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Unemployment Exits Before and During the Crisis

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Abstract

Using administrative data from Spanish Social Security, we compare the pattern and the determinants of individual unemployment durations and the stability of jobs found after unemployment before and during the recent crisis. We find particularly strong effects of the crisis on the hazards in the beginning of the unemployment spell. The groups hit hardest by the crisis are men, immigrants, older workers, and individuals with lower levels of education. The disadvantage of men is mainly due to the more pro-cyclical nature of men´s jobs. Decompositions show that the increase in average unemployment duration and the decrease in average duration of the new job during the crisis are not explained by changing characteristics of the individuals who become unemployed.

Keywords: Unemployment durations; Job durations, Business cycle; Re-employment probability

JEL classification: J64, C41, E32

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1. Introduction

The recent economic recession in Spain has led to important adjustments in the labour market, with a reduction of working hours and the dismissal of many workers. The case of Spain is particularly dramatic compared to many other countries that suffered from the crisis. According to the Spanish Labour Force Survey (SLFS), the unemployment rate in Spain rose from 8.5% in 2006 to 25% in 2012. Young workers were strongly affected, with youth unemployment reaching 55% by the end of 2012. The long-term unemployment rate rose from 2% in 2006 to 14% by the end of 2012. This is specifically worrying because long-term unemployment implies a loss of human capital, reducing welfare and increasing the risk of social exclusion. Unemployment also has important consequences for the sustainability of the Social Security system, reducing contributions and increasing the amount of benefits to be paid.

In order to understand the nature of unemployment it is important to consider both entry into and exit out of unemployment. In this study we analyse transitions from unemployment to any job before and during the Great Recession as well as the duration of the jobs found after an unemployment spell, which can be seen as an indicator of the quality of the match. We contribute to the microeconomic literature on the impact of the crisis on labour market outcomes in several ways. We focus on identifying the groups of workers who suffered most from the crisis, in terms of longer unemployment durations or less stable subsequent jobs. In addition, we analyse how the effects of the Great Recession relate to the cyclical variation in local labour market conditions. Finally, we decompose the changes in the transition rates into sample composition effects and residual changes induced by changing economic conditions.

We compare unemployment and subsequent job duration patterns and their determinants in a period of expansion (2005-2007) and the recent recession (2009-2011). The differential effects of the crisis across socio-economic groups will show which unemployed workers suffered most from the crisis, in terms of a larger decrease in re-employment probability or in stability of the job found after an unemployment spell. Workers' Protection systems and active labour market policies can target such vulnerable groups. This makes our results useful to address some of the remaining challenges stated in European Commission (2015), such as reducing unemployment

and promoting stable employment. The Europe 2020 strategy explicitly recommends special attention for vulnerable groups (European Commission, 2010).

Our data come from the Longitudinal Working Lives Sample (LWLS), based upon administrative records from the Spanish Social Security Administration. The use of an administrative dataset in the analysis of transitions avoids the bias associated with misreported transitions and the detailed information on durations in days avoids the time aggregation bias, which may be quite relevant in a country with high mobility like Spain.

LWLS contains detailed information on employment and unemployment transitions, and individual and job characteristics. Following Arulampalam and Stewart (1995), we compare two inflow samples at different points in time. We construct two separate comparable and homogeneous samples that include all unemployment spells with or without benefits starting in the calendar years 2005 and 2009. We include multiple spells of the same individuals (following Imbens and Lynch, 2006). We observe individuals who enter unemployment in these years until the end of their subsequent job or the end of the observation period - 31 December 2011 for the 2009 data and, for comparability, 31 December 2007 for the 2005 data.[3] This procedure avoids sample selection problems (see Ham and Lalonde, 1996).[4] It implies that we can interpret the results as differences between unemployment and subsequent job spells that started in different economic contexts. Our samples are representative for unemployed workers who lost their job in 2005 or 2009, not for labour market entrants looking for their first job (and certainly not for the complete workforce).

For both samples, following Ham and Lalonde (1996), we estimate a bivariate continuous time hazard rate model for unemployment to any job and for subsequent job to unemployment, with potentially correlated unobserved heterogeneity terms. In the benchmark model, the explanatory variables only include individual characteristics and the regional unemployment rate. An extended model also includes (previous or current) job characteristics.

We find particularly strong effects of the crisis on the hazards in the first year (and specifically the first few months) of the unemployment spell. The groups most

[3] Unemployment figures started to rise substantially in 2009.
[4] We do have right censored spells; this is accounted for in the empirical analysis.

negatively affected by the crisis are men, immigrants, older workers, and individuals with low education level. The results of the extended model with job characteristics show that the difference between men and women is mainly due to the more pro-cyclical nature of men´s jobs. The average characteristics of workers who become unemployed in 2005 and 2009 differ significantly. Still, decompositions show that only a small part of the increase in average unemployment duration and the decrease in stability of the new job during the crisis are explained by changing characteristics of the individuals who become unemployed.

The remainder of the paper is organized as follows. Section 2 positions the paper in the existing literature. Section 3 briefly explains the relevant characteristics of the Spanish unemployment benefit system. Section 4 describes the data. Section 5 introduces the econometric framework of unemployment and subsequent job durations. Section 6 provides the main results and conclusions are drawn in Section 7.

2. Background

Job search theory gives an ambiguous prediction of the relationship between the business cycle and the duration of unemployment. Increases in unemployment will reduce the reservation wage but also the probability of receiving a job offer. Lynch (1989) and Dynarski and Sheffrin (1990) found that higher unemployment results in lower re-employment probabilities. On the other hand, the models of Meyer (1990) and Solon (1985) suggest that the average duration of unemployment falls in a recession. Similarly, the quality of new matches and the duration of jobs found by the unemployed can also vary with the business cycle in different ways. During a recession, the employer may recruit better workers because of the larger applicant pool but workers may also be willing to accept poorer matches. The roles of entry and exit to explain the changes in the aggregate unemployment rate have been extensively studied in the macroeconomic literature, with diverging conclusions. For instance, for the US, Shimer (2012) found that the job finding rate explains three quarters of the volatility in the unemployment rate. Sahin et al. (2010) and Elsby et al. (2010) concluded that both an increasing inflow into unemployment and a decline in outflow explain the recent upswing in unemployment rate. The latter two papers also found that specific groups

(men, young, low-skilled and minority workers) experienced greater cyclicality of unemployment mainly due to a more pronounced increase of the inflow rate.

For 14 OECD countries, Elsby et al. (2013) found that fluctuations in both inflow and outflow rates contribute substantially to the variation in unemployment, with varying weights across countries. For Spain, Silva and Vázquez-Grenno (2013) found that entry is more important at the early stage of the crisis while lower job finding rates drive high unemployment later on.

In the current paper, we do not focus on the business cycle variation in aggregate durations, but on the heterogeneity across different socio-economic groups. The literature suggests that three key mechanisms may explain differences in the impact of the recession across groups (Cho and Newhouse, 2013): occupational segregation, employment decisions of firms, and labour supply decisions. The nature of the shock, labour market adjustments, and institutions may also play a role.[5] The groups hit hardest by the crisis may therefore differ from one crisis to the next and across countries.

One strand of the literature analysing heterogeneity across socio-economic groups focuses on the ratios of unemployment, employment and non-participation (Cho and Newhouse, 2013, Hoynes et al. 2012). Other studies, more in line with our study, analyse transitions between labour market states, mainly the probability to find a job (Bachman et al. 2015, Bergin et al. 2015). Generally, these studies find that youth, men,[6] and racial and ethnics minorities are the groups most severely affected by a recession, while the effect of education is ambiguous.

Hoynes et al. (2012), analysing employment and unemployment of socio-economic groups for 1979-2011 in the US, found that the Great Recession is deeper and longer than the recessions in the 1980s, but the same groups were hit hardest: youth, low-educated workers, men, African-American, and Hispanic workers. Differences across groups are mainly explained by variation in the exposure to cycles across industries and occupations. Cho and Newhouse (2013), regressing the change in the trends in (un-)employment for 17 middle-income countries and comparing 2006-2008 with 2008-2009, found that men, youth, and workers with medium levels of education were most affected by the crisis.

[5] The period analysed did not have major labour market reforms that could contaminate the results.
[6] Estimates of ILO (2010) imply that the downturn has affected male and female outcomes in virtually identical ways.

Bachmann et al. (2015), using European Union data for the period 2005-2010, explored the effects of the crisis on the labour market transitions focusing on differences between countries, socio-demographic groups (gender, age, level of skills), and type of employment contract. They found important differences across countries and concluded that men and young persons were particularly hit by the crisis, both in terms of unemployment exit and entry. Medium-skilled workers were more affected by an increase in the unemployment entry probability, while high skilled workers experienced a relatively large fall in the job finding rate. On the other hand, Bell and Blanchflower (2011), using microdata from three Eurobarometer surveys, found that the more educated were relatively less likely to enter unemployment during recessions.

Bergin et al. (2015) explored the impact of socioeconomic and job characteristics on transitions from unemployment to employment and vice versa in Ireland, with separate models for pre- and post-Great Recession data. Controlling for job characteristics, they found that immigrants, men, and low educated workers were affected most severely by the crisis. Young people became much less likely to exit unemployment but also got a lower risk of losing their job.

Earlier studies on the length of the unemployment spells usually control for the business cycle including the current local unemployment rate (Van den Berg, 2001), not interacted with other characteristics. An exception is Arulampalam and Stewart (1995), who looked at the impact of the business cycle in a comparable way as we do, using two inflow cohorts at very different points in time.

Some existing studies analysed the impact of the Great Recession on unemployment and employment durations in Spain. Carrasco and García-Pérez (2015) found that immigrants are more sensitive to the business cycle than natives, in terms of unemployment as well as employment hazards. De la Rica and Rebollo (2015) and Nagore (2016) explored gender differentials in labour market transitions, using a macro and a micro approach, respectively. They found pro-cyclicality in the gender gap due to the more procyclical jobs of men. Nagore and van Soest (2015) analysed the stability of all new job matches and found that the increase in the probability of job loss during the crisis was particularly large for males, young workers, low educated workers, manual occupations, and workers in construction. The main novelty of the current study is that we analyse both the job finding rate and the stability of the new job of those who leave unemployment in a joint framework.

3. The Unemployment Benefit System in Spain

Since individuals receiving unemployment benefits are a substantial part of our final sample, it is relevant to summarize the main aspects of Spanish unemployment for the period under study, 2005-2011.[7] The system provides coverage to wage workers (excluding civil servants and domestic employees)[8] who lost their job, are willing to work, and have a minimum period of contributions to the Social Security System. There are two levels of protection: contributory (Unemployment Insurance Benefit, UIB) and assistance (Unemployment Assistance Benefit, UAB). UIB is based on actuarial and financial principles and covers unemployed workers who contributed for at least 12 months in the last six years preceding unemployment. On the other hand, UAB is a means-tested benefit available to unemployed workers not or no longer entitled to UIB. The minimum period of contribution required in this case is three months in the last six years. UIB duration increases with the contribution record, with approximately one month of benefits for three months of contributions, a minimum of four and a maximum of 24 months. The UIB amount includes contributions to old age pensions (largely paid by the Public Employment Service, SPEE) and is equal to 70% (during the first 180 days) or 60% (from the 181st day) of the average daily contributory base, calculated on contributions made during the 180 days prior to unemployment. The benefit level is related to the wage level prior to unemployment, with maximum and minimum depending on the number of dependants below age 26. For instance, the monthly UIB amount in 2005 was between €438.48 (no dependent children) and €1,233.23 (two or more dependent children). The amount of UAB is not related to the previous wage; it was €376 in 2005 for everyone. UAB duration depends on the family responsibilities, the age of the recipient, and the length of the contributory period in the last six years.

Table 1 summarizes the (maximum) unemployment benefit duration for UIB and UAB. For instance, the unemployment benefit duration is between 3 and 60 months for an unemployed worker more than 44 years old with family responsibilities. If the unemployed worker is older than 52 years old and entitled to an old age pension, the unemployment benefit may last until the age of retirement.[9]

[7] The main legislation reference for the period under study is the Royal Legislative Decree No 1/1994 of 20 June.
[8] Some groups (i.e., workers in agriculture) are covered but have different rules.
[9] Several changes were made in the reform of July 2012 (after our observation window), making the system less generous.

Table 1: Duration of unemployment benefits (UIB and UAB)

Nr. Of months contributed in the last six years (tenure)	Contributory Unemployment Benefits (months)	Assistance Benefits			
		With family responsibilities		Without family responsibilities	
		Younger than 45	Older than 44	Younger than 45	Older than 44
3	-	3	3	-	-
4	-	4	4	-	-
5	-	5	5	-	-
6-11	-	21	21	6	6
12-17	4	18	24	-	6
18-71	2 x integer(tenure/6)= 6,8,10…22	24	30	-	6
72	24	24	36	-	6
Older than 52	-	Until the age of retirement			
Others (*)	-	6, 12 or 18			

Source: Own elaboration from *Toharia et al. (2010)*
(*) returning emigrants, released from prison, disabled but able to work.

4. Data and Descriptive Statistics

Our data come from the Longitudinal Working Lives Sample[10] (LWLS) based upon administrative records of the Spanish Social Security Administration (SSA). LWLS is collected annually since 2004 and contains information on a four percent random sample of the population who ever had any relationship with the SSA in the sample period, paying contributions or receiving benefits. It has approximately one million people. Individuals in the 2004 LWLS remain in the sample as long as they have a relationship with SSA. LWLS contains information on the labour market histories of adults who ever did paid work. This database is useful for our study because of its longitudinal design and the rich information on employment and unemployment transitions and individual and job characteristics, e.g. gender, age, nationality, firm size, sector of activity, type of contract, and information on contributory and non-contributory benefits.

We constructed two samples that include all unemployment spells of workers who lost their job in 2005 (the expansion period) and 2009 (the recession), with or without unemployment benefits (UIB and UAB)[11] and including multiple spells of the

[10] For a detailed description of this data set, see Duran (2007), García-Perez (2008) and Lapuerta (2010).
[11] Unemployed individuals who have never worked are not included in LWLS since they have no relation (yet) with SSA. We only consider unemployment spells without benefits longer than 15 days for individuals with a certain degree of labour market attachment - UB receipt in the year of reference or at least 30 days of paid work experience.

same individual. Individuals in these samples who exit from unemployment to a paid job[12] remain observed until the end of either this job spell or the observation period. The latter is 31 December 2011 for the 2009 data and, for comparability, set to 31 December 2007 for the 2005 data. The data are obtained by merging the datasets LWLS 2005-2006-2007 and LWLS 2009-2010-2011. We applied several filters to our samples (see Table A1 in the Appendix). For instance, our samples are restricted to workers aged 16-65 in 2005 or 2009 who exclusively worked in jobs falling under the General Social Security Regime,[13] because of differences in benefit arrangements. We removed individuals with incomplete information and recoded overlapping spells[14] and dropped observations from Ceuta and Melilla (two small Spanish enclaves in Africa).

The length of an unemployment spells is measured as the difference (in days) between the date of ending a job and the date of starting a new one, irrespective of benefit receipt. Unemployment spells may therefore also include periods without job search (e.g. of discouraged workers)[15] or periods of emigration (e.g. younger workers looking for better job opportunities abroad). Exits from unemployment to self-employment or other labour market states than paid work are treated as right-censoring, as well as unemployment spells exceeding the end of the observation period. Analogously, the subsequent job duration is defined as the difference (in days) between the termination and starting dates of the job. Other exits, such as retirement or death, are considered right-censored.

Descriptive analysis

Our sample for 2005 consists of 87,950 individuals with 133,413 unemployment spells and 124,132 job spells. The 2009 sample has 133,197 individuals with 181,010 unemployment spells and 139,885 job spells. The difference between the two years reflects the large increase of the number of transitions into unemployment between 2005 and 2009 and the decline in the job finding rates.

The Kaplan Meier survival functions in the top panel of Figure 1 show the probability of not having found a job as a function of spell duration t. The median unemployment duration has increased from 72 days in the 2005 sample to 147 days in

[12] Those who become self-employed are no longer observed.
[13] Other Social Security Regimes include, for instance, Self-employment, Agriculture and Household Special Regimes.
[14] We keep the spells with 1) the highest part-time coefficient; 2) if these are equal, the longest duration, or 3), if these are both equal, the highest contributory base.
[15] Brandolini et al. (2006) recognise the need to include the group of non-participants when looking at labour market dynamics. They argue that participants and non-participants do not differ substantially in their job search activity.

2009. On the other hand, the bottom panel shows that the median duration of the job after an unemployment spell has decreased from 149 days in 2005 to 113 days in 2009. This suggests the impact of the crisis is larger for unemployment exit than for entry.

Figure 1: Kaplan Meier Survival estimates; exits from –unemployment to any job (top panel) and from job to unemployment (bottom panel). 2005 and 2009 samples. Durations in days.

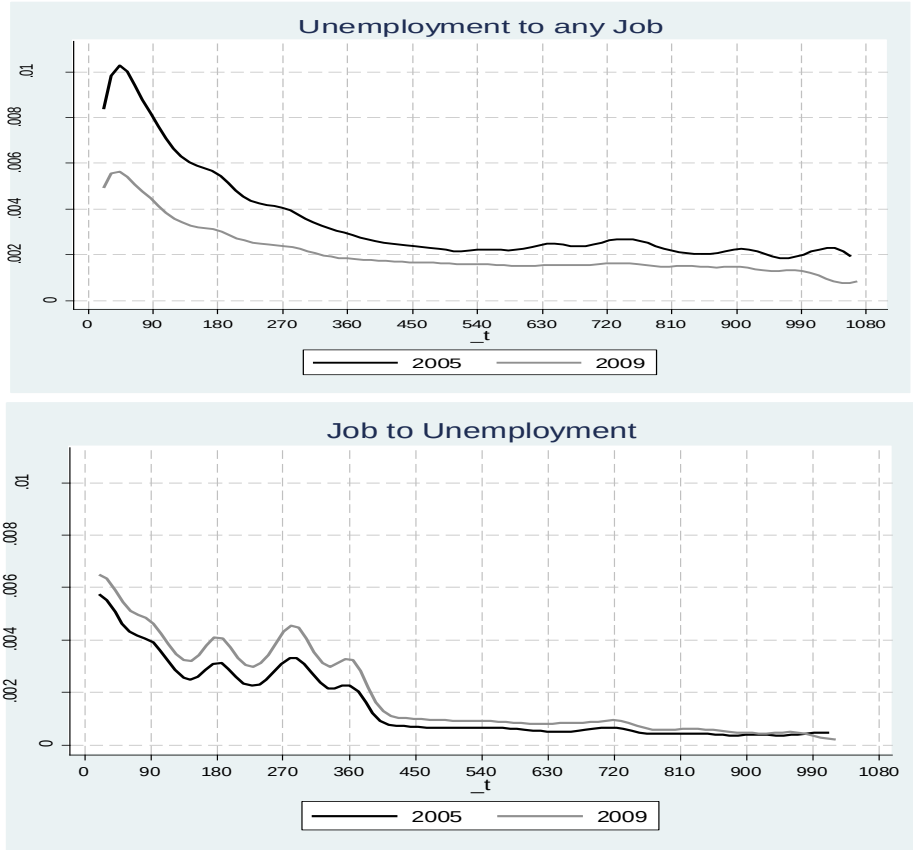
Source: Own elaboration from LWLS.

The corresponding hazard rates are sketched in Figure 2.[16] Note that the negative associations in Figure 2 (both top and bottom panel) may reflect genuine negative state dependence, but may also be due to heterogeneity and the changing nature of the pool of the unemployed and re-employed over time. These explanations will be disentangled in the econometric model. The top panel shows that the crisis reduced the unemployment exit hazard mainly in the first year of unemployment. There is a negative association between each hazard rate and the duration of the spell, and it is

[16] The estimates use Kernel smoothing; the empirical hazard rate at time *t* is the proportion of individuals unemployed (employed) for *t* days that find (loss) a job on day *t+1*.

stronger in 2005. The bottom panel shows that job loss patterns are similar in both periods with declining hazards until about 400 days of tenure. Some local peaks in the hazard are found at 180, 270 and 360 days. These peaks are also found in previous studies and correspond to the usual duration of temporary contracts. During the crisis the likelihood to re-enter unemployment rises, particularly during the first year. But the difference is smaller than in the top panel, suggesting that the probability to find a new job is more sentitive to the business cycle than the probability to lose that job.

Figure 2: Kaplan Meier smoothed hazard functions; exits from –unemployment to any job (top panel) and from job to unemployment (bottom panel). 2005 and 2009 samples. Durations in days.



Note: Durations in days.
Source: Own elaboration from LWLS.

According to job search theory, the probability to exit from unemployment into employment depends, on the one hand, on variables affecting the probability of receiving a job offer, such as the local unemployment rate and the level of education, and on the other hand on variables driving the probability to accept an offer, such as family circumstances. We therefore include individual and family characteristicscs as explanatory variables, as well as the (quarterly) regional unemployment rate. Table 2 provides some descriptive statistics of the explanatory variables in both samples. For all these variables, the sample means in the two samples are significantly different from each other. During the financial crisis the composition of unemployment has changed. In 2005, about 49% of the sample were males, but in 2009 this proportion had risen to 54%, reflecting the larger growth in unemployment of males compared to females due to the crisis. The average age at the time of becoming unemployed rises from 33 years in the 2005 sample to 37 years in the 2009 sample, mainly due to the large growth of the share of individuals older than 45 and the decline of the proportion younger than 30. During our observation window, unemployed workers older than 51 years who satisfy all the requirements for a retirement pension were elegible to receive UAB until retirement age. We therefore expect a lower probability to find a job for this group. Only 20% of the unemployed in the two samples have children. The proportion of non-Spanish-speaking unemployed immigrants increased from 4.41% in 2005 to 8.2% in 2009, while the fraction of Spanish-speaking immigrants only increased from 3% to 4%.

Table 2: Descriptive statistics for the 2005 and 2009 samples.

	2005		2009	
Variable	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**
Male (*)	0.486		0.541	
Age at the start of the unemployment spell	33.3	10.20	37.0	11.23
Children below 4 (*)	0.058		0.060	
Children ages 4-15 (*)	0.154		0.150	
Nationality				
Spanish native (*)	0.924		0.874	
Spanish speaking immigrant (*)	0.032		0.045	
Non-Spanish speaking immigrant (*)	0.044		0.082	
Level of education				
Primary (*)	0.189		0.214	
Lower secondary (*)	0.401		0.410	
Upper secondary (*)	0.239		0.226	
Post-secondary (*)	0.143		0.126	
Unemployment rate (quarterly regional)	9.33%	0.04	18.2%	0.05
Male unemployment rate (quarterly regional)	7.37%	0.02	17.8%	0.05
Female unemployment rate (quarterly regional)	11.18%	0.05	18.6%	0.05
Inhabitants > 40,000 (*)	0.451		0.470	

Notes: Variable definitions are given in Table A2 in the Appendix. Descriptive statistics refer to the first observation of the first unemployment spell of each individual; (*): Dummy variable

About 40% of unemployed individuals in both samples have lower-secondary level of education. The share of unemployed workers with primary education increased from 19% in 2005 to 21% in 2009, mainly due to the high proportion of low educated workers in construction. The share with post-secondary level fell from 14% to 12%. This is in line with Rosholm (2001) who found that the quality of those becoming unemployed is higher during a boom.

To account for regional economic conditions we use the quarterly unemployment rate by region and gender. The average unemployment rate in the crisis period (18.2%) is on average almost twice as high as during the expansion (9.3%). Moreover, unemployment rates vary substantially across regions, possibly reflecting inefficiencies and lack of flexibility in the labour market (Fernandez-Kranz, 2014). Degree of urbanization is captured by a dummy for living in a larger municipality. Around 45% (47%) of workers live in a municipality with more than 40,000 inhabitants in the 2005 (2009) sample.

To sum up, there are substantial differences in the sample composition of the workers becoming unemployed in 2009 versus their counterparts in 2005. These differences could explain changes in unemployment length and job stability between the two periods.[17]

[17] In an extended model, we will also control for variables related to the individual's labour market history and characteristics of the previous job (see Appendix Table A2 for variable definitions and A3 for descriptive statistics). The sample compositions in 2005 and 2009 also substantially differ in terms of some of these variables; for example, the 2009 sample has workers with more experience who are more often entitled to unemployment benefits.

5. Econometric Framework

To analyse the unemployment and subsequent job duration, we estimate a bivariate hazard rate model for the two potentially correlated transitions: unemployment to any job, and subsequent job to unemployment. Transitions to other states, such as retirement or self-employment, are treated as right-censored. Since all durations are measured in days, we consider the duration of each spell as a continuous random variable.

The unemployment (job) hazard rate at duration t is the probability of leaving unemployment (job) at spell length t conditional on not leaving unemployment (job) earlier. Formally the hazard rate is defined as h(t) = f(t)/S(t), where f(t) is the density function of the unemployment (job) duration and S(t) is the survival function (S(t) =1-F(t), where F(t) is the cumulative density). The hazard rate can be interpreted as the conditional probability of leaving unemployment (job) in a short (one-unit) time interval.

Conditional on observed and unobserved heterogeneity, the two consecutive risks are assumed to be independent. We specify the following Multivariate Mixed Proportional Hazard (MMPH) model with gap-time representation; time is reset to zero after each unemployment and consecutive job spell (see, e.g., van den Berg, 2001) with hazards $h_m(t|X_i(t), V_i^m)$ for the two types of transitions – unemployment to any job (m=uj), and job to unemployment (m=ju) – of individual i, conditional on observed and unobserved characteristics:

$$h_m(t|X_i(t), V_i^m) = h_0^m(t) \cdot \exp(X_i(t)'\beta^m) \cdot \exp(V_i^m) \qquad (1)$$

The proportional hazard assumption implies that the shape of the duration dependence is the same for all individuals, but the level of the hazard may vary across individuals. The hazard rate for the transition process m=uj, ju evaluated at spell duration t for spell s of individual i is given by the product of the baseline hazard, $h_0^m(t)$, an observed heterogeneity factor, $X_i(t)'\beta^m$ including time-varying and time-invariant covariates (and excluding the intercept, as a normalization needed to identify the model) and an unobserved heterogeneity ("frailty") component V_i^m.

The baseline hazard, $h_0^m(t)$, follows an exponential distribution with piecewise constant duration dependence, using (mainly quarterly)[18] cut-points $\tau_l, l = 0, \dots, L$:

[18] We tried to estimate the model with a piece-wise constant specification of the baseline hazard using monthly and weekly cut points to allow for greater flexibility, but the estimation algorithm did not converge.

$$h_0^m(t) = \bar{h}_l,\ t \in (\tau_{l-1}, \tau_l), l = 1, \ldots, L$$

(2)

This baseline hazard specification has the advantage of not imposing a particular functional form, thus allowing for a flexible shape of duration dependence. The parameters of main interest are the vectors $\beta^m, m = uj, ju,$ that determine how the hazards vary with individual characteristics and regional unemployment. A positive coefficient of a covariate implies that, keeping other observed variables and the unobserved heterogeneity constant, an increase in the covariate raises the probability to find a job (m=uj) or to lose that job (m=ju). A way to interpret the size of the coefficients is through the percentage change in the hazard produced by a change in the covariate by one unit, obtained as $(e^{\beta^m} - 1) \cdot 100$.

The unobserved heterogeneity terms are V_i^m. Following Heckman and Singer (1984), we use discrete frailty and allow the unobserved heterogeneity components of the two transitions: from unemployment to any job (V_i^{uj}) and from job to unemployment (V_i^{ju}) to be correlated. Ignoring this correlation could create a sample selection problem and bias the results (Ham and Lalonde, 1996). This discrete distribution is a computationally attractive way to allow for correlation between unobserved heterogeneity terms of different exits. It is computationally easier than a bivariate continuous distribution and allows for a more flexible distribution if the number of mass points grows large. Moreover, it is very common in the literature on unemployment dynamics; see, e.g., Uhlendorff and Zimmermann (2014), Caliendo et al. (2013), or Rebollo and García-Pérez (2015).

Under this discrete frailty distribution, the population consists of several subpopulations with different risks. For instance, one group of motivated individuals could have higher exit probabilities for the transition from unemployment to any job but low chances to lose that job, another group might have low chances of finding a job and bad prospects for keeping that job, etc. The group to which an individual actually belongs is never observed. The population fractions of the groups are unknown parameters pk.[19] with $\sum_{k=1}^{K} p_k = 1$; K is the number of groups (the number of mass points of the distribution of (V_i^{uj}, V_i^{ju})).

[19] To ensure the probability is between zero and one we assume $p_k = \frac{\exp(a_k)}{(1+\sum_{l=1}^{K-1}\exp(a_l))}$

We assume that unobserved heterogeneity is constant over time (within and across spells of the same individual) and independent of observed characteristics, the standard assumptions in this kind of models (van den Berg, 2001). Moreover, since we do not impose a normalization on the baseline hazard or on $X_i(t)'\beta^m$, we need to impose E(Vm) = 0: $\sum_{k=1}^{K} p_k V^m$ for m=uj, ju.

Ignoring unobserved heterogeneity may lead to biases in the estimates of β^m and would make the estimated duration dependence more negative (Nickell, 1979). The flexible baseline hazard and inclusion of frailty make it possible to analyse genuine duration dependence before and during the crisis.

The parameters are estimated jointly by Maximum Likelihood. The likelihood function is, under the independence assumption, the product of the Likelihood function of all the individuals (i), $L = \prod_i L_i$. The likelihood contribution L_i of individual i for two consecutive risks (m=uj, ju) can be written as the expected value of the conditional likelihood given (V_i^{uj}, V_i^{ju}): $L_i = \sum_{k=1}^{K} P_k \cdot L_i(V^k)$, where $L_i(V^k)$ is the conditional likelihood contribution given (V_i^{uj}, V_i^{ju}) is equal to the kth mass point $V^k = (V_k^{uj}, V_k^{ju})$. This conditional likelihood contribution is a standard likelihood contribution in a model without unobserved heterogeneity; it includes the conditional density function for the observed exits of the completed spells and the conditional survival function for right-censored spells at each transition.

$$L_i(V^k) = \prod_{s=1}^{s_{uj}} h_{uj,s}(t_i|X_i(s), V_{uj}^k)^{d_{i,uj,s}} \cdot S_{uj,s}\left(t_i\middle|X_i(s), V_{uj}^k\right) \cdot \prod_{s=s_{uj+1}}^{s_{uj}+s_{ju}} h_{ju,s}(t_i|X_i(s), V_{ju}^k)^{d_{i,uj,s}\cdot d_{i,ju,s}} \cdot S_{ju,s}(t_i|X_i(s), V_{ju}^k)^{d_{i,uj,s}} \quad (3)$$

Here s=1,…,S are the spells of individual i, and $d_{i,uj,s}$ is a dummy that is 1 if there is a transition from unemployment to any job at the spell s. Analogously, $d_{i,ju,s}$ is 1 if there is a transition from job to unemployment. Our Stata code for estimation is largely based upon that of Bijwaard (2014).

6. Estimation Results

Table 3 presents the results for our benchmark model. Estimates for an extended model that adds characteristics of the job (sector, type of contract, etc.) and benefit entitlement are presented in the appendix. We refer to this extended specification when there are interesting differences. The best likelihood for the benchmark model is obtained using a discrete unobserved heterogeneity distribution with three points of support.[20]

Parameter estimates

One of the main determinants of unemployment and job durations is the (quarterly) local unemployment rate. Table 3 shows that for the expansion sample, the probability of getting a job is smaller in regions with high unemployment. This is in line with Arranz and Muro (2004), Alba et al. (2012), Arranz et al. (2010) and Bover et al. (2002) who also found that in Spain, unemployment durations are longer in regions with higher unemployment rates. The effect reverses, however, in the recession sample. This may be related to increasing regional mobility during the crisis (cf. Sala and Trívin, 2014). In the extended model (Appendix Table A4) the relation with regional unemployment during the crisis is weaker. This suggests that it is not the unemployment rate as such that drives job finding rates during the recession, but the fact that in certain regions unemployed workers more often come from (and will want to go back to) sectors and job types where mobility is larger. Moreover, we find that subsequent job stability is lower in regions with higher unemployment rates, especially in the 2009 sample. This effect diminishes in the extended model and remains significant for the 2009 sample only, suggesting that regions with large unemployment are regions with more short-term jobs that are at risk during the recession.

Men's job finding rates exceed women's (by 5.3%) during the expansion period but are lower than those of women during the recession (6.1%), ceteris paribus. For earlier periods, Arranz and Muro (2004), Arranz et al. (2010) and Alba et al. (2012) found higher exit probabilities to new jobs (not recalls) for men. Once a job is found, the jobs last longer for men than for women, but this advantage has fallen substantially

[20] The model with three mass points is significantly better than the model with two mass points. We did not estimate the model with four mass points because the estimated probability of the additional mass point is extremely small for the two samples.

during the recession (from 10.5% in 2005 to 4.5% in 2009. The results for the extended model (Appendix Table A4) suggest that this can be because men and women have different types of jobs.

Age patterns for transitions from unemployment to any job and vice versa are similar for both periods. Unemployment exit probabilities fall with age. This is in line with the literature (Bover and Gómez, 2004; Arranz et al. 2010; Bover et al., 2002). This may reflect the reluctance of employers to recruit older workers (see, e.g., Taylor and Walker, 1998). Older workers may also have higher reservation wages due to more labour experience (Folmer and van Dijk, 1988) or more difficulties to adapt to a new job (Narendranathan and Nickell, 1985). Once a new job is accepted, this job is shorter for younger workers and longer for the middle-aged workers compared to those aged 52 to 65. During the crisis, the decreasing age pattern in the job finding rate is even steeper than before the crisis and workers in the age group 52-65 are still much worse off. The age pattern in job exit rates is much less pronounced. Only the youngest age groups have much less stable jobs, and this difference decreases somewhat during the crisis. Controlling for job characteristics (Table A4) reduces the differences across age groups. For example, part of the explanation why young workers have less stable jobs is that they often have non-permanent contracts. If type of contract and other observed characteristics of jobs are kept constant, the jobs of the young are still less stable than those of older workers but the difference is much smaller.

Unemployed immigrants suffer much more from the crisis than Spanish natives. In the expansion period immigrants and particularly Spanish speaking immigrants experienced shorter unemployment periods than comparable natives,[21] and the subsequent job duration did not differ from the one of an otherwise identical native. During the downturn, however, job finding probabilities of immigrants fell more than for comparable natives, and subsequent job stability fell as well, especially for Spanish-speaking immigrants. This might be because of more competition with other unemployed – perhaps natives are willing to accept jobs during the recession that they would not have accepted in the expansion period. Both of these findings are in line with Carrasco and García-Perez (2015) who argue that this is because immigrants have lower capital labour complementarity than natives. Return migration or discrimination might also be the explanation.

[21] Rebollo-Sanz (2012, Table A1) also found that the hazard of re-employment is higher for immigrants.

Table 3: Estimation results of bivariate continuous-time hazard rate model for unemployment to any job and for subsequent job to unemployment; 2005 and 2009 samples

	2005 sample				2009 sample			
	Unemployment to any Job		Job to Unemployment		Unemployment to any Job		Job to unemployment	
	Coeff.	SE	Coeff.	SE	Coeff.	SE	Coeff.	SE
Regional unemployment rate	-1.653***	(0.0859)	0.562***	(0.120)	0.554***	(0.0523)	1.086***	(0.0629)
Male	0.0515***	(0.00679)	-0.111***	(0.00997)	-0.0627***	(0.00576)	-0.0457***	(0.00714)
Age 16-19	0.674***	(0.0188)	0.423***	(0.0254)	0.982***	(0.0242)	0.312***	(0.0303)
Age 20-24	0.686***	(0.0145)	0.257***	(0.0191)	0.930***	(0.0121)	0.189***	(0.0144)
Age 25-29	0.605***	(0.0145)	-0.0125	(0.0194)	0.805***	(0.0117)	0.00536	(0.0141)
Age 30-34	0.503***	(0.0151)	-0.106***	(0.0203)	0.680***	(0.0120)	-0.0742***	(0.0146)
Age 35-39	0.481***	(0.0157)	-0.0551***	(0.0210)	0.627***	(0.0125)	-0.0778***	(0.0152)
Age 40-44	0.506***	(0.0161)	-0.0241	(0.0214)	0.606***	(0.0129)	-0.0509***	(0.0156)
Age 45-51	0.429***	(0.0159)	-0.0108	(0.0209)	0.481***	(0.0123)	0.00222	(0.0147)
Children below 4	-0.320***	(0.0137)	-0.136***	(0.0196)	-0.315***	(0.0124)	-0.105***	(0.0158)
Children 4-15	-0.0893***	(0.00908)	-0.0199	(0.0127)	-0.110***	(0.00842)	-0.0233**	(0.0104)
Spanish speaking imm.	0.157***	(0.0163)	-0.0301	(0.0241)	0.141***	(0.0133)	0.184***	(0.0162)
Non-Spanish speaking imm.	0.103***	(0.0142)	0.00231	(0.0201)	0.0506***	(0.0104)	0.0426***	(0.0128)
Lower secondary	0.00256	(0.00791)	-0.0737***	(0.0111)	0.0908***	(0.00756)	-0.0433***	(0.00919)
Upper secondary	-0.0224**	(0.00891)	-0.0887***	(0.0127)	0.0909***	(0.00870)	-0.144***	(0.0109)
Post-secondary	0.0346***	(0.0102)	-0.0147	(0.0145)	0.267***	(0.0101)	-0.202***	(0.0128)
Inhabitants>40,000	-0.0153***	(0.00583)	-0.0713***	(0.00818)	0.0418***	(0.00562)	0.00686	(0.00698)
V1	0.528***	(0.0392)	0.467***	(0.0149)	0.858***	(0.0304)	0.885***	(0.0155)
V2	2.793***	(0.0741)	1.349***	(0.0342)	2.573***	(0.0511)	1.085***	(0.0251)
a1	-5.666***	(0.154)			-5.273***	(0.0921)		
a2	-3.069***	(0.0610)			-2.980***	(0.0513)		
Number of individuals	87,950		79,912		133,197		97,518	
Number of spells	133,413		124,132		181,010		141,776	
Number of exits	124,413		75,831		141,776		101,507	
Observations	768,017				1,174,101			
Log Likelihood	-1.22E+06				-1.545e+06			

Note 1: Correlated Competing risks estimation: piecewise baseline and discrete distribution of unobserved heterogeneity with three mass points.E[V]=0 for both samples.
Note 2: For 2005 estimation, Pr(Type I)= 0.35%; Pr(Type II)=4.63%; Pr(Type III)= 95%; V3(U-J)=-0.14; V3(J-U)=-0.07; Rho=0.99. For 2009 estimation, Pr(Type I)= 0.51%; Pr(Type II)=5.05%; Pr(Type III)=94.4%; V3(U-J)=-0.14; V3(J-U)=-0.06; Rho=0.98.
Note 3: Reference categories:female, Aged_52_65, Primary level of education.Unemployment rate is a time-varying variable. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Unemployed with young children (particularly, children younger than 4 years old) face more difficulties to find any job than unemployed without children in both periods, but the jobs they get are more stable. There are no large differences between the two time periods in this respect.

Search theory implies that a higher level of education is associated with more productivity (Toharia and Cebrián, 2007), implying a higher arrival rate and job stability but also a higher reservation wage. Arranz and Muro (2004) find no significant effect of education, while according to Bover and Gómez (2004) having a university degree reduces the hazard to a temporary job but increases the hazard to a permanent job. We find that before the crisis, the effects of education on unemployment exits are small, possibly due to high demand for low educated workers during the building boom and an oversupply of graduates. The stability of jobs found by workers with secondary education is somewhat higher than for low or post-secondary education levels. In contrast, during the crisis a higher level of education substantially increases the probability of getting a job. During the recession, employers are able to select more on skills given that workers with high education are more often willing to accept low skilled jobs. Moreover, job stability increases monotonically and substantially with education level. This might be explained by the destruction of low-skilled jobs during the recession, for example in construction. Thus, low educated individuals suffer more from the crisis than those with higher level of education. The education effects during the crisis in the extended model have the same sign but are smaller in magnitude. This shows that a large part of the disadvantage of low educated workers can be attributed to the nature of their jobs. For instance, low educated men tend to work in construction sector, heavily affected by the crisis.

Municipality size influences access to training, networks that are useful for finding a job, etc., and may therefore affect the job exit rate. Moreover, municipality size may be associated with specific industries and types of jobs. During the expansion period, those living in a municipality with more than 40,000 inhabitants need somewhat more time than others to get a job, but the jobs they get are more stable than in smaller municipalities. During the crisis on the other hand, the unemployed in larger municipalities have better chances of getting a job, and there is no significant difference in the stability of that job. This suggests that the crisis hits smaller towns harder in terms of job opportunities, but the reverse is true for job stability.

Unobserved heterogeneity is significant in both samples and in both processes, demonstrating the importance of unobserved characteristics such as motivation, job search effort, social pressure etc., for the chances to find a new job and retain that job. The population is divided into one very large group (group III: 95% of the population in 2005 and 94.4% in 2009) and two small groups with much higher chances for both

exits. The estimated correlation between the two unobserved heterogeneity terms is high (0.99 for the 2005 sample, 0.98 for the 2009 sample). This implies that someone who is likely to find a job also has higher chances of re-entering into unemployment. It confirms the importance of allowing for a correlation between the two processes.[22]

Coefficients of the labour market variables in the extended model (Appendix Table A4) capture time-persistent heterogeneity or causal effects. Most results are as expected. In short, we find that, for the two samples, employees with a temporary contract have higher job turnover (shorter unemployment and job spells), jobs in non-manual occupations are more stable and workers receiving unemployment benefits have longer unemployment spells. Part-time workers have higher job finding rates and lower chances to lose their new job. As expected the receipt of UB is negatively associated with the probability to find a job. This effect is stronger for the UA than for the UI. As emphasized by Cockx and Picchio (2012), these coefficients cannot be given a structural interpretation and might reflect unobserved productivity differences rather than a disincentive for work effect. All these effects are similar for the two samples.

For unemployed workers in the construction sector, stability of the new job deteriorated disproportionally during the crisis, whereas unemployed workers in manufacturing got a much harder time to get a new job. A positive relation between job stability and firm size already existed in the 2005 sample but became much stronger during the crisis.

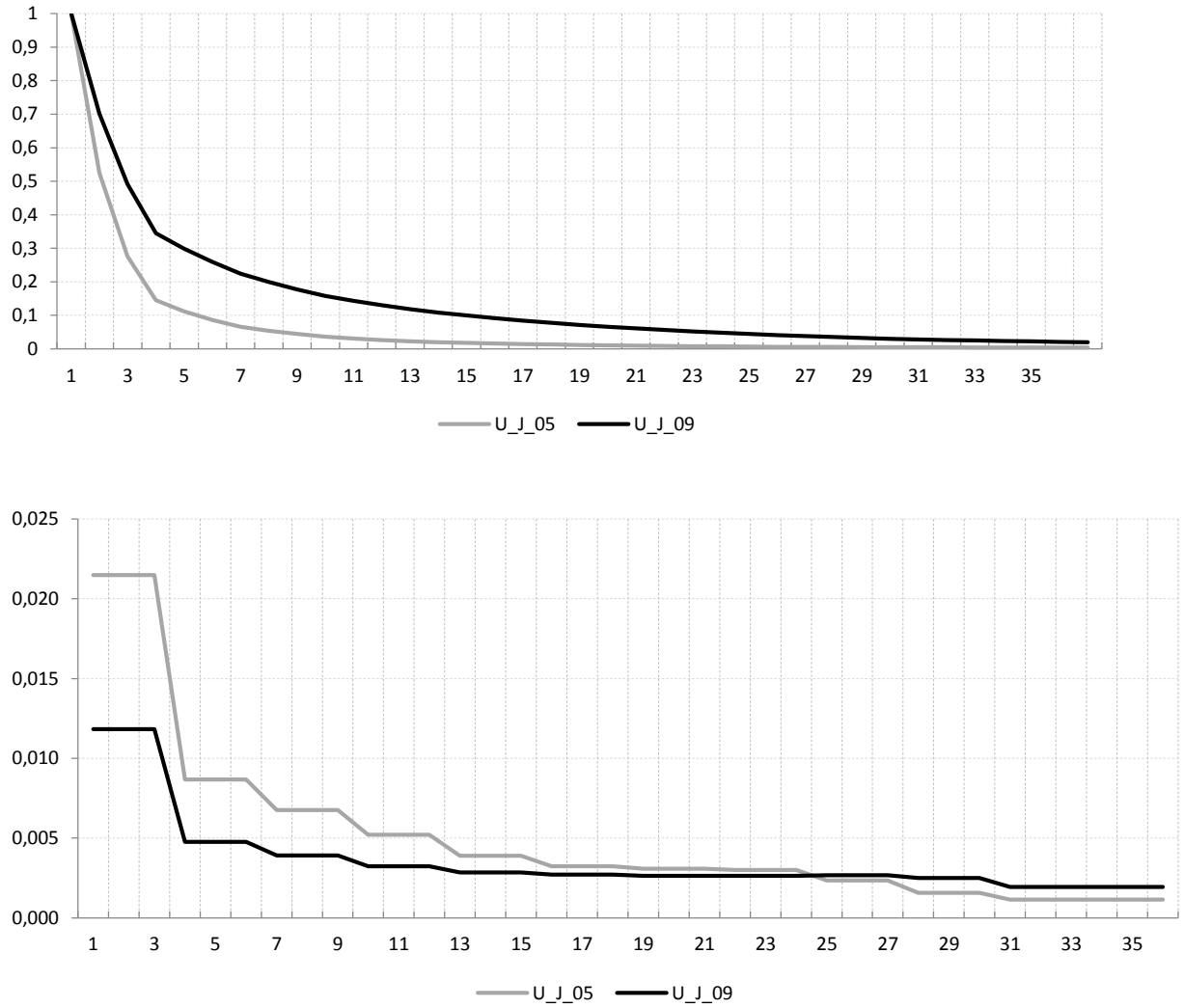
Baseline Hazards

Figure 3 shows the estimated survival functions and hazards for transitions from unemployment to any job for a benchmark person for both samples. The top panel shows, for example, that in the benchmark group in 2005, the probability of getting a job within six months is 91.4%. In 2009, this probability has fallen to 74.1%. The bottom panel shows the corresponding hazard rates. For the 2009 sample, we find much smaller hazards in the first year of the unemployment spell, and particularly during the first three months. Since observed and unobserved heterogeneity are controlled for, the negative slopes can be interpreted as true negative state dependence (Heckman and Singer, 1984). Duration dependence is stronger in 2005 than in 2009, possibly due to a

[22] In the extended model, convergence was obtained only for a model with two mass points. Unobserved heterogeneity is again significant in both processes, and the correlation between the unobserved heterogeneity terms is 1. (It is automatically 1 or -1.)

negative stigma effect that makes employers reluctant to hire the long-term unemployed before the crisis. This stigma largely disappears during the crisis when more productive workers also remain unemployed. This result is in line with a theoretical study of Lockwood (1991), arguing that negative duration dependence is weaker the higher the unemployment rate. In contrast, Blanchard and Diamond (1994) get the opposite result. Our result differs from those of Rosholm (2001) and Kalwij (2010) who (not controlling and controlling for unobserved heterogeneity, respectively) found no significant evidence of negative duration dependence.

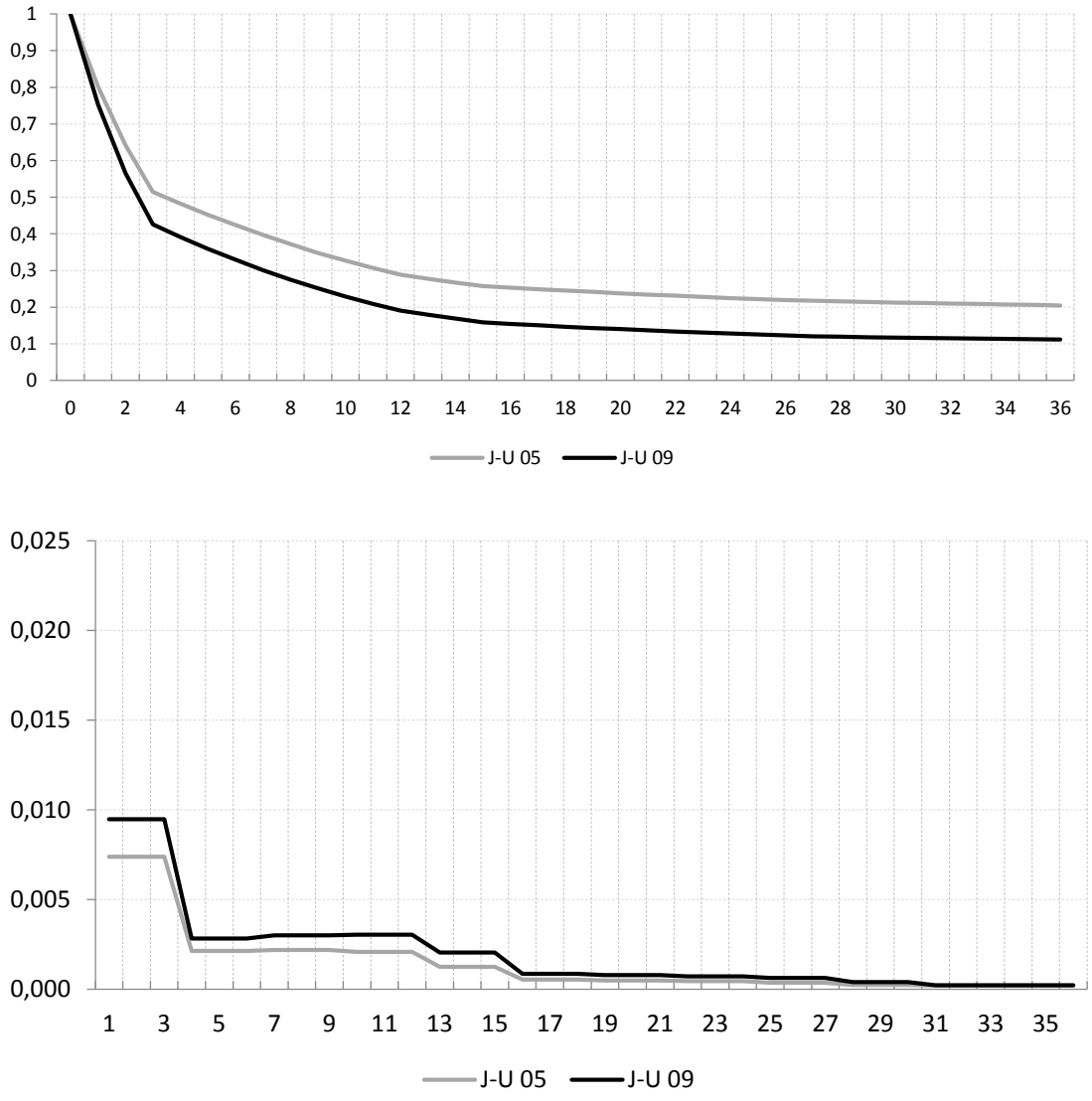
Figure 3. Survival functions (top panel) and hazard rates (bottom panel) benchmark person for Unemployment to any Job; 2005 and 2009 samples.



Source: Own elaboration.
Notes: Durations in months. Benchmark: man, age group 25 - 29, primary level of education, Spanish native, without children, municipality with more than 40.000 inhabitants.

Figure 4 shows the estimated survival and hazard functions for transitions from the new job to unemployment for a benchmark person. The top panel shows, for instance, a 71% chance to lose the job and become unemployed within one year for the 2005 sample. This probability has increased by 10% points in the 2009 sample. The bottom panel shows that the corresponding hazard rates are decreasing for both time periods. The hazard rates into unemployment are slightly higher during the recession, but the differences are less pronounced than in Figure 3 and the hazards are almost identical at durations of more than 15 months.

Figure 4. Survival functions (top panel) and hazard rates (bottom panel) benchmark person for job to unemployment; 2005 and 2009 samples; competing risks model.



Source: Own elaboration.

Notes: Durations in months. Benchmark: man, age group 25 - 29, primary level of education, Spanish native, without children, municipality with more than 40,000 inhabitants.

Decompositions

Table 4 shows the results of decompositions of the difference between the survival probabilities after 360 days in the periods before and during the crisis, in the spirit of, for example, Rosholm (2001) or Verho (2014). The first rows give the average survival probabilities for the two samples according to the model estimates and the difference between these two. The average predicted probability of not finding a job within a year was 15% in the 2005 sample[23] and increased to 31.6% in the 2009 sample. On the other hand, the average probability of keeping the new job decreased from 26% in the 2005 sample to 19.7% for the 2009 sample.

The remaining rows decompose these differences. First, we use the 2005 estimates and the 2005 regional unemployment rates by gender, but compute the average probabilities for the individuals in the 2009 sample. Comparing with the 2005 probabilities in row 3 gives the composition effect: the difference explained by the differences in individual characteristics in the two samples. The composition effect explains only one tenth of the reduction in the probability to find any job (1.7 out of the 16.6%-points). The decrease in stability of the new job is not explained at all by differences in sample characteristics - in fact, the difference in sample composition would predict a change in the other direction (of 0.75%-points). These results are in line with Bergin et al. (2015) for Ireland, who find that changes in the composition of the population from 2006 to 2011 explain very little of the changes in unemployment and employment transition rates.

Table 4 Decomposition analysis for correlated consecutive events from the benchmark model.

	Unemployment to any job		Job to Unemployment	
Total Effect: Difference between:	**16.62%**	**100%**	**-6.28%**	**100%**
S09 09	31.62%		19.71%	
S05 05	15%		26.00%	
Composition effects	**1.74%**	**10%**	**0.75%**	**-12%**
Business cycle effects	**14.88%**	**90%**	**-7.04%**	**112%**

Note: Evaluated using the 2005 model. S09 09 (S05 05) is the average survival probability at month 12, using the model of 2009 (2005) for the sample of 2009 (2005).
Source: Own elaboration from LWLS.

[23] The similarity between the average survival probabilities from the estimation model and from raw data (section 3, Descriptive analysis) confirms the goodness of fit of our model.

In the extended model, a similar decomposition leads to a larger role of the composition effect for unemployment exits (Appendix Table A5): this now captures one third of the change in the probability to remain unemployed after 12 months between the two samples (5.74 %-points). This is because the composition effect now also captures differences in previous employment sector, job characteristics, and labour market history and benefit entitlement. The main reason is the latter: Compared to the 2005 sample, the unemployed in the 2009 sample have very different labour market histories with different types of contract and larger entitlement to UB or UA benefits (cf. Table A3). Since benefit entitlement is associated with a lower unemployment exit rate (Table A4), this explains part of the difference. Still, the conclusion remains that the largest part of the change in the probability to remain unemployed for more than 12 months is not explained by sample composition effects. For job stability, the results for the extended model are similar to those of the benchmark model: composition effect hardly explain the difference in stability of the jobs found by those who became unemployed in 2005 or 2009.

7. Conclusions

We have analysed transitions from unemployment to any job during an expansion period (2005-2007) and during the recent recession (2009-2011) and the stability of the new jobs found by the unemployed. We have modelled the two transition processes jointly, estimating bivariate continuous-time hazard rate models, using a rich administrative data set from the Spanish Social Security Administration.

Our results confirm the pro-cyclicality of unemployment to employment transitions and the counter-cyclicality of exits from employment back into unemployment. The effects of the crisis were much stronger for exits out of unemployment than for the hazard to lose the new job, emphasizing the importance of policies that help individuals to find jobs during the recession. Negative duration dependence of the unemployment hazard was much stronger in the expansion period than during the crisis, supporting the arguments of Lockwood (1991) that stigma effects dissapears during the crisis.

A decomposition analysis shows that changes in characteristics of the pool of unemployed explain only a limited part of the reduction in exit probabilties and do not

explain the rising instability of the new jobs. Instead, unemployed individuals with given characteristics become less likely to find a job, particularly a stable job. Our estimates show that there is substantial heterogeniety in this respect: the chances to find a job deteriorate much more for some socio-economic groups than for others. Groups particularly affected in terms of the highest increase in unemployment duration during the downturn are men, older workers (ages 45-65), non-Spanish speaking immigrants, lower educated workers, and those living in smaller towns. Analogously, those most affected in terms of a reduction of job stability are Spanish-speaking immigrants, those with lower level of education, middle-aged workers, and workers in urbanized areas.

Comparing with an extended model that not only controls for individual characteristics but also for job characteristics and benefit entitlement shows that part of the explanation is that different socio-economic groups tend to have different types of jobs. For example, men's unemployment exit rates decrease much more during the crisis than those of women. The difference largely disappears, however, when controlling for sector and job characteristics, implying that the more pro-cyclical nature of male employment is due to the gender segregation in employment.

To reintegrate the targeted groups into the labour market and give them stable jobs, it is necessary to involve them into continuous training and educational programs. Active labour market policies in Spain currently include dual training and employment programs provided to unemployed workers combining employment and training in a training centre.[24] Qualified dual vocational training programs are boosted and Spain is using the European Social Fund to create employment by funding vocational training. Previous evidence shows that targeted schemes aimed at disadvantaged groups are more effective in raising employment than broad training programmes combined with job search interventions. According to Cho and Newhouse (2013), "One key lesson is that traditionally disadvantaged groups of workers may not necessarily be most vulnerable to labor market disruptions during a crisis. So, programs to mitigate the impacts of the recession, rather than serving traditionally disadvantaged groups of people, could consider targeting most affected workers by the crisis."

Our results show that the importance of education rises during the downturn, increasing unemployment exit chances and securing job stability. Moreover, the position of older unemployed workers during the crisis seems particularly concerning.

[24] Royal Decree 1529/2012, 9 November 2012.

Unemployment exit rates already fell strongly with age before the crisis, and the slope has even become much more negative during the recession. Only part of these effects are due to the differences between older and younger workers in job characteristics and benefit entitlements. On the other hand, the stability of the jobs found by older and middle-aged unemployed workers is not very different, either before or during the crisis. Unemployment among older workers during the recession therefore seems to require special attention from policymakers, in line with findings for other countries. See, e.g., Gielen and van Ours (2006) for the Netherlands, who mention on the job training and wage subsidies as potential policy measures focused on older workers.

From a policy point of view, the current study has the limitation that demand and supply factors could not be disentangled, implying that the mechanisms explaining why some groups suffered more from the recession than others are not always clear. Separating declining job arrival rates from changes in reservation wages and job acceptance rates in a more structural analysis (probably also involving data on job search that are not available in our administrative records) seems a useful topic of future analysis.

Appendix

Table A1. Sample selection

Filters 2005	Number of individuals	
	2005 sample	**2009 sample**
Number of individuals starting any non-employment spell in the year of reference	**116,777**	**163,198**
Number of individuals starting any UB spell in the period of reference	67,244	132,049
Number of individuals starting any non-employment spell in the year of reference (excluding those starting an UB spell)	49,533	31,149
Drop individuals after merging consecutive unemployment spells, drop spells starting before the year of reference	17,987	26,110
Drop individuals from agriculture	1,128	697
Drop individuals for lack of relevant (individual) information	7,956	1,295
Drop individuals from Ceuta and Melilla	312	442
Drop individuals with any degree of disability	1,444	1,899
Final sample (number of individuals)	**87,950**	**132,755**
Number of individuals starting an UB spell in the year of reference	55,934	114,534
Number of individuals starting a non-employment spell following a job spell	40,724	27,510
Those who did not started an UB spell	32,016	18,221
Number of individuals starting a non-employment spell following an UB spell	23,919	45,859

Notes:
Individuals between 16 and 65 years old who did not work in a Social Security Regime other than General ("Regimen General de la Seguridad Social") since 1996. Self-employed workers are therefore not in the sample.
Source: Own elaboration from 2005-2007 LWLS and 2009-2011 LWLS.

Table A2. Definitions of explanatory variables

Individual characteristics	
Male	1 if male
Age	Dummies for ages 16-19; 20-24; 25-29;30-34;35-39;40-44;45-51;52-60<u>5</u>; (time-varying)
Spanish native	1 if Spanish citizenship
Spanish-speaking immigrants	1 if immigrant from a Spanish-speaking country
Non-Spanish speaking immigrants	1 if immigrant from a non-Spanish-speaking country
Children below 4	1 if the individual has children younger than 4 years old. It is a time-varying covariate
Children 4-15	1 if the individual has children between 4 and 15 years old. It is a time varying covariate
Primary education	1 if none and elementary education level
Lower secondary	1 if lower secondary education (middle school)
Upper secondary	1 if upper secondary education level (high school)
Post-secondary	1 if tertiary education level
Regional unemployment rate	Quarterly unemployment rate by gender and region (time-varying); source: Economically Active Population Survey (EPA)
Inhabitants>40,000	1 if the number of inhabitants of the municipality where the individual is living is greater than 40.000
(current or previous) Job characteristics	
Unemployment Benefit	1 if the unemployed is receiving contributory unemployment benefits
Unemployment Assistance	1 if the unemployed is receiving assistance benefits
Sector of activity	Dummies for sector of activity in which the individual has been working the longest: Manufacturing, construction or services
Industry	Dummies for manufacturing, construction and services industries.
High Technology	1 if sector of activity in high technology according with the classification of industries by technologic level.
Type of contract	Permanent, open-ended, on-call temporary, temporary Permanent: contract with indefinite duration. Open-ended: A type of permanent contracts to develop works that have the character of fixed discontinuous. Temporary: is an employment contract for a fixed or uncertain term. On-call temporary: is an employment contract to substitute employees with a reserved right to their job.
THA (Temporary Help Agency)	1 if the employment is signed through a temporary help agency
Size of the firm	Dummies for 0 (missing), 1-19, 10-19,20-49,50-249, >250
Non-manual occupation	1 if non-manual occupation
Part-time coefficient	Hours worked as a fraction of full time work (1 is a full time job)
Public Sector	1 if the employer is Public Sector.

Source: Own elaboration from LWLS and INE.

Table A3. Descriptive statistics for the 2005 and 2009 samples: Additional regressors extended model

	Unemployment				Job			
	2005		2009		2005		2009	
	Mean	**Std. Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**	**Mean**	**Std. Dev.**
Missing firm size	14.2%	0.35	4.2%	0.20	9.1%	0.29	3.8%	0.19
Size 1-9	30.8%	0.46	37.5%	0.48	32.1%	0.47	34.3%	0.47
Size 10-19	9.1%	0.29	10.5%	0.31	9.5%	0.29	10.3%	0.30
Size 20-49	11.6%	0.32	12.9%	0.34	12.1%	0.33	13.3%	0.34
Size 50-249	16.8%	0.37	18.0%	0.38	17.6%	0.38	19.1%	0.39
Size ≥250	17.6%	0.38	16.9%	0.37	19.6%	0.40	19.3%	0.39
THA	5.4%	0.23	3.2%	0.17	8.5%	0.28	5.8%	0.23
Public	10.0%	0.30	9.0%	0.29	10.3%	0.30	12.0%	0.32
Industry								
Construction	16.8%	0.37	19.2%	0.39	17.2%	0.38	16.6%	0.37
Manufacturing	12.9%	0.34	15.8%	0.36	10.0%	0.30	8.7%	0.28
Services	70.3%	0.46	64.7%	0.48	72.7%	0.45	74.2%	0.44
High technology	3.1%	0.17	5.5%	0.23	2.6%	0.16	2.9%	0.17
Type of contract								
Open ended	5.8%	0.23	6.1%	0.24	6.3%	0.24	7.9%	0.27
Permanent	14.8%	0.35	29.1%	0.45	14.4%	0.35	14.1%	0.35
Temporary	71.5%	0.45	57.7%	0.49	70.3%	0.46	68.3%	0.47
On call temporary	7.9%	0.27	7.2%	0.26	9.1%	0.29	9.8%	0.30
Non-manual	36.1%	0.48	36.5%	0.48	36.9%	0.48	37.4%	0.48
Part-time	21.3%	0.41	20.3%	0.40	21.6%	0.41	26.3%	0.44
UI	48.0%	0.50	65.8%	0.47				
UA	9.8%	0.30	16.1%	0.37				

Notes:
See Table A2 for variable definitions.
Descriptive for the first observation of each individual in each event and sample; for unemployment event, job characteristics refer to the previous job spell.

Table A4. Estimation results of extended model; 2005 and 2009 samples.

	2005 sample				2009 sample			
	U-J		J-U		U-J		J-U	
	Coeff.	SE	Coeff.	SE	Coeff.	SE	Coeff.	SE
Unemployment rate	-1.570***	(0.0872)	-0.0530	(0.118)	0.131**	(0.0518)	0.394***	(0.0605)
Male	0.0339***	(0.00738)	-0.0807***	(0.0104)	0.0364***	(0.00628)	-0.0551***	(0.00781)
Age 16-19	0.0461**	(0.0193)	0.259***	(0.0248)	0.294***	(0.0236)	0.165***	(0.0293)
Age 20-24	0.158***	(0.0151)	0.151***	(0.0185)	0.361***	(0.0122)	0.0928***	(0.0141)
Age 25-29	0.218***	(0.0150)	-0.0536***	(0.0187)	0.413***	(0.0118)	-0.0497***	(0.0137)
Age 30-34	0.214***	(0.0154)	-0.123***	(0.0195)	0.399***	(0.0120)	-0.105***	(0.0141)
Age 35-39	0.212***	(0.0159)	-0.0846***	(0.0202)	0.370***	(0.0124)	-0.0952***	(0.0146)
Age 40-44	0.220***	(0.0163)	-0.0606***	(0.0205)	0.354***	(0.0128)	-0.0885***	(0.0149)
Age 45-51	0.184***	(0.0161)	-0.0484**	(0.0200)	0.267***	(0.0122)	-0.0299**	(0.0140)
Children below 4	-0.171***	(0.0138)	-0.0589***	(0.0188)	-0.180***	(0.0122)	-0.0610***	(0.0153)
Children 4-15	-0.0305***	(0.00916)	-0.00685	(0.0122)	-0.0354***	(0.00829)	-0.00348	(0.0100)
Spanish speaking imm.	0.0551***	(0.0165)	-0.0514**	(0.0236)	0.0402***	(0.0129)	0.0964***	(0.0156)
Non-Spanish sp. Imm.	-0.000267	(0.0143)	-0.00753	(0.0197)	-0.0608***	(0.0102)	-0.000355	(0.0124)
Lower secondary	0.0400***	(0.00794)	-0.0482***	(0.0107)	0.0852***	(0.00739)	-0.0142	(0.00882)
Upper secondary	0.0176*	(0.00928)	-0.000125	(0.0127)	0.0834***	(0.00877)	-0.0337***	(0.0108)
Post-secondary	-0.00621	(0.0109)	0.0624***	(0.0151)	0.122***	(0.0105)	-0.0767***	(0.0133)
Inhabitants>40,000	-0.00745	(0.00591)	-0.0917***	(0.00799)	0.0315***	(0.00553)	-0.0419***	(0.00681)
THA	0.115***	(0.0113)	0.327***	(0.0144)	0.155***	(0.0130)	0.342***	(0.0149)
Public	-0.0437***	(0.0116)	-0.200***	(0.0156)	-0.0634***	(0.0107)	-0.124***	(0.0120)
Missing firm size	-0.0816***	(0.00985)	0.191***	(0.0148)	-0.123***	(0.0142)	0.425***	(0.0172)
Size 10-19	0.0269**	(0.0109)	-0.0374***	(0.0145)	0.0382***	(0.00961)	-0.0613***	(0.0118)
Size 20-49	0.0369***	(0.00998)	-0.00272	(0.0132)	0.0471***	(0.00889)	-0.0970***	(0.0109)
Size 50-249	0.0221**	(0.00904)	-0.0152	(0.0119)	0.0273***	(0.00823)	-0.0754***	(0.00995)
Size ≥250	0.0416***	(0.00949)	-0.0381***	(0.0125)	0.0409***	(0.00906)	-0.127***	(0.0110)
Construction	0.0836***	(0.00944)	-0.349***	(0.0132)	0.0271***	(0.00847)	-0.0153	(0.0106)
Manufacturing	-0.0358***	(0.0102)	-0.184***	(0.0143)	-0.307***	(0.0104)	-0.148***	(0.0133)
High tech	-0.0509***	(0.0182)	-0.0321	(0.0256)	-0.225***	(0.0163)	-0.0116	(0.0228)
Open-ended	0.165***	(0.0140)	-0.139***	(0.0159)	0.451***	(0.0118)	-0.195***	(0.0129)
Permanent	-0.405***	(0.0108)	-2.049***	(0.0193)	-0.651***	(0.00836)	-1.777***	(0.0155)
On call temp	0.155***	(0.0117)	0.166***	(0.0152)	0.256***	(0.0108)	0.227***	(0.0125)
Part-time coef.	0.0960***	(0.0147)	-0.0982***	(0.0186)	0.0889***	(0.0136)	-0.0646***	(0.0153)
Non-manual	-0.0251***	(0.00723)	-0.147***	(0.00989)	-0.00569	(0.00678)	-0.135***	(0.00856)
UI	-0.639***	(0.00693)			-0.693***	(0.00638)		
UA	-0.916***	(0.0125)			-0.909***	(0.00836)		
V1	0.261***	(0.0178)	1.719***	(0.0302)	0.517***	(0.0178)	1.642***	(0.0263)
a1	-3.747***	(0.0635)			-3.858***	(0.0525)		
Number of individuals	84,728		79,402		125,633		97,061	
Number of spells	129,545		121,721		171,063		137,967	
Number of exits	121,721		75,286		137,967		101,237	
Observations	715,577				1,111,724			
Log Likelihood	-1,174,273				-1,494,204			

Notes:

Correlated Competing risks estimation: piecewise baseline and discrete distribution of unobserved heterogeneity with two mass points.E[V]=0 for both samples .

For 2005 estimation, Pr(Type I)=2.3 %; Pr(Type II)= 97.7%; V2(U-J)=-0.01; V2(J-U)=-0.041; For 2009 estimation, Pr(Type I)=2.07%; Pr(Type II)=97.9%; V2(U-J)=-0.011; V2(J-U)=-0.035;

Firm characteristics correspond to the previous job spell for the unemployment and to the current spell for the employment

For the reference category from personal characteristics, see table 3. Size_1_9, Services, temporary. Regional

unemployment rate and part-time coefficient are continuous variables.

Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table A5. Decomposition analysis for correlated consecutive events from the extended model.

	Unemployment to any job		Job to unemployment	
Total Effect: Difference between:	**16.72%**	**100%**	**-6.60%**	**100%**
S09 09	31.98%		21.23%	
S05 05	15.26%		27.83%	
Composition effects	**5.74%**	**34%**	**-0.09%**	**1.3%**
Business cycle effects	**10.98%**	**66%**	**-6.52%**	**99%**

Note: Evaluated using the 2005 model. S09 09 (S05 05) is the average survival probability at month 12, using the model of 2009 (2005) for the sample of 2009 (2005).
Source: Own elaboration from LWLS.

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