared to the contribution of the strong rise in the disaggregated LTU shares for most cohorts. In Bentolila *et al.* (2017) we construct a counterfactual series for the overall LTU share in which the weights of different groups (formed by crossing age, gender, education, and nationality) in unemployment are allowed to change in accordance with the actual data, while the group-specific LTU shares are fixed at their pre-crisis values in 2008Q3. By construction, the counterfactual LTU share for the reference period therefore coincides with the pre-crisis minimum of the actual LTU share. While this is a rough exercise that excludes the role of time-varying industry shares, it is telling that composition effects account only for a small part of the increase in the LTU share from 2008Q3 to 2016Q2 (our computations say that composition effects only account for about 4 of the 38 pp observed increase).

2.3 Labor market flows during the crisis

In the previous sections we have provided evidence on the evolution of the stocks of employed (E), unemployed (U), and non-participating (N) workers. The purpose of this section is to describe the evolution of the flows among these states. To be consistent with the empirical analysis in the next section, we distinguish between permanent and temporary jobs and we restrict our sample to male prime-age workers (25-54 years old). We use aggregates constructed from LFS flow microdata using population weights and measured as three-quarter centered moving averages, so as to reduce seasonal volatility.

We begin with the initial impulse to the rise in unemployment, i.e. the increase in the transition rates between employment and unemployment at the start of the two recent recessions. In the left-hand panel of Figure 3 we observe a sharp increase in the termination of temporary jobs at the start of the first recession and another jump in the second recession. The spikes in dismissals of permanent workers are even more apparent in the right-hand panel, though it should be noted that the separation rate of permanent jobs is almost a factor of magnitude smaller than the one for temporary jobs. Overall, between 2008 and 2013, the quarterly transition rate between temporary employment and unemployment more than doubled –and even today it is still substantially higher than in 2008. By contrast, the transition rate between permanent employment and unemployment is currently close to its pre-crisis level.

Our next figure depicts the transition rates in the opposite direction, namely between unemployment and employment, on either permanent or temporary jobs. Both transition rates fell dramatically during the crisis and are still far below their pre-crisis levels. In the case of temporary jobs the average quarterly transition rate fell by almost 20 pp, from a pre-crisis level of around 32% to a minimum of 12% and its current level lies around 17%. The quarterly transition rate between unemployment and permanent employment follows a similar pattern, but once again the rates are almost an order of magnitude smaller than in the case of temporary employment. Before the crisis, around 5% of unemployed males managed to transit from unemployment to a permanent job within a quarter, while currently only close to 2% do. The above-mentioned rates are

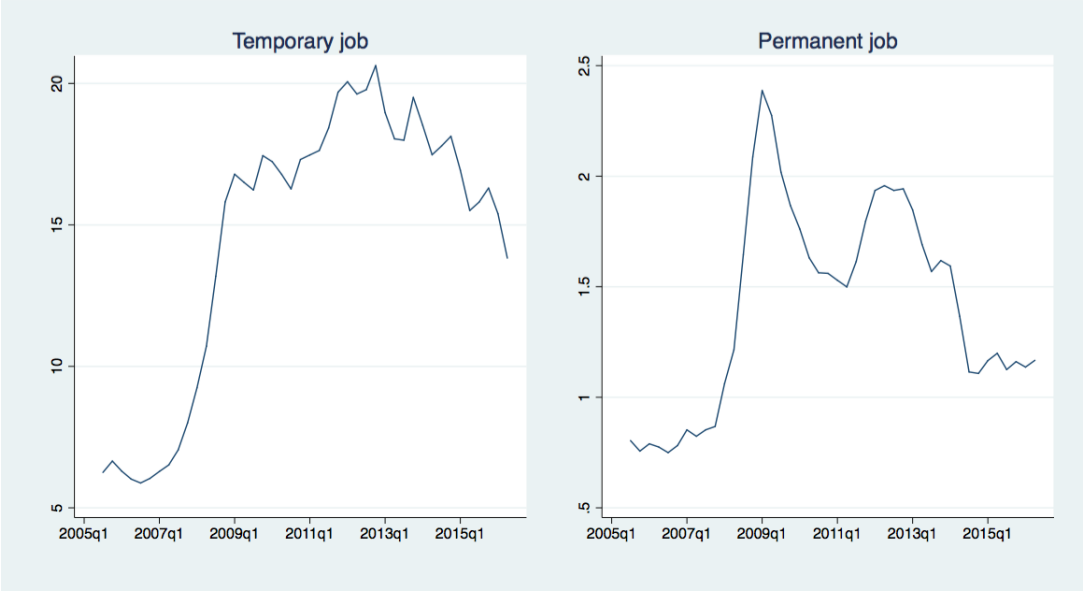


Figure 3: Quarterly flow rates into unemployment by contract type in Spain, 2005Q3-2016Q2 (%)

Note: Rates are computed with respect to the reference population in the preceding quarter. Source: INE, Encuesta de Población Activa.

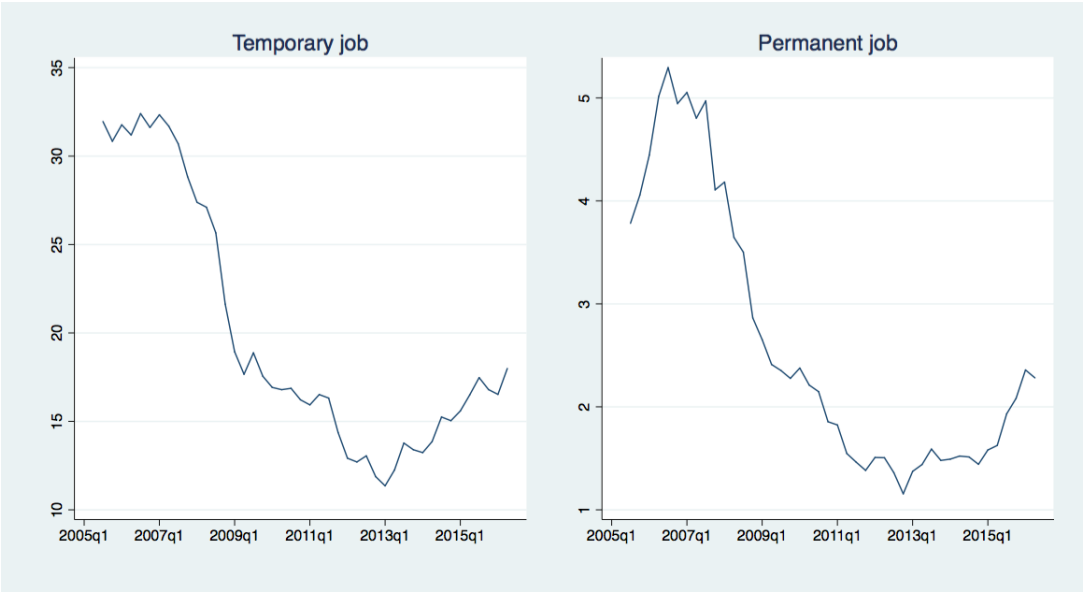


Figure 4: Quarterly flow rates into employment by contract type in Spain, 2005Q3-2016Q2 (%)

Note: Rates are computed with respect to the reference population in the preceding quarter. Source: INE, Encuesta de Población Activa.

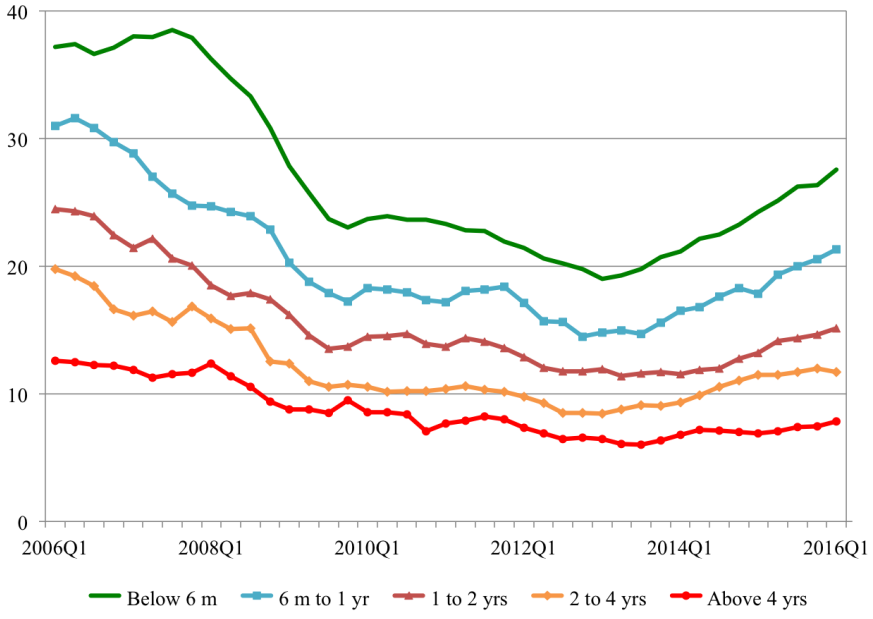


Figure 5: Unemployment duration-specific transitions to employment in Spain, 2006Q1-2016Q1 (%)

Source: INE, Encuesta de Población Activa.

averages and do not account for changes in the pool of unemployed, but they provide an indication of the dramatic increase in the average length of unemployment spells and thereby the risk of LTU during the crisis.

In view of our focus on LTU, it is also worth checking whether the flow rate from unemployment to employment falls with duration. To this aim, Figure 5 represents quarterly exit rates to employment according to the duration of unemployment. There is a clear gradient, with the exit rate falling with the spell length. And the differences are huge: at the end of the expansion around 37% of workers who had been unemployed for up to 6 months left within a quarter, whereas only 14% of those who had been unemployed for 4 years or more did so. During the recession all exit rates fell, but differences across duration groups shrank significantly. Then, when the recovery arrived, differences widened again, with the exit rates of the long-term unemployed increasing by less than those of the short-term unemployed, a situation which has reinforced their chances of remaining stuck in unemployment. This is *prima facie* evidence of the presence of duration dependence, but at this stage we cannot exclude the alternative explanation of dynamic selection. Namely, that over time the composition of the pool of the unemployed may worsen as the most employable workers leave first and this process of selection also generates a negative relationship between the average job finding rate and duration. We address this issue rigorously in Section 3.

The foregoing evidence highlights the dramatic changes that took place during the

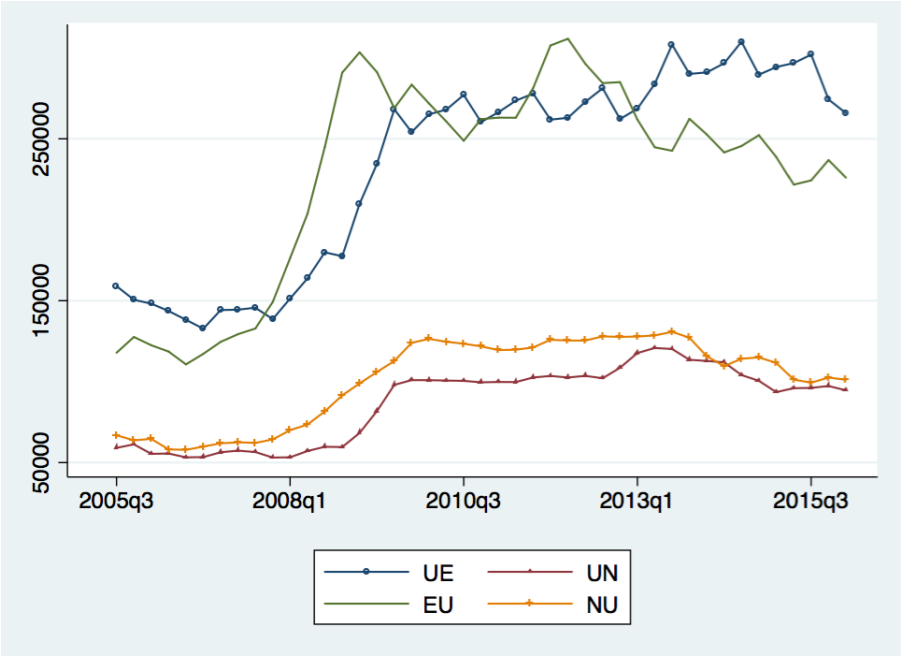


Figure 6: Quarterly flows to/from unemployment in Spain, 2005Q3-2016Q1 (persons)

Note: E, employment; U, unemployment; and N, nonparticipation. Source: INE, Encuesta de Población Activa.

crisis in the flows between employment and unemployment. It also reveals the overwhelming dominance of temporary jobs vis-à-vis permanent jobs in labor market flows, as well as a clear deterioration in the quality of new jobs. Now an even larger share of the newly created jobs are temporary and these jobs last less than before the crisis. It is important to take this aspect into account, because it is not clear *a priori* to what degree relatively short-duration temporary jobs may help a long-term unemployed person to restore his or her working career after several years in unemployment.[7]

Another relevant issue relates to the transitions into and out of nonparticipation. The participation rates of prime age males are fairly constant during the crisis. Nonetheless, these net rates hide frequent movements into and out of non-participation. Figure 6 depicts absolute gross flows to and from unemployment from 2005 to 2016. It is apparent that the recession brought about a very large jump in absolute flows between employment and unemployment, but also that many people enter or leave nonparticipation from one quarter to the next. These flows are equivalent to at least one-third of those between employment and unemployment. According to LFS definitions, workers are considered to be: (a) employed if they have worked for at least one hour in the survey reference week, (b) unemployed if they are not employed, they have searched for a job during the previous month, and they are also available to start work within two

[7]Nagore and van Soest (2016a) analyze the change in the behavior of labor market flows during the crisis using the Spanish Social Security data.

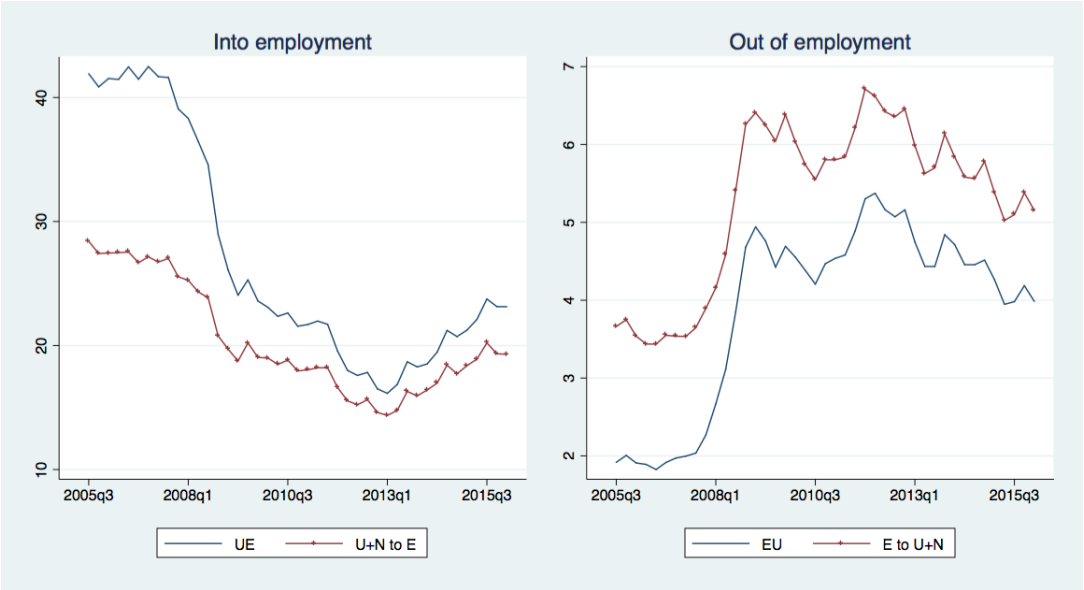


Figure 7: Quarterly flow rates into employment and unemployment in Spain, 2005Q3-2016Q3 (%)

Note: Rates are computed with respect to the reference population in the preceding quarter. Source: INE, Encuesta de Población Activa.

weeks, and (c) nonparticipants otherwise. These definitions reveal that the dividing line between unemployment and nonparticipation is quite permeable. People without a job who stop searching for one quarter go from unemployment to nonparticipation but if they look for work again in the following quarters they will go through the reverse transition. These transient flows are frequent, indicating that nonparticipation is not necessarily persistent, especially within this age bracket. Thus, as has been pointed out for the US by Elsby *et al.* (2015), flows involving non-participation can be a crucial element in explaining the dynamics of unemployment.

In view of this finding –and also due to data limitations– in Section 3 we study the transitions between nonemployment and employment. It is therefore worth checking their cyclical behavior vis-à-vis the traditional ones involving unemployment rather than nonemployment. These are represented in Figure 7, where the left panel depicts the employment inflow rates and the right panel the employment outflow rates. It is apparent that the transition rates from/to nonemployment follow very similar pattern as the traditional transition rates between employment and unemployment.

3 What affects the probability of entering and leaving LTU?

This section presents our empirical results on the probability of entering and exiting LTU. Our key goal is to estimate the impact of individual characteristics and duration dependence on the profile of the job finding rates of the long-term unemployed. Throughout the analysis we control for unobserved heterogeneity to avoid bias due to selection on unobservables. We begin this section by explaining our empirical approach, we then move on to describe the key characteristics of the sample, and we end the section with a discussion of our main estimation results.

3.1 Empirical approach

We assume that individuals move between two mutually exclusive states, employment and unemployment (or, more rigorously, nonemployment). Furthermore, in view of the dual nature of the Spanish labor market, we distinguish between temporary and permanent, or open-ended, employment. Thus, since we are interested in analyzing the transitions between these labor states, our empirical model will consider the joint estimation of the exit or hazard rates out of both employment and unemployment.

We specify these hazard rates using a discrete-time duration model (Jenkins 1995), where they are given by the following conditional probability:

$$h(t) = \Pr(T = t \mid T \geq t)$$

and where T is a discrete random variable denoting either employment or unemployment duration. The hazard $h(t)$ therefore measures the probability of a transition for a person who has remained in the same state for exactly t periods. We will also allow for multiple destinations from each state, i.e. we employ a competing risks model for each state (Lancaster 1990).

Our framework is similar to the one in Carrasco and García-Pérez (2014) and very close to that in Rebollo and García-Pérez (2015). The hazard rate of unemployed individuals depends on unemployment duration, t, and on a vector of variables $x(t)$ which includes a set of individual, sectoral, and aggregate variables described below. It also depends on the person's unemployment benefit entitlement, which is captured by two variables. The first one is a dummy variable, $b(t)$, that takes the value one when the worker is receiving an unemployment subsidy in period t. The second one is a discrete variable $e(t)$ that measures the remaining months of entitlement for persons who are eligible to contribution-based unemployment benefits and receiving them. These two benefits variables and some variables in $x(t)$ are allowed to have heterogeneous effects over the unemployment spell through their time-varying coefficients, $\alpha_i(t)$, through an interaction with log duration. In the case of $e(t)$ this is modelled by allowing $\alpha_3(t)$ to be a cubic polynomial in log duration.[8] Accordingly, the unemployment hazard rate

[8]This is a departure from the setup in Rebollo and García-Pérez (2015), who assume linearity.

has the following structure:

$$h_u^j(t) = \Pr(T_u = t \mid T_u \geq t, x(t), b(t), e(t), \eta^u) = F(\alpha_0^j(t) + \alpha_1^j(t)x(t) + \alpha_2^j(t)b(t) + \alpha_3^j(t)e(t) + \eta^u)$$

where $j = e_T, e_P$ denotes, respectively, the two alternative exits from unemployment, i.e. employment with a temporary contract and with a permanent contract, as in Bover and Gómez (2004). The last term in the formula, η^u, is discussed below.

The exit from employment is also estimated as a competing risks model and all coefficients in these hazard rates are allowed to differ between temporary and permanent employees. The two competing risks for employed workers are moving to another job, e, without going through unemployment, and moving to unemployment, u. Hence, the employment hazard rate has the following specification:

$$h_e^k(t) = \Pr(T_e = t \mid T_e \geq t, x(t), \eta^e) = F(\beta_0^k(t) + \beta_1^k(t)x(t) + \eta^e)$$

where $k = e, u$ denotes the two alternative exits. This hazard rate is simpler than the one in Rebollo and García-Pérez (2015), where it also depends on previous receipt of unemployment benefits. The variables in $x(t)$ are the same as in the unemployment hazard.

Following Bover *et al.* (2002) and García-Perez and Muñoz-Bullón (2004), we use a logistic distribution to model all hazard rates.[9] Moreover, as we are considering competing risks models for both employment and unemployment, exit from a given state needs to be specified as a multinomial logit model with two alternative risks for each state:

$$h_u(t) = h_u^{e_T}(t) + h_u^{e_P}(t)$$

$$h_e(t) = h_e^e(t) + h_e^u(t)$$

Lastly, in order to avoid spurious duration dependence in the hazard rate generated by the presence of unobserved factors (van den Berg 2001), we control for unobserved heterogeneity affecting the flows both to and from unemployment. This is captured by the terms η^u and η^e in the expressions for the hazards. We exploit the fact that we observe multiple spells for the same individual and we estimate the unemployment and employment hazard rates simultaneously, assuming that unobserved heterogeneity follows a discrete distribution function with different mass points, as in Heckman and Singer (1984). In particular, we allow a four-mass-point distribution function, namely two different points for each state, η_1^u and η_2^u for unemployment, and η_1^e and η_2^e for employment, so that four different types may emerge with joint probabilities: (η_1^u,

[9] We could have alternatively used the extreme value distribution. As explained in van den Berg (2001), this distribution allows the model to verify the mixed proportional hazard assumption. Our approach departs from the proportionality assumption, at the cost of imposing more structure, because we want to allow the potential impact of duration and of both observed and unobserved heterogeneity on the exit from employment and unemployment not to be proportional.

η_1^e), (η_1^u, η_2^u), (η_2^u, η_1^e), and (η_2^u, η_2^e). Standard errors for the estimated coefficients are computed using the delta method. The existence of repeated employment and unemployment spells and, more importantly, of some time-varying covariates allows non-parametric identification (Abbring and van den Berg 2004, Gaure *et al.* 2007).

Before presenting our the results it is worth noting that our empirical approach only yields reduced-form estimates, resulting from the interplay of labor demand and supply, so that they do not have a causal interpretation. They are nevertheless informative on the characteristics which are associated with a higher risk of LTU.

3.2 The sample and the set of control variables

Our initial data set is a 20% random sample of the prime age males aged 25 to 54 whose records appear in the nine waves of the Continuous Sample of Working Lives (Muestra Continua de Vidas Laborales or MCVL) corresponding to the years 2006-2014. All individuals must have experienced at least one spell of nonemployment between 2001 and 2014. Individuals only appear in the records when they pay Social Security contributions –roughly, if they are either employed or nonemployed and receiving unemployment benefits. No information about job search activity is available in the data set, hence we cannot distinguish between unemployed and nonparticipating individuals. Moreover, nonemployed individuals drop out of the records if they stop receiving benefits. This is however not a problem, since the length of completed spells of nonemployment can be reconstructed using the information on the subsequent job. Furthermore, to exclude persistent nonparticipants, we limit nonemployment duration to three years, after which spells are treated as censored. With this caveat in mind, as already indicated, we take the license of referring to individuals without employment as unemployed rather than nonemployed individuals.

Against this potential drawback, the MCVL data have crucial advantages vis-à-vis the flow data from the LFS. They allow us to follow workers since the start of their working careers –whereas the LFS only follows individuals for six quarters– and they have a daily frequency, which permits the observation of all employment spells –while many labor market transitions are lost in the LFS due to its quarterly frequency. We can also construct a worker's entitlement to benefits from the MCVL, whereas the LFS only allows us to know whether the worker is receiving benefits or not. By contrast, using this data source implies that we lose information on people living in the same household as the unemployed worker, that is available in the LFS.

In order to avoid problems concerning attrition, we only analyze employment spells in the general Social Security regime, thus excluding special regimes such as agriculture, public employment, and self-employment, and treat exits from unemployment to these states as right-censored (García-Pérez 2006). Moreover, in the case of workers who are recalled to the same firm, a feature which has become increasingly important in Spain (Arranz and García-Serrano 2014), we only consider intervening unemployment spells lasting more than 30 days.[10] Lastly, to maximize the probability of observing

[10] Accordingly, two employment spells with the same firm which have an unemployment spell lasting

the individual's complete work history, we exclude both immigrants and people who appear for the first time in the sample being 30 years old or older (who may have an unrecorded work history).

Our set of control variables includes both individual characteristics and aggregate variables. The former comprise: (a) Age, grouped into three ten-year intervals. (b) Education, measured by dummy variables for the highest degree attained.[11] (c) Skill, divided into high, medium, and low, computed from grouped Social Security tax categories.[12] (d) Actual experience, measured by the number of months employed divided by the number of months of potential experience computed since the person entered the labor market. (e) A dummy that captures whether the worker was fired from his previous job. And (f) our controls for benefit entitlement. As explained before, entitlement to contributory benefits is measured by the remaining months of entitlement in each month (Meyer 1990). The latter is computed from each individual's employment and insurance claim history (since residual benefits not claimed in one unemployment spell can be claimed again in a later spell).[13] This entitlement is entered as a cubic polynomial in order to capture nonlinear effects.

The aggregate variables included in the model are as follows: (a) Employment growth, captured by the monthly growth rate of the number of employees by province. (b) The quarterly national unemployment rate. (c) 17 region dummies. (d) 6 industry dummies. (e) 12 monthly dummies. And (f) two step dummy variables for the labor reforms in June 2010 and February 2012, that take the value one from those dates onwards.

Finally, the following variables are interacted with log duration: age, education, skill, unemployment insurance entitlement (linear and quadratic), receipt of unemployment assistance benefits, monthly growth rate of employees by province, quarterly national unemployment rate, and industry. In an extension we also interact the type of contract (see below).

3.3 Descriptive statistics

Table 3 shows the main characteristics of workers in the sample when they enter unemployment, separately for the expansion and the recession. It should be stressed that

less than 30 days in between are considered as a single employment spell.

[11]The data on educational attainment in the MCVL is officially revised using information from the Spanish Continuous Census of Population (Padrón Continuo). It has however been improved since 2009 with data from the Ministry of Education, and so we use the latter information, imputing it backwards (see De la Roca and Puga 2016, footnote 7).

[12]High skill includes college and junior college graduates, and top and middle managers (groups 1 to 6 in the Social Security classification), medium skill includes administrative assistants and so-called first- and second-level officers (groups 7 and 8), and low skill includes third-level officers and unskilled workers (groups 9 and 10), see García-Pérez (1997).

[13]Workers having access to two different sets of benefit entitlements must choose between them. We assume that they choose the one with the higher length. For more information see Rebollo and García-Pérez (2015).

Table 3: Descriptive statistics of sample spells, 2001-2014 (%)

	Expansion		Recession	
	Mean	St. dev.	Mean	St. dev.
Age				
25-34 years old	88.3	(32.2)	60.7	(48.8)
35-49 years old	10.7	(30.9)	35.6	(47.9)
45-54 years old	1.0	(10.1)	3.7	(19.0)
Education				
Primary or less	11.9	(32.3)	13.5	(34.2)
Secondary, 1st stage	48.7	(50.0)	51.0	(50.0)
Secondary, 2nd stage	25.3	(43.5)	23.2	(42.2)
College	14.2	(34.9)	12.3	(32.8)
Skill				
High	31.1	(46.3)	41.4	(49.3)
Low	24.7	(43.1)	11.6	(32.0)
Medium	44.2	(49.7)	47.0	(49.9)
Experience				
Fraction of potential	76.9	(25.7)	77.3	(24.6)
Dismissal from previous job				
Dismissed	80.8	(39.4)	91.5	(27.9)
Not dismised	19.2	(39.4)	8.5	(27.9)
Industry of previous job				
Manufacturing	13.6	(34.2)	12.7	(33.3)
Construction	29.6	(45.7)	25.7	(43.7)
Non-market services	7.5	(26.3)	9.5	(29.3)
Trade	11.4	(31.7)	11.5	(31.8)
Hospitality	9.1	(28.8)	11.4	(31.8)
Other services	28.9	(45.3)	29.3	(45.5)
Unemployment benefits				
Contributory	26.2	(44.0)	29.0	(45.4)
Assistance	17.7	(38.2)	35.6	(47.9)
No benefits	56.1	(49.6)	35.3	(47.8)
Previous job contract type				
Permanent	16.4	(37.0)	20.3	(40.3)
Temporary	83.6	(37.0)	79.7	(40.3)
Number of spells	37,399		62,045	

Note: The sample is made of males aged 25-54 years old. The expansion corresponds to the period 2001-2007 and the recession to 2008-2014. The characteristics correspond to individuals in their first month in unemployment. Columns add up to 100 by characteristic.

Table 4: Unemployment duration, benefit duration, and raw hazard rates

	Expansion	Recession
A. Unemployment duration (months)		
Exit to a temporary job		
Median	3.0	4.0
Third quartile	6.0	8.0
Mean	4.5	6.3
Share of spells (%)	76.6	69.9
Exit to a permanent job		
Median	3.0	4.0
Third quartile	6.0	8.0
Mean	5.1	6.2
Share of spells (%)	7.9	8.1
Censored spells		
Median	5.0	8.0
Third quartile	11.0	16.0
Mean	7.6	10.7
Share of spells (%)	15.5	22.0
B. Unemployment benefit duration (months)		
All		
Median	7.0	10.0
Mean	9.8	11.2
Temporary previous job		
Median	6.0	8.0
Mean	8.9	10.0
Permanent previous job:		
Median	16.0	20.0
Mean	15.6	16.8
C. Hazard rates out of unemployment (%)		
Exit to a temporary job		
No benefits	16.9	9.3
Contributory benefits	9.7	7.2
Assistance benefits	11.3	7.9
Exit to a permanent job		
No benefits	1.6	1.0
Contributory benefits	1.3	1.2
Assistance benefits	1.2	0.7
Number of spells	37.399	62,045

Note: The sample is made of males aged 25-54 years old. The expansion corresponds to the period 2001-2007 and the recession to 2008-2014. Shares of spells add to 100 by column.

they correspond to workers involved in inflows rather than to worker stocks. The majority are younger than 35 years old, have completed at most compulsory secondary education, and enter unemployment from a temporary job. These facts confirm the remarks made above regarding Table 1, which is based on LFS data. The education breakdown matches well the skill structure measured via occupations. In the expansion, close to one-half of the unemployed do not qualify for benefits. There are sizeable changes in the composition of inflows from expansion to recession, with workers becoming on average younger and less educated. They are also more likely to have been dismissed (as quits fall) and to come from permanent jobs. Accordingly, they are more likely to be entitled to unemployment benefits. The share of construction workers falls, which is unexpected, but it simply reflects a reduction in turnover in that industry, given the scarcity of new jobs.[14]

Next, Table 4 presents some descriptive statistics of spells. The vast majority (91%) of non-censored exits from unemployment are to temporary jobs, but the average length is similar for spells ending in both types of contracts.[15] Completed spells are quite lengthy, around 5 months on average in the expansion and 6 months in the recession, though, as shown in the Table for the third quartile, they are much lengthier in the upper part of the distribution. Moreover, the recession is characterized by a strong rise in both the share and the duration of censored spells. Average benefit entitlement periods last 10-11 months, but they are much smaller for workers coming from temporary jobs than from permanent jobs, for whom the mean is around 17 months in the recession. Lastly, average monthly hazard rates from unemployment to permanent jobs are around one-tenth of those to temporary jobs and they are typically smaller for workers who receive unemployment benefits.

3.4 Empirical results

We now present our estimation results. In order to save space, we restrict our attention to the hazard of leaving unemployment (the estimates for the hazard of leaving employment are included in Bentolila *et al.* 2017).

3.4.1 Hazards

Before introducing our main results for the inflows to and outflows from LTU, it is instructive to analyze the average hazard rates. The hazard rate for the recession is shown in Figure 8. It corresponds to a hypothetical individual with the average characteristics in our sample and it takes into account all interactions between these characteristics and duration. The underlying estimated coefficients are reported in Appendix Table A.1.

[14] According to the LFS, in 2014 flows into construction from nonemployment were equal to one-third of the level in 2008.

[15] Overall, 18.4% of spells are censored, which correspond to true censoring (13.3%) and exit after 36 months (5.1%).

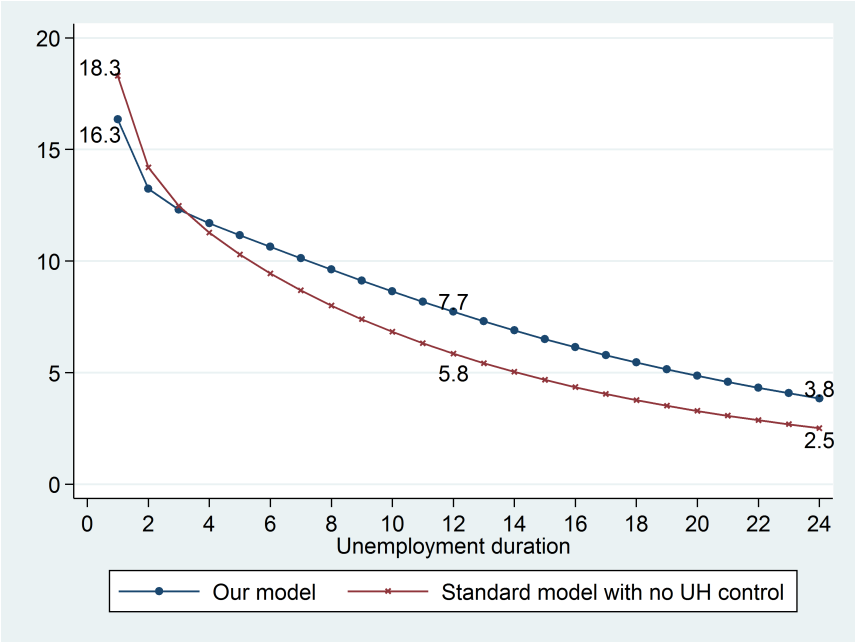


Figure 8: Duration dependence in the baseline model and the model without unobserved heterogeneity (%)

Note. Source: own computations.

The time profile of the average hazard rate provides our best estimate of the magnitude of duration dependence. During the first twelve months of unemployment the monthly hazard of the average individual drops by 55% or 9 pp –from 16.4% to 7.4%– and it halves again during the second year. For comparison, the first figure is equal to two-thirds of the difference between the hazard rates in the first month for an average worker with no benefits and for one with a 12-month entitlement to contributory benefits (which amounts to 13.5 pp). It is also equal to 47% of the difference between the former hazard rate and the hazard rate for an average worker with a 24-month entitlement (which amounts to 19.2 pp). Moreover, the impact of duration dependence is much higher than the 5.6 pp difference implied by the movement from the 25th to the 75th percentile in the distribution of actual work experience and the meager 2.8 pp difference between the initial hazard rates of 25-34 year-old workers and 45-54 year-old ones.

It is interesting to see how cycle affects hazard rates depending on unemployment duration. To this end, we have computed the average hazard rate in the recession (2008-2014) in two scenarios. The good scenario is one in which provincial employment growth is set at its third quartile and the national unemployment rate at its first quartile in the sample (4.8% and 18.6%, respectively), while in the bad scenario each variable is set at the opposite quartile (first and third, respectively, with values of -0.44% and 24.9%). The predicted hazard rates for unemployment durations of 1, 12, and 24

months, respectively, are: 18.6 vs. 14.6, 8.2 vs. 7.4, and 4.0 vs. 3.7. In other words, upon an improvement in the labor market, those who have been unemployed for one month leave 27.3% faster (4 pp), those unemployed for one year leave 11.7% faster (0.8 pp), and those unemployed for 24 months leave only 7% faster (0.25 pp). These figures clearly illustrate that a cyclical improvement helps the short-term unemployed much more than the long-term unemployed, and confirms the profile of exit rates previously shown in Figure 5.

Figure 8 also confirms the need to control for unobserved heterogeneity. Failure to do so biases the estimation of duration dependence upwards, as indicated by the steeper slope of the curve which depicts the average hazard for our baseline model in that case (which now falls by 68%, from 18.3% to 5.8%, during the first year).[16] While their interpretation is not straightforward, it is still worth to briefly comment the results. According to our results, there are two prevailing types of workers: the (η^u_{fast}, η^e_{slow}) type, i.e. those with a strong labor-market attachment, who leave unemployment fast and employment slowly (32% of the individuals in the expansion), and the (η^u_{low}, η^e_{slow}) type, i.e. those who leave both unemployment and employment slowly (50% of the individuals). The changes in the shares from the expansion to the recession indicate that part of the increase in unemployment duration is due to a fall in the share of fast unemployment leavers (by 13 pp) and an offsetting increase in slow leavers, especially among those with low exit rates from both unemployment and employment.

3.4.2 The predicted probability of entering LTU or VLTU

Both the initial hazard rate and the slope of the hazard function differ across individuals. In order to assess the marginal impact of individual characteristics on the probability that a workers enters LTU or VLTU, we report in Table 5 the survival rates in unemployment at 12 and 24 months −i.e. the probability that an individual does not manage to find a job during the first or second year of unemployment. Apart from the average survival rates reported in the top row, these survival rates are constructed by varying one characteristic at a time and setting all other characteristics at the values used to construct the average hazard rates. Moreover, in order to construct the survival rates at 24 months we have rescaled the survival rate at 12 months to 100. The survival rate at 24 months therefore measures the conditional probability that an individual who is still unemployed after one year remains unemployed during the second year of unemployment.

Inspection of the table reveals that the chances of entering LTU have almost doubled in recent years, as indicated by the rise in the survival rate at 12 months from 13.8% in the expansion to 25.5% in the recession. The difficulty to escape LTU is captured by the survival rate at 24 months. In the expansion two-thirds of the long-term unemployed managed to leave unemployment in the next twelve months, but in the recession this figure has dropped to one-half. Hence, the recent recession is characterized by a strong

[16] The estimates related to unobserved heterogeneity appear in Table A.2 of the Appendix.

Table 5: Survival rates in unemployment at 12 and 24 months (%)

	Expansion		Recession	
	12 months	24 months	12 months	24 months
Overall	13.8	33.2	25.5	51.3
Age				
25-34 years old	13.3	32.2	23.7	48.9
35-44 years old	16.0	38.6	27.5	53.9
45-54 years old	27.5	49.8	34.9	62.9
Education				
Primary or less	13.4	31.0	28.4	51.8
Secondary, 1st stage	11.2	27.7	24.2	47.1
Secondary, 2nd stage	14.7	32.0	26.6	50.8
College	19.6	35.9	27.6	50.7
Skill				
High	14.9	36.0	26.0	52.6
Medium	12.5	31.6	23.4	49.0
Low	14.5	31.9	30.7	54.0
Experience				
P75	10.7	28.3	18.5	43.4
P50	12.2	30.7	22.3	47.9
P25	16.2	36.6	30.8	56.5
P10	20.6	42.2	40.6	64.9
Industry				
Manufacturing	12.3	30.0	23.7	47.3
Construction	11.8	31.8	23.6	48.3
Non-market services	18.4	39.7	30.9	58.3
Trade	14.8	30.6	27.7	50.2
Hospitality	12.3	27.5	22.1	46.8
Other services	14.7	36.2	25.7	54.7
Unemployment insurance				
No benefits	5.7	19.9	12.1	36.0
6 months	13.0	19.9	20.6	36.0
12 months	35.0	19.9	41.9	36.0
18 months	45.3	30.9	52.6	46.9
24 months	53.3	56.1	60.7	65.6
Unemployment assistance				
No	5.7	19.9	12.1	36.0
Yes	31.3	51.7	44.3	67.3
GDP growth				
High	12.6	30.7	22.2	49.5
Low	15.2	35.8	28.4	52.6

Note: The sample is made of males aged 25-54 years old. The expansion corresponds to the period 2001-2007 and the recession to 2008-2014. The probability for the 24th month is computed after resetting it to 100 at 12 months.

rise in the inflow rate to LTU and a substantial drop in the outflow rate, leading to a higher incidence and a stronger persistence of LTU.

In line with our previous results, we find that individuals over 45 years old are especially prone to enter LTU, with estimated survival probabilities at 12 and 24 months of, respectively, 35% and 63% in the recession. The vulnerability of these workers is not new, however, as the survival probabilities for this group were also quite high in the expansion. Given our controls for other worker characteristics, this finding points at structural problems for older workers in rebuilding their working careers after job loss.

In the case of education we find surprisingly small and non-monotonic effects. For the expansion, we find a U-shaped pattern in the relation between education and the survival rate. The individuals with at most mandatory education enjoy the lowest survival rates while college graduates face the largest ones. The difference in survival rates is around 8 percentage points. This counter-intuitive result is likely due to the relatively intensive growth of low-skilled jobs during the expansion, which was fueled by the housing boom (Bonhomme and Hospido 2017). Indeed, for the recession period, characterized by a steep drop in the demand for low-skill labor, the differences across all educational attainment groups are much smaller. For example, the difference in the chances of becoming long-term unemployed between college and compulsory education graduates is reduced to 3 pp. The U-shaped relationship is also present for our occupation-based measure of skill, with less skilled workers showing higher survival rates in the recession. Apart from labor demand, another factor underlying this finding could be that higher job opportunities for more skilled workers are hampered by their higher reservation wages. On the other hand, in our companion estimation of employment hazards we do find that more educated and more skilled workers have significantly longer employment durations (Bentolila *et al.* 2017).

Table 5 reveals that higher work experience does significantly reduce the chances of becoming long-term unemployed. Moving from the 10th to the 75th percentile of the distribution of actual work experience reduces the survival rate at 12 months by around 50% in both the expansion and the recession. This accords with more experienced workers having higher expected productivity.

As expected, we find that workers from construction have the lowest survival rate (at twelve months) in the expansion. What is somewhat more surprising is that they still enjoy the second-lowest survival rates in the recession, despite the bursting of the housing bubble. The explanation for this finding seems to be the comparatively high degree of turnover in construction. Almost 60% of jobs in this industry last less than 3 months, compared to 44% for the other industries. Thus, short-duration temporary contracts seem to act as an informal work-sharing arrangement, allowing a relatively large share of unemployed construction workers to find work and thus avoid entering LTU. The alternative explanation of a high degree of occupational mobility is belied by the data. Indeed, inspection of post-unemployment outcomes reveals that unemployed construction workers are less likely to take up jobs in a different industry than workers from the other industries. And those who do mostly relocate to non-knowledge intensive

service industries (see Jansen *et al.* 2016 for details).

At first sight these results may seem inconsistent with the fact that almost 20% of the very long-term unemployed used to work in Construction (see Table 2). However, it should be recalled that this industry made up around 13% of total employment before the crisis and, contrary to our initial descriptive analysis, we are now controling for worker characteristics. The prevalence of high-school dropouts and other unfavorable characteristics among construction workers shows up in low average exit rates, but differences with other industries become much less hampering once we control for those characteristics.

The largest adverse effects on unemployment exit rates are associated with the receipt of unemployment benefits.[17] As shown in Figure 9, during the recession a worker with a 12-month entitlement to contributory benefits has an almost 42% chance of entering LTU. By contrast, for a similar worker with no benefits this probability is 30 pp lower. Similarly, conditional on having entered LTU, a 24-month benefit entitlement raises the probability of entering VLTU by a further 30 pp vis-à-vis a comparable worker without UI benefits (0.656–0.36), or 35 pp unconditionally (0.607×0.656–0.121×0.36).[18] In terms of entry into LTU, the effect for a worker without benefits of getting just a 6-month UI entitlement is larger than the effect of the economy moving from expansion to recession –namely, from 5.7% to 13% in the first column vs. going from the first figure to 12.1% in the third column. And much higher differences are found for higher entitlement lengths. The size of the change in the survival rate attached to unemployment assistance benefits is similar. These differences in survival rates suggest that workers with unemployment benefits either exert a relatively low search effort and/or have a high reservation wage, and that, when they step up their effort to find a job, they are affected by strong duration dependence, so that at that point their chances of leaving unemployment are much lower than the initial ones. In the next Section we take a look at declared reservation wages.

In sum, we find that the probabilities of entering LTU and VLTU are quite large and that they increased significantly in the recession. The determinants of both states are quite similar and we obtain evidence of strong duration dependence. While individual characteristics like educational attainment and skill are not associated with large changes in the rate of exit from unemployment, there are larger changes attached to mature age, low experience, and receipt of unemployment benefits.[19] The latter finding points at the lack of active labor market policies.

[17]Large effects of unemployment benefits on unemployment duration have been found in many articles, like those cited regarding Spain in the Introduction. Recent research has moved towards exploiting natural experiments, see Moffit (2014) for an overview. For Spain, Rebollo-Sanz and Rodríguez-Planas (2016) have found a strong impact on the job-finding probability of the reduction in the benefit replacement rate that took place in Spain in 2012.

[18]Note that the rates of survival to 24 months shown in the table are the same for all benefit entitlements up to 12 months since, by construction, upon exhaustion of their benefits, the monthly hazard rate for workers with UI benefits converges to the hazard rate for workers without benefits.

[19]Several of these results echo those in Nagore and van Soest (2016b), who analyze unemployment exits during the crisis using the same data source but different methods.

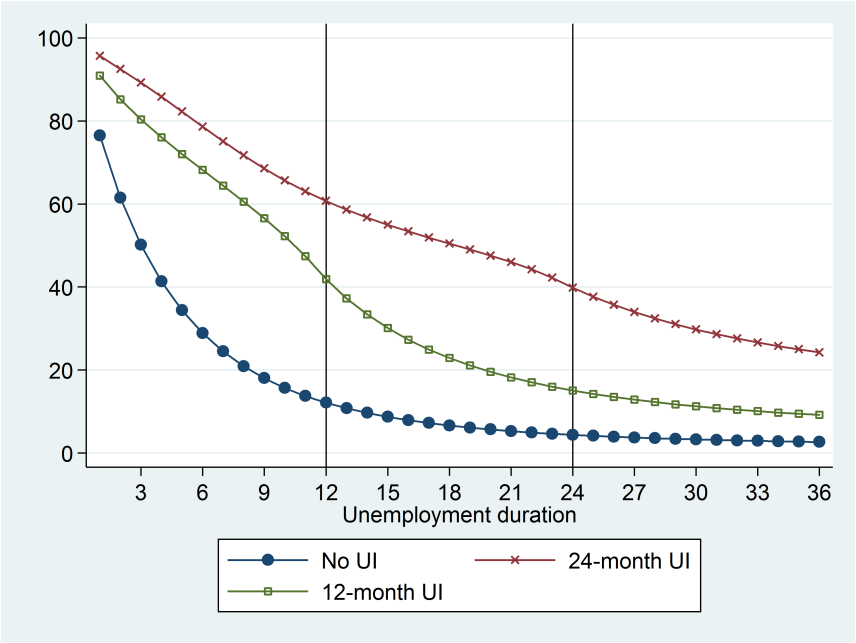


Figure 9: Effect of unemployment insurance (UI) on the survival rate in unemployment

Source: own computations.

3.5 Alternative specifications

Given the prevalence of temporary labor contracts in Spain, many of the jobs that unemployed workers find are very short-term. We may therefore wonder what types of workers mostly benefit from the availability of those short-duration jobs and whether these jobs act as stepping-stones to more stable employment. One way to look at this issue is to treat exits from unemployment to employment spells that last less than 30 days as right-censored rather than as standard exits.[20] We can then examine the resulting changes in the estimated coefficients associated with worker characteristics, which we again do through the lens of survival rates to 12 and 24 months in unemployment.

The full results are reported in Appendix Table A.3, while the most relevant changes are presented in Table 6. By construction, the average survival rates turn out to be larger than in the baseline, by 17.4 and 14.5 pp, respectively, for LTU and VLTU. But the important point is that the effect of ignoring short-duration employment spells varies across population groups. To see this, notice that the increase in the LTU rate for college graduates is 6.7 pp less than for workers without secondary education. Similarly, the survival rate in LTU for workers at the first decile of the distribution of experience increases by 9.8 pp more than for those at the 75th percentile. Furthermore, in this alternative specification the estimated survival rates at both 12 and 24 months are monotonically decreasing in the educational attainments of the unemployed. These

[20] We also adjust the estimation of the employment hazard accordingly.

Table 6: Survival rates in unemployment at 12 and 24 months in the recession. Censored model (%)

	Survival rate		Change vs. baseline	
	12 months	24 months	12 months	24 months
Overall	42.9	65.7	17.4	14.5
Education				
Primary or less	47.6	67.8	19.2	16.0
Secondary, 1st st.	42.7	63.6	18.5	16.5
Secondary, 2nd stage	42.8	64.0	16.3	13.2
College	40.1	61.4	12.5	10.7
Experience				
P75	31.4	56.0	13.0	12.6
P50	37.8	61.6	15.5	13.8
P25	50.7	71.6	19.9	15.1
P10	63.4	80.0	22.7	15.1
Unemployment insurance				
No benefits	28.3	53.4	16.2	17.5
6 months	36.4	53.4	15.7	17.5
12 months	54.8	53.4	12.9	17.5
18 months	63.7	61.8	11.0	15.0
24 months	71.4	75.5	10.7	9.9
Unemployment assistance				
No	28.3	53.4	16.2	17.5
Yes	59.5	76.6	15.1	9.3

Note: The sample is made of males aged 25-54 years old. The recession corresponds to the period 2008-2014. The probability for the 24th month is computed after resetting it to 100 at 12 months. The last two columns report the change in the survival rate in pp vis-à-vis the baseline estimates given in Table 5.

results indicate that the somewhat counterintuitive U-shape pattern of the survival rates in our baseline specification is a reflection of the rotation of relatively low-skilled and inexperienced workers on short-duration contracts of less than a month.

One objective of future labor market reforms should be to improve access to stable jobs, especially for workers at the bottom of the skill distribution. The latter may require restrictions on the use of temporary contracts, but policymakers should take into account that such policy might have adverse effects on the current long-term unemployed. A short-duration temporary contract may be the only viable option after years of unemployment and in some cases such spells may improve the subsequent job-finding rate of the long-term unemployed. The latter issue is extremely important for policy purposes, but an analysis of the causal impact of short-duration spells or placements on the subsequent job-finding rates of the long-term unemployed is left for future research, as it requires different techniques.[21]

Next, in an attempt to test the robustness of our results, we have also estimated an alternative model that includes a control for the type of contract in the previous job –temporary or permanent– and its interaction with the linear term of UI entitlement. The objective is to minimize the possibility that our controls for benefit entitlements capture unobserved traits of the individuals that are not captured by our controls for time-invariant unobserved heterogeneity. For example, some workers may wish to reenter unemployment periodically once they qualify for UI benefits. These persons may be more inclined to accept temporary jobs and their propensity to exhaust their benefit entitlements is likely to be relatively strong. The estimated effects of specific worker characteristics indeed differ with respect to our baseline specification, but in terms of the overall impact on entry to LTU or VLTU the estimates are very similar, both qualitatively and quantitatively. The largest difference appears in the effect of having a 24-month entitlement to UI benefits on the survival rate at 12 months, which is 12 pp lower in this alternative specification.[22]

4 Reservation wages during the crisis

Given the reduced-form nature of our estimation, we cannot distinguish between demand and supply factors. This raises the question of whether exit rates are low because

[21]This impact could be estimated by means of a natural experiment. For example, the public employment services could introduce variation in the payment structure of private placement agencies. Some agencies could be rewarded for placements of any length, while others would only receive a reward for placements with a minimum duration. Currently, the minimum requirement is typically six months of continuous employment, but it is questionable whether this requirement is optimal for the hard-to-place long-term unemployed. The estimation of duration models with alternative definitions of LTU, that either ignore certain short-duration employment spells like we do here or that consider the accumulated length of all unemployment spells during a pre-specified period, is also an option. But in this case it is challenging to identify truly causal effects. The best that one could do would be to compare the predictive power of the different specifications using out-of-sample observations.

[22]These results are available upon request.

labor demand is low or because the long-term unemployed have too high reservation wages. To answer this question we analyze the adjustment of self-reported reservation wages during the recession.

The contents of this section are related to an ongoing discussion on wage cyclicality in Spain. Real wage rigidity is well-known to be high in Spain.[23] Reentry wages of unemployed workers have however fallen considerably during the crisis and this process started as early as 2010. For example, for the prime-age males in our sample, the average reentry real wage dropped by 15.3% between 2009 and 2014 (while it grew by 5.2% between 2005 and 2009). This suggests a relatively strong adjustment in reservation wages. However, the evolution of reentry wages depends on the composition of the unemployed who reenter employment and their decisions to accept job offers. Moreover, the difference between the reentry wages of the short- and long-term unemployed in our sample is very small. This could point at the need for further reductions in the reservation wages of the latter, but it might also reflect the fact that only the best-qualified long-term unemployed manage to find a job. For this reason, we pursue an alternative route, by analyzing the behavior of self-reported reservation wages.

4.1 Self-reported reservation wages in the crisis

We use data from the Spanish Survey of Family Finances (Encuesta Financiera de las Familias, EFF). This survey asks unemployed respondents the following question: "At what gross monthly wage would you be willing to work?". We keep only people who report a positive nominal reservation wage up to 4,000 euros per month in any year.[24] To avoid having people with a low job search effort, we exclude respondents who report being unemployed for more than five years.

We pool together four waves, corresponding to years 2002, 2005, 2008, and 2011. The sample size is small, 2,816 observations, of which 1,538 are for men. The data constitute an incomplete panel, but the panel dimension is small, since around 80% of individuals are only observed once. For this reason, we estimate by pooling the data. The descriptive statistics of the variables are presented in Table A.4. The average monthly reservation wage is around 1,200 euros (at 2011 prices), the average duration is high –2 years–, one-third of individuals are receiving unemployment insurance benefits and 10% are getting unemployment assistance benefits. By household, average gross real income is around 40,500 euros per year and average real net wealth is around

[23] According to the estimates with MCVL data for 1987-2013 by Font *et al.* (2015), a 1 pp increase in unemployment leads to a fall in real wages of -0.24 pp in the recession, and between -0.38 and -0.48 in the expansion. Real wage elasticity is found to be about 70% higher for job movers (see also De la Roca 2014). Izquierdo and Puente (2015) suggest that there has been a structural change after the 2012 labor reform, which changed the regulation of wage bargaining (García-Pérez and Jansen 2015), finding a 0.13 pp increase in the real wage elasticity for 2012Q3-2013Q4. Nonetheless, these elasticities are small in comparison with the estimates above unity in Pissarides (2009) for the US and several European countries. Moreover, the elasticity is countercyclical, i.e. employment destabilizing.

[24] There are only 11 observations above this threshold; keeping them slightly raises the estimated elasticity to unemployment duration.

390,000 euros.

Our empirical specification is as follows:

$$log(\omega_{it}) = \alpha_t + \beta log(Dur_{it}) + \gamma UI_{it} + \delta UA_{it} + X'_{it}\mu + u_{it}$$

where subindex i denotes an individual, t is time, α_t is a year fixed effect, Dur is unemployment duration measured in years, UI is an indicator for the receipt of unemployment insurance benefits, UA is an indicator for the receipt of assistance unemployment benefits, and X is a vector that contains the following variables: age, age squared, marital status (single, married, unmarried partner, separated, divorced), education (secondary first stage or less, secondary second stage, college), household-head status, household size (number of members), annual household gross real income (in the preceding year), real non-financial wealth, real financial wealth, and real debt. All monetary variables are deflated with 2011 as the base year. The EFF provides five imputations for missing values in most variables. We estimate the wage equation with each of the five data sets and then compute simple averages of the coefficients and corrected standard errors (Banco de España 2015). Standard errors are clustered at the household level.

In view of the small sample size we report estimates for the total sample, including a dummy variable for females, and then for males alone. The first column of Table 7 shows that women have significantly lower reservation wages, but we should recall that, on average, women also earn lower wages even after controling for observable characteristics.

Reservation wages are higher for male workers who are older, more educated, and married or cohabiting. They are also larger the higher are the household's size, income, and financial wealth. Surprisingly, real assets are not significant and debt attracts a positive sign, so that it is not net wealth that matters. Receipt of contributory benefits is significant for the full sample but not for males, and assistance benefits are not significant for any sample.

An important coefficient is the elasticity of the reservation wage to unemployment duration, which is equal to –1.6% overall and –2.0% for males. These are very small elasticities, which suggest that reservation wages do not respond very strongly to unemployment duration. Nevertheless, this finding cannot be directly linked to the probability of leaving unemployment, which also depends on the actual level of the reservation wage, job search effort, and the job-offer arrival rate (García-Pérez 2006).

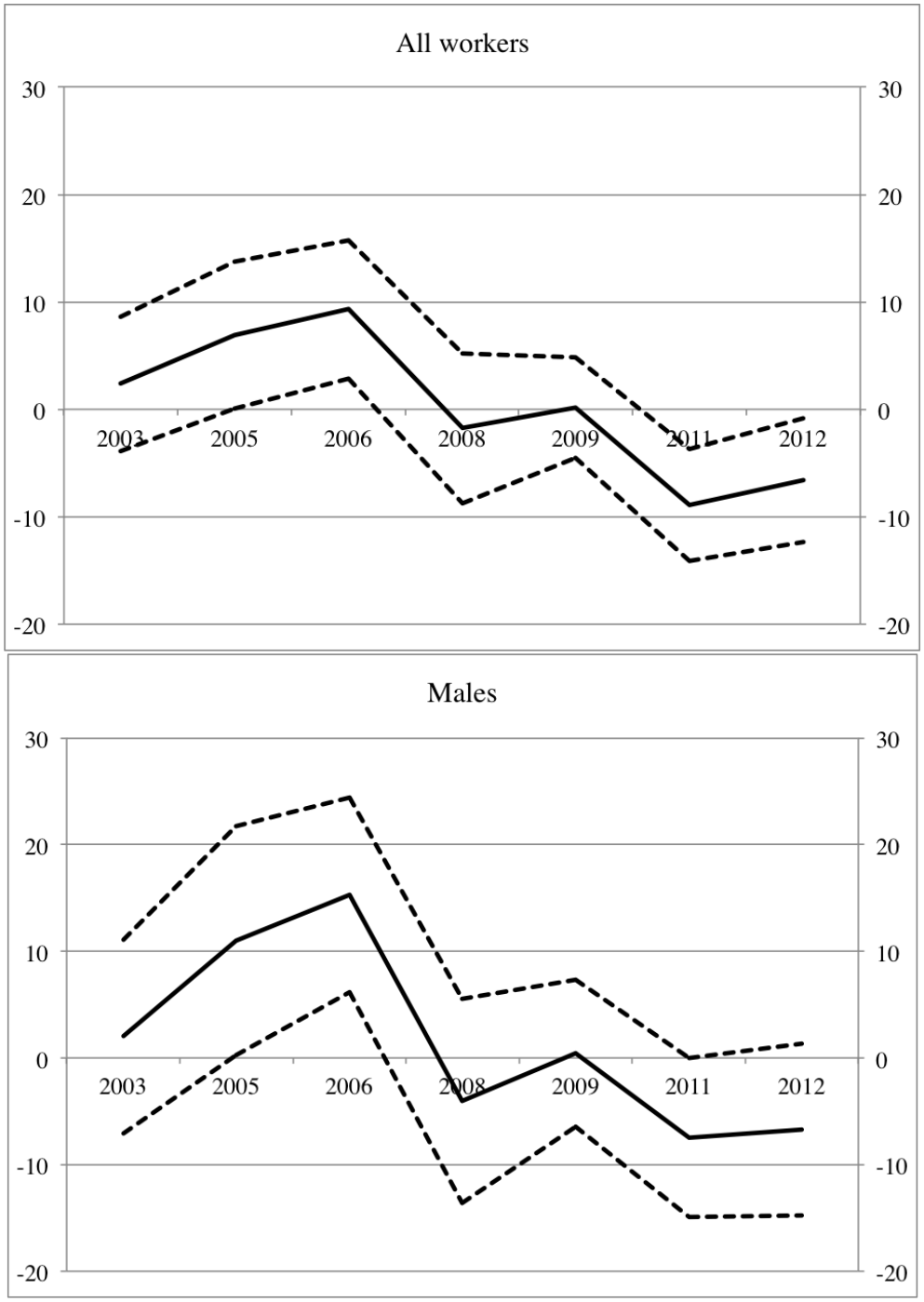
The profile of the coefficient on the yearly dummy variables, depicted in 10, shows that declared reservation wages do indeed trace the business cycle. With respect to 2002, there are significant increases for all workers in the boom years of 2005 and 2006 and large reductions in 2011 and 2012. The pattern is even stronger for males, although it is estimated with less precision due to the smaller number of observations.

Compared to the existing literature on self-reported reservation wages, we provide better control for the financial situation of households. Our findings are at variance with the results in Koenig *et al.* (2016), who use data on self-reported reservation

Table 7: Reservation wages
Dependent variable: Log monthly reservation wage

	All		Males	
Log duration	-0.016	***	-0.020	**
	(0.006)		(0.008)	
Contributory benefits	0.055	***	0.021	
	(0.015)		(0.020)	
Assistance benefits	0.013		-0.022	
	(0.021)		(0.029)	
Age	0.004	***	0.005	***
	(0.001)		(0.001)	
Female	-0.159	***		
	(0.013)			
Married	0.059	***	0.070	***
	(0.018)		(0.026)	
Unmarried partner	0.057	**	0.107	***
	(0.028)		(0.041)	
Separated	0.033		0.011	
	(0.040)		(0.055)	
Divorced	0.076	*	0.054	
	(0.048)		(0.062)	
Household head	-0.014	**	-0.011	
	(0.006)		(0.007)	
Secondary education, 2nd stage	0.031	**	0.048	**
	(0.014)		(0.020)	
College	0.070	***	0.056	***
	(0.014)		(0.018)	
Household size	0.224	***	0.194	***
	(0.022)		(0.032)	
Total income	0.014	**	0.024	**
	(0.007)		(0.012)	
Real assets	0.001		0.002	
	(0.002)		(0.002)	
Financial assets	0.013	***	0.011	***
	(0.002)		(0.003)	
Debt	0.005	***	0.005	**
	(0.001)		(0.002)	
Observations	2,816		1,358	
R^2	0.251		0.262	

Note: The specification includes also year dummy variables. Standard errors are clustered at the household level. * $p<0.10$, ** $p<0.05$, *** $p<0.01$.



Source: own computations. Dashed lines represent 95% confidence intervals.

Figure 10: Annual profile of reservation wages, deviations from 2002 (%)

wages for the United Kingdom and Germany –from 1991 to 2009 and from 1984 to 2010, respectively– and find low and borderline significant elasticities to the aggregate unemployment rate. Outside the self-reported reservation wage literature, however, the strong and volatile cyclical behavior, that we associate with the coefficients on the time fixed effects, is also found by Chodorow-Reich and Karabarbounis (2016), who compute, for the US from 1961 to 2012, the flow value of the opportunity cost of employment, dividing it into a benefit component and the forgone value of nonworking time. While the first component is small and countercyclical, the second one –which is closer to our measure, since we are controling for the receipt of unemployment benefits– is procyclical and volatile over the business cycle.

The low response of reservation wages to the receipt of unemployment benefits has also been found by Krueger and Muller (2016). These authors collected a Survey of Unemployed Workers in New Jersey, in 2009-2010, and cannot find a significant relationship between reservation wages and benefits. Like in our sample, however, these wages are positively correlated with financial variables like severance payments and savings. Lastly, our results partially differ from those in Addison *et al.* (2009) using data from the European Community Household Panel over 1994–1999. Only in 7 out of 13 countries in their sample do self-reported reservation wages respond significantly to the receipt of unemployment benefits, but Spain is one of them, with a significant coefficient of 0.092. However, their equation is less informative than ours, since it only includes gender, schooling, age, and time fixed effects as controls. On the other hand, they do not find a significant elasticity of the reservation wage to unemployment duration in most countries, but in the Spanish case it is significant and low, around 1%, which is smaller but close to our estimates.

5 Conclusions

In this paper we have analysed the determinants of the buildup and the persistence of the exceptionally high levels of LTU in Spain. Our empirical analysis indicates that mature age, low experience, and entitlement to UI benefits are the main risk factors. Low educational attaintments and low skill levels also raise the chances of entering LTU, but to a lesser extent. Moreover, two of the most striking findings are the relatively low risk of LTU for workers from construction and the pervasive impact of negative duration dependence. The job-finding hazard of the average worker drops by 55% during the first twelve months and halves again during the next twelve months. These effects are much larger in size than the cross-sectional differences in job-finding rates across unemployed workers with the same duration. Moreover, over time the job-finding rates become less responsive to improvements in the aggregate labor market conditions.

The overall conclusion of our analysis is that the current levels of LTU entail a substantial risk of social and economic exclusion. Growth alone will not be sufficient to eradicate this risk and the room for further wage adjustments seems very limited due to the strong adjustment in reentry wages that is matched by a strong decline is

self-reported reservation wages. Spain will therefore have to step up its effort to design effective active labor market policies that help to improve the employment prospects of the long-term unemployed. Moreover, with a view to the future, Spain should intensify the early activation of the unemployed in order to avoid that the receipt of benefits reduces job search intensity.

Our study offers no clear prescriptions for the design of efficient active labor market policies for the long-term unemployed. Nonetheless, the recent meta-analysis of Card *et al.* (2015a,b) indicates that these policies can make a significant contribution in reducing unemployment and especially LTU. Investments in training and hiring incentives that are carefully targeted at the LTU deliver the best long-term results while public employment programs tend to deliver the poorest results, but the impact estimates vary considerably between different studies. The appropriate design of the programs is therefore key and the interventions need to be tailored to the need of each participant. If these conditions are satisfied, the programs for the long-term unemployed are often both effective and efficient (Csilag and Fertig 2015).

After a long period of inaction, Spanish policymakers are slowly recognising the need to develop effective tools to fight LTU. Following a recent recommendation of the European Council, Spanish authorities have approved a three-year program –the Programa de Acción Conjunta para Desempleados de Larga Duración– to offer individualised support to one million long-term unemployed people. This support involves the assignment to a personal tutor and the preparation of an individual integration agreement. However, this is nothing but the first step. Spain has a poor track record in the field of active labor market policies and its public employment services are outdated and play at best a marginal role as intermediaries on the labour market (Jansen 2016a,b). These problems need to be addressed before we may expect positive results from the recently announced plan.

Appendix

A Additional empirical results

This Appendix provides additional empirical results on the exit rate from unemployment and reservation wages.

A.1 Hazard rates and survival probabilities

Table A.1: Estimates of the hazard of leaving unemployment

	Expansion				Recession			
	Exit to temporary contract		Exit to permanent contract		Exit to temporary contract		Exit to permanent contract	
	Coeff.	z	Coeff.	z	Coeff.	z	Coeff.	z
Age 35-44 years old	-0.035	-0.89	0.180	2.05	-0.106	-4.69	0.232	4.36
Age 45-54 years old	-0.586	-4.16	0.442	2.04	-0.295	-5.00	0.438	3.79
Age 35-44 y.o.×log Dur	-0.064	-2.51	-0.038	-0.65	-0.018	-1.55	-0.135	-4.34
Age 45-54 y.o.×log Dur	-0.006	-0.07	-0.127	-0.93	-0.067	-2.06	-0.210	-3.07
Secondary education, 1 st.	0.107	2.66	-0.033	-0.32	0.160	4.86	-0.185	-2.41
Secondary education, 2 st.	-0.079	-1.75	-0.093	-0.82	0.103	2.72	-0.111	-1.31
College education	-0.348	-6.64	-0.117	-0.92	-0.024	-0.53	0.101	1.05
Secondary ed 1 st.×log Dur	0.008	0.33	0.147	1.96	0.003	0.19	0.166	3.43
Secondary ed 2 st.×log Dur	0.004	0.14	0.172	2.17	-0.035	-1.83	0.180	3.45
College education×log Dur	0.018	0.60	0.191	2.25	-0.027	-1.16	0.193	3.33
High skill	0.093	2.64	0.224	2.67	0.273	7.38	0.015	0.18
Medium skill	0.187	6.29	-0.079	-0.98	0.347	10.11	-0.155	-1.92
High skill×log Dur	-0.088	-4.09	0.020	0.37	-0.093	-5.01	0.173	3.48
Medium skill.×log Dur	-0.071	-3.82	0.130	2.40	-0.071	-4.15	0.171	3.49
Dismissal	0.259	12.67	-0.004	-0.08	0.601	26.33	-0.469	-12.19
Experience	0.471	12.89	1.730	18.47	1.031	31.63	1.596	22.48
ΔEmployment	2.405	3.37	-5.286	-3.06	4.142	14.25	2.904	3.78
ΔEmployment×log Dur	-0.663	-1.50	1.122	0.96	-0.711	-4.49	-0.002	0.00
Unemployment rate	0.010	1.19	-0.220	-14.25	-0.017	-5.63	-0.066	-10.51
Unempl. rate×log Dur	-0.021	-3.63	0.053	4.13	0.006	4.94	0.019	6.57
Labor reform 2010					-0.072	-3.25	-0.073	-1.27
Labor reform 2012					0.009	0.51	0.158	3.47
log Dur	-0.378	-4.43	-0.991	-5.21	-0.693	-13.77	-0.850	-7.06
$(\log \text{Dur})^2$	0.255	6.22	-0.052	-0.47	0.337	11.37	-0.032	-0.39
$(\log \text{Dur})^3$	-0.066	-7.28	0.000	-0.01	-0.104	-16.79	-0.029	-1.72
Unemployment insurance	-0.436	-33.82	-0.381	-11.92	-0.380	-37.52	-0.157	-7.04
Unemployment insurance^2	0.042	22.32	0.031	6.84	0.035	24.83	0.014	4.55
Unemployment insurance^3	-0.001	-17.91	-0.001	-4.76	-0.001	-20.30	0.000	-3.80
U. insurance×log Dur	-0.026	-3.55	-0.039	-2.51	-0.024	-4.95	-0.035	-3.28
U. insurance×$(\log \text{Dur})^2$	0.006	1.63	0.021	2.80	0.009	3.95	0.009	1.85
U. assistance	-1.083	-32.14	-1.086	-11.38	-1.107	-46.11	-1.287	-18.78
U. assistance×log Dur	0.036	1.44	0.111	1.63	0.030	1.96	0.187	4.29
No. of spells	37,399				62,045			
No. of observations	1,346,016				1,641,889			
Log likelihood	-352,712.09				-446,462.87			

Note: The sample is made of males aged 25-54 years old. The expansion corresponds to the period 2001-2007 and the recession to 2008-2014.

Table A.2: Unobserved heterogeneity estimates, 2001-2014

	Males			
	Expansion		Recession	
Types	Coeff.	z	Coeff.	z
η_1^u	-2.19	-17.98	-1.83	-23.20
η_2^u	-0.96	-7.93	-3.17	-40.97
η_1^e	-2.68	-14.37	-3.59	-23.58
η_2^e	-1.73	-9.24	-2.59	-16.94
Probabilities				
$Pr(\eta_{fast}^u, \eta_{slow}^e)$	0.32	17.93	0.23	10.18
$Pr(\eta_{fast}^u, \eta_{fast}^e)$	0.15	–	0.11	3.68
$Pr(\eta_{slow}^u, \eta_{slow}^e)$	0.50	30.48	0.58	–
$Pr(\eta_{slow}^u, \eta_{fast}^e)$	0.03	6.63	0.08	11.99

Note: The sample is made of males aged 25-54 years old.

A.2 Descriptive statistics of the reservation wage sample

Table A.3: Survival rates in unemployment at 12 and 24 months with censoring of short jobs (%)

	Expansion		Recession	
	12 months	24 months	12 months	24 months
Overall	26.7	48.0	42.9	65.7
Age				
25-34 years old	26.2	47.2	40.7	63.6
35-44 years old	28.4	52.0	45.3	68.0
45-54 years old	40.1	63.7	52.5	75.5
Education				
Primary or less	27.5	47.8	47.6	67.8
Secondary, 1st st.	24.1	44.2	42.7	63.6
Secondary, 2nd stage	27.5	46.6	42.8	64.0
College	30.8	48.2	40.1	61.4
Skill				
High	27.9	50.3	43.6	66.8
Medium	24.6	46.1	40.9	64.2
Low	28.4	48.0	46.8	67.1
Experience				
P75	19.0	39.2	31.4	56.0
P50	22.6	43.5	37.8	61.6
P25	32.8	54.1	50.7	71.6
P10	43.1	63.4	63.4	80.0
Industry				
Manufacturing	22.1	42.2	38.4	61.0
Construction	23.3	46.2	40.1	62.4
Non-market services	31.2	52.0	47.4	71.1
Trade	26.1	44.9	42.2	63.8
Hospitality	24.5	42.6	38.3	61.9
Other services	30.8	53.2	46.4	70.1
Unemployment insurance				
No benefits	16.0	35.5	28.3	53.4
6 months	25.1	35.5	36.4	53.4
12 months	46.2	35.5	54.8	53.4
18 months	55.2	45.8	63.7	61.8
24 months	63.4	65.9	71.4	75.5
Unemployment assistance				
No	16.0	35.5	28.3	53.4
Yes	43.8	62.1	59.5	76.6
GDP growth				
High	24.6	44.6	39.2	64.1
Low	28.9	51.6	46.1	67.0

Note: The sample is made of males aged 25-54 years old. The expansion corresponds to the period 2001-2007 and the recession to 2008-2014. The probability for the 24th month is computed after resetting it to 100 at 12 months.

Table A.4: Descriptive statistics for reservation wage sample

	Mean	Std. dev.
Reservation wage (monthly)	1,177.3	511.5
Unemployment duration	2.0	1.4
Contributory benefits	33.1	51.6
Assistance benefits	9.8	32.6
Age	36.6	13.8
Female	45.4	54.5
Single	50.4	54.8
Married	36.6	52.8
Unmarried partner	5.1	24.2
Separated	3.5	20.1
Divorced	3.3	19.4
Household head	35.5	52.4
Secondary education, 1st stage or less	45.8	54.6
Secondary education, 2nd stage	35.7	52.5
College	18.5	42.5
Household size	3.5	1.5
Total income	40,532.6	59,082.0
Real assets	350,236.4	1,613,022.5
Financial assets	77,356.3	700,874.5
Debt	38,007.4	360,404.6

Note: 2,816 observations. All variables are percentage shares except monetary variables, which are in 2011 euros, age and unemployment duration which are in years, and household size which is the number of members.

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