Nova School of Business and Economics

Universidade Nova de Lisboa

Dissertation, presented as part of the requirements for the Degree of Doctor of Philosophy in Economics

Essays on Job and Unemployment Protection: The Impacts on Unemployment Duration, Wages, and Fertility

Marta Cardoso Lopes, Nr. 669

A Dissertation carried out on the PhD in Economics, under the supervision of Professor Pedro Portugal

August 26 , 2019

Abstract

This thesis analyses the impact of job and unemployment protection on unemployment duration, wages, and fertility. The first chapter provides an extensive literature review on the impact of potential duration of unemployment benefits on unemployment duration, which is compared to a survival analysis that accounts for unobserved heterogeneity. The second chapter makes use of control function approach, with selection on treatment and outcome variables, to evaluate the impact of unemployment duration on re-employment wages. The third chapter uses both reduced-form and structural modelling approaches to evaluate the impact of job security on labour supply and fertility decisions.

Keywords: unemployment benefits, wages, fertility, job security

Acknowledgements

I am glad there is a physical component of this Dissertation that helps me to believe this PhD came to an end. From the start, or even before that, there was a considerably large number of people that encouraged me to achieve one of the greatest accomplishments of my life. Here, I would like to publicly acknowledge their importance and apologize for the unnamed that also occupy an important space in my heart.

First and foremost, I must thank my advisor, Pedro Portugal. He was one of the reasons I enrolled in the PhD in Economics program, and definitely one of the most important factors that contributed to my success in it. Since I first met him in 2011, Pedro offered me the necessary tools, especially his Econometrics and Labour knowledge, but also some of the challenging times, namely when it came to the actual writing of this dissertation. For both of them, I must recognize I am thankful.

Alongside with Pedro, other faculty members at Nova School of Business Economics (Nova SBE) were important contributions. I want to thank Pedro Pita Barros and Paulo M. M. Rodrigues for the extraordinary MSc teaching and for helping me to have a better view of what an Academic Economist is. I also want to thank Pedro Martins for the help provided with the Social Security data, and to Susana Peralta for all the fruitful discussions on our joint project on tax competition that enriched my skills portfolio, and for the help in taking some academic decisions too. I must not end the acknowledgement to Nova SBE without mentioning the opportunity to embrace teaching activities, the financial support on the presentation of some of this work at international conferences, and all the administrative assistance, especially in the job market.

I must also acknowledge the university that hosted me for the last three and a half years. At University College London (UCL) I was exposed to excellent lectures and to uncountable seminars provided by top Economists in the world. For this, I am deeply thankful to Pedro Carneiro and Hugo Reis that made this visiting possible. I also want to give my special thanks to Lars Nesheim, for his support on my last chapter of this dissertation, which was used as job market paper. Even without the official supervisor title, he took the time to teach me dynamic structural models, to discuss the contents of my paper, to attend all my presentations at UCL, and to push the quality of the paper to another level. Just as important were Aureo de Paula, Cormac O'dea, Elaine Kelly, Marcos Vera Hernandez, Michela Tincani, Mónica Costa Dias, and Sarah Cattan. Any good result from the job market is owed to you. I am also grateful to my colleagues at the Centre for Microdata Methods and Practice (especially Alan Crawford, Cristina Gualdani, and Priscilla Fialho) and at the Institute for Fiscal Studies, especially to the ones belonging to the Tag Rugby team (in particular Christine Farquharson and George Stoye), for the encouragement and support provided both in and out of the office.

Last but not least. To my brother and my sister-in-law that made living in London a lot easier. To my parents, my aunt, my uncles, and my grandmothers for backing me up, psychologically and financially, in all challenges, even when they did not understood their deep meaning. To my grandfathers that I hope to have the best seats in the sky to observe the presentation of a project with impact in the Portuguese economy, as they have always done through their lives, at Trinhão and Maria Gomes. To all the inhabitants and descendants of both villages that inspired me to work for the greater good. To the extraordinary companionship of Bruno Carvalho, Claudia Garcia, Gonzalo Paz, Guillermo Uriz, João Vaz, Pedro Freitas, Pedro Luís Silva, and Rui Costa. Probably you are the ones that know best how hard it was and how meaningful it is, this PhD. Finally, to the support and encouragement of all my friends, in Lisbon and in London, who always struggled to understand why have I extended my studies for six additional years. At this point I can tell you it was worth it, and this is only the start!

This work had the support of Fundação para a Ciência e Tecnologia (FCT) under the grant SFRH/BD/121190/2016.

Contents

Introduction

In response to the rapid increase in unemployment rates during the 1970s, the Economics literature focused on the study of job and unemployment protection. During that period, job search theory emerged with Mortensen (1977) developing a framework that also accounted for the limited duration of unemployment benefits, and Nickell (1978) developed a dynamic labour demand model which accounted for the costs of employment protection. The consideration for the labour market institutions has intensified in the 1980s, when most of the industrialized western countries were achieving historic records of unemployment rate since the end of World War II – United States reached 11% in 1983, United Kingdom, 12% in 1984, and Spain passed 20% for the first time in 1985.

The heterogeneity in the (un)employment adjustment to the aggregate (oil) shocks triggered a new literature in the 1990s, devoted to study which institutional features and their differences across labour markets act as main drivers of the differential rates of adjustment. Bentolila and Bertola (1990), Nickell (1997), and Blanchard and Portugal (2001) concluded that some characteristics of the European labour markets were important determinants to explain the lower flows into job creation and job destruction, or the so called *Eurosclerosis* (Giersch, 1985), which resulted in longer unemployment duration (Layard et al., 1991). Ljungqvist and Sargent (2008) nominated the two most important institutional drivers very clearly: "Europe has continued to have stronger employment protection and more generous unemployment insurance than the United States".

As documented by Nickell et al. (2005), during the 1980s, Europe has increased further the generosity of the benefit level and duration of unemployment insurance, as well as the protection of employment. In an apparently opposite direction, the introduction of fixed-term contracts with low severance payments was aimed at increasing the rates of job hiring (Cahuc and Postel-Vinay, 2002). This setting lead to the emergence of dual labour markets, especially in Southern European countries, where the permanent (or open-end) contracts with high severance payments contrast with fixed-term (or temporary) contracts with restrictions on duration and renewals (Booth et al., 2002a).

Portugal was no exception in following the European trends. Until the major labour market reforms during *Troika*[1] supervision, the country was constantly occupying the leading position in the OECD employment protection legislation (EPL) index regarding dismissals, especially due to the high severance payments for short tenure duration. For all the contracts starting in 2012, the minimum 3-month payment was eliminated and the 1-month severance for each full year of seniority was reduced to 20 days. In 2013, the severance payment was further reduced to 12 days.

The high protection of employment until recent years has motivated the increasing use of temporary contracts. Since early 2000s, the share of temporary contracts has been consistently around 20%, ranking among the top 5 in Europe. Even with such trial of increasing the flows out of unemployment, the long-term unemployment in Portugal reached its peak in 2013, with 9.3%. In consistency with the empirical and theoretical literature, Portugal had indeed the longest (limited) unemployment benefit potential duration until 2012 – up to 3 years and 2 months, excluding potential subsequent

[1]Term attributed to the team, composed of Economists of European Central Bank, European Commission and the International Monetary Fund (IMF), which was responsible to monitor the countries that received financial loans provided for by the European Union and the IMF in early 2010s.

unemployment assistance.

Considering the labour market institutions in Portugal, it is with no doubt an interesting case of study. This thesis focuses on the study of two institutional features that, albeit analysed separately, are both important in shedding light about the design of optimal joint policies that have already been discussed (Blanchard and Tirole, 2008). I analyse their consequences on labour market outcomes, such as unemployment duration, re-employment wages, and (female) labour supply, but also the unintended consequence of dual labour markets on fertility decisions.

In chapter 1, I do an extensive literature review on empirical studies that studied the impact of unemployment insurance, and its potential duration, on the hazard rate to employment and on the unemployment duration. I then evaluate the same effects using administrative data from the Social Security in Portugal between 2005 and 2012. I estimate complementary log-log survival models with unobserved heterogeneity. The identification relies on a policy reform introduced in 2007 that induced exogenous variation of potential durations across time age, career history and behaviour during the insured spells. The main results indicate that an increase of one month in potential duration of unemployment insurance leads to an average decrease of 9% in re-employment probability. The unobserved heterogeneity estimates indicate that about 20% of the individuals in the sample are characterized as long-term unemployed. Finally, recurring to a temporary non-anticipated extension of the potential duration of subsequent unemployment assistance in 2009, I estimate that eligible individuals have a 4% lower re-employment probability.

The main goal of unemployment insurance is to help individuals to find a better match in the job market. As the potential duration of unemployment benefits leads to longer unemployment duration, it is therefore important to also evaluate the consequences that longer spells of unemployment might have on re-employment outcomes.

In chapter 2, I estimate the impact of unemployment duration on re-employment wages. Using the same source of data (administrative data from the Social Security in Portugal between 2005 and 2012), I take a control function approach to overcome the difficulty raised by the simultaneity issues that arise in naive approaches. In the first stage I consider three alternative instruments to account for the endogeneity of the unemployment duration: the first, is based in the potential duration of unemployment benefits, the second, explores the age discontinuity in the potential duration rules of unemployment benefits; and the third, takes advantage of the change in the rules of the potential duration introduced by a reform in 2007. The main results show that each additional month in unemployment duration is expected to decrease the re-employment wage by around 0.5%. The effect seems to be robust when I control for selectivity on unemployment durations, and also when I control for selectivity on re-employment wages, either by considering the conventional Heckman two-step approach or the augmented inverse probability of censoring weighted estimation.

In both previous studies I exclude the women that took advantage of long duration entitlement to unemployment insurance to take maternity leave and come back to a constant inflow of income. Thinking of these dynamics into a broader sense, I redirect my research towards the different types of job contracts in Portugal, which differ significantly in terms of job security. This research is motivated by the facts: 1) the restrictions on duration and number of allowed renewals implied by fixed-term contracts generate lower job security; 2) in many countries, these contracts are particularly common among young women.

In chapter 3, I study the impact of job security on fertility decisions. Using a policy reform, which took place in Portugal in 2003, I show that the lower job security associated with fixed-term contracts decreases the likelihood of giving birth. The negative effect is particularly strong for shorter contracts. To identify the different channels that explain these results, I build and estimate a dynamic life-cycle structural model where women decide both labour supply and fertility, conditional on the characteristics of the job contract. I then simulate two different labour market policies that have been discussed in the public debate. Imposing an automatic conversion into permanent contracts, with higher job security, at the end of the fixed-term contract limit, decreases the number of childless women by 8.3%. In contrast, applying contract-specific tax rates, penalizing fixed-term contract wages and subsidizing permanent contract wages, induces 20% of the women who are already mothers to have their second child. These results corroborate the reduced-form evidence, showing that job security is especially important at first birth, whereas income is relatively more important for subsequent birth decisions.

Chapter 1

Revisiting the Impact of Potential Duration of Unemployment Benefits on Joblessness Duration

1.1 Introduction

Under imperfect information (Stigler, 1961, 1962), the optimal strategy for a job seeker consists of refusing all proposals for which the value is lower than her reservation wage, thus accepting all the others. According to job search theory (Mortensen, 1977), unemployment benefits affect individuals search behaviour by lowering the opportunity cost of being unemployed, thus increasing the reservation wage of insured individuals. This reaction decreases the job search effort and the number of acceptable jobs for the beneficiaries and, as a consequence, increases the probability of remaining unemployed for a longer period of time (Nickell, 1979; Lancaster, 1979).

To foster re-employment, unemployment benefits have a pre-determined potential duration associated.[1] Therefore, when the unemployment benefits approach the point of exhaustion, the reservation wage decreases sharply generating spikes in the unemployment escape rates which may triple, when compared with the rest of the spell (Moffitt, 1985). Since the 1980s, several authors have devoted their research to investigate the impact of potential duration of unemployment benefits on the spikes at the end of the unemployment spells.

In the first part of this chapter, I revisit the literature and present an exhaustive survey of the empirical evidence on the impact of potential duration of unemployment benefits on the unemployment duration (and on the hazard rate into re-employment). Filges et al. (2015) attempted a similar literature review but the authors adopted a different strategy, by replicating the results of 12 studies, following the Campbell Collaboration guidelines. This chapter complements this attempt by exposing a careful comparison of 48 empirical studies in 5 different groups: the impact of i) unemployment insurance (UI) recipiency, ii) potential duration of UI benefits, iii) time to exhaust the potential duration of UI benefits, and iv) exhaustion of potential duration of UI benefits, on the hazard rate into re-employment, and v) the impact of potential duration of UI benefits on unemployment duration.

The main contributions of this literature review are twofold. Firstly, it provides the reader with an important source of information regarding the current stand point of the empirical literature. Secondly, the inclusion of detailed information on the country, period, data and restrictions on each selected sample, and the indication of the estimation model adopted in each study, make it a useful tool for future researchers in understanding how different sample and estimation choices might lead to different results.

[1] Only Belgium puts no limit on the duration of unemployment benefits but compensates the incentive with large cuts on the amount of the benefit along the spell.

In the empirical part of this chapter, I use administrative data from Portuguese Social Security records, between 2005 and 2012. The dataset contains very rich information on the career history of the individuals and the different types, potential duration and amount of unemployment benefits. Additionally, the period of the data includes two reforms that changed the potential duration of unemployment insurance. These reforms are a crucial source of variation to estimate the impact of the unemployment benefits recipiency and potential duration on the hazard rate into re-employment.

Hence, this study also adds up to the literature by analysing a country with generous unemployment benefits (Stovicek et al., 2012), during a period which has not been the focus of any study at present, as it is more recent than what was used in other published studies. Addison and Portugal (2008a,b), and Portugal (2008) found a substantial negative impact of UI recipiency on the hazard rate into re-employment in Portugal in the 90s and early 2000s, but very little is know on the impact of the potential duration. Only Addison and Portugal (2008b) have studied such impact and the estimate is remarkably lower than other studies that have used the same empirical specification. In this chapter I extend the analysis of the authors by studying a more recent period and by including the impact of subsequent unemployment assistance (UA), which has also been many times neglected in the literature.

In this study I use variation across time, age, and working experience as exogenous shifters of potential duration, hence enabling to identify and estimate the causal impact of potential duration on re-employment rate. I explore two policy reforms that changed the potential duration of unemployment insurance and unemployment assistance in 2007 and 2009/2010, respectively. By strengthening the link between unemployment insurance potential duration and working experience, the policy change in 2007 has increased the potential duration of some of the new beneficiaries, decreased the potential duration for others, while still leaving a (control) group unaffected with respect to its potential duration. Furthermore, the policy change in 2009/2010 increased the potential duration of unemployment assistance with no additional means testing, age or experience requirements.

The main results from the estimation of complementary log-log hazard models, with two mass points of unobserved heterogeneity (Heckman and Singer, 1984), reveal that UI recipiency decreases the hazard rate of re-employment by 73%, and that an increase of one month in potential duration of unemployment insurance leads to an average decrease of 9% in the hazard rate. Both results are larger than what was previously found in the literature (the upper bound estimates were on a magnitude of 50% and 8% for changes in recipiency and duration, respectively). When unobserved individual heterogeneity is taken into account, the estimates indicate that around 20% of the individuals in the sample have a substantially lower hazard rate to employment. This result is below of what has been reported in the literature (around 40%).

Taking advantage of a non-anticipated temporary extension of 6 months (at a lower benefit amount) on the potential duration of subsequent unemployment assistance in 2009 and 2010 I observe that recipiency of such an extension decreased the hazard rate into employment by 10%. Overall, the results show that both unemployment insurance and subsequent unemployment assistance have an important impact on the shape of the re-employment rates, especially in scenarios with large long-term unemployment rates.

The remainder of the paper is structured as follows. In Section 1.2 I provide the literature review of 48 studies. Section 1.3 explains the rules of potential duration of unemployment benefits in Portugal during the period of study. Section 1.4 describes the data. Section 1.5 provides descriptive evidence on the research question. Section 1.6 details the identification model and the estimation procedure. Section 1.7 presents the main results and discusses the empirical findings of the hazard models including the results for the temporary policy change. Finally, section 1.8 concludes.

1.2 Literature Review

In this section I provide a literature review of the empirical studies that evaluated the impact of unemployment insurance (UI) and its potential duration on the transition rate to employment and on the unemployment duration.[2] The selection of studies is based on the availability of estimates' interpretation, given that the usually non-linear estimation requires additional computation to obtain the average effects.

Even though most of the studies were already mentioned on the review by Filges et al. (2015), only 12 of them were actually analysed, as the authors wanted to re-compute the effects using the same methodology. The survey in this chapter provides a clear comparison of the original estimates on 48 studies, thus offering the reader an important source of information regarding the current stand point of the empirical literature in this topic.

1.2.1 Overview of Data and Methodologies

Table 1 reviews the studies that provide an estimate for the average effect on the transition rate to re-employment. Each panel focuses on each independent variable – UI recipiency, potential duration of UI, time to exhaust UI, and the moment of UI exhaustion. Table 2 reviews the average effect of potential duration of UI on the duration of unemployment. In each table/panel, the studies are organized by chronological order.

In total, 15 countries were analysed by the studies in this review. There is a major

[2]Katz and Meyer (1990a); Hunt (1995); Addison and Portugal (2008b) also studied the impact on the hazard rate to other exits than employment. This literature review focuses on the hazard rate into re-employment to be in line with the research question of this paper.

emphasis in the United States (14 out of the 48 studies), but Austria, Germany, and Spain, have also a reasonable share (6, 5 and 4 studies, respectively).

Not surprisingly, in this topic, 72% of the studies used administrative data. When studying unemployment benefits it is important to get precise data on the length of their potential duration. However, almost all studies that evaluate the impact of UI recipiency use surveys – in order to compare recipients with non-recipients. Only Vodopivec (1999), in this group of studies, uses administrative data which, in the case of Slovenia, contains information about both recipients and non-recipients of UI.

Most of the studies included both genders in their analysis. However, it should be emphasized that no study focused solely on women and some of studies (16) restricted their sample to men only, due to gender differences in job search behaviour. The period of the analysis used in the male-only studies started, on average, roughly 6 years before than the studies that included both genders, which indicates the literature is accommodating for the increasing female labour force participation in more recent years.

Four main types of estimations were used: semi-parametric hazard models, parametric hazard models, the linear probability model (LPM), and OLS. In the first case, I distinguish between continuous (Cox (1972) or piecewise) and discrete (cloglog or logistic) proportional hazards. In the second case, I present the distribution that was assumed for unemployment duration (exponential, logistic, normal, or weibull). In the last case, I distinguish the studies that adopted an regression discontinuity design (RDD).[3].

[3]Note that Lalive (2007, 2008) and Schmieder et al. (2012) also presented robustness to the RDD estimates by testing different bandwidths on the local linear regressions

Table 1: Literature Review - The impact on transition rate to employment

Study	Country	Period	Data	Sample		Hazard Model[a]	Unob. Het[b]	Avg. Effect[c]	Change
UI Recipiency									
Adamchik (1999)	Poland	1994-1996	survey	both	15-60	Cox PH	no	-27%	
Vodopivec (1999)	Slovenia	1990-1992	admin	both	15-64	Cox PH	no	-42%	
Addison and Portugal (2008a)	Portugal	1992-1997	survey	both	29-55[d]	Cox PH	two mass points	-50%	
Addison and Portugal (2008b)	Portugal	1992-1996	survey	male	16-64	piecewise	gamma	-42%	
Portugal (2008)	Portugal	1998-2008	survey	male	15-64	piecewise	no	-38%	
UI Potential Duration (for each month)									
Katz and Meyer (1990a)	USA[e]	1978-1983	admin	male	17-54	Cox PH	gamma	-8%	3
Hunt (1995)	Germany	1983-1988	survey	both	44-48	Cox PH	gamma	-5%	10
Lalive and Zweimüller (2004)	Austria	1986-1998	admin	male	45-54	Cox PH	no	-0.1%	45
Lalive (2007)	Austria	1987-1991	admin	male	46-53	LPM	no	-0.1%	43
	Austria	1987-1991	admin	female	46-53	LPM	no	-1.4%	43
Van Ours and Vodopivec (2008)	Slovenia	1997-2001	admin	male	15-64	LPM	no	[0.8,1]%	[-9,-3]
	Slovenia	1997-2001	admin	female	15-64	LPM	no	[-1.7,-0.4]%	[-9,-3]
Caliendo et al. (2013)	Germany	2001-2007	admin	male	44-46	logistic	five mass points	-3.8%	6
	Germany	2001-2007	admin	female	44-46	logistic	four mass points	-5.7%	6
Farber and Valletta (2015)	USA	2000-2005	survey	both	20-64	normal	no	-2.6%	3
	USA	2007-2012	survey	both	20-64	normal	no	n.s.	10
Le Barbanchon (2016)	France	2000-2002	admin	both	16-50	Cox PH	no	-3.2%	8
Time to exhaustion (for each month)									
Gonzalo (2002)	Spain	1989-1991	admin	male	25-55	normal	two mass points	-13.7%	
Arranz and Muro (2004)	Spain	1987-1989	admin	both	18-59	exponential	three mass points	-11.2%	
Addison and Portugal (2004)	USA	1995-1998	survey	both	20-60	piecewise	gamma	-13.2%	
Addison and Portugal (2008b)	Portugal	1992-1996	survey	male	16-64	piecewise	gamma	[-9.7;-4.3]%	

[a] Cox PH - Cox Proportional Hazard Model; piecewise - Piecewise-constant Hazard Model, LPM - Linear Probability Model

[b] Unobserved Heterogeneity is either parametric (gamma distributed) or non-parametric (in which case the number of mass points is indicated)

[c] When needed, the average effect is re-scaled as if the change was 1 month. To scale back to the original size of the effect use the last column.

[d] Ages 29,30,34,35,39,40,44,45,49,50,54, and 55 only. [e] States: GA, ID, LA, MO, NM, NY, NC, PA, SC, VT, WA, and WI only.

Table 1: Literature Review - The impact on transition rate to employment (continuation)

Study	Country	Period	Data	Sample		Hazard Model[a]	Unob. Het[b]	Avg. Effect	Measured at
At the Time of Benefit Exhaustion									
Han and Hausman (1990)	USA	1980-1981	survey	both	20-65	extreme	gamma	up to 2.2	1 month
Katz and Meyer (1990b)	USA[f]	1979-1980	admin	both	17-54	Cox PH	gamma	2.36	1 month
Meyer (1990)	USA	1978-1983	admin	male	17-54	cloglog	gamma	3.45	1.5 months
Fallick (1991)	USA	1984	survey	both	15-64	Cox PH	no	n.s.	1 month
Narendranathan and Stewart (1993)	UK	1978-1979	admin	male	<64	logistic	two mass points	n.s.	1 month
Carling et al. (1996)	Sweden	1991	admin	both	<55	cloglog	gamma	1.69	2 months
Lubyova and Van Ours (1997)	Slovakia	1991-1992	admin	both	15-64	Cox PH	no	n.s.	1 month
	Slovakia	1994-1995	admin	both	15-64	Cox PH	no	1.54 to 4	1 month
Rogers (1998)	USA[g]	1980-1986	admin	male	<55	logistic	no	1.3 to 5.4	1 month
Vodopivec (1999)	Slovenia	1990-1992	admin	both	15-64	Cox PH	no	2.14	1 month
Puhani (2000)	Poland	1994	survey	male	18-55	logistic	two mass points	1.88	3 months
	Poland	1994	survey	female	18-55	logistic	two mass points	1.93	3 months
Bover et al. (2002)	Spain	1987-1994	survey	male	20-64	logistic	two mass points	up to 1.47	1 month
Jurajda and Tannery (2003)	USA[g]	1980-1985	admin	both	15-64	logistic	two mass points	3.51 to 6	1 week
Røed and Zhang (2003)	Norway	1990-1999	admin	male	<60	cloglog	five mass points	1.4	2 months
	Norway	1990-1999	admin	female	<60	cloglog	four mass points	1.5	2 months
Addison and Portugal (2004)	USA	1998	survey	both	20-61	piecewise	gamma	3.4	1 week
Jenkins and García-Serrano (2004)	Spain	1987-1991	admin	male	20-59	logistic	gamma	up to 1.3	2 months
Pellizzari (2006)	Europe[h]	1994-2001	survey	both	15-64	cloglog	normal	n.s.	2 months
Van Ours and Vodopivec (2006)	Slovenia	1997-2001	admin	male	18-45	Cox PH	no	2.27	1 month
	Slovenia	1997-2001	admin	female	18-45	Cox PH	no	2.48	1 month
Card et al. (2007b)	Austria	1981-2001	admin	both	20-50	Cox PH	no	3.16	1 week
Schmitz and Steiner (2007)	Germany	1995-2004	survey	male	< 58	logistic	two mass points	up to 1.67	2 months
	Germany	1995-2004	survey	female	< 58	logistic	two mass points	up to 1.58	2 months
Rebollo-Sanz (2012)	Spain	2000-2007	admin	male	18-55	cloglog	three mass points	up to 2.5	1 month
	Spain	2000-2007	admin	female	18-55	cloglog	three mass points	up to 1.7	1 month
Geerdsen et al. (2018)	Denmark	1998-2003	admin	male	25-44	cloglog	two mass points	1.1	2 months

[f] Missouri [g] Pennsylvania [h] Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, United Kingdom

Table 2: Literature Review - The impact of potential duration of unemployment benefits on unemployment duration

Study	Country	Period	Data	Sample		Hazard Model[a]	Unob. Het[b]	Variability[c]	Avg. Effect
UI Potential Duration (for each month)									
Holen (1977)	USA[d]	1969-1970	admin	both	>17	OLS	no	S	0.8
Moffitt and Nicholson (1982)	USA	1976	survey	both	≥ 17	OLS	no	P,S	0.1
Moffitt (1985)	USA	1978-1983	admin	male	17-82	Cox PH	no	P,S	0.15
Ham and Rea Jr (1987)	Canada	1975-1980	admin	male	18-64	logistic	three mass points	P	0.33
Katz and Meyer (1990a)	USA	1978-1983	survey	both	20-65	Cox PH	gamma	P,S	[0.16,0.22]
Gritz and MaCurdy (1992)	USA	1978-1985	survey	male	young	logistic	no	P,S	0.1
Lindeboom and Theeuwes (1993)	Netherlands[e]	1982-1984	admin	both	17-63	piecewise	three mass points	E	1.3
Winter-Ebmer (1998)	Austria	1986-1991	admin	male	< 65	weibull	gamma	P,S,A	0.03
Terrell and Sorm (1999)	Czech Repub.	1992-1993	admin	male	15-64	weibull	gamma	P,E	0.38
	Czech Repub.	1992-1993	admin	female	15-64	weibull	gamma	T,E	0.28
Card and Levine (2000)	USA	1995-1997	admin	both	18-65	logistic	no	T,S	0.31
Belzil (2001)	Canada	1972-1984	admin	male	18-25	weibull	four mass points	T,E	[0.14,0.21]
Lalive and Zweimüller (2004)	Austria	1986-1998	admin	both	45-54	Cox PH	no	P,S,A	0.055
Lalive et al. (2006)	Austria	1987-1991	admin	both	35-54	Cox PH	two mass points	P,A,E	[0.05,0.1]
Van Ours and Vodopivec (2006)	Slovenia	1997-2001	admin	both	18-45	Cox PH	no	P,E	0.19
Lalive (2007)	Austria	1987-1991	admin	both	46-53	OLS-RDD	no	P,S,A	0.02
Lalive (2008)	Austria	1986-1998	admin	male	46-53	OLS-RDD	no	P,S,A	0.09
	Austria	1986-1998	admin	female	46-53	OLS-RDD	no	P,S,A	0.32
Schmieder et al. (2012)	Germany	1975-2004	admin	both	40-43	OLS-RDD	no	P,A	0.21
Farber and Valletta (2015)	USA	2000-2012	survey	both	20-64	normal	no	P,S	0.06
Le Barbanchon (2016)	France	1999-2004	admin	both	< 50	Cox PH	no	A,E	0.2

[a] Cox PH - Cox Proportional Hazard Model; piecewise - Piecewise-constant Hazard Model, LPM - Linear Probability Model

[b] Unobserved Heterogeneity is either parametric (gamma distributed) or non-parametric (in which case the number of mass points is indicated)

[c] P - policy; S - space; A - age; E - experience

[d] The sample comes from an experiment that was conducted in San Francisco, Boston, Phoenix, Seattle and Minneapolis-St. Paul.

[e] Leiden only.

Most of the studies opted for the estimation of hazard models as the lack of large longitudinal datasets leads to high proportion of censored observations. Throughout time, the proportional hazards (PH) models were the mostly used ones in these studies. The PH method is popular because it does not require any assumption on the baseline hazard function. The continuous version (Cox PH) is even more handy as it also does not demand the hazard to be constant in each interval of time, which is required when estimating discrete proportional hazard models or piecewise-linear models. Among the parametric models, the logistic distribution is the most popular. As Jenkins and García-Serrano (2004) justify, this distribution facilitates the estimation of duration models with large datasets.

In a separate column, I indicate whether the study has tested for the presence of unobserved heterogeneity and if so, which type was used.[4] In most of the cases, a mixing discrete distribution of mass points (Heckman and Singer, 1984) was applied, and in that case I also report the number of mass points chosen. 16 out of the 27 selected studies, that have controlled for unobserved heterogeneity, have chosen this methodology. More than half of the studies have chosen two mass points, which were typically associated with short- and long-run unemployment. 10 studies have opted to use the gamma distribution and the 20 studies that have not controlled for it at all display a "no" in the respective column of the table.

For Table 2, I have also included the exogenous sources of variation in the potential duration of unemployment benefits that were used in each study. This column shows the letter P, if the study used a policy reform, S, if the study used spacial variation, A, if the study used age variation, and E, if the study used experience variation. Most of the studies relied on policy reforms. The second mostly used source of variation was space variation as many policies, like in the USA and in Austria, are state- or

[4]Note that the average effects, however, are not necessarily the ones that accounted for unobserved heterogeneity as I have rather chosen the most preferred specification reported by the authors.

county-specific. Finally, some other studies also used variation in age and experience, as in many countries, the potential duration varies (discontinuously) with these two determinants.

1.2.2 Review on the Impact of UI Potential Duration

Before analysing the intensive margin of the impact unemployment benefits (potential duration), it is important to know what is the extensive margin of unemployment benefits (recipiency) on the transition from unemployment to employment. According to the first panel of Table 1 the average effect ranges between -27% and -50%, where studies that account for unobserved heterogeneity, on average, present a higher magnitude of the estimated effect.

In the second panel of Table 1, I have listed the studies that study the impact of one additional month of potential duration on the transition rate into employment. Note that different studies report different measures of time, either because of differences in the frequency of the data or because of different policy changes that implied different changes in the potential duration of UI. For these reasons, I have scaled the average effects to the impact of one month, but maintained the original changes in potential duration used by the authors in the last column of the table.

The impact of each additional month of potential duration of UI on the transition into employment is most of the times negative. The larger (negative) effects seem to be around the period when the unemployment rates were higher – in particular the -8% found by Katz and Meyer (1990a). The estimates of more recent studies, between 2000 and 2012, lie within a narrower range, between -5.7% and -2.6%. Only Farber and Valletta (2015) finds a non-significant effect (n.s.) for the period between 2007 and 2012. This result might in part be explained by the type of data (Displaced Workers Survey from the Current Population Survey), which presents some measurement error in the duration and transitions out of joblessness. Van Ours and Vodopivec (2008) presents a positive effect for men but note that the outcome variable in the study is the transition rate into a permanent job rather than a temporary one. As it will be discussed in the third chapter of this thesis, males are more likely to get permanent jobs than females, so this fact can help to explain such results. Consistently, both Lalive (2007) and Caliendo et al. (2013) also find larger negative results for female samples.

Most of the studies that study the impact of the potential duration of UI on the transition rate to employment take a different perspective and analyse how much time is left until the exhaustion of the potential duration. In the third panel of Table 1, I present the 4 studies that have studied the linear impact of time to exhaustion on the transition rate to employment. The average effects reported in the studies about Spain and USA report similar magnitudes (between 11.2% and 13.7%), whereas the study about Portugal reports smaller estimates (-4.1% for exiting into dependent employment and -9.7% for exiting into self-employment) . The range of estimates can be explained by different types of exits that the authors study, being the exits into self-employment and part-time the ones with the largest effects and the exit to full-time contracts the ones with the smallest effects. As discussed in these studies, most of the individuals that leave unemployment for the self-employment or part-time reveal different job search preferences. The authors emphasize that individuals leaving unemployment to self-employment or part-time at the very last months of the potential duration of unemployment benefits typically search for full-time dependent employment until they approach the end of the benefit span – either because the reservation wages are much larger than part-time wages, or because they never engaged into self-employment before, as this type of job is not entitled to UI.

The vast majority of studies explores the time-varying dynamics of the unemployment benefits, as there is wide evidence that the transition rate into employment is not uniform along the spell but rather U-shaped along the potential duration of the unemployment benefits. That explains why the last panel of Table 1 contains the largest number of studies in this literature review. As the potential duration of UI varies substantially between countries and periods of study, not all points along the UI spell would be comparable across studies. In this table I give preference to the last measured effect before exhaustion, which still presents a wide heterogeneity, between 1 week and 3 months before exhaustion. The spike in the hazard to employment ranges from non-significance (n.s.) to 6 times more, relative to the last point at where it is measured. In general, the closer to exhaustion is the point, at which it is measured, the larger is the average effect. The average across all studies with significant estimates predicts the hazard rate to more than double (2.5) near the benefits exhaustion.

Once again, in the panel for the spike analysis, the larger effects are associated with studies that analyse the period of the 1980s, when unemployment was the highest in most of the countries.[5] Note also that, once again, in 4 out of the 6 studies that presented results by gender, the spike in benefit exhaustion is larger for female than for male – these studies included Slovakia, Poland, Norway, and Slovenia. For the studies about Germany and Spain, actually the male sample seems to present a higher spike in the transition to employment at the point of exhaustion, which might be related to the lower female attachment to the labour market – especially in Germany, where the incidence of full-time employment on women is much lower than that of men, according to the OECD.

Finally, Table 2 presents the impact of potential duration of unemployment benefits on unemployment duration. The range of the average effect is widely spread, and lies between 0.02 and 1 additional months of unemployment duration, for each additional month of unemployment benefits.

[5]Note that the studies in this panel did not include the latest crisis yet, as the latest date in the data to be used was 2007.

A few things should be highlighted in the last table. Most of the studies report an estimate below 0.4. The outliers have two factors in common: both did not use any policy reform, selected only specific regions in the country of study, and applied alternative estimation methods. Holen (1977) did not account for the censoring in the data, by using an ordinary least squares model. Lindeboom and Theeuwes (1993) applied a piecewise model, which assumes a constant hazard in each interval of time (a week in this case). On the other hand, it should also be noted that the studies that reported the smallest estimates have used survey data – also the case in Table 1.

1.3 Institutional Background

I dedicate this section to the rules of the potential duration of unemployment benefits in Portugal during the period of the analysis, January 2005 to March 2012. The description of the remaining rules (funding, entitlement requirements, monetary amounts and training) is presented in the Appendix of this chapter.

Until January 2007, the potential duration of Unemployment Insurance (UI) benefits in Portugal was solely determined by age, except for individuals at age of 45 and over. Since January 2007, the entitlement for individuals below age 45 also depended, to a minor extent, on contributions to Social Security (since last spell of unemployment) and working experience.[6] As a consequence of the reform, some individuals were entitled to a longer potential duration, others to a shorter potential duration, and others had no changes in their entitlement. Note that individuals at age of 45 and over were also affected by the reform. In Table 1.4 I summarize the potential duration for each group of individuals.

[6]The number of months of contributions to Social Security (SS months) might not correspond to number of months with labour market experience as, for example, receiving disability benefits also counts for contribution purposes.

Table 1.4: Potential Duration Rules of UI (all values in months except age)

	Before Jan 2007		After Jan 2007		
Age at UI Start	Potential Duration	Extra Potential Duration †	Social Security Contributions‡	Potential Duration	Extra Potential Duration †
<30	12	-	≤24	9	-
			>24	12	1
[30,40[	18	-	≤48	12	-
			>48	18	1
[40,45[	24	-	≤60	18	-
			>60	24	1
≥45	30	2	≤72	24	-
			>72	30	2

† for each 5 years of working experience in the 20 years preceding the date of unemployment
‡ Social Security contributions since last spell of unemployment include both the recipiency of wages as well as other benefits such as the ones to insure parental leave and disability.

Before January 2007, beneficiaries aged less than 45 years old were entitled to a fixed potential duration of UI benefits, regardless of the number of months they have contributed to Social Security. Beneficiaries that were younger than 30 years old would be entitled to 12 months; beneficiaries aged between 30 and 40 years old would be entitled to 18 months; beneficiaries aged between 40 and 45 years old would be entitled to 24 months; and beneficiaries aged 45 or more would be entitled to at least 30 months. Only for the last group of beneficiaries the working experience mattered. If they had working experience of at least 5, 10, 15 or 20 years, then the potential duration of UI would increase to 32, 34, 36, or 38 months, respectively.

For beneficiaries unemployed after January of 2007 the working experience became more important in the determination of their UI potential duration. A beneficiary younger than 30 years old would be entitled to: 9 months if she worked 2 years or less; 12 months if she worked between 2 and 5 years; 13, 14, 15, or 16 months if she worked at least 5, 10, 15, or 20 years, respectively.[7] A beneficiary aged between 30 and 40

[7]Note that even though the legal minimum age to work in Portugal is 16 years, it is however allowed to start working before, as long, as the individual is supervised by an adult. In the dataset there are very few individuals that are entitled to 16 months, and this should be no surprise as in the Portuguese

years old would be entitled to: 12 months if she worked 2 years or less; 18 months if she worked between 4 and 5 years; 19, 20, 21, or 22 months if she worked at least 5, 10, 15, or 20 years, respectively. A beneficiary aged between 40 and 45 years old would be entitled to: 18 months if she worked 5 years or less; 24 months if she worked more than 5 non years ; 25, 26, 27, or 28 months if she worked at least 5, 10, 15, or 20 years, respectively. A beneficiary aged 45 years old or more would be entitled to: 24 months if she worked 6 years or less; 30 months if she worked more than 6 years; 32, 34, 36, or 38 months if she worked at least 5, 10, 15, or 20 years, respectively.

During the entire period in analysis, the unemployed in Portugal could also be entitled to Unemployment Assistance (UA) in two cases. First, if the individuals did not fulfil the minimum requirements to receive UI benefits, especially in terms of working experience – Initial Unemployment Assistance. Second, if the individuals exhausted the potential duration of UI benefits – Subsequent Unemployment Assistance. Both types of benefits provided a lower amount than that of UI benefits.[8]

Before 2007, the subsequent UA beneficiaries were entitled to half the potential duration of UI, according to the working experience they had when they started the UI, and to their age at the end of UI. After the reform, the potential duration of UA beneficiaries also changed. The changes were applied to all the individuals that claimed UA after January 2007, regardless of when they started UI. Therefore, the reform affected both individuals that were affected in the potential duration of UI, and individuals that were not affected by the reform in the potential duration of UI, but were affected in the potential duration of the subsequent UA. In the Appendix of this chapter I provide extensive tables on all the possible entitlement situations according to

Labour Force Survey of the first quarter of 2005, when the dataset in this study starts, there were 7.3% of the individuals aged below 30 years old responding they started their first job before they were 16 years old.

[8]Once again I focus here on the description of the potential duration of the benefits and I write the remaining rules in the Appendix of this chapter. The analysis on initial UI is out of the scope of this chapter.

the age and starting date of UI and UA. The following table summarizes the potential duration of UA recipients for each group of individuals, according to their age and experience.

Table 1.5: Potential Duration Rules of subsequent UA (all values in months except age)

	Before Jan 2007		After Jan 2007		
Age at UA Start	Potential Duration	Extra Potential Duration †	Social Security Contributions‡	Potential Duration	Extra Potential Duration †
<30	6	-	≤ 24	4.5	-
			>24	6	0.5
[30,40[	9	-	≤ 48	6	-
			>48	9	0.5
[40,45[	12	-	≤ 60	9	-
			>60	12	0.5
≥ 45	15	1	≤ 72	12	-
			>72	15	1

† for each 5 years of working experience in the 20 years preceding the date of unemployment
‡ Social Security contributions since last spell of unemployment include both the recipiency of wages as well as other benefits such as the ones to insure parental leave and disability.

In an unprecedented year (2009) in terms of long-term unemployment in Portugal – 4.4% of the active population was unemployed for one year or more –, the Portuguese Government, decided to extend the UA potential duration (both initial and subsequent) for six months in order to give extra support to the long-term unemployed. Contrary to the job experience and age pre-requisites necessary to be a UI or UA beneficiary, the extended benefit program only required the beneficiaries to exhaust their UA potential duration, without any type of means-test. In the subsequent year, when the long-term unemployment rate kept rising to 5.8%, the same Government renewed the measure, under the same conditions. However, considering the large expenditure in 2009, the policy renewal was forcibly cancelled in July 2010. Such decision was carried out under the *Stability and Growth and Program 2010-2013*, by the tripartite committee responsible for the financial assistance programme in Portugal.

1.4 Data

1.4.1 Data source

The database used in this chapter is a 1% representative sample (39 972 individuals) of the panel data entitled *Microdados do Sistema de Informação da Segurança Social*, which contains monthly administrative records of the unemployment beneficiaries inflow between January 2005 and March 2012.

Regarding the unemployment information, this dataset contains the reason for unemployment, the type of benefit, the daily amount received, the potential and actual duration of each spell, and the reason for the end or interruption of the unemployment benefit when the spell is not censored.

The same dataset also contains information on employment registers. For each individual's working experience, the dataset includes the type of job (employee or employer) and the duration of the job spell. In case the employment register occurred between January 2005 and March 2012, the dataset also includes the earnings and the economic activity of that job.

Finally, the dataset also includes some demographic variables, such as gender, age, region of residence and nationality.

1.4.2 Sample construction

From the 39 972 individuals available, 105 are excluded due to irregularities in their process.[9] Additionally, I exclude those individuals (8 283) that benefited from training, occupational programs or parenthood benefits during the period of unemployment

[9]Those individuals were in one of the two situations: either not following the rules for receiving the unemployment benefit and therefore had to give back the benefit; or without information during some months of the period of unemployment being impossible to distinguish if the two moments observed were disjoint unemployment spells or not.

insurance, as I do not model those decisions in this study.[10]

This chapter focuses on the study of Unemployment Insurance (UI) and Subsequent Unemployment Assistance (UA). Therefore, I exclude unemployment spells of alternative unemployment insurance / assistance benefits (initial unemployment assistance, partial benefits, benefit associated to wage arrears, benefit of public school professors, benefit of emigrant workers, etc.) due to differences on the eligibility conditions.[11]

Finally, I exclude the individuals that had no complete information on all the variables of interest. The final sample accounts for 22 000 beneficiaries. From this group I select 3 623 that were also observed as subsequent UA beneficiaries and create another sample. The second sample is particularly useful to study the impact of the temporary increase in the potential duration of unemployment assistance in Portugal during the whole year of 2009 and first half of 2010.

1.4.3 Descriptive Statistics

Descriptive statistics on the variables of interest and some demographic characteristics for all individuals that benefited from UI are provided in Table 1.6.

The first panel shows that joblessness duration is, on average, 5 months longer than the UI subsidized duration. In fact, 38% of the individuals that received UI have experienced some joblessness period without UI benefits. Less than half (43%) of these individuals also received UA after exhausting UI potential duration. Therefore, in the estimation I use the joblessness duration (i.e., time until re-employment) as measure of duration of the unemployment spell.

The second panel shows that the average level of benefits matches roughly the general rule of 65% of gross wage. The same panel shows that approximately 50%

[10] For more details on the interruptions see the sample selection note in the Appendix of this chapter.

[11] These types of benefits account for 30% of the unemployment spells recorded in the sample.

Table 1.6: Descriptive statistics of UI beneficiaries

	Mean	Std	Min	Max
Benefit duration (UI) (in months)	13.51	10.85	1	38
Joblessness duration (in months)	18.48	18.28	1	87
Benefit dur. = Joblessness dur.	0.62			
Level of benefits (€/month)	496.15	207.18	43.5	1257.6
Reference wage (€/month) [1]	745.59	584.54	48.9	14407.5
$\mathbb{1}$ < minimum guaranteed benefits	0.20			
$\mathbb{1}$ Minimum guaranteed benefits	0.29			
$\mathbb{1}$ Maximum level of benefits	0.01			
Tenure (in years)	4.74	5.22	1	45
Number of jobs	2.30	1.85	1	28
Reason of unemployment				
end of fixed-term contract	0.38			
redundancy	0.23			
mutual agreement	0.17			
plant closure or mass lay-off	0.10			
others	0.12			
Year of unemployment				
2005	0.15			
2006	0.16			
2007	0.10			
2008	0.12			
2009	0.13			
2010	0.12			
2011	0.18			
2012 (until March)	0.04			
Age (in years)	39.18	12.10	17	74
Male	0.49			
Region of Residence				
North	0.43			
Center	0.17			
Lisbon	0.25			
Alentejo	0.06			
Algarve	0.04			
Azores and Madeira	0.05			
Re-employed (any job)	11 335 (10 762 as employees)			
Retired, disabled, dead	903			
Censored spells	9 762			
Transited to UA	3 626			
Sample	22 000			

[1] Reference wage is calculated as the sum of the earnings received within the first 12 months of the 14 months previous to the date of unemployment, divided by 12.

of the beneficiaries received either the minimum guaranteed amount or below that if their net wage was already below that defined amount, and that only 1% received the maximum allowed amount. The first fact is in line with the average reference wage being below the average monthly-basis gross wage in Portugal during the period in analysis, approximately 893€. The second fact might be explained by the low average tenure length in the sample – less than 5 years – and by the previous switch between jobs, which potentially left scars on their previous wages, in case of non-consecutive job spells.

In line with a labour market characterised by a high share of temporary contracts and a low conversion into permanent contracts, which are heavily protected, it is not surprising that the termination of a fixed-term contract is the main reason for job loss, representing 38% of the individuals.[12] After that, redundancy (i.e., the employer claimed the job no longer existed or the individual was not suited for the job anymore) and mutual agreement are the most representative ones with 23% and 17%, respectively.

The distribution of unemployment inflows along time is approximately uniform with the exception of 2007, when the inflow decreased, which might be explained by the enforcement of the new rules. However, during the following years, with the onset of the financial crisis, the inflow quickly caught up the previous level, reaching the peak of inflows in 2011, 0.18 (note that 2012 is not fully contained in this dataset). The high average age (39 years old) should be related with the low attachment to the labour market of older workers.[13] Finally, the spatial distribution of unemployed individuals along the different regions of Portugal, matches roughly the one of employed individuals.[14]

[12]The share of temporary contracts in Portugal is 22% of dependent employment, double of the OECD average, and the yearly conversion rate is less than 15%, which is one of the lowest in European Union, as pointed out by (Boeri et al., 2011).

[13]Even eliminating the individuals whose age is above the minimum required for retirement the average age of the sample is still high.

[14]If any difference, the North has a higher proportion of unemployed relative to employed and Lisbon

In the last panel of Table 1.6 I present descriptive statistics on the last observed status for each individual in the sample. Roughly half of the individuals in the sample are observed in a new job before March 2012, whereas 44% of the spells of joblessness are censored. Note that such a high rate of censoring is both explained by the high proportion of long-term unemployment in Portugal (around 51% during the period in analysis), as well as by the onset of the crisis that increased the flows into unemployment in the last years of the sample. As 95% of the individuals that are observed to transit into employment find a dependent employment job, I restrict the analysis, of hazard rates to employment, to this type of work only.

Table 1.7 presents similar descriptive statistics for UA beneficiaries. As only the first two panels are substantially different from the full sample, the remaining of the descriptive statistics for UA beneficiaries are presented in the Appendix of this chapter.

Table 1.7: Descriptive statistics of subsequent UA beneficiaries

	Mean	Std	Min	Max
Benefit Duration (UI)	21.85	8.89	9	38
Benefit Duration (UA)	9.20	4.87	1	19
Joblessness Duration	39.48	18.27	10	87
UI benefits (€/month)	14.91	5.36	4.18	41.92
UA benefits (€/month)	12.38	1.62	4.18	13.97
Reference wage (€/month)	614.6	369.9	132.3	5 943.3
Re-employed (any job)	1 702 (1 584 as employees)			
Retired, disabled, dead	166			
Censored spells	1755			
Sample	3 623			

Note: Additional descriptive statistics are provided in the Appendix of this chapter.

In comparison with the previous table, the individuals that benefit from subsequent UA have, on average, much longer subsidized (UI) spells and still experience some

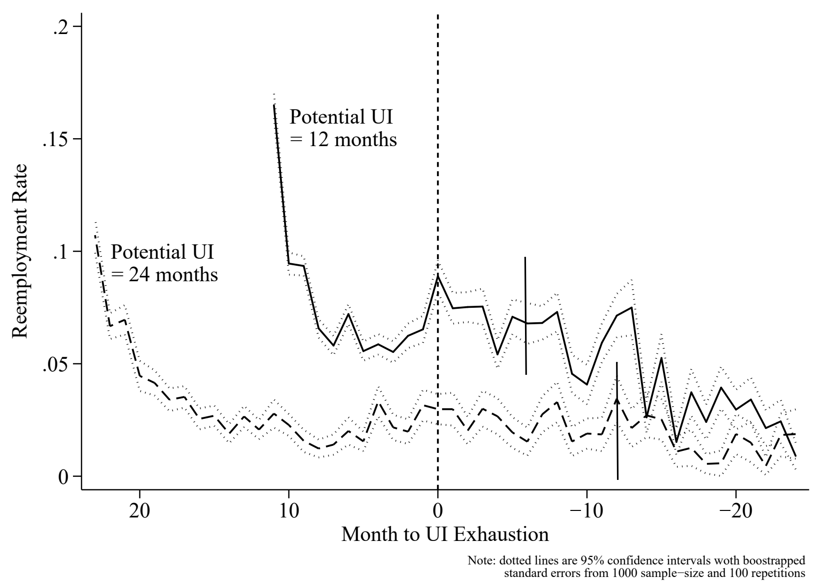
has a lower proportion, which might be in line with the fact that in the North a third of the unemployed had their jobs extinct, whereas 43% of the unemployed individuals in Lisbon lost their job due to the end of a fixed-term contract.

period with no unemployment income support. Note also that not only the average UA benefit amount if lower, but also that the previous wage is roughly 100€less, which combined leads to a lower average amount of UI benefits for these individuals, relatively to the full sample.

1.5 Descriptive Evidence on the Impact of the Potential Duration of Unemployment Benefits

Preceding the regression analysis, it is helpful to firstly analyse the descriptive evidence of outflows into employment along the spell of unemployment. In the following figures I plot the average behaviour of two groups of beneficiaries, the ones entitled to 12 and 24 months of unemployment insurance, as these are two of the most common entitlements in the data.[15]

Figure 1.1: Re-employment rates, by Potential Duration of UI



[15]The behaviour for the other potential duration entitlements follows the same trend and is available upon request.

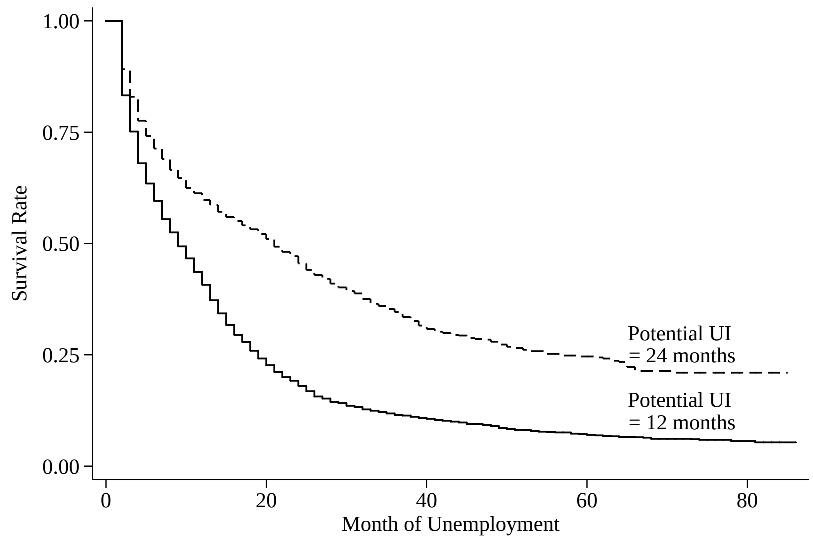
In figure 1.1, until the moment of UI exhaustion, two patterns stand out. The first one is that, irrespective of the potential duration, re-employment rates fall sharply after the first months of unemployment. The second one is that the spike at benefit exhaustion is larger for beneficiaries entitled to the shorter potential duration. [16]

In both series there is also a (smaller) spike at the point of subsequent UA exhaustion – 6 months after UI exhaustion for the series of 12 months entitlement, and 12 months after UI exhaustion for the series of 24 months entitlement, marked with short vertical lines. The spike is more evident on the shorter series as a larger proportion of individuals had also benefited from subsequent UA (19% vs 16%). After the point of subsequent UA exhaustion, a decay is observed in both series. However, the re-employment rate for the beneficiaries entitled to 24 months does not decrease as much as for the counterpart series, due to the statistically significant larger proportion of individuals that have access to RSI (*Rendimento Social de Inserção*, or *minimum guaranteed income*, an income support provided to Portuguese families with severe economic distress), after the exhaustion of subsequent UA (25% vs 20%).

Even though the previous figure shows a clear difference in the transition rate into employment for the different groups of individuals, the transition rate statistic does not incorporate the fact that many individuals are still jobless at the end of the dataset. The censoring may either come from the long unemployment spells or by the late entry into the dataset. In Figure 1.2, I plot the empirical Kaplan-Meier survival functions for the same two groups of potential duration entitlement, as before (12 and 24 months).

[16]The fact that there is a spike in both series is in line with the optimization strategy of the beneficiaries that rationally delay the starting date of re-employment in order to benefit from subsidized leisure (Boone and van Ours, 2012).

Figure 1.2: Kaplan-Meier Survival Rates, by Potential Duration of UI



The empirical survival function, at a certain month t is given by the following:

$$\hat{S}(t) = \prod_{j|t_j<t} \left(1 - \frac{d_j}{n_j}\right) \tag{1.1}$$

where d_j stands for the number of exits to employment on month t_j, n_j indicates the number of unemployed "at risk" of becoming re-employed on that month, and j indexes the month number for individual i. In other words, this survival rate can be interpreted as the probability that an individual did not find a job in each month t_j, included in the period until t.

The pattern illustrated in the graph shows that, even after 80 months of unemployment, almost 25% of the individuals entitled to two years of UI potential duration did not find a job, whereas only 5% of those entitled to one year of potential duration were in that situation.

This graphical analysis shows evidence that the potential duration of unemployment benefits (both UI and subsequent UA) is a strong determinant of the shape of the joblessness duration distribution.

1.6 Estimation and Identification

1.6.1 Discrete Proportional Hazard Model

To estimate the impact of potential duration of unemployment benefits on re-employment rates I use an hazard model, accounting for the fact that many of the unemployment spells in the data are censored. In particular, I choose the discrete time representation of the proportional hazard model (Cox, 1972) – a complementary log-log (c-log-log) hazard model. Its discrete feature is particularly convenient in the case of this study as the data is grouped in months (Jenkins, 2005).[17] Accordingly, it is assumed that only one event occurs at each month (from t_{j-1} to t_j). To account for that, the c-log-log (extreme value) function is asymmetrical in the sense that it approaches zero slowly and converges to one quickly, opposed to the balanced distributions of the logit and probit alternatives. In other words, the probability of an event to occur at a specific period in time is either very large or negligible.

In this model, the probability of surviving until t, i.e., staying unemployed until t, is given by the following:

$$S(t|\boldsymbol{x'_{it}\beta}) = \exp\left[-\exp(\boldsymbol{x'_{it}\beta})\int_0^t h_0(u)du\right] \tag{1.2}$$

where t denotes the time (in months) elapsed since the beginning of the spell of unemployment; $h_0(t)$ is defined as the baseline hazard function capturing the duration dependence; $\boldsymbol{x'_{it}}$ is the vector of explanatory variables (both constant and time varying) for individual i in time t; and $\boldsymbol{\beta}$ is the corresponding regression coefficient vector.

[17]The data is actually provided in days, but given the long spells of unemployment (in the limit, more than 30.000 days of duration), I convert them to a monthly basis for purposes of estimation.

The discrete conditional probability that an individual i gets out of unemployment in month t, given that she was unemployed up to that point, is written as:

$$\begin{aligned} h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta}) &= \frac{S(t-1|\boldsymbol{x}'_{it}\boldsymbol{\beta}) - S(t|\boldsymbol{x}'_{it}\boldsymbol{\beta})}{S(t-1|\boldsymbol{x}'_{it}\boldsymbol{\beta})} = 1 - \frac{S(t|\boldsymbol{x}'_{it}\boldsymbol{\beta})}{S(t-1|\boldsymbol{x}'_{it}\boldsymbol{\beta})} \\ &= 1 - \exp\left[-\exp(\boldsymbol{x}'_{it}\boldsymbol{\beta})\left(\underbrace{\int_0^{t-1} h_0(u)du}_{H_{t-1}} - \underbrace{\int_0^{t} h_0(u)du}_{H_t}\right)\right] \end{aligned} \tag{1.3}$$

The complementary log-log transformation, is applied into two steps (logs), firstly:

$$\begin{aligned} \log[1 - h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta})]) &= \log[1 - (1 - \exp[-\exp(\boldsymbol{x}'_{it}\boldsymbol{\beta})(H_{t-1} - H_t)])]) \\ &= \exp(\boldsymbol{x}'_{it}\boldsymbol{\beta})(H_{t-1} - H_t) \end{aligned} \tag{1.4}$$

and secondly:

$$\log(-\log[1 - h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta})]) = \log(-\exp(\boldsymbol{x}'_{it}\boldsymbol{\beta})(H_{t-1} - H_t) = \boldsymbol{x}'_{it}\boldsymbol{\beta} + \log(H_t - H_{t-1}) \tag{1.5}$$

Note that instead of $H_{t-1} - H_t$ there is now $H_t - H_{t-1}$.

Rewriting the last expression:

$$h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta}) = 1 - \exp[-\exp(\boldsymbol{x}'_{it}\boldsymbol{\beta} + \gamma_t)] \tag{1.6}$$

where γ_t stands for $\log(H_t - H_{t-1})$, which summarizes the duration dependence pattern. The likelihood for estimation is based on the general likelihood for right-censored discrete data (Allison, 1982):

$$\mathcal{L} = \prod_{i=1}^{N} [\Pr(T_i = t)]^{\delta_i} [\Pr(T_i > t)]^{1-\delta_i} \tag{1.7}$$

where δ_i is an individual censor indicator, which takes the value 0 if the duration spell of individual i is complete, and 1 otherwise, and T_i is the month in which individual i finds a job. By definition, the hazard rate defined in 1.6 represents $\Pr(T_i = t | T_i \geq t, \boldsymbol{x_{it}})$. Using the properties of conditional probabilities, the likelihood can be written as a function of the hazard rate:

$$\mathcal{L} = \prod_{i=1}^{N} \left[h(t|\boldsymbol{x'_{it}\beta}) \prod_{j=1}^{t_{i-1}} (1 - h(t|\boldsymbol{x'_{it}\beta})) \right]^{\delta_i} \left[\prod_{j=1}^{t_i} (1 - h(t|\boldsymbol{x'_{it}\beta})) \right]^{1-\delta_i} \tag{1.8}$$

$$= \prod_{i=1}^{N} \left(\frac{h(t|\boldsymbol{x'_{it}\beta})}{1 - h(t|\boldsymbol{x'_{it}\beta})} \right)^{\delta_i} \prod_{j=1}^{t_i} (1 - h(t|\boldsymbol{x'_{it}\beta})) \tag{1.9}$$

which is estimated by maximum likelihood.

1.6.2 Unobserved Heterogeneity

Even controlling for age, gender, reason of unemployment among other covariates, the econometric model may be misspecified if some explanatory variables are inappropriately omitted. Following Heckman and Singer (1984), I use a semi-parametric estimator with respect to the distribution of the individual-specific effects. Namely, to complete the specification of the likelihood function presented above, I include a binomial-mixture distribution function which approximates the unknown probability distribution for two support points, whose location and probability mass associated are estimated according to the data.

With this approach I consider that a sub-population of the unemployed will take much longer to get a job (*slower exiters*[18]), which ends up being more realistic than assuming that all individuals, with similar observed characteristics, exit into re-employment at the same rate.[19] This generates two different intercepts in the equation to estimate,

[18] Adopting the terminology by (Jenkins, 2005)

[19] This type of model is usually denoted as *split-population model* (Schmidt and Witte (1989). It is also associated to the concept of "defective risks" - see Addison and Portugal (2003) for a detailed explanation and Portugal (2008) for another empirical application, with more recent data.

one for each group. The group with the larger intercept leaves unemployment at a faster rate than the other group (*faster exiters*). The hazard rate becomes:

$$h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta}) = \begin{cases} 1 - \exp[-\exp(\boldsymbol{x}'_{it}\boldsymbol{\beta} + \gamma_t)], & \text{if } i \text{ is a } \textit{slower exiter} \\ 1 - \exp[-\exp(m_z + \boldsymbol{x}'_{it}\boldsymbol{\beta} + \gamma_t)], & \text{if } i \text{ is a } \textit{faster exiter} \end{cases} \tag{1.10}$$

with $m_z > 0$. In this context likelihood function becomes:

$$\mathcal{L} = \pi \times \mathcal{L}_{slowerexiter} + (1 - \pi) \times \mathcal{L}_{fasterexiter} \tag{1.11}$$

where $\pi \in (0, 1)$ is the proportion of unemployed in group of *slower exiters*, and $\mathcal{L}_{slowerexiter}$ is the expression in (1.8) evaluated at the hazard rate for *slower exiters*, and $\mathcal{L}_{fasterexiter}$ is the analogous for *faster exiters*.

1.6.3 Identification

When estimating the models described above, one should pay attention to common factors that might determine both the potential duration of unemployment benefits and the re-employment rates. The period of time used in this chapter allows the exploitation of three different sources of variation in the potential duration of both UI and UA, which helps to identify the causal impact of an additional month of unemployment benefits on the re-employment probability.

First, there is variation across ages. For example, before 2007, an individual that was 45 years old and had 20 years of working experience, was entitled, to 14 additional months of potential duration of UI than an individual with the same characteristics except 1 year younger.

Second, there is variation across working experience. For example, after 2007, an individual that was 40 years old, but had 60 months or less months of contributions to Social Security, was entitled to 18 months of UI potential duration, whereas in the same period, for an individual with the same age, one additional month of contributions would entitle her to 24 months of UI potential duration.

Third, there is variation across time. For example, an individual that was 40 years old, with less than 5 years of working experience, and registered in the employment center in December 2006, was entitled to 24 months of UI potential duration, whereas an identical individual registered in the employment center in January 2007, was entitled to only 18 months of UI potential duration. Note that unlike the previous sources of variation, starting to receive UI benefits after the reform might yield less, the same, or additional months of potential duration, when compared to pre-reform rules.

The three types of variation also apply in the same directions to the potential duration of subsequent UA, with the caveat that the entitlement in this case is defined as half of the corresponding UI potential duration. Additionally, there is another source of time variation, as the individuals that exhausted the potential duration of UA in 2009, or first half of 2010, were also entitled to 6 additional months.

Using these sources of variation, seven different specifications are estimated.[20] Specification 1.12 includes the potential duration of the unemployment insurance benefits at the start of the unemployment spell, which does not vary over time.

$$h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \text{UIPotentialDuration}_i + \boldsymbol{x}'_{it}\boldsymbol{\beta} + \ln(t))] \tag{1.12}$$

Considering the different sources of variation mentioned before, the initial potential duration of unemployment benefits, together with the necessary controls, should identify the impact of the institutional setting in the probability of re-employment. Then, in

[20]Note that all specifications are defined here without unobserved heterogeneity, but that will also be taken into account in all of them in the Results section.

1.13, I take advantage of the reform in 2007 and study the impact of the change in the potential duration of the unemployment insurance benefits on the hazard rate to employment.

$$h(t|\boldsymbol{x}'_{it}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \Delta_{2007}\text{UIPotential Duration}_i + \boldsymbol{x}'_{it}\boldsymbol{\beta} + \ln(t))] \quad (1.13)$$

If the joblessness duration of the individuals that were affected by the reform moves in the same direction as the change in the potential duration of the unemployment benefits, then this must be a good proxy for the potential duration of unemployment benefits. Table 1.8 summarizes the average joblessness duration by each age group and by whether the individuals were entitled to more, less, or the same potential duration of unemployment insurance benefits, in comparison to what they would receive had they started receiving the benefits before January 2007.

Table 1.8: Average joblessness duration (in months), by age group and difference between the pre- and post-reform rules of potential duration of unemployment insurance

Age group	Negative change	No change in rules	Positive change
<30	9.07	13.58	13.97
[30,40[	9.81	12.84	16.11
[40,45[	12.08	15.28	18.43
$\geq$ 45	16.04	22.25	NA

Note: the reform of 2007 did not extend the potential duration of unemployment benefits for individuals with age 45 or more.

As Table 1.8 shows, the individuals that were entitled to a longer (shorter) potential duration of unemployment insurance benefits after 2007, had also a longer (shorter) spell of joblessness. The advantage of using this proxy is that, unlike the potential duration of unemployment insurance benefits, the effect of age and working experience is less present in the values of this variable. For example, an individual might have been entitled to one additional month of potential duration and either be 30 years old or 50 years old; or an individual might have been entitled to 6 less months of potential duration and either have 20 or 55 months of Social Security contributions.

Specifications 1.14 and 1.15 are similar to 1.12 and 1.13, but also include the potential duration, and the change, of unemployment assistance benefits. In these two models, the sample was restricted to all individuals that have benefited from both unemployment insurance and unemployment assistance, to evaluate the additional impact of the latter.

$$h(t|\boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \text{UIPotentialDur}_i + \alpha_2 \times \text{UAPotentialDur}_i + \boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta} + \ln(t))] \tag{1.14}$$

$$h(t|\boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \Delta_{2007}\text{UI Potential Dur}_i + \alpha_2 \times \Delta_{2007}\text{UAPotentialDur}_i + \boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta} + \ln(t))] \tag{1.15}$$

Finally, specifications 1.16, 1.17, and 1.18 include the time-varying effects of unemployment insurance and unemployment assistance on the hazard rate to employment. Specification 1.16 includes two mutually exclusive dummies, UI and UA, which allow to compare between the two states of insurance with the base group state being the absence of both benefits.

$$h(t|\boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \mathbb{1}\text{UIRecipient}_{it} + \alpha_2 \times \mathbb{1}\text{UARecipient}_{it} + \boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta} + \ln(t))] \tag{1.16}$$

Specification 1.17 includes the potential duration of unemployment benefits (UI and UA) that was attributed to the individual at the beginning of the benefit recipiency, for as long as the individual receives the unemployment benefits, thus combining both recipiency and potential duration effects.

$$h(t|\boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta}) = 1 - \exp[-\exp(\alpha_1 \times \text{UIPotentialDur}_{it} + \alpha_2 \times \text{UAPotentialDur}_{it} + \boldsymbol{x}'_{\boldsymbol{it}}\boldsymbol{\beta} + \ln(t))] \tag{1.17}$$

Finally, specification 1.18 includes the time that is left to spend from the potential duration of unemployment benefits that was attributed at the beginning of the recipiency period, i.e., the number of months until UI/UA exhaustion.

$$h(t|\boldsymbol{x'_{it}\beta}) = 1 - \exp[-\exp(\alpha_1 \times \text{TimetoExhaustUI}_{it} + \alpha_2 \times \text{TimetoExhaustUA}_{it} + \boldsymbol{x'_{it}\beta} + \ln(t))] \qquad (1.18)$$

The interpretation of the associated coefficients allows to evaluate the impact of the limited benefit duration feature in a non-stationary environment of job search (Van den Berg, 1990). Using 12 months of potential UI duration as an example, the following table illustrates the definition of *time to exhaustion*:[21]

Table 1.9: Definition of time to exhaustion variables

Month	1	2	3	4	...	11	12	13	14	...	17	18	19	20	21
UI	12	11	10	9	...	2	1	0	0	...	0	0	0	0	0
Subsequent UA	0	0	0	0	...	0	0	6	5	...	2	1	0	0	0

For each additional calendar month, the UI entitlement period decreases linearly until the point of exhaustion when *time to exhaust UI* first assumes the value zero. In that same month, in case the individual is entitled to subsequent UA, *time to exhaust subsequent UA* assumes the value of the potential duration and then decreases linearly until exhaustion. Therefore, in the case the individual does not transit to subsequent UA, the second variable is always zero. As both benefits get exhausted, by month 19, both variables assume the value zero from that point onwards.

All specifications account for a logarithmic duration dependence function ($\ln(t)$), assumed to be common to all individuals. Different specifications for the baseline hazard function, such as dummies and splines, were also tried and the results were not perturbed. When unobserved heterogeneity is taken into account, the baseline hazard changes discontinuously for the *faster exiters*.

[21]For the sake of the example, I assume that the individual age has not changed significantly such that the potential subsequent UA duration is more than half of that for the UI

1.6.4 Control Variables

To control for the age effects, and also for the age discontinuities in the rules for the potential duration of unemployment benefits, I use linear splines with cuttoffs at ages 30, 40 and 45. Note that the age that is used to define to potential duration of unemployment benefits is the age at the beginning of the unemployment episode, therefore this is the one I assume for the splines, which are not time-varying.

The replacement ratio is also defined using the first UI amount in the spell. The reasons behind this choice are twofold: first, the UI benefit amount was constant in Portugal, until the most recent law of 2012; and second, the amount earned in the UA is linked to the one received in UI.[22] Therefore, the amount of UA is excluded from these specifications. Note that I use the logarithm of the replacement ratio, so that the coefficient can be interpreted as an elasticity. To be coherent with the concept of replacement ratio used by the Social Security, the denominator of the replacement ratio is the *reference wage*, instead of being the last wage observed in the data before the individual become unemployed.[23] This is defined as the average salary received during the first 12 months of the 14 months preceding the unemployment rate. This option brings two advantages: first, unlike the last wage before unemployment, I do observe the *reference wage* for all the individuals; second, using this measure avoids large penalizations to individuals that did not receive wages, or received large wage cuts, in the last months of employment (as it is typical in plant closures).

Additional demographic controls are included in all specifications. A gender dummy

[22]See the Appendix of this chapter for the rules defining the UA monetary amount.

[23]See the Appendix of this chapter for a precise definition of the concept of *reference wage*. In order to measure the balance between the benefit amount and the previous wage I run a different regression considering the individuals for which I had access to the actual previous wage for the purpose of robustness. The results on the other variables were roughly the same and the elasticity on the replacement ratio has slightly decreased, showing evidence for the fact that with the current rule, the individuals that have not worked during the whole reference period end up receiving less than a counterpart with the same wage but working in all of these 12 months.

(equal to one if the individual is male), a foreign dummy (equal to one if the individual does not have Portuguese nationality), the quarterly unemployment rate, which is time-varying, five regional dummies (four regions in mainland – Alentejo, Algarve, Centro and Norte – and Azores and Madeira, which were included together due to the smaller number of observations in these two regions), four reasons of unemployment dummies (end of a fixed-term contract, redundancy, mutual agreement, and plant closure or mass lay-off), tenure in the last job (measured in years), and the number of previous jobs with contributions to the Social Security.

1.7 Empirical Results

1.7.1 Transitions from UI or Subsequent UA to employment

This subsection presents the estimates of the impact of unemployment benefits and their potential duration on the hazard rate to (dependent) employment. All the following tables account for unobserved heterogeneity with two mass points.[24] The Appendix A4 of this chapter presents the estimates without unobserved heterogeneity in the same order. The estimates in Table 1.10 refer to specifications 1.12 and 1.13.

The first column of Table 1.10 shows that, for each additional month of potential duration of unemployment insurance benefits, the hazard rate is expected to decrease by approximately 9.1%. This estimate is close to the upper bound of the range that was previously estimated in the literature – between 0.1% and 8% (for the studies that reported a negative impact), summarised in the second panel of Table 1.

The large magnitude of the estimate of interest in column (1) of Table 1.10 might be explained by two facts. First, like Katz and Meyer (1990a), this study uses a

[24]The specifications that account for unobserved heterogeneity were estimated using the Stata user-written code `hshaz` by Stephen Jenkins.

Table 1.10: Estimated c-log-log regression, with unobserved heterogeneity – the impact of potential duration of unemployment benefits on the hazard rate to employment

Variables	(1)		(2)	
	Coeff.	(SE)	Coeff.	(SE)
UI Potential Duration (in months)	-0.0955	(0.0035)		
Δ_{2007} **UI Potential Duration (in months)**			-0.1203	(0.0035)
Ln(Joblessness Duration) †	0.1218	(0.0163)	0.1112	(0.0168)
Age spline (in years)				
min(age,30)	0.0128	(0.0052)	-0.0244	(0.0049)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	0.0080	(0.0051)	-0.0417	(0.0048)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	0.1816	(0.0133)	-0.0372	(0.0113)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1123	(0.0053)	-0.1438	(0.0055)
Ln(Replacement Ratio) ‡	-0.4331	(0.0659)	-0.4299	(0.0654)
Past Career Information:				
Tenure (in years)	-0.0150	(0.003)	-0.0187	(0.003)
Number of jobs	0.1364	(0.0071)	0.1363	(0.007)
$\mathbb{1}$ Industry	Yes		Yes	
$\mathbb{1}$ Reason of Unemployment				
end of fixed-term contract	0.0745	(0.0398)	0.0571	(0.0397)
redundancy	-0.1791	(0.0430)	-0.1829	(0.0429)
mutual agreement	-0.3245	(0.0495)	-0.3341	(0.0494)
plant closure or mass layoff	0.0986	(0.0534)	0.1000	(0.0531)
Other Controls:				
$\mathbb{1}$ Gender (Male=1)	0.2094	(0.0247)	0.2127	(0.0246)
$\mathbb{1}$ Nationality (Foreign=1)	-0.0674	(0.0342)	-0.0701	(0.034)
Quarterly Unemployment Rate †	-0.2847	(0.0065)	-0.2853	(0.0065)
$\mathbb{1}$ Region of Residence				
North	-0.1018	(0.0309)	-0.1037	(0.0307)
Center	0.3092	(0.036)	0.3114	(0.0358)
Alentejo	0.5511	(0.0506)	0.5461	(0.0508)
Algarve	0.4573	(0.0584)	0.4458	(0.0583)
Azores and Madeira	-0.2353	(0.0588)	-0.2348	(0.0585)
Unobserved heterogeneity:				
Δ intercept (*faster exiters*)	1.4603	(0.0774)	1.4239	(0.0774)
Intercept (*slower exiters*)	-0.6156	(0.2054)	-0.8256	(0.2059)
Proportion of *faster exiters*	0.7915	(0.0299)	0.7899	(0.0339)
Proportion of *slower exiters*	0.2085	(0.0299)	0.2101	(0.0339)
Log-likelihood	-43462.6190		-43527.5360	
Nr of observations	406 626		406 626	

† Time-varying covariates ‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).
Note: column (2) only includes observations of individuals that also received unemployment assistance.

period/country when the unemployment rate was particularly high.[25] Second, like the studies in the literature that reported the largest effects, this study also accounts for unobserved heterogeneity, which is proven to be non-negligible in all the specifications.[26]

In column (2) of Table 1.10, the total unemployment benefits potential duration is replaced by the change in the potential duration of unemployment benefits that occurred due to the reform in 2007. The main estimate of interest shows that each additional month of potential duration for the affected individuals is expected to decrease the hazard rate to employment by 12%.

The estimates in Table 1.11 include all the individuals that were also observed to receive unemployment assistance. These specifications also account for the potential duration of unemployment assistance, in a separate coefficient. The results reveal that the impact of unemployment insurance in re-employment is larger than that for unemployment assistance but the two coefficients are not statistically different from each other, in both specifications.

The estimates that use the change in the potential duration of unemployment benefits due to the policy reform (in columns (2) and (4)) are slightly larger than the ones that use directly the potential duration of unemployment benefits (in columns (1) and (3)). This result highlights that one should be careful when interpreting the main coefficients in columns (1) and (3), given that the potential duration is defined by age and experience, which also influence the hazard into re-employment.

The estimates in columns (5), (6), and (7), in Table 1.12, take into account the time-varying impact of unemployment benefits. The first specification shows that the recipiency of unemployment insurance is expected to decrease the hazard rate to employment by approximately 73%.

[25]The unemployment rate in Portugal during the period in analysis increased from 7.5% in the first quarter of 2005 to 14.8% in the first quarter of 2012.

[26]Note that the estimates for the main coefficients of interest tend to be of smaller magnitude when unobserved heterogeneity is not taken into account. See Appendix A4 of this chapter.

Table 1.11: Estimated c-log-log regression, with unobserved heterogeneity – the impact of changes in the rules of UB potential duration, on the hazard rate to employment

Variables	**(3)** Coeff.	(SE)	**(4)** Coeff.	(SE)
UI Potential Duration (in months)	-0.1288	(0.0172)		
UA Potential Duration (in months)	-0.1270	(0.0332)		
Δ_{2007} **UI Potential Duration (in months)**			-0.1115	(0.0455)
Δ_{2007} **UA Potential Duration (in months)**			-0.1373	(0.0910)
Ln(Joblessness Duration) †	3.6414	(0.1002)	3.5324	(0.0992)
Age spline (in years)				
min(age,30)	-0.0812	(0.0151)	-0.1476	(0.0136)
(age-30) $\times \mathbb{1}$ (age $\in$ [30,40[) + 10 $\times \mathbb{1}$ (age $\geq$ 40)	-0.0402	(0.0171)	-0.1558	(0.0143)
(age-40) $\times \mathbb{1}$ (age $\in$ [40,45[) + 5 $\times \mathbb{1}$ (age $\geq$ 45)	0.0418	(0.0503)	-0.4146	(0.0451)
(age-45) $\times \mathbb{1}$ (age $\geq$ 45)	-0.1036	(0.0247)	-0.1402	(0.0263)
Ln(Replacement Ratio) ‡	-1.6398	(0.2974)	-1.6471	(0.2437)
Past Career Information:				
Tenure (in years)	-0.0146	(0.0097)	-0.0230	(0.0096)
Number of jobs	0.0989	(0.0351)	0.1083	(0.0355)
$\mathbb{1}$ Industry	Yes		Yes	
$\mathbb{1}$ Reason of Unemployment				
end of fixed-term contract	0.0350	(0.1186)	0.0799	(0.1177)
redundancy	-0.1137	(0.1273)	-0.0632	(0.1225)
mutual agreement	-0.0984	(0.1399)	-0.0617	(0.1395)
plant closure or mass layoff	0.3717	(0.1575)	0.4014	(0.1553)
Other Controls:				
$\mathbb{1}$ Gender (Male=1)	-0.0767	(0.0802)	-0.0888	(0.0754)
$\mathbb{1}$ Nationality (Foreign=1)	-0.0208	(0.1090)	-0.0142	(0.1137)
Quarterly Unemployment Rate †	-0.2608	(0.0186)	-0.2634	(0.0188)
$\mathbb{1}$ Region of Residence				
North	-0.1407	(0.0934)	-0.0928	(0.0939)
Center	-0.1800	(0.1280)	-0.0301	(0.1284)
Alentejo	0.6335	(0.2002)	0.6917	(0.1757)
Algarve	0.0301	(0.2473)	0.1441	(0.2421)
Azores and Madeira	-0.4199	(0.1518)	-0.3472	(0.1560)
Unobserved heterogeneity:				
Δ intercept (*faster exiters*)	4.4643	(0.1306)	4.3485	(0.1313)
Intercept (*slower exiters*)	-11.3045	(0.5861)	-11.8064	(0.6040)
Proportion of *faster exiters*	0.6594	(0.0133)	0.6559	(0.0138)
Proportion of *slower exiters*	0.3406	(0.0133)	0.3441	(0.0138)
Log-likelihood	-7135.7812		-7188.5358	
Nr of observations	141 183		141 183	

† Time-varying covariates ‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).
Note: column (2) only includes observations of individuals that also received unemployment assistance.

Once again, the estimates reported in this study lie above the upper bound of estimates reported for UI recipiency in the first panel of Table 1.[27] This is not surprising as the base group in this study is the group of individuals that have exhausted the unemployment benefits, and therefore it does not include individuals that are not eligible to UI (as these are not identified in the records of administrative data). The group of individuals that exhausted the benefits have, by construction, longer joblessness spells, and therefore are expected to have a lower hazard rate into employment. Even though the estimate for unemployment assistance is larger, once again that is not statistically different than the estimate for unemployment insurance.

The main differences between column (6) and column (2) are twofold: the latter is estimated for the entire sample, and the values of the initial potential duration of unemployment benefits are zero whenever the individual is not receiving the unemployment benefit. Therefore, the estimates in column (6) should lie between that of column (2) and column (7). The coefficient estimates reveal a larger effect of the initial potential duration of unemployment assistance, that is statistically different from that of unemployment insurance. Such difference is exacerbated in the coefficient estimates of column (7), which contains the time to exhaust unemployment benefits as main variables of interest. The estimated effect for each month further away from exhaustion is a decrease of approximately 2.4% for unemployment insurance, and more than triplicates for unemployment assistance. These estimates are in line with the results from Addison and Portugal (2008b), which are summarized in the last panel of Table 1.

Summarising the coefficient estimates on the logarithmic baseline hazard, three aspects should be highlighted. For the full sample (columns (1), (2) and (5) to (7)), when unobserved heterogeneity is not taken into account, the negative estimated coefficient

[27] Note however that when the dummy for UA receivers is omitted from column (5), the coefficient estimate for the dummy for UI receivers becomes much closer to the ones reported in the literature for UI recipiency.

Table 1.12: Estimated c-log-log regression, with unobserved heterogeneity – the time-varying impact of unemployment benefits on the hazard rate to employment

Variables	**(5)**		**(6)**		**(7)**	
	Coeff.	(SE)	Coeff.	(SE)	Coeff.	(SE)
$\mathbb{1}$ **Recipiency** †						
Unemployment Insurance	-1.3179	(0.0437)				
Unemployment Assistance	-1.2390	(0.0524)				
UB Initial Potential Duration †						
Unemployment Insurance			-0.0549	(0.0016)		
Unemployment Assistance			-0.1119	(0.0049)		
Time to Exhaust UB (in months) †						
Unemployment Insurance					-0.0240	(0.0025)
Unemployment Assistance					-0.0925	(0.0074)
Ln(Joblessness Duration) †	0.0323	(0.0157)	0.0480	(0.0161)	-0.0462	(0.0223)
Age spline (in years)						
min(age,30)	-0.0163	(0.0056)	0.0003	(0.0054)	-0.0183	(0.0048)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	-0.0364	(0.0054)	-0.0113	(0.0053)	-0.0287	(0.0048)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	0.0061	(0.0126)	0.1109	(0.0128)	0.0268	(0.0117)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1567	(0.006)	-0.1485	(0.0058)	-0.1392	(0.0054)
Ln(Replacement Ratio) ‡	-0.3979	(0.0722)	-0.4061	(0.0701)	-0.4110	(0.0638)
Past Career Information:						
Tenure (in years)	-0.0254	(0.0033)	-0.0208	(0.0032)	-0.0228	(0.003)
Number of jobs	0.1610	(0.0086)	0.1540	(0.0077)	0.1359	(0.007)
$\mathbb{1}$ Industry	Yes		Yes		Yes	
$\mathbb{1}$ Reason of Unemployment						
end of fixed-term contract	0.1382	(0.0446)	0.1256	(0.0433)	0.1252	(0.0382)
redundancy	-0.1496	(0.0486)	-0.1546	(0.0468)	-0.1179	(0.0413)
mutual agreement	-0.3498	(0.0551)	-0.3341	(0.0537)	-0.2993	(0.0479)
plant closure or mass layoff	0.1559	(0.0598)	0.1668	(0.0578)	0.1387	(0.0517)
Other Controls:						
$\mathbb{1}$ Gender (Male=1)	0.2529	(0.0277)	0.2350	(0.0267)	0.2255	(0.0239)
$\mathbb{1}$ Nationality (Foreign=1)	-0.0139	(0.0382)	-0.0352	(0.0369)	-0.0150	(0.0329)
Quarterly Unemployment Rate †	-0.3578	(0.0072)	-0.3393	(0.007)	-0.2974	(0.0064)
$\mathbb{1}$ Region of Residence						
North	-0.1132	(0.0347)	-0.1031	(0.0335)	-0.1072	(0.0298)
Center	0.3559	(0.0394)	0.3415	(0.0385)	0.3042	(0.0348)
Alentejo	0.5799	(0.0544)	0.5809	(0.0529)	0.5202	(0.0497)
Algarve	0.5045	(0.0632)	0.4957	(0.0617)	0.4587	(0.0569)
Azores and Madeira	-0.2287	(0.0686)	-0.2279	(0.0639)	-0.2015	(0.0569)
Unobserved heterogeneity:						
Δ intercept (*faster exiters*)	2.1958	(0.0531)	1.8974	(0.0554)	1.3876	(0.1102)
Intercept (*slower exiters*)	0.3096	(0.208)	-0.6002	(0.2035)	-0.4711	(0.2232)
Proportion of *faster exiters*	0.7339	(0.0115)	0.7538	(0.0157)	0.8316	(0.0357)
Proportion of *slower exiters*	0.2661	(0.0115)	0.2462	(0.0157)	0.1684	(0.0357)
Log-likelihood	-43331.3170		-43236.9730		-43748.0230	
Nr of observations	406 626		406 626		406 626	

† Time-varying covariates

‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).

expresses the negative duration dependence that is usually associated to longer joblessness spells, that is, the longer the individual stays without a job, the less likely the individual is to transit to employment.[28] In almost all specifications, this coefficient is revealed to be positive when unobserved heterogeneity is taken into account. Note that this is a mechanic effect that is compensated by the change in the intercept. Finally note that the baseline hazard is always positive for the restricted sample of individuals that are observed to receive UA as these will have, by construction, longer joblessness spells, and therefore exit to employment at a later stage.

To provide a clearer interpretation of the age effects, I plot the marginal effects in the hazard rate, by age, for each of the five specifications that are estimated for the full sample, taking unobserved heterogeneity into account.

Figure 1.3: Age Effects, by specification

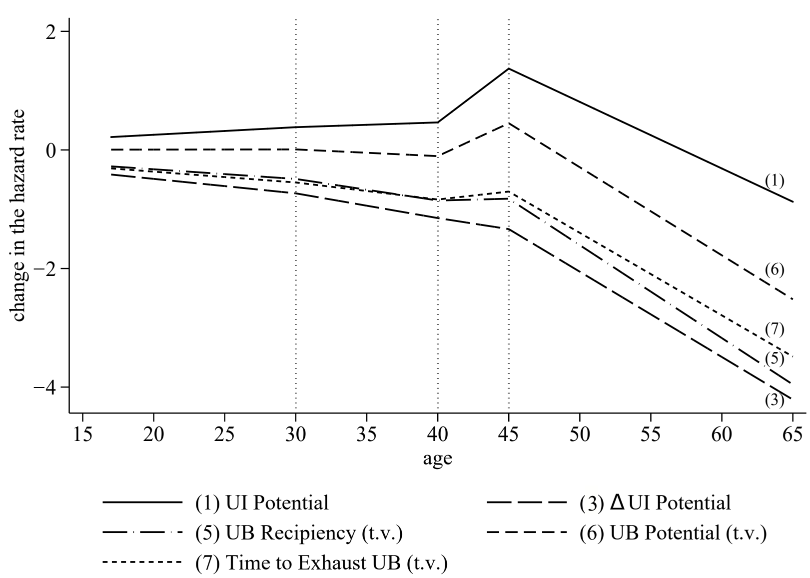


Figure 1.3 illustrates three main patterns: the age effect is, in most of the specific-

[28]According to Steiner (2001) the negative effect might be explained by human capital depreciation, employer stigmatization, or sorting.

ations, expected to decrease the hazard rate; the hazard rate to employment might increase with age for the group of individuals between 40 and 45 years old; and, after age 45, each additional year of age decreases the hazard rate dramatically. The first and the last pattern are also verified for the individuals that benefited from UA but the age effect during 40 to 45 years old is either negative or non-significant.

The remainder of the coefficient estimates are briefly described here.[29] Higher values of replacement ratio decrease the hazard rate, as the opportunity cost of being unemployed decreases. In terms of the career impact, tenure reduces the hazard rate, while the previous number of jobs increases the hazard rate, as verified by Addison and Portugal (2008b). However, unlike the results in the same study, the reasons of unemployment seem to play an important role shaping the hazard rate. While end of the fixed-term contract and plant closure/mass layoffs contribute positively to the hazard rate, redundancy and mutual agreement have a negative impact on the hazard rate. Males have, on average, higher hazard rates than females, and foreigners have, on average, lower hazard rates than nationals, but both coefficients are non-significant for individuals that received unemployment assistance. As expected, the current quarterly unemployment rate has an important impact on the hazard rates. The regional dummies shape the average hazard rates throughout the country, being the individuals in the North and in the islands, the ones that take longer to transit to employment, and the individuals in the Center the ones with higher hazard rates.

Finally, in terms of unobserved heterogeneity, the coefficient estimates reveal that between 17% and 27% of the individuals have hazard rates that are close to zero – either due to lack of acceptable offers or by avoiding the transition into inactivity (Addison and Portugal, 2008a). This result is unexpectedly large for individuals that have benefited

[29]The industries that lead to larger decreases in the hazard rate are "Financial Intermediation", "Real Estate and Renting", "Scientific and Technical Consulting Activities", and " Recreational, Cultural and Sporting Activities". The coefficients for the industry dummies are available upon request.

from unemployment assistance (34%). These numbers lie in the range of results reported in the literature that also assumed two mass points for the distribution of unobserved heterogeneity in the estimation (Narendranathan and Stewart (1993) reported 41%, Bover et al. (2002) reported 4%, Schmitz and Steiner (2007) reported 27.4%, Addison and Portugal (2008a) reported 45%, and Geerdsen et al. (2018) reported 12%).

1.7.2 Evidence on a short-term extended UA benefit program

In this section I analyse the impact of the program that extended the potential duration of unemployment assistance in Portugal between January 2009 and July 2010. The short-term nature of the policy measure enables the construction of two control groups: one that captures the pre-reform period, with a lower unemployment rate, and another for the post-reform period, with a higher unemployment rate, when compared with the treatment period. This helps to disentangle the impact of the reform from the increasing trend in unemployment rate. To study the impact of this measure I only consider individuals that received the subsequent UA. In this subsection, I estimate two different specifications.

The first specification identifies the change in the hazard rates to employment when individuals are receiving UA and are eligible to extra UA potential duration (1095 individuals), given by α_1; or receiving UA but non-eligible to extra UA potential duration (2528 individuals), given by α_2; or receiving the extra potential duration (508 individuals), given by α_3, in comparison with individuals that have already exhausted the benefits.[30]

[30]Remember that an individual is eligible to the extension by exhausting the unemployment assistance potential duration between January 2009 and July 2010.

$$h(t|\boldsymbol{x'_{it}\beta}) = 1 - \exp[-\exp(\alpha_1 \times \mathbb{1}\text{UARecipientNonEligible}_{it} + \alpha_2 \times \mathbb{1}\text{UARecipientEligible}_{it} + \alpha_3 \times \mathbb{1}\text{ExtraUARecipient}_{it} + \boldsymbol{x'_{it}\beta} + \ln(t))] \tag{1.19}$$

The second specification follows the same identification as specification 1.16 in the previous subsection, by analogously measuring the effect of the limited duration of unemployment assistance and the extra unemployment assistance.

$$h(t|\boldsymbol{x'_{it}\beta}) = 1 - \exp[-\exp(\alpha_1 \times \text{TimetoExhaustUA}_{it} + \alpha_2 \times \text{TimetoExhaustExtraUA}_{it} + \boldsymbol{x'_{it}\beta} + \ln(t))] \tag{1.20}$$

As in the previous subsection, when unobserved heterogeneity is taken into account, the coefficient on recipiency of the extra potential duration of UA becomes larger and statistically significant. This evidence supports the importance of accounting for unobserved heterogeneity when individuals have long unemployment spells, as it is the case for subsequent UA beneficiaries. Table 1.13 shows that approximately 40% of the individuals have a substantially lower hazard rate to employment, and that individuals that received the extra potential duration of UA had an hazard rate that was 30% lower than that for individuals that have exhausted their UA benefits.

In the second specification, which includes the *time to exhaust* variables, the coefficient estimate for unemployment assistance is still large and significant, as in previous subsection, but the coefficient for the extra unemployment assistance is not statistically significant. This means that even though recipiency of the extra benefit is estimated to decrease the hazard rate into employment, the number of months (6) of the potential duration of the extra benefit did not significantly affected the hazard rate into employment.

Table 1.13: Estimated c-log-log regression, with unobserved heterogeneity – the impact of unemployment assistance eligibility, recipiency, and potential duration on the hazard rate to employment

Variables	(8)		(9)	
	Coeff.	(SE)	Coeff.	(SE)
Eligibility to Extra UA†				
Non-eligible UA Beneficiary	-0.5922	(0.0983)		
Eligible UA Beneficiary	-0.7178	(0.137)		
𝟙 **Recipiency of Extra UA**†	-0.3532	(0.1418)		
Time to Exhaust UB (in months) †				
Unemployment Assistance			-0.0546	(0.0205)
Extra Unemployment Assistance			-0.0345	(0.0319)
Ln(Joblessness Duration) †	0.4270	(0.0675)	0.5112	(0.0919)
Age spline (in years)				
min(age,30)	-0.0577	(0.0164)	-0.0601	(0.0163)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	-0.0671	(0.0149)	-0.0664	(0.0147)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	-0.0702	(0.0375)	-0.0671	(0.0375)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1070	(0.0186)	-0.1052	(0.0184)
Ln(Replacement Ratio) ‡	-0.6418	(0.1113)	-0.6414	(0.108)
Past Career Information:				
Tenure (in years)	-0.0114	(0.0099)	-0.0114	(0.0097)
Number of jobs	0.1090	(0.0362)	0.1108	(0.0343)
𝟙 Industry	Yes		Yes	
𝟙 Reason of Unemployment				
end of fixed-term contract	0.0534	(0.1381)	0.0240	(0.1290)
redundancy	-0.1322	(0.1400)	-0.116	(0.1320)
mutual agreement	-0.1559	(0.1570)	-0.1374	(0.1496)
plant closure or mass layoff	0.3130	(0.1722)	0.3162	(0.1639)
Other Controls:				
𝟙 Gender (Male=1)	-0.0285	(0.0824)	-0.0262	(0.0785)
𝟙 Nationality (Foreign=1)	0.1397	(0.1173)	0.1299	(0.1132)
Quarterly Unemployment Rate †	-0.2203	(0.0181)	-0.2153	(0.0174)
𝟙 Region of Residence	Yes		Yes	
Unobserved heterogeneity:				
Δ intercept (*faster exiters*)	2.5633	(0.1309)	2.5326	(0.1668)
Intercept (*slower exiters*)	-4.1577	(0.2054)	-4.4893	(0.7238)
Proportion of *faster exiters*	0.5911	(0.0299)	0.6296	(0.0275)
Proportion of *slower exiters*	0.4089	(0.0299)	0.3704	(0.0275)
Log-likelihood	-6601.8425		-6620.7192	
Nr of observations	63 142		63 142	

† Time-varying covariates

‡ The replacement ratio is defined as ln(monthly UA) - ln(reference wage).

1.8 Conclusions

In this chapter I use a rich administrative dataset from the Portuguese Social Security records between 2005 and 2012. The data includes very detailed information on the career history of the unemployed individuals and precise information on the potential duration of both unemployment insurance and unemployment assistance benefits. Throughout this chapter, I show that a longer potential duration of unemployment benefits dramatically decreases the hazard rate into employment.

Firstly, a descriptive analysis shows evidence that the transition rate into employment spikes at the point of benefit exhaustion, being the spike lower for unemployment assistance exhaustion than for that of unemployment insurance.

Secondly, I explore different specifications in the estimation of complementary log-log hazard models, with two mass points of unobserved heterogeneity. I analyse both the impact of the recipiency and the potential duration of unemployment benefits. The identification is based on the variation across time, age, and working experience that define the rules of the potential duration of unemployment benefits. I explicitly explore the policy change of 2007 which exogenously changed the potential duration of unemployment benefits. The results reveal that UI recipiency decreases the hazard rate of employment by 73%, and that an increase of one month in potential duration of unemployment insurance leads to an average decrease of 9% in the hazard rate.

Comparing these results with an extensive and careful literature review provided in the beginning of this chapter, I conclude that both effects are larger than what was previously found (up to 50%, and up to 8%, respectively). The unobserved heterogeneity estimates indicate that about 20% of the individuals in the sample have a substantially lower hazard rate to employment. This result is lower than what has been reported in the literature (around 40%).

Finally, in order to evaluate the impact of an unexpected benefit duration extension

I use a reform that was applied in 2009 and 2010. According to this policy, all the individuals which had/or would have exhausted the unemployment assistance benefit in 2009 were entitled to an additional 6 months, regardless of their age, career or any kind of means test. Eligibility seems to decrease the hazard rates of UA beneficiaries but the estimate is not statistically significant. However, once unobserved heterogeneity is taken into account, the recipiency of the extra potential duration is estimated to decrease the hazard rate to employment in 30%, when compared to individuals that have exhausted all the UA benefits they were entitled to.

Appendix of Chapter 1

A1 Unemployment Benefits in Portugal

A1.1 Funding and Entitlement Requirements

Following the European standards, the unemployment benefits' system in Portugal is composed of two subsystems: unemployment insurance (UI) and unemployment assistance (UA). The former is funded by contributions from workers and employers, while the latter is funded by general government revenues. An individual is entitled to receive the UA in two cases: either it does not have the necessary requirements to get access to the UI (initial UA beneficiaries), or it has exhausted the potential duration of UI (subsequent UA beneficiaries). In both situations the individual is means tested – the aggregate income of the household should not overcome the 80% of IAS. [31]

In order to be entitled to either of the two unemployment benefits subsystems, an individual must have been fired as an employee. [32] The discern rule between UI and initial UA is the number of monthly contributions to the Social Security during the two years before the date of unemployment. The contributions requirement to be eligible for UI was eighteen months between January 2005 (beginning of the period in analysis) and

[31] IAS - *Indexantes de Apoios Sociais* - is a minimum monthly guaranteed amount of money that has evolved in the following way: 374.70 €(2005); 385.90 €(2006); 397.86 €(2007); 407.41 €(2008); 419.22 €(henceforth)

[32] Note that voluntary dismissals, unless fair, are not entitled to any type of unemployment insurance. Additionally, until the last major change in in April 2012, self-employed workers were not entitled to UI

January 2007, fifteen months between January 2007 and January 2010, twelve months between January 2010 and July 2010 and fifteen months again until March 2012 (ending of the period in analysis). The contributions requirement to be eligible for initial UA was six months only, throughout the entire period.

A1.2 Monetary Amount

The UI benefit amount is defined according to the past remunerations (including at most one Holidays' subsidy and one Christmas's subsidy) received during the first 12 of the 14 months previous to the date of unemployment.[33]

Table 1.14: Benefit amount rules - Unemployment Insurance

<table>
<tr><th rowspan="2"></th><th rowspan="2">net wage < IAS</th><th colspan="2">net wage ≥ IAS</th><th rowspan="2">ceiling</th></tr>
<tr><th>65% gross wage ≤ IAS</th><th>65% gross wage > IAS</th></tr>
<tr><td>Before July 2010</td><td rowspan="2">net wage</td><td rowspan="2">IAS</td><td>65% gross wage</td><td rowspan="2">3xIAS</td></tr>
<tr><td>After July 2010</td><td>75% net wage</td></tr>
</table>

Notes: by gross wage is meant the sum of the base wage with, if so, one Holidays and one Christmas subsidy. The net wage is calculated as the gross wage minus contributions to Social Security and income taxes.

Due to the lighter contributions requirement, the remunerations' period shrinks to the first 6 during the 8 months previous to the date of unemployment, in the case of initial UA. Although the period of subsidized unemployment contributes for the calculus of the old-age pensions, the unemployment benefits are exempted from income taxes. [34] The definition of the UI amount (once again, in place during the period of analysis) is

[33]In Portugal the annual remunerations are defined in 14 months because in June and December an employee receives two monthly salaries. In June,the regular salary is complemented by the "Holidays' subsidy", and in December the regular salary is complemented by the "Christmas's subsidy".

[34]Under a crisis environment, the Portuguese Government has implemented a 6% tax in the unemployment benefits in 2013, which has been refused by the Constitutional Court and given back to all the beneficiaries

summarized in Table 1.14. If the average net wage during the first 12 of the 14 months previous to the date of unemployment is lower than the IAS in place .

The UA amount is solely determined by the household composition. If the unemployed lives alone, then it receives the IAS. If not, then it receives 80% of that amount. In case these amounts are higher than the previous net wage of the unemployed, that amount replaces the previously mentioned values. During the extensions provided in 2009 and 2010 the daily amount received in this program was 8.38€, i.e., 60% of the IAS, plus 1.4€ per each child until the maximum of 13.97€ daily benefit which corresponds to the IAS.

A1.3 Training

Both potential durations and benefit amounts paid by Social Security may change due to training or occupational programs, during the period of UI/UA. Besides actively looking for a job, the subsidized unemployed is obliged to do training provided by IEFP (*Instituto de Emprego e Formação Profissional*) which helps not only the re-integration into the job market but also the pursuing of higher degree studies. During the training, the IEFP provides food and travel allowance. To guarantee that the individual does not earn less than the initial benefit amount (as an incentive to do training), Social Security provides the difference between the initially set up benefit amount and the allowances provided by the IEFP. During the period of training, the potential duration of the unemployed is frozen. At the end of that period, in order to know for how much longer the unemployed is still entitled to receive the unemployment benefit, the Social Security sums all the money given in compensation and divides by the initially set daily benefit amount: subtracting the final result to the potential duration of the individual before starting training yields the number of days, if any, for which the individual is still entitled to receive the initially set benefit amount. Regarding the occupational program, only the daily amount (potential duration remains unchanged) is different if the occupation is remunerated, i.e., the unemployed may receive more than the amount that was set at the beginning of the insurance/assistance period.

A2 Sample Exclusion Note

The individuals that receive training from the training centres in Social Security are eligible to receive the unemployment benefits for a longer period than what was initially defined. During the training period, they only receive part of the unemployment benefits in order to complement the transports and food allowances that are provided in training. Dividing the smaller benefit amount by the period of training at the end of it yields the actual number of days receiving the unemployment benefits, which is smaller than the actual time period, thus extending the potential duration of unemployment benefits. Those in occupational programs receive an extra amount on top of the unemployment benefits. According to the job search theory, a higher benefit yields a higher reservation wage and thus a smaller probability of leaving unemployment. On the other hand, the occupational program works as a signal to the employer, hence yielding a larger probability of leaving unemployment. In order to avoid bias on the estimates, I opt to exclude these individuals. Finally, as for the parenthood benefits I verify that those interrupting the unemployment benefit to receive this alternative insurance have used 84% of the potential duration of unemployment benefits, contrasting with the average of 57% when considering the whole initial sample. Moreover, 40% of those are actually right-censored spells, i.e., no transition from the unemployment benefit to employment is observed.

A2 Potential Duration of Unemployment Benefits in Pre- and Post- 2007 Reform

Table 1.15: Potential Duration of Unemployment Benefits at age< 30, by experience and start of benefits recipiency

Age at UI entry	Age at UA entry	UI start	UA start	SS months	Experience (in years)	Potential Duration (in months) UI	Δ_{2007} UI	UA	Δ_{2007} UA
<30	<30	<2007	<2007	–	–	12	–	6	–
			≥2007	≤24	–	12	–	4.5	-1.5
				>24	<5	12	–	6	0
					[5,10[			6.5	0.5
					[10,15[			7	1
					[15,20[			7.5	1.5
					≥ 20			8	2
		≥2007	≥2007	≤24	–	9	-3	4.5	-1.5
				>24	<5	12	0	6	0
					[5,10[	13	1	6.5	0.5
					[10,15[	14	2	7	1
					[15,20[	15	3	7.5	1.5
					≥ 20	16	4	8	2
	[30,40[	<2007	<2007	–	–	12	–	9	0
			≥2007	≤48	–	12	–	6	-3
				>48	<5	12	–	9	0
					[5,10[			9.5	0.5
					[10,15[			10	1
					[15,20[			10.5	1.5
					≥ 20			11	2
		≥2007	≥2007	≤24	–	9	-3	6	-3
				]24,48]	–	12	0	6	-3
				>48	<5	12	0	9	0
					[5,10[	13	1	9.5	0.5
					[10,15[	14	2	10	1
					[15,20[	15	3	10.5	1.5
					≥ 20	16	4	11	2

Note: Age at entry in Unemployment Insurance (UI) and Unemployment Assistance (UA) are defined according to the thresholds in the law. The number of months of contributions to Social Security (SS months) might not correspond to number of months with labour market experience as, for example, receiving disability benefits also counts for contribution purposes. Δ_{2007} stands for the change in the potential duration of the unemployment benefit that comes from the fact that the individual claimed that benefit after January 2007.

Table 1.16: Potential Duration of Unemployment Benefits at age]30, 40], by experience and start of benefits recipiency

Age at UI entry	Age at UA entry	UI start	UA start	SS months	Experience (in years)	Potential Duration (in months) UI	Δ_{2007} UI	UA	Δ_{2007} UA
			<2007	–	–	18	–	9	–
				≤48	–	18	–	6	-3
					<5			9	0
		<2007	≥2007		[5,10[			9.5	0.5
				>48	[10,15[	18	–	10	1
					[15,20[			10.5	1.5
	[30,40[				≥ 20			11	2
				≤48	–	12	-6	6	-3
					<5	18	0	9	0
		≥2007	≥2007		[5,10[	19	1	9.5	0.5
				>48	[10,15[	20	2	10	1
					[15,20[	21	3	10.5	1.5
					≥ 20	22	4	11	2
[30,40[			<2007	–	–	18	–	9	–
				≤60	–	18	–	9	-3
					<5			12	0
		<2007	≥2007		[5,10[			12.5	0.5
				>60	[10,15[	18	–	13	1
					[15,20[			13.5	1.5
	[40,45[				≥ 20			14	2
				≤48	–	12	-6	9	-3
				]48,60]	–	18	0	9	-3
					<5	18	0	12	0
		≥2007	≥2007		[5,10[	19	1	12.5	0.5
				>60	[10,15[	20	2	13	1
					[15,20[	21	3	13.5	1.5
					≥ 20	22	4	14	2

Note: Age at entry in Unemployment Insurance (UI) and Unemployment Assistance (UA) are defined according to the thresholds in the law. The number of months of contributions to Social Security (SS months) might not correspond to number of months with labour market experience as, for example, receiving disability benefits also counts for contribution purposes. Δ_{2007} stands for the change in the potential duration of the unemployment benefit that comes from the fact that the individual claimed that benefit after January 2007.

Table 1.17: Potential Duration of Unemployment Benefits at age $]40, 45]$, by experience and start of benefits recipiency

Age at UI entry	Age at UA entry	UI start	UA start	SS months	Experience (in years)	Potential Duration (in months)			
						UI	Δ_{2007} UI	UA	Δ_{2007} UA
			<2007	–	–	24	–	12	–
				≤60	–	24	–	9	-3
					<5			12	0
		<2007			[5,10[			12.5	0.5
			≥2007	>60	[10,15[	24	–	13	1
					[15,20[			13.5	1.5
	[40,45[				≥ 20			14	2
				≤60	–	18	-6	9	-3
					<5	24	0	12	0
		≥2007	≥2007		[5,10[	25	1	12.5	0.5
				>60	[10,15[	26	2	13	1
					[15,20[	27	3	13.5	1.5
					≥ 20	28	4	14	2
					<5			15	
					[5,10[			16	
[40,45[			<2007		[10,15[	24	–	17	–
					[15,20[			18	
					≥ 20			19	
		<2007		≤ 72	–	24	–	12	-3
					<5			15	
			≥2007		[5,10[			16	
	≥45			>72	[10,15[	24	–	17	–
					[15,20[			18	
					≥ 20			19	
				≤60	–	18	-6	12	-3
				]60,72]	–	24	0	12	0
					<5	24	0	15	
		≥2007	≥2007		[5,10[	25	1	16	
				>72	[10,15[	26	2	17	0
					[15,20[	27	3	18	
					≥ 20	28	4	19	

Note: Age at entry in Unemployment Insurance (UI) and Unemployment Assistance (UA) are defined according to the thresholds in the law. The number of months of contributions to Social Security (SS months) might not correspond to number of months with labour market experience as, for example, receiving disability benefits also counts for contribution purposes. Δ_{2007} stands for the change in the potential duration of the unemployment benefit that comes from the fact that the individual claimed that benefit after January 2007.

Table 1.18: Potential Duration of Unemployment Benefits at age≥ 45, by experience and start of benefits recipiency

Age at UI entry	Age at UA entry	UI start	UA start	SS months	Experience (in years)	Potential Duration (in months) UI	Δ_{2007} UI	UA	Δ_{2007} UA
					<5	30		15	
					[5,10[	32		16	
			<2007		[10,15[	34	–	17	–
					[15,20[	36		18	
					≥ 20	38		19	
		<2007		≤72	–	30	0	12	-3
					<5	30		15	
					[5,10[	32		16	
≥45	≥45		≥2007	>72	[10,15[	34	0	17	0
					[15,20[	36		18	
					≥ 20	38		19	
				≤72	–	24	-6	12	-3
					<5	30		15	
					[5,10[	32		16	
		≥2007	≥2007	>72	[10,15[	34	0	17	0
					[15,20[	36		18	
					≥ 20	38		19	

Note: Age at entry in Unemployment Insurance (UI) and Unemployment Assistance (UA) are defined according to the thresholds in the law. The number of months of contributions to Social Security (SS months) might not correspond to number of months with labour market experience as, for example, receiving disability benefits also counts for contribution purposes. Δ_{2007} stands for the change in the potential duration of the unemployment benefit that comes from the fact that the individual claimed that benefit after January 2007.

A3 Additional Descriptive Statistics of Subsequent UA Beneficiaries

Table 1.19: Descriptive statistics of subsequent UA beneficiaries (continuation)

	Mean	Std	Min	Max
Reason of unemployment				
end of fixed-term contract	0.41			
redundancy	0.33			
mutual agreement	0.16			
plant closure or mass lay-off	0.10			
others	0.10			
Year of unemployment				
2005	0.22			
2006	0.21			
2007	0.13			
2008	0.19			
2009	0.18			
2010	0.06			
2011	0.01			
Beginning of UA spell				
2006	0.11			
2007	0.15			
2008	0.16			
2009	0.17			
2010	0.20			
2011	0.18			
2012 (until March)	0.03			
Age (in years)	39.18	12.10	17	74
Male	0.49	0.50		
Region of Residence				
North	0.56			
Center	0.12			
Lisbon	0.21			
Alentejo	0.03			
Algarve	0.02			
Azores and Madeira	0.06			
Re-employed (any job)	1 702 (1 584 as dependent employees)			
Retired, disabled, dead	166			
Censored spells	1755			
Sample	3 623			

A3 Empirical Results Without Unobserved Heterogeneity

Table 1.20: Estimated c-log-log regression – the impact of potential duration of unemployment insurance benefits on the hazard rate to employment

Variables	(1)		(2)	
	Coeff.	(SE)	Coeff.	(SE)
UI Potential Duration (in months)	-0.0819	(0.0030)		
Δ_{2007} **UI Potential Duration (in months)**			-0.1037	(0.0030)
Ln(Joblessness Duration) †	-0.0424	(0.0096)	-0.0445	(0.0096)
Age spline (in years)				
min(age,30)	0.0061	(0.0044)	-0.0249	(0.0041)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	0.0073	(0.0043)	-0.0372	(0.0041)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	0.1523	(0.0113)	-0.0277	(0.0098)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1022	(0.0048)	-0.1334	(0.0050)
Ln(Replacement Ratio) ‡	-0.4498	(0.0555)	-0.4446	(0.0553)
Past Career Information:				
Tenure (in years)	-0.0136	(0.0027)	-0.0166	(0.0027)
Number of jobs	0.1165	(0.0056)	0.1167	(0.0056)
$\mathbb{1}$ Industry $\mathbb{1}$ Reason of Unemployment				
end of fixed-term contract	0.0563	(0.0329)	0.0400	(0.0329)
redundancy	-0.1288	(0.0362)	-0.1377	(0.0362)
mutual agreement	-0.2789	(0.0426)	-0.2852	(0.0426)
plant closure or mass layoff	0.1047	(0.0458)	0.1051	(0.0458)
Other Controls:				
$\mathbb{1}$ Gender (Male=1)	0.1868	(0.0209)	0.1859	(0.0209)
$\mathbb{1}$ Nationality (Foreign=1)	-0.0586	(0.0287)	-0.0686	(0.0288)
Quarterly Unemployment Rate †	-0.2654	(0.0058)	-0.2676	(0.0058)
$\mathbb{1}$ Region of Residence				
North	-0.0973	(0.0264)	-0.0991	(0.0264)
Center	0.2139	(0.0302)	0.2170	(0.0302)
Alentejo	0.4284	(0.0414)	0.4176	(0.0413)
Algarve	0.3551	(0.0489)	0.3408	(0.0489)
Azores and Madeira	-0.2271	(0.0506)	-0.2216	(0.0506)
Intercept	0.5600	(0.1501)	0.3261	(0.1475)
Log-likelihood	-43581.9280		-43635.5839	
Nr of observations	406 626		406 626	

† Time-varying covariates ‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).
Note: column (2) only includes observations of individuals that also received unemployment assistance.

Table 1.21: Estimated c-log-log regression – the impact of changes in the rules of UB potential duration on the hazard rate to employment

Variables	**(3)** Coeff.	(SE)	**(4)** Coeff.	(SE)
UI Potential Duration (in months)	-0.0625	(0.0132)		
UA Potential Duration (in months)	-0.0557	(0.0239)		
Δ_{2007} **UI Potential Duration (in months)**			-0.0611	(0.0302)
Δ_{2007} **UA Potential Duration (in months)**			-0.0227	(0.0567)
Ln(Joblessness Duration) †	1.4672	(0.0499)	1.4538	(0.0499)
Age spline (in years)				
min(age,30)	-0.0453	(0.0111)	-0.0800	(0.0098)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	-0.0283	(0.0122)	-0.0819	(0.0105)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	0.0250	(0.0395)	-0.1820	(0.0345)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1116	(0.0225)	-0.1409	(0.0235)
Ln(Replacement Ratio) ‡	-1.1698	(0.1574)	-1.1962	(0.1580)
Past Career Information:				
Tenure (in years)	-0.0002	(0.0072)	-0.0054	(0.0075)
Number of jobs	0.1655	(0.0242)	0.1657	(0.0244)
$\mathbb{1}$ Industry	Yes		Yes	
$\mathbb{1}$ Reason of Unemployment				
end of fixed-term contract	-0.0601	(0.0844)	-0.0666	(0.0845)
redundancy	-0.0692	(0.0909)	-0.0843	(0.0910)
mutual agreement	-0.3040	(0.1044)	-0.2973	(0.1043)
plant closure or mass layoff	0.2495	(0.1152)	0.2456	(0.1155)
Other Controls:				
$\mathbb{1}$ Gender (Male=1)	0.1095	(0.0560)	0.1038	(0.0561)
$\mathbb{1}$ Nationality (Foreign=1)	0.1596	(0.0788)	0.1584	(0.0791)
Quarterly Unemployment Rate †	-0.2405	(0.0146)	-0.2388	(0.0147)
$\mathbb{1}$ Region of Residence				
North	-0.1342	(0.0699)	-0.1417	(0.0700)
Center	-0.1943	(0.0930)	-0.1889	(0.0928)
Alentejo	0.1387	(0.1448)	0.1042	(0.1447)
Algarve	-0.1153	(0.1898)	-0.1342	(0.1898)
Azores and Madeira	-0.2503	(0.1227)	-0.2324	(0.1227)
Intercept	-4.0066	(0.4429)	-4.2205	(0.4390)
Log-likelihood	-7573.0759		-7596.5625	
Nr of observations	141 183		141 183	

† Time-varying covariates ‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).
Note: column (2) only includes observations of individuals that also received unemployment assistance.

Table 1.22: Estimated c-log-log regression – the time-varying impact of unemployment benefits on the hazard rate to employment

Variables	**(5)**		**(6)**		**(7)**	
	Coeff.	(SE)	Coeff.	(SE)	Coeff.	(SE)
$\mathbb{1}$ **Recipiency** †						
Unemployment Insurance	-0.6338	(0.0334)				
Unemployment Assistance	-0.6973	(0.0457)				
UB Initial Potential Duration (in months) †						
Unemployment Insurance			-0.0403	(0.0014)		
Unemployment Assistance			-0.0838	(0.0046)		
Time to Exhaust UB (in months) †						
Unemployment Insurance					-0.0276	(0.0024)
Unemployment Assistance					-0.0883	(0.0073)
Ln(Joblessness Duration) †	-0.1851	(0.0124)	-0.1875	(0.011)	-0.1894	(0.0163)
Age spline (in years)						
min(age,30)	-0.0229	(0.0042)	-0.0099	(0.0042)	-0.0176	(0.0042)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	-0.0288	(0.0041)	-0.0124	(0.0041)	-0.0233	(0.0042)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	-0.0001	(0.0099)	0.0795	(0.0103)	0.0339	(0.0105)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.1309	(0.005)	-0.1230	(0.0049)	-0.1286	(0.0050)
Ln(Replacement Ratio) ‡	-0.4667	(0.0555)	-0.4690	(0.0553)	-0.4328	(0.0557)
Past Career Information:						
Tenure (in years)	-0.0208	(0.0027)	-0.0172	(0.0027)	-0.0203	(0.0027)
Number of jobs	0.1254	(0.0056)	0.1250	(0.0055)	0.1200	(0.0056)
$\mathbb{1}$ Industry	Yes		Yes		Yes	
$\mathbb{1}$ Reason of Unemployment						
end of fixed-term contract	0.1028	(0.0328)	0.0993	(0.0328)	0.0934	(0.0328)
redundancy	-0.0845	(0.0362)	-0.0818	(0.0362)	-0.0947	(0.0362)
mutual agreement	-0.2639	(0.0426)	-0.2601	(0.0426)	-0.2671	(0.0426)
plant closure or mass layoff	0.1418	(0.0459)	0.1600	(0.0459)	0.1297	(0.0459)
Other Controls:						
$\mathbb{1}$ Gender (Male=1)	0.2002	(0.0208)	0.1983	(0.0208)	0.1974	(0.0208)
$\mathbb{1}$ Nationality (Foreign=1)	-0.0097	(0.0286)	-0.0333	(0.0286)	-0.0146	(0.0286)
Quarterly Unemployment Rate †	-0.2923	(0.0059)	-0.2921	(0.0059)	-0.2792	(0.0058)
$\mathbb{1}$ Region of Residence						
North	-0.1010	(0.0264)	-0.0950	(0.0264)	-0.0970	(0.0264)
Center	0.2223	(0.0302)	0.2154	(0.0302)	0.2269	(0.0302)
Alentejo	0.4124	(0.0413)	0.4157	(0.0413)	0.4135	(0.0413)
Algarve	0.3580	(0.0489)	0.3564	(0.0489)	0.3631	(0.0489)
Azores and Madeira	-0.2142	(0.0506)	-0.2148	(0.0506)	-0.2048	(0.0506)
Intercept	1.2323	(0.1520)	0.8005	(0.1492)	0.7258	(0.1489)
Log-likelihood	-43734.2790		-43523.9778		-43821.3340	
Nr of observations	406 626		406 626		406 626	

† Time-varying covariates

‡ The replacement ratio is defined as ln(monthly UI) - ln(reference wage).

Table 1.23: Estimated c-log-log regression – the impact of unemployment assistance eligibility, recipiency, and potential duration on the hazard rate to employment

Variables	**(8)**		**(9)**	
	Coeff.	(SE)	Coeff.	(SE)
Eligibility to Extra UA†				
Non-eligible UA Beneficiary	-0.8057	(0.0967)		
Eligible UA Beneficiary	-0.9296	(0.1328)		
$\mathbb{1}$ **Recipiency of Extra UA**†	-0.1270	(0.1342)		
Time to Exhaust UB (in months) †				
Unemployment Assistance			-0.1772	(0.0166)
Extra Unemployment Assistance			-0.0257	(0.0307)
Ln(Joblessness Duration) †	-0.1956	(0.0450)	-0.2960	(0.0476)
Age spline (in years)				
min(age,30)	-0.029	(0.0113)	-0.0231	(0.0114)
(age-30) $\times\mathbb{1}$ (age $\in$ [30,40[) + 10 $\times\mathbb{1}$ (age $\geq$ 40)	-0.0420	(0.0105)	-0.0315	(0.0106)
(age-40) $\times\mathbb{1}$ (age $\in$ [40,45[) + 5 $\times\mathbb{1}$ (age $\geq$ 45)	-0.0229	(0.0276)	0.0120	(0.0278)
(age-45) $\times\mathbb{1}$ (age $\geq$ 45)	-0.0937	(0.0158)	-0.0823	(0.0157)
Ln(Replacement Ratio) ‡	-0.3532	(0.0789)	-0.3677	(0.0790)
Past Career Information:				
Tenure (in years)	-0.0014	(0.0073)	0.0012	(0.0073)
Number of jobs	0.1357	(0.0251)	0.1259	(0.0251)
$\mathbb{1}$ Industry	Yes		Yes	
$\mathbb{1}$ Reason of Unemployment				
end of fixed-term contract	-0.0291	(0.0903)	-0.0444	(0.0903)
redundancy	-0.0714	(0.0967)	-0.0843	(0.0967)
mutual agreement	-0.1932	(0.1091)	-0.1955	(0.1092)
plant closure or mass layoff	0.2273	(0.1199)	0.2187	(0.1198)
Other Controls:				
$\mathbb{1}$ Gender (Male=1)	0.0859	(0.0580)	0.0783	(0.0579)
$\mathbb{1}$ Nationality (Foreign=1)	0.1027	(0.0800)	0.0758	(0.0799)
Quarterly Unemployment Rate †	-0.1847	(0.0141)	-0.1742	(0.0142)
$\mathbb{1}$ Region of Residence	Yes		Yes	
Intercept	-1.3217	(0.5373)	-1.4801	(0.5355)
Log-likelihood	-6688.1420		-6664.1371	
Nr of observations	63 142		63 142	

† Time-varying covariates

‡ The replacement ratio is defined as ln(monthly UA) - ln(reference wage).

Chapter 2

Unemployment Duration and Re-employment Wages

2.1 Introduction

Theory predicts that unemployment duration can affect re-employment wages through three main channels: skill depreciation, job searching, and stigmatization. First, the longer the unemployment spell, the longer the individual is prevented from accumulating work experience, thus deteriorating both general and job-specific human capital (Becker, 1962). Second, in an opposite direction, job search models predict that a longer spell of unemployment allows individuals to search for a better match, generating positive wage returns from unemployment (Mortensen, 1988). Altonji and Shakotko (1987) argued that job shopping is one of the main contributors for wage growth over a career. Finally, a stigma effect might arise in labour markets with imperfect information (Vishwanath, 1989; Belzil, 1995). As the employer cannot distinguish whether a longer spell of unemployment represents low productivity or a low rate of job offer arrivals, the wage offered might be reduced, at least during the screening period in the new job, as part of a "ranking" hiring strategy that gives preference to individuals with shorter unemployment duration (Blanchard and Diamond, 1994).

While the theoretical literature seems to agree on the existence of an impact of unemployment duration on re-employment wages, the empirical literature so far has not reached a consensus. The main difficulty is the simultaneity issue present in the estimation. As reservation wages are not usually observed, it is hard to distinguish whether the unemployed individual accepted the new job at a certain point in time because that was when the offered wage was above the reservation wage that was set in the beginning of the job search, or whether the unemployment duration up until that moment lead the individual to decrease the threshold rule to be re-employed faster.

To overcome this endogeneity problem, Addison and Portugal (1989) recurred to instrumental variables (IV) estimation. In the first stage, the authors used previous job duration and reason for unemployment as potential sources of exogenous variation to explain unemployment duration. The authors estimated this step using both complete and incomplete spells to improve prediction, keeping the estimation sample as representative of the unemployed population. Results from the second stage indicated that a 10% increase in unemployment duration lowered accepted wages by about 1% for the United States between 1979 and 1984.

Other studies resorted to regression discontinuity designs, relying on the institutional settings of unemployment insurance. They do not evaluate directly at the impact of unemployment duration on re-employment wages but rather at the channel of unemployment insurance, which is assumed to affect re-employment wages through the primary effect on unemployment duration. Using such methodology the direction of the results is unclear.Lalive (2007) and Card et al. (2007a) found no significant impact on wages. Both studies used at Austrian data. The former explored the discontinuity in benefit duration at the age of 50 whereas the latter explored the discontinuity in severance pay at 36 months of tenure. Nekoei and Weber (2017) were the only ones to find a positive effect. A nine-week extension in the potential duration of unemployment benefits in Austria, was expected to increase unemployment duration by 2 days, but predicted to increase re-employment wages by 0.5%. Centeno and Novo (2009) and Schmieder et al. (2016) have reported statistically significant negative impacts. The two studies explored age discontinues as sources of quasi-experimental variation in unemployment duration and achieved quite similar results. For each additional month of unemployment duration, the former study reported an effect of 1% using Portuguese data between 1999 and 2002, and the latter study reported a reduction in wage offers of 0.8% in Germany between 1975 and 2008.

In this chapter, I adopt an empirical approach that has not been used before, in the literature of unemployment duration and re-employment wages. I employ a control function approach, which is essentially an instrumental variables estimation (Wooldridge, 2015). This approach has two main advantages in the context of the research question in this chapter. First, to estimate the first stage in a non-linear fashion (accelerated time failure model) and to use both complete and incomplete spells of unemployment duration. Second, by plugging the residuals from the first stage into the re-employment wage equation, to test and control for endogeneity.

This chapter also distinguishes from the previous literature in the fact that, the source of exogenous variation I use, affects the beneficiaries in a non-monotonic fashion. Most of the previous studies recurred to either extensions or shortenings in the entitled duration. I use a reform, implemented in Portugal in January 2007, which changed the potential unemployment benefits' duration in different directions. Around 41% of the individuals got longer potential duration of unemployment benefits, 23% got shorter potential duration of unemployment benefits, and approximately 36% of them have not faced any change in the potential duration of unemployment benefits.

Exploring the exogenous variation in the reform, I construct three alternative instrumental variables: the potential duration of unemployment duration; the age discontinuities in the potential duration of unemployment duration; and the changes in the potential duration of unemployment duration that were generated by the reform. Among the three different alternatives, the last instrument is arguably more convincing. In comparison with the first one, it is less affected by other factors such as age and labour market experience which might also directly impact on re-employment wages. In comparison with the second one, it uses a larger number of observations, thus providing more precision in the results.

The main result of this chapter is: one additional month in unemployment duration leads to a decrease of approximately 0.5% in re-employment wages. Even though the magnitude is smaller, the result is in line with Addison and Portugal (1989) and Schmieder et al. (2016).

Finally, I test for both self-selection into treatment (longer unemployment duration) and self-selection into re-employment. First, I include in the outcome equation the interaction between the residuals and the treatment variable (unemployment duration) to test for self-selection into treatment. Second, I employ the traditional Heckit two-step approach accounting, in the selection equation, for the period during which the individual is observed to be unemployed. Third, I recur to an augmented inverse probability of censoring weighted estimation where the weights are the inverse probability of the individual to be censored at the end of the sample, March of 2012. According to these three tests, there is no indication that self-selection is a serious problem.

The chapter is organized as follows. Section 2.2 describes the institutional framework. Section 2.3 presents the data. Section 2.4 explains the empirical methodology. Section 2.5 describes each of the three instruments. Section 2.6 presents the main results. Section 2.7 performs a robustness analysis. Section 2.8 concludes.

2.2 Institutional Framework

This section describes the rules of the potential duration of unemployment benefits insurance (UI) system in Portugal during the period of the analysis, i.e., January 2005 to March 2012. Particular emphasis is given to the changes originated by the reform of 2007. The description of the remaining rules (funding, entitlement requirements, monetary amounts and training) is presented in the Appendix of Chapter 1.

The potential duration of unemployment benefits in Portugal currently depends on both age and career history, but this was not always the case. Before January 2007, only the individuals that were 45 years old or older could be entitled to extra potential duration, according to the number of years of Social Security contributions.[1] After the major reform in 2007, the rules for all individuals included both age and career history. The differences in entitlements are summarized in Table 2.1.

Table 2.1: Potential Duration Rules of UI (all values in months except age)

	Before Jan 2007		After Jan 2007		
Age at UI Start	Potential Duration	Extra Potential Duration †	Social Security Contributions‡	Potential Duration	Extra Potential Duration †
<30	12	-	≤24	9	-
			>24	12	1
[30,40[	18	-	≤48	12	-
			>48	18	1
[40,45[	24	-	≤60	18	-
			>60	24	1
≥45	30	2	≤72	24	-
			>72	30	2

† for each 5 years of working experience in the 20 years preceding the date of unemployment
‡ Social Security contributions since last spell of unemployment include both the recipiency of wages as well as other benefits such as parental leave and disability insurance.

[1] The new law was published in November of 2006, but the changes in potential duration of unemployment benefits have only taken place after January of 2007. Further changes took place in the end of 2006 and are briefly described at the end of this section.

Before the reform, the potential duration of unemployment benefits for individuals younger than 45 years old was solely determined by their age. Individuals younger than 30 years old were entitled to 12 months, individuals between 30 and 40 years old were entitled to 18 months, and individuals between 40 and 45 years old were entitled to 24 months of unemployment benefits. Individuals with 45 years old or more were entitled to 30 months but could extend this potential duration for additional 2, 4, 6, or 8 months, in case they have worked for at least 5, 10, 15 or 20 years, respectively.

After the reform, for individuals of all ages, the longer careers were awarded with an increased potential duration, while the shorter careers were entitled to a shorter potential duration of unemployment benefit. Some individuals, however, were entitled to the same potential duration as defined by the previous rules. An individual younger than 30 years old was entitled to: 9 months if she had 24 months or less of contributions to the Social Security; 12 months if she had between 24 months and 5 years of contributions; 13, 14, 15, or 16 months if she had, respectively, at least 5, 10, 15, or 20 years of contributions. An individual aged between 30 and 40 years old was entitled to: 12 months if she had 48 months or less of contributions to the Social Security; 18 months if she had between 48 months and 5 years of contributions; 19, 20, 21, or 22 months if she had, respectively, at least 5, 10, 15, or 20 years of contributions. An individual aged between 40 and 45 years old was entitled to: 18 months if she had 60 months or less of contributions to the Social Security; 24 months if she had between 60 months and 5 years of contributions; 25, 26, 27, or 28 months if she had, respectively, at least 5, 10, 15, or 20 years of contributions. Finally, the individuals with 45 years old or more were entitled to the same potential duration as before with the exception of those with 72 months or less of contributions to the Social Security - these individuals were entitled to 24 months rather than to 30 months as in the previous rules.

In addition to the change in potential duration of unemployment benefits, two additional measures were implemented in the unemployment benefits system. First, the required working period to guarantee eligibility to the unemployment benefit has decreased from 18 to 15 months.[2] Second, early retirement became more easily accessible to some of the unemployed.[3] These changes could potentially affect the results but the number of individuals affected by them is negligible.

2.3 Data

2.3.1 Portuguese Social Security Data

The database I use is a 1% representative sample from the panel of unemployment beneficiaries inflows registered in the Portuguese Social Security between January 2005 and March 2012.[4] The observations are recorded on a daily basis for 39 972 unemployment beneficiaries. It includes basic demographic information, career history, and all the necessary details about the unemployment benefits.

At the individual level, there is information on gender, date of birth, and district of residence. In terms of career history, the database goes back to the first registered job of the oldest individual (1941). Even though monthly earnings (split into base wage, regular, and non-regular benefits) are only available between 2005 and 2012, the database includes information on whether the individual is a dependent employee or self-employed, some details about the firm where the individual is employed (region and industry) and the starting/ending date of each job, for all paid jobs the individual ever

[2]The required working period is counted during the 24 months preceding the date of unemployment.

[3]The legal age for early retirement has increased from 60 to 62 years old for all the individuals aged at least 57 at the date of unemployment and with at least 15 years of contributions. For all the individuals aged at least 55 at the date of unemployment and with at least 22 (instead of 30) years of contributions, the legal age has decreased from 58 to 57 years old. Note that the individuals under the first conditions may opt for the second scheme as long as they have 22 (and not only 15) years of contributions.

[4]Entitled as *Microdados do Sistema de Informação da Segurança Social*

had.

The unemployment information comprises the start and ending dates of the unemployment spells (or potential ending date, in case of uncompleted spells), the amount and type of benefit, and the participation on any type of training or occupational program. Finally, there is also information on other types of exit besides re-employment, such as illness, disability, maternity, income support and old-age pension.

2.3.2 Sample Construction

The original sample contains 39 972 unemployment beneficiaries. In this chapter, I focus the analysis on the individuals that were covered by the unemployment insurance (UI) system described in Section 2.2. Therefore, I exclude the beneficiaries of unemployment initial assistance that have, by definition, less attachment to the labour market (8 862 individuals). I further exclude from the analysis all the individuals whose potential duration was extended due to training, parenthood, disease, or disability benefits as these events might raise endogeneity issues. Such decision resulted in the exclusion of 19% of the 29 760 individuals selected in the first step. For the same reason, I also exclude the individuals that have entered into early retirement because of the change in the rule that also occurred since January 2007. Finally, I also exclude all the individuals for which there is lack of information on any of the explanatory variables used in the regressions. Such options lead to a sample comprised by 18 543 individuals. Descriptive statistics are briefly described in the next subsection.

2.3.3 Descriptive Statistics

Table 2.2, presents the descriptive statistics for the constructed sample, which includes both individuals that were re-employed and individuals with censored unemployment duration, because they did not find a job before the last observed date in the data.

The average benefit duration of unemployment insurance benefit is around 12 months. However, the average individual remains jobless for approximately 18 months. In the empirical analysis, I opt to use the jobless duration rather than the benefit duration as the time for which the individual receives no benefit may also contribute to the loss of skills, and therefore may also impact the re-employment wage.

Table 2.2: Descriptive statistics - Full Sample

	Mean	Std.D.	Min	Median	Max
Benefit Duration (UI) (in months)	12.34	10.63	1	9.5	38
Jobless Duration (in months)	17.56	17.21	1	11	83
Level of benefits (€/month)	508.31	210.85	43.50	419.10	1257.60
Previous wage (€/month)	784.42	570.78	43.50	612.92	14208.65
Tenure (in years)	5.43	6.40	0.08	2.83	47.92
Reason of unemployment					
end of fixed-term contract	0.40				
redundancy	0.25				
mutual agreement	0.17				
plant closure or mass lay-off	0.11				
others	0.07				
Year of start unemployment spell					
2005	0.07				
2006	0.17				
2007	0.11				
2008	0.13				
2009	0.14				
2010	0.13				
2011	0.20				
2012 (until March)	0.05				
Age at unemployment (in years)	38.42	11.51	17	37	69
Male	0.49				
Region of Residence *					
North	0.37				
Center	0.22				
Lisboa	0.30				
Alentejo	0.03				
Algarve	0.05				
Azores and Madeira	0.04				
Re-employed (any job)	9 027 (8 423 as dependent employees)				
Social Benefit, disabled, dead	446				
Incomplete spells	9 070				
Sample	18 543				

The average monthly benefit is approximately 65% of the average previous monthly gross wage, which corresponds to the rule applied to most beneficiaries. However, the replacement ratio in the data actually ranges between 10% and 100%. A low replacement ratio is associated with high previous earnings, as the maximum level of benefits defined during the period of the analysis was 1257.60€. A high replacement ratio is associated with low previous earnings, as during the period of the analysis, the minimum benefit was set to be the individual's net wage if 65% of the monthly gross wage was lower than 419.22€.[5]

On average, each individual has worked for 5 years and 6 months in the job she had previous to the unemployment spell. However, the sample also includes workers with little experience (less than a year), as well as workers with almost 50 years of tenure.

Almost half of the sample lost their job because the fixed-term contract they had was not (or could not be) renewed. After that, redundancy (i.e., the employer claimed the job no longer existed or the individual was not suited for the job anymore) and mutual agreement are the most representative ones with 25% and 17%, respectively. Only 11% of the individuals lost their job due to plant closure or mass lay-off.

The distribution of individuals by year of start of the unemployment spell is roughly uniform between 2007 and 2010. The peak in 2011 illustrates the *catastrophic job destruction*, as described by Carneiro et al. (2014). The smaller number of individuals in 2005 is mostly related with the lack of complete information on all variables of interest. As the database ends in March of 2012, the number of individuals starting the spell in that year is also smaller.

The average age at the start of the unemployment spell is around 40 years old. However, the distribution is wide enough to guarantee that it is possible to evaluate

[5]Note that the previous wage is defined here the reference value used by the Social Security to calculate the benefit value. This value can vary between 12 and 14 monthly wages summed over the first 12 of the 14 months preceding the date of unemployment.

the discontinuities in the potential duration around the thresholds of 30, 40, and 45 years old. In terms of gender, the sample is evenly distributed. Lastly, the distribution of unemployed individuals in the sample is proportional to the regional distribution of the active population in Portugal.

To get an idea of the characteristics that might contribute to a faster exit to re-employment, Table 2.3 presents some descriptive statistics for the individuals that found a job until March 2012.

Table 2.3: Descriptive Statistics - Re-employed Sample

	Mean	Std.D.	Min	Median	Max
Benefit Duration (UI) (in months)	7.94	7.34	1	5.77	38
Jobless Duration (in months)	11.83	11.77	2	7	75
Level of benefits (€/month)	483.48	193.43	65.10	413.7	1257.60
Previous wage (€/month)	746.01	512.67	116.99	607.18	10463.61
re-employment wage (€/ month)	575.23	359.75	10.57	495.23	5602.57
Age at re-employment (in years)	34.36	9.49	17	33	66
Male	0.49	0.50	0	0	1
Sample			8 423		

In comparison with the whole sample, the re-employed individuals have substantially smaller duration of both unemployment protection and joblessness. They also receive, on average, a lower level of unemployment benefits, which might be correlated with the fact that they are younger, on average. Finally, the difference between the distribution of re-employment wages and the distribution of previous wages provides suggestive evidence of the wage loss experienced after an unemployment spell.

2.4 Methodology

The impact of the unemployment duration on the re-employment wage is given by the parameter α_1, in the following wage equation (in log form):

$$Ln(PostW_i) = \alpha_0 + \alpha_1 Ln(UD_i) + \boldsymbol{x_i'}\boldsymbol{\beta} + u_i \qquad (2.1)$$

where $PostW_i$ is the re-employment wage of individual i, and UD_i is the duration of the unemployment spell until re-employment. The vector of control variables, $\boldsymbol{x_i}$, includes pre-unemployment wage, gender, age, tenure, reasons of unemployment, previous industry, region, and the starting quarter of the unemployment spell. The error term, u_i, is assumed to be normally distributed.

When estimating the equation 2.1, one should have into account two potential issues: simultaneity bias and selection bias.

The first is related to the possibility that unemployment duration is rather determined by (instead of being a determinant of) the wage. According to job search theory (Mortensen, 1977), individuals should reject all the wage offers until the accepted wage equals the reservation wage.[6] Therefore, low reservation wages (and their associated accepted wages) might imply shorter unemployment duration spells.

The second issue is related to the possibility that there is selection in the sample of individuals for which the accepted wage is observed. The selection may either be positive if high-skilled individuals (earning higher earnings) are employed faster, and therefore are more likely to be observed. The selection may also be negative in case the accepted wage is so low that the individual has many offers to choose from, and therefore, higher likelihood to be re-employed at a faster pace, and be observed.

[6] Addison et al. (2013) have shown that reservation wages are not stationary along the course of the unemployment or jobless spell but also that the elasticity of the reservation wage with respect to elapsed unemployment duration is small.

The next two subsections present different alternative specifications to estimate the main outcome equation, which account for these two potential issues.

2.4.1 Accounting for Simultaneity Bias

To get around the potential simultaneity bias, I use a control function approach, which is inherently an instrumental variables estimation (Wooldridge, 2015). In both methods the estimation is done in two stages, with the first stage providing exogenous estimates for the endogenous variable ($Ln(UD_i)$) in the second stage.

The fundamental difference is the way the estimates from the first stage enter in the second stage (also called as main outcome equation). In the standard instrumental variables approach the fitted values of the first stage substitute for the endogenous variable in the second stage. In the control function approach, the generalized residuals obtained in the first stage ($\hat{\nu}_i$) are plugged in the second stage, without removing the endogenous variable from there, as it is displayed below:

$$Ln(PostW_i) = \alpha_0 + \alpha_1 Ln(UD_i) + \alpha_2 \hat{\nu}_i + \boldsymbol{x_i'\beta} + u_i \tag{2.2}$$

In terms of bias correction the two methods are equivalent, as the exogenously predicted outcome for the endogenous variable equals the difference between the estimated residual and the endogenous variable itself. In this set-up, α_1 measures the exogenous impact of unemployment duration on re-employment wages and α_2 indicates both the direction and significance of endogeneity, like an Hausman test (Hausman, 1978). The remaining variables and coefficients of equation 2.2 are defined as in 2.1.[7]

The main advantage of the control function approach is that the first stage can

[7]Even after this correction, there is another concern related with the validity of the usual standard errors. Since the two stages are estimated in separate steps, I follow Wooldridge (2015) and bootstrap the sample in 1000 repetitions for all the estimations.

be estimated with an accelerated failure time model, which accounts for the censoring in some of the unemployment spells present in the data.[8] By doing so, I am using information on both complete and incomplete unemployment spells, thus improving the estimation of the unemployment duration.[9] As I assume normality of the error term in the main outcome equation (following the standard in the literature), I have chosen a parametric estimation in the first stage, namely a lognormal distribution to estimate an accelerate failure time hazard model (Kalbfleisch and Prentice, 2011), where the covariates are multiplicative on time, as expressed in the following equation:

$$Ln(t_i \exp(-z_i'\delta)) = \nu_i \tag{2.3}$$

where t_i is the elapsed unemployment duration, z_i' is a vector of covariates for individual i that include the instrumental variable, and $\nu_i = \sigma u_i$ is a normally distributed generalised error term (u_i is an error term and σ is a scale factor). Note that, even though this step is estimated in a non-linear fashion, the generalized residuals are still obtained in a linear way and can therefore be directly plugged in the second stage equation (2.2).

2.4.2 Accounting for Selection Bias

Even after addressing the simultaneity bias, one might be concerned with selection bias, which may arise in two ways – selection into the treatment and/or selection into re-employment. Selection into the treatment occurs if there are unobserved characteristics that lead individuals with longer duration spells to have larger impact (α_1) on re-employment wages. Selection into the re-employment occurs if the individuals that got a job, and for which the re-employment wage is observed, do not have the same

[8] A spell is censored on the last date the individual is observed in the database if the individual remains jobless by then.

[9] Incomplete spells are defined to be such that the end-date of the spell is not observed in the data.

(unobserved) characteristics of the individuals who remain unemployed.

To overcome the first type of selection, I follow Wooldridge (2015) and include an interaction term between the residuals of the first stage and the treatment variable (unemployment duration) in the second stage. Therefore, the main outcome equation to be estimated is:

$$Ln(PostW_i) = \alpha_0 + \alpha_1 Ln(UD_i) + \alpha_2 \hat{\nu}_i + \alpha_3 \hat{\nu}_i \times Ln(UD_i) + \boldsymbol{x_i'\beta} + u_i \tag{2.4}$$

If selection into treatment exists, then α_3 must be significant. In case the coefficient has the same sign as the estimated α_1, then individuals with longer duration have a larger impact of the unemployment duration on the re-employment wage.

To overcome the selection into employment I present two alternatives based on the same tool. As the sample is truncated at March of 2012 but the start of the unemployment spell might be anywhere between January 2005 and the sample end date, one should have in mind that the probability of observing a complete spell is not the same for all individuals.

First, I employ the traditional two-step approach proposed by Heckman (1976). The instrument I use to estimate the selection equation into employment (E_i) is the (log) duration of the time window ($Ln(T_i)$) in which the individual can be observed. Conditional on the covariates defined before, x_i, this variable should only affect the re-employment wage via the (expected positive) impact on the probability of observing the individual to get a job, given by γ_1 in equation 2.5

$$\Pr(E = 1|x_i) = \Phi(\gamma_0 + \gamma_1 Ln(T_i) + \boldsymbol{x_i'\beta}) = \Phi(z_i'\gamma) \tag{2.5}$$

In this set-up, the expected value of the (log) re-employment wage, conditional on the individual getting a job and on the values of x_i, is:

$$\begin{aligned}\mathbb{E}(Ln(PostW_i)|E_i = 1, x_i) &= \alpha_0 + \alpha_1 Ln(UD_i) + \alpha_2 \hat{\nu}_i + \boldsymbol{x_i'\beta} + \mathbb{E}(u_i|E_i = 1) \\ &= \alpha_0 + \alpha_1 Ln(UD_i) + \alpha_2 \hat{\nu}_i + \boldsymbol{x_i'\beta} + \alpha_\lambda \lambda(-z_i'\gamma) \\ &= \alpha_0 + \alpha_1 Ln(UD_i) + \alpha_2 \hat{\nu}_i + \boldsymbol{x_i'\beta} + \alpha_\lambda \frac{\phi(-z_i'\gamma)}{(1 - \Phi(-z_i'\gamma)}\end{aligned} \tag{2.6}$$

where λ is the inverse mills ratio evaluated at $z_i'\gamma$, i.e., and α_λ, to be estimated, is $\rho_{u\epsilon}\sigma_u$, i.e., the product between the correlation between the error of the outcome equation, u_i, and the error of the selection equation, ϵ_i, and the standard deviation of the error of the outcome equation (σ_u). Therefore, the sign of the α_λ indicates the sign of the correlation between the unobservables in the selection equation and the unobservables in the outcome equation.

Alternatively, I adopt the augmented inverse probability of censoring weighed estimation suggested by Rotnitzky et al. (1998). The idea is basically the same that is employed in the construction of sample surveys, i.e., to weight each contribution by the inverse probability of being observed. In this case I weight each observation by the probability of re-employment, i.e., by the probability that the complete duration of the unemployment spell is shorter than the total window for which I observe the individual, either unemployed or employed. As I do not observe the complete duration of the unemployment spell for some individuals (because their unemployment spells are censored), I use instead predicted unemployment duration ($\widehat{UD}$) from the first stage 2.3. The probability of interest is the following:

$$\text{Prob}\left(\widehat{Ln(UD)} < Ln(T_i)\right) = \Phi\left(\frac{\widehat{Ln(UD)} - Ln(T_i)}{\hat{\sigma}}\right) \tag{2.7}$$

where $\hat{\sigma}$ is the standard deviation parameter of each lognormal estimation in the first stage. Finally, I apply the inverse of this probability as weights in the re-employment wage equation, which is estimated by weighted least squares.

2.5 Identification

Here, I present three alternative instruments that are used to obtain an exogenous estimate of the unemployment duration in equation 2.3.

2.5.1 First Instrumental Variable: Potential Duration of Unemployment Benefits

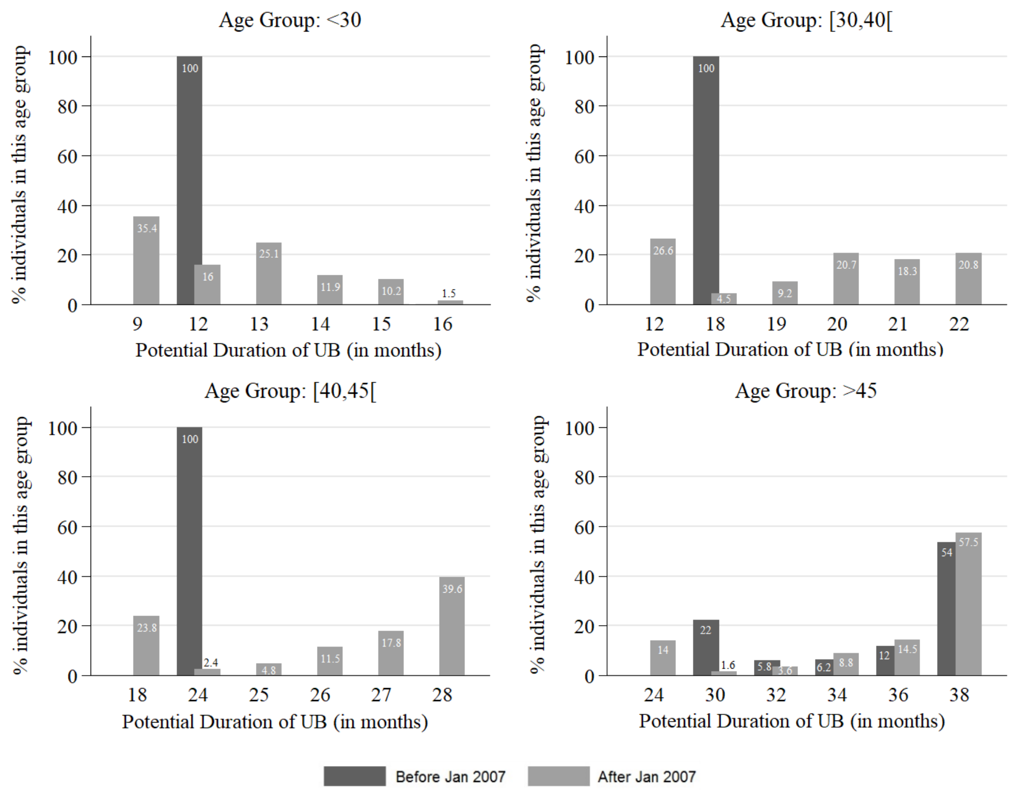
There is a vast literature reporting a strong (positive) relation between the potential duration of unemployment benefit and the joblessness duration.[10] Methodologies vary between the use of cross sectional variation – exploring age, work experience and regional differences – and the use of time series variation – exploring policy changes that affected the potential duration of unemployment benefits.[11] The sample used in this study includes both sources of variation. The former is obtained trough definition of potential duration of unemployment benefits, which is defined as a function of age and labour market experience. The latter is obtained through the 2007 policy change, which affected the relationship between the potential duration and its determinants.

To provide a better understanding of the actual change in the distribution of the potential duration of unemployment benefits, Figure 2.1 presents the distributions before and after the law change for the different age groups, according to the entitlement regulation.

[10] See Chapter 1 of for an extensive literature review. Recent evidence is provided by (Schmieder et al., 2012; Farber and Valletta, 2015; Le Barbanchon, 2016)

[11] See section 2.2 for a detailed description on the policy change.

Figure 2.1: Percentage of individuals by age group and potential duration of unemployment benefit, before and after the 2007 reform



In the top left panel, all the unemployed beneficiaries below 30 years old were entitled to 12 months of potential duration of unemployment benefit before 2007, and after the change in the rules, the distribution has spread between 9 and 16 months. Given the age of this group, it is not surprising that the individuals entitled to 9 months, because of few social security contributions, represents the largest share (35,4%). However, within this group, there is also a considerable proportion (47.2%) that was awarded with longer potential duration of unemployment benefits due to their previous work experience.

In the top right and the bottom left panels two patterns stand out. The percentage of beneficiaries entitled to the same or less potential duration as before 2007 decreases, relative to the first panel, and the percentage of individuals entitled to longer extensions increases with age. This reflects that labour market experience among individuals aged between 40 and 45 is larger than for those between 30 and 40. In both groups, around 70% of the beneficiaries were entitled to a longer potential duration in comparison to what they would have been offered if unemployed before January 2007.

Contrary to the previous age groups the last panel presents no major changes in what concerns the potential duration attributed before and after the law change. Before January 2007 this group was already entitled to 2 potential months of extension for each 5 years of work. However, the law change introduced a penalty to this age group. As a result, 14% of the individuals were only entitled to 24 instead of 30 months of potential duration of unemployment benefit, because the number of Social Security contributions between the previous and current spell of unemployment was less than 6 years.

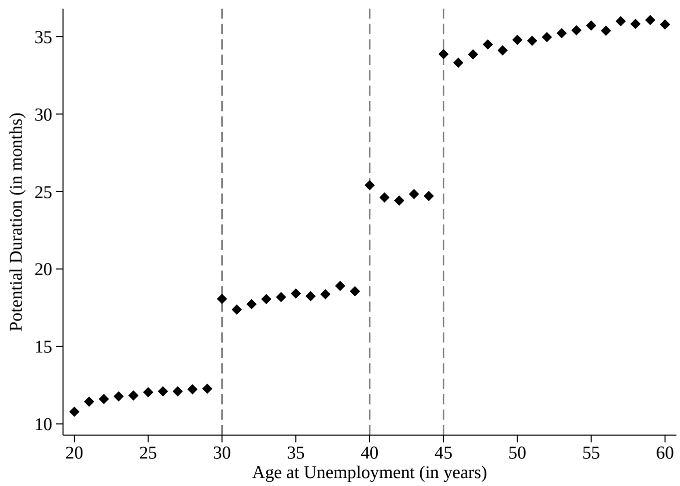
2.5.2 Second Instrumental Variable: Age Discontinuity in Potential Duration of Unemployment Benefit

As the figure in the previous subsection suggests, the age discontinuities are also an important source of exogenous cross-sectional variation in the potential duration of unemployment insurance. As argued in Addison and Portugal (2008a), one additional year of age, keeping everything else constant, does not make two individuals sufficiently different in terms of observed characteristics evaluated by the employer. Therefore, for two individuals with a one year age difference, but similar in all other dimensions, if the one that is entitled to a longer potential duration of unemployment benefits has a longer spell of unemployment, then the two spells must be related.

To illustrate these discontinuities in the potential duration of unemployment benefits for the individuals in the sample, I plot the average potential duration by age.[12]

[12] Note that, within the same age group, individuals might still be entitled to different potential duration of unemployment benefits as this also varies with work experience – see Table 2.1

Figure 2.2: Binscatter of average potential duration of unemployment benefits by age at unemployment



The discontinuities in the average potential duration between individuals of different consecutive age groups are substantial (between 5 and 8 months). To validate this instrument, I also test for the continuity of density on the number of individuals around each threshold, and whether sorting might be associated with pre-unemployment wages. Both figures,2.4 and 2.5 in the Appendix of this chapter, demonstrate no major concerns in both tests – in terms of continuity of the number of individuals and pre-unemployment wages around the age thresholds.

When using this instrument, I use only individuals whose age is around the threshold. The idea is to define a binary variable that assumes the value 1 if the individual is 30, 40 or 45 years old, and 0 if the individual is one year younger (29, 39 or 44 years old, respectively) at the date of unemployment. For the individuals, unemployed before January 2007, that are 29, 39 or 44 years old, one additional year of age increases the potential duration by six months. However, for the individuals unemployed after January 2007, one additional year of age may actually lead to no change in the potential duration of the unemployment benefit. This happens for three groups of individuals: i) 29 years old unemployed with work experience between 24 and 48 months; ii) 39 years old unemployed with work experience between 48 and 60 months; iii) 44 years old unemployed with work experience between 60 and 72 months. In all other cases, the individual could be entitled to at most 10 additional months of unemployment benefits, if the unemployment spell started after January 2007.

2.5.3 Third Instrumental Variable: Change in Potential Duration of Unemployment Benefit

Here, I use the change in the potential duration of unemployment benefit generated by the law of 2007. This policy did not change the age thresholds but rather the size of their discontinuities for different individuals, according to their work experience.

For each individual, I define *change* as the difference, in months, between the current potential duration of unemployment benefit and the one she would be entitled to, according to the rules in place before January 2007. Therefore, *change* is equal to 0 for all the individuals that entered the system before that month. For all the unemployed that entered the system after the reform, *change* may either assume positive, negative or a null value, according to their previous work experience. Table 2.4 shows that

Table 2.4: Change, in potential duration of unemployment benefits after 2007

Change	Individuals	%
-6	1959	13.95
-3	1259	8.97
0	5042	35.90
+1	1354	9.64
+2	1475	10.50
+3	1452	10.34
+4	1503	10.70

Notes: Changes are displayed in months for all individuals in the sample that became unemployed after January 2007. The third column displays the percentage of individuals affected by each change.

approximately 36% of the individuals, who became unemployed after January 2007, were not affected by the new rules. Additionally, about 23% were entitled to a shorter potential duration, as they have not reached the minimum period of contributions necessary, to be entitled to at least the same potential duration as before. Finally, roughly 41% has actually benefited from the institutional change applied due to their longer job histories.

The main goal of this alternative instrument is to set a clear cut between age and potential duration effects.[13] For the sake of example, take an individual that is eligible to 13 months. Using the definition in the first instrumental variable, the value of the instrument (13) automatically indicates the individual is 30 years old. Using the *change* definition, the instrument will assume the value "+1", which is the same value in case the individual would be entitled 19 or 25 months of potential duration, as well, therefore leaving uncertain to which age category the individual belongs. Therefore, the correlation between *change* and age should be much smaller than that of the first instrument (potential duration of unemployment benefits).

This is a convenient instrument because the individuals that benefited from the new rules have actually had longer spells of unemployment in comparison with those not affected by the reform.

Table 2.5: Average duration of unemployment benefits, by age group and difference between 2007 and 1999 potential duration of unemployment benefit rules

Age group	Negative change	No change in rules	Positive change
<30	9.64	13.71	13.87
[30,40[	10.79	15.17	16.73
[40,45[	13.22	16.53	18.13

Notes: Average duration is displayed in months and age displayed in years.

Table 2.5 shows that, for all age groups, the actual duration of unemployment

[13] The correlation between potential duration of unemployment benefits and age is 0.90 whereas that for this instrumental variable is -0.03. The correlation off both instrumental variables with work experience is 0.52 and 0.18, respectively.

benefits has followed the direction of the change in the rules. Individuals that have benefited from 1 to 4 additional months of unemployment have had an actual average duration that was 0.16 to 1.6 months longer, when compared to their age counterparts that were not entitled to any change. On the other hand, individuals with a penalty of -3 to -6 in potential duration of unemployment benefits have had an actual unemployment duration that was 3.31 to 4.38 months shorter, when compared to their age group counterparts.

2.6 Results

In the next subsection I present the OLS estimates for the re-employment wage equation and provide a brief discussion of the results. For robustness, I also estimate the equation on wage loss/gain, by restricting the coefficient on the previous wage equal to 1. By doing so, instead of having the log wage as the outcome variable in the main equation, I have the difference between the post- and the pre-log wages. In the remaining subsections, where I always use the re-employment wage as the outcome variable, I show the results of the control function approach, for each of the three instruments described in Section 2.5. First stages of the control function approach are provided in Table 2.15 in the Appendix of this chapter. Subsection 2.6.3 presents the estimates that also control for the different types of selectivity (as described in subsection 2.4.2). Estimates for the probit estimates are presented in Table 2.16 in the Appendix of this chapter.

2.6.1 OLS results

Table 2.6 provides the OLS estimates for the re-employment wage and loss/gain equations.

Table 2.6: Re-employment wage equations

	Full Sample		Restricted Sample	
Outcome: Ln(PostW)	(1)	(2)	(3)	(4)
Ln(Duration)	-0.059***	-0.056***	-0.069***	-0.062***
	(0.008)	(0.007)	(0.020)	(0.022)
Ln(Previous Wage)	0.567***	[1]	0.599***	[1]
	(0.012)		(0.036)	
Constant	2.915***	0.151**	2.519***	0.010
	(0.105)	(0.073)	(0.358)	(0.293)
R^2	0.278	0.058	0.320	0.056
N	8 423		1017	

Notes: standard errors in parenthesis below the coefficient estimates. The equations also include a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies. The second equation restricts the coefficient on the log of previous wage to be equal to unity, which is equivalent to estimate the effect of log duration on the difference of log wages.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

The first two columns include all individuals for which the post-unemployment wages is observed, whereas the last two columns are restricted to 29, 30, 39, 40, 44, or 45 years old individuals, for later comparison with the second instrument estimation. The coefficient estimate of column (1) shows that an increase of 10% in the joblessness duration leads to a decrease of around 0.6% in the post-unemployment wage. The negative effect is in line with the theories of skill depreciation and stigma, described in Section 2.1. The result does not change significantly when the pre-unemployment wage is restricted to 1 (columns 2 and 4), but is slightly larger, in absolute value, in the restricted sample.

2.6.2 Control Function Approach Results

Table 2.7 presents the control function approach results for the two variables of main interest in the outcome equation, i.e., the logarithm of joblessness duration

(Ln(Duration)) and the residual of the first stage ($\hat{\nu}_i$). Each column is associated with a different instrumental variable, defined in Section 2.5.[14]

Table 2.7: Re-employment wage equations using three different instrumental variables

Instrument in the first stage:	Potential Duration	Age Discontinuities	Δ Potential Duration
Outcome: Ln(PostW)	(1)	(2)	(3)
Ln(Duration)	-0.047***	-0.096***	-0.049***
	[0.010]	[0.027]	[0.013]
$\hat{\nu}_i$	-0.071***	0.054	-0.068***
	[0.020]	[0.068]	[0.028]
R^2	0.267	0.311	0.267
N	8 423	1 017	8 423

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equations also include the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

The results in columns (1) and (3) show that, if the unemployment duration increases by 10%, the re-employment wage is expected to decrease by 0.5%.[15] The larger estimate presented in column (2) should not be subject to direct comparisons with the other two columns, as the sample is substantially restricted. In fact, applying the first and third instruments in the restricted sample yields coefficients of -0.92% and -0.82%, respectively. The relative difference between the control function estimates and the OLS estimates is reflected in the sign of the coefficient associated to the generalized residual. When the coefficient of $\hat{\nu}_i$ positive, the coefficient of main interest becomes closer to zero, thus attenuating the effect found before. When the coefficient of $\hat{\nu}_i$ is negative, the negative effect of $Ln(Duration)$ on wages is even larger. The significance in the first and third instruments suggest there is evidence of endogeneity in that sample.

[14]The results for the first stage estimates using each of the three alternative instruments are presented in Appendix, in Table 2.15.

[15]Using the mean of joblessness duration in the sample, this means that each additional month implies an average decrease of 0.4% in re-employment wage

2.6.3 Control Function Results - accounting for self-selection

As discussed in Subsection 2.4.2, two types of self-selection might be present: into treatment, and in the observability of the outcome. To test for self-selection into treatment, each outcome equation includes the interaction between the logarithm of joblessness duration and the residual of the first stage. The first stage associated with each instrument is left unchanged relative to the previous Subsection. Table 2.8 presents the results associated with each instrument, accounting for this type of selection.

Table 2.8: Re-employment wage equations accounting for both endogeneity and self-selection into treatment

Instrument in the first stage:	Potential Duration	Age Discontinuities	Δ Potential Duration
Outcome: Ln(PostW)	(1)	(2)	(3)
Ln(Duration)	-0.048***	-0.071	-0.052***
	[0.013]	[0.053]	[0.013]
$\hat{\nu}_i$	0.056	-0.390	0.075
	[0.063]	[0.310]	[0.090]
$\hat{\nu}_i \times$ Ln(Duration)	-0.035**	0.107*	-0.037*
	[0.017]	[0.064]	[0.022]
R^2	0.268	0.310	0.267
N	8 423	1 017	8 423

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equations also include the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

The sign of the coefficient on the interaction term indicates the direction of the impact of selection on the effect of unemployment duration on re-employment wages. The estimates indicate evidence of negative selection into treatment in columns (1) and (3), which suggests that individuals with unobserved characteristics that lead to longer unemployment duration have a larger losses in the re-employment wage. In accordance, the coefficient associated to Ln(Duration) goes down, relative to Table 2.7. On the other hand, in column (2), where the interaction coefficient is positive, the coefficient associated to Ln(Duration) increases, thus attenuating the effect, relative to Table 2.7.

To test for self-selection into re-employment, I estimate two alternative models, as explained in Subsection 2.4.2. The first model is the conventional Heckit selection model (referenced as Heckit in the following tables), where the time window for which each individual can be observed is used as exclusion restriction. This instrument relies on the fact that the database ends exogenously at March 2012. Therefore, individuals that are firstly observed on a date that is close to the end of the database should be, on average, less likely to be observed in re-employment. The selection equation estimates are presented in the Appendix of this chapter, in Table 2.16, and the estimates for the main outcome equation are presented in Table 2.9.

The second model corresponds to the inverse probability (of censoring) weighted equation (referenced as IPW in the following tables). The weights are the inverse of the probability of censoring, which is computed from the first stage estimates that are kept unchanged. The estimates for the main outcome equation are also presented in Table 2.9.

Table 2.9: Re-employment wage equations accounting for endogeneity and self-selection into employment (Heckit two-step approach and augmented inverse probability of censoring weighted estimation)

Instrument:	Potential Duration		Age Discontinuities		Δ Potential Duration	
Outcome: Ln(PostW)	Heckit	IPW	Heckit	IPW	Heckit	IPW
Ln(Duration)	-0.045***	-0.047***	-0.102***	-0.081***	-0.047***	-0.048***
	[0.011]	[0.015]	[0.045]	[0.055]	[0.013]	[0.016]
$\hat{\nu}_i$	-0.077***	-0.046	0.069	0.034	-0.071***	-0.044
	[0.028]	[0.034]	[0.099]	[0.120]	[0.028]	[0.037]
$\hat{\lambda}$	-0.060		-0.166		-0.050	
	[0.041]		[0.130]		[0.043]	
N	8 423		1 017		8 423	

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equations also include the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

Accounting for selection, on top of endogeneity, does not change substantially the estimates on the coefficient of $Ln(Duration)$, relative to Table 2.7. In line with that, the coefficient on the inverse mill ratio (λ) is non significant in all three Heckit equations. In parallel, the IPW estimates on the first stage residual became non-significant.

Finally, in Table 2.10, the estimates account for both self-selection into unemployment duration and into re-employment. The sign and significance of the coefficient estimates of $\hat{\nu}_i$ and its interaction with the logarithm of unemployment duration are the same as in Table 2.8, with the exception of the non-significance of the later in the second instrument, i.e., even accounting for self-selection into re-employment, the results associated with the first and third instrument suggest that individuals with unobserved characteristics that lead to longer unemployment duration have a larger losses in the re-employment wage.

Table 2.10: Re-employment wage equations accounting for endogeneity,sef-selection into treatment, and self-selection into the observability of the outcome (Heckit two-step approach and augmented inverse probability of censoring weighted estimation)

Instrument	Potential Duration		Age Discontinuities		Δ Potential Duration	
Outcome: Ln(PostW)	Heckit	IPW	Heckit	IPW	Heckit	IPW
Ln(Duration)	-0.045***	-0.047***	-0.057	-0.023	-0.050***	-0.053***
	[0.013]	[0.015]	[0.064]	[0.051]	[0.012]	[0.016]
$\hat{\nu}_i$	0.041	0.008	-0.502	-0.473	0.066	0.041
	[0.062]	[0.087]	[0.425]	[0.337]	[0.071]	[0.093]
$\hat{\nu}_i\times$ Ln(Duration)	-0.032*	-0.015	0.129	0.109	-0.035**	-0.021
	[0.017]	[0.022]	[0.129]	[0.077]	[0.017]	[0.023]
$\hat{\lambda}$	-0.051		-0.100		-0.044	
	[0.043]		[0.174]		[0.048]	
N	8 423		1 017		8 423	

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equations also include the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

Like in Table 2.9, the alternative estimation methods (Heckit and IPW) do not significantly change the estimates on the impact of unemployment duration on re-employment wages for the first and third instruments. However, with the age discontinuity instrument, controlling for both types of selection yields a coefficient estimate that is remarkably closer to the one obtained in the other two instruments.

2.7 Robustness Analysis

For purposes of robustness, I assess the sensitivity of the results along four dimensions: period of the analysis, pre-unemployment wage quartiles, transition to a different industry, and the definition of re-employment wages. Firstly, I check whether the effect of joblessness duration on re-employment wage varies across time in two ways: 1) by excluding the economic crisis and 2), by narrowing it to the period around of the policy change in potential duration of unemployment benefits. Secondly, I test for the robustness of the results with respect to different pre-unemployment wage quartiles. Thirdly, I test for heterogeneous effects in individuals that have been re-employed in a different industry. Finally, I analyse the impact of joblessness duration in the re-employment wage earned one, three, and five years after the date of re-employment.

For the sake of brevity, all the estimates presented in this section use the third instrumental variable, i.e., the change in the potential duration of unemployment benefits generated by the policy change in 2007. This is my preferred estimate, both for the methodological reasons presented in Subsection 2.5.3, and for the stability of the results presented in the previous section.

2.7.1 Period of the Analysis

A major concern regarding the main estimates could be the fact that the period in analysis includes the most recent economic crisis in Portugal, which generated *cata-*

strophic job destruction (Carneiro et al., 2014). [16]

To test whether the wage losses are explained by the increase in dismissals after 2009, I perform the same analysis for the individuals dismissed during the period 2005-2008, thus excluding the most critical period of the economic crisis. Additionally, to ensure that the effect of joblessness duration on re-employment wage is being identified by the reform, I also conduct the same exercise for the individuals in 2006-2007. The results for both tests are shown in Table 2.11.

Table 2.11: Re-employment wage equations, by year of unemployment

	2005-2008	2006-2007
Variable	(1)	(2)
Ln(Duration)	-0.048***	-0.064***
	[0.012]	[0.018]
$\hat{\nu}_i$	-0.053**	-0.054
	[0.025]	[0.033]
R^2	0.310	0.311
N	6 478	3 992

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equations also include the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

The estimates presented in the first column of Table 2.11 do not significantly differ from those reported in Table 2.7. Hence, I conclude the inclusion of the years 2009-2012 (March) are not driving the results. When I concentrate the period of the analysis to the years before and after the implementation of the reform, the effect of unemployment duration on wages is even larger than the one presented in Table 2.7. This suggests that, if something, the main results may be interpreted as a lower bound, which reassures the validity of the third instrumental variable.

[16] According to Couch and Placzek (2010) and Couch et al. (2011), earnings losses of displaced workers are larger during recessionary times.

2.7.2 Distribution of Pre-unemployment Wages

In this subsection I investigate if there is a particular portion of the pre-unemployment wage distribution for which the effect of unemployment duration on re-employment wages is larger. To do so, I interact the duration effect with each quartile of the pre-unemployment wage distribution. Table 2.12 shows that the impact of joblessness duration on the re-employment wage gets wider along the pre-unemployment wage distribution, that is, unemployed individuals with larger pre-unemployment wages get larger wages losses in re-employment.

Table 2.12: Re-employment wage equation, by pre-unemployment wage quartiles

	Second Stage
Variable	(1)
Ln(Duration)	-0.005
	[0.014]
Ln(Duration) $\times \mathbb{1}$ (506€$<$ Previous Wage $\leq$ 613 €)	-0.006
	[0.015]
Ln(Duration) $\times \mathbb{1}$ (613€$<$ Previous Wage $\leq$ 841 €)	-0.051***
	[0.014]
Ln(Duration) $\times \mathbb{1}$ (Previous Wage $>$ 841 €)	-0.104***
	[0.015]
$\hat{\nu}_i$	-0.073***
	[0.024]
$\mathbb{1}$ (506 €$<$ Previous Wage $\leq$ 613 €)	0.081***
	[0.031]
$\mathbb{1}$ (613 €$<$ Previous Wage $\leq$ 841 €)	0.268***
	[0.028]
$\mathbb{1}$ (Previous Wage $>$ 841€)	0.749***
	[0.033]
R^2	0.221
N	8 423

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equation also includes the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

This result may be explained by two characteristics of the labour market in Portugal.

The existence of a national minimum wage explains the non-negative sign observed for the first quartile. The fact that this interaction is not significant is consistent with Centeno and Novo (2014), which found that low-wage workers tend to react less to benefit extensions. On the other hand, the evidence for nominal wage rigidity in Portugal (Blanchard, 2007; Portugal, 2006; Raposo et al., 2019) supports the fact that workers with higher pre-unemployment wages face larger wage losses after being unemployed. As workers are fired with wages above their productivity, the wage loss after displacement is higher than expected.[17] A similar result was found by Burda and Mertens (2001) for Germany.

2.7.3 Job Change Between Industries

As argued in the theoretical literature, part of the decrease in wages after unemployment might be related to the loss of human capital, and in particular, to the loss of industry specific human capital (Neal, 1995). In the re-employed sample used in here, 63% of the individuals are observed to move to a different industry. To test whether this is driving the main results, I test whether the effect of $Ln(duration)$ changes for those whose re-employment job is in a different industry, by interacting this variable with a dummy which is equal to one if the individual transited to another industry. The results are presented in Table 2.13.

The estimates suggest lower re-employment wages for individuals that move to another industry after the unemployment episode but no statistically different losses due to longer spells of unemployment. Given that the change in industry may be endogenous, I follow the suggestion of Addison and Portugal (1989) and test whether this difference varied by tenure in the previous firm (as a proxy for tenure in the previous industry) and no statistically difference is found.

[17] See Shapiro and Stiglitz (1984) for a theoretical explanation.

Table 2.13: Re-employment wage equation, by pre-unemployment industry movement

	Second Stage
Variable	(1)
Ln(Duration)	-0.046***
	[0.012]
Ln(Duration) $\times\mathbb{1}$ (New industry $\neq$ Last industry)	0.001
	[0.011]
$\hat{\nu}_i$	-0.069***
	[0.023]
$\mathbb{1}$ (New industry $\neq$ Last industry)	-0.047**
	[0.024]
R^2	0.268
N	8 423

Notes: bootstrapped standard errors in square brackets below the coefficient estimates. The equation also includes the log of previous wage, a dummy for gender, age dummies, a quadratic function of tenure, dummies for the reason of unemployment, seven year dummies and three quarter dummies, according to the date of re-employment, industry dummies, and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

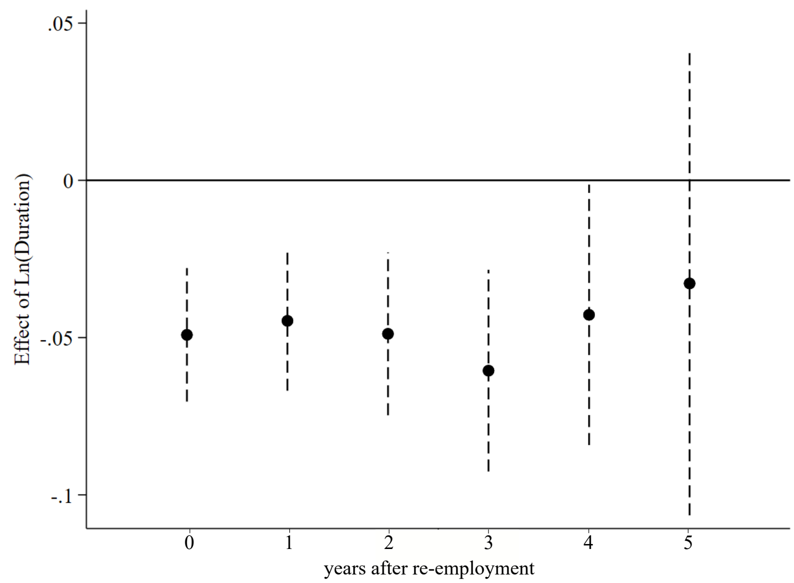
2.7.4 Wage Loss Persistence

Finally, I also estimate the same specification for wages along the first 5 years in re-employment. By doing so, I want to test the persistence of the negative impact of joblessness duration on re-employment wages.

Figure 2.3 shows that the negative effect of unemployment duration on wages is persistent until, at least, the fourth year after re-employment. In the fifth year the estimate is still negative, but the drop in observations due to the relatively short database time window implies much larger confidence intervals.

One of the possible explanations for this persistence could be the relatively short tenure in the first re-employment job. In case the recently re-employed fall into unemployment again or transit to another job, any returns to job-specific tenure are lost. In fact, I observe that roughly 45% of the re-employed face a new unemployment spell in the first six years after re-employment. Moreover, after one year of re-employment only

Figure 2.3: Percentage of individuals by age group and potential duration of unemployment benefit, before and after the 2007 reform



37% of the workers remain in the same company that hired them after unemployment. However, I find no evidence for the relationship between unemployment duration and re-employment tenure, which is consistent with the results found by Nekoei and Weber (2017).

2.8 Conclusions

A broad range of methodologies have been applied to estimate the causal effect of unemployment duration on re-employment wages. Without information on the reservation wage (which is the case when using administrative data), one cannot directly identify whether the re-employment wage is being determined by the unemployment duration (through stigma or skill depreciation), or the re-employment wage is actually determining the unemployment duration. This study adds two main contributions to this literature.

First, I take advantage of an institutional reform that changed the potential duration of unemployment benefits, in January of 2007, in Portugal, to adopt two alternative instruments to predict the unemployment duration. While the age discontinuity in entitlement to different potential duration of unemployment benefits has already been explored in the literature, little is known about a policy change that created heterogeneous effects in the entitlement.

The first alternative instrument is the potential duration of unemployment benefits. While this variable yields a strong positive effect on the joblessness duration, it is strongly correlated with age and work experience. I overcome this problem by using the change in the potential duration of unemployment benefits between the the pre- and post- reform period (second alternative instrument). Whereas this kind of policies usually rely on a general increase (or a general decrease) of the potential duration of the beneficiaries, this policy contains a unique feature. Around 41% of the individuals have benefited from longer potential duration of unemployment benefits, 23% of the individuals were entitled to smaller potential duration of unemployment benefits, and approximately 36% of the individuals were not affected by the reform. I also perform the same analysis using the age discontinuities present in the rules of unemployment benefits for purposes of comparison with other studies.

The second main contribution is methodological and related to the estimation procedure, which incorporates three main features: 1) the use of control function approach, which accounts for endogeneity in the outcome equation, with a non-linear first stage 2) the consideration of the information on the incomplete spells of unemployment in the first stage of the estimation, which allows me to get more accurate estimates on the predicted duration 3) the use of the probability of re-employment to account for selectivity in the results, which is employed in the inverse probability of censoring weighted estimation suggested by Rotnitzky et al. (1998).

The main estimates reveal that each additional month of unemployment duration leads to a statistically significant decrease of approximately 0.5% in re-employment wages. The estimates are robust to the different instruments in use, and also when self-selection, in both unemployment duration and in the re-employment wages, is accounted for.

The main result is in line with Addison and Portugal (1989) and Schmieder et al. (2016), even though at a smaller scale. Such difference may be associated with the decline in the responsiveness of real wages to unemployment changes over the last decade (Robalo Marques et al., 2010).

The contributions of this chapter are not restricted to the estimate of the causal effect of unemployment duration on re-employment wages. Actually, they suggest important policy implications. As pointed in Nekoei and Weber (2017), a positive effect of unemployment duration in re-employment wages would suggest that unemployment insurance subsidizes search and not just leisure, which is not the case here. Instead, the negative impact of unemployment duration on re-employment wages seems suggests that additional job search assistance programs may be justified in Portugal (Raposo et al., 2019).

Appendix of Chapter 2

Table 2.14: Descriptive Statistics - Re-employed Sample

	Mean	Std.D.	Min	Median	Max
Benefit Duration (UI) (in months)	7.94	7.34	1	5.77	38
Jobless Duration (in months)	11.83	11.77	2	7	75
Level of benefits (€/month)	483.48	193.43	65.10	413.7	1257.60
Previous wage (€/month)	746.01	512.67	116.99	607.18	10463.61
re-employment wage (€/ month)	575.23	359.75	10.57	495.23	5602.57
Age at re-employment (in years)	34.36	9.49	17	33	66
Male	0.49	0.50	0	0	1
Sample			8 423		

B2.1 Instruments validity

Figure 2.4: Validity of second instrumental variable - continuity of density

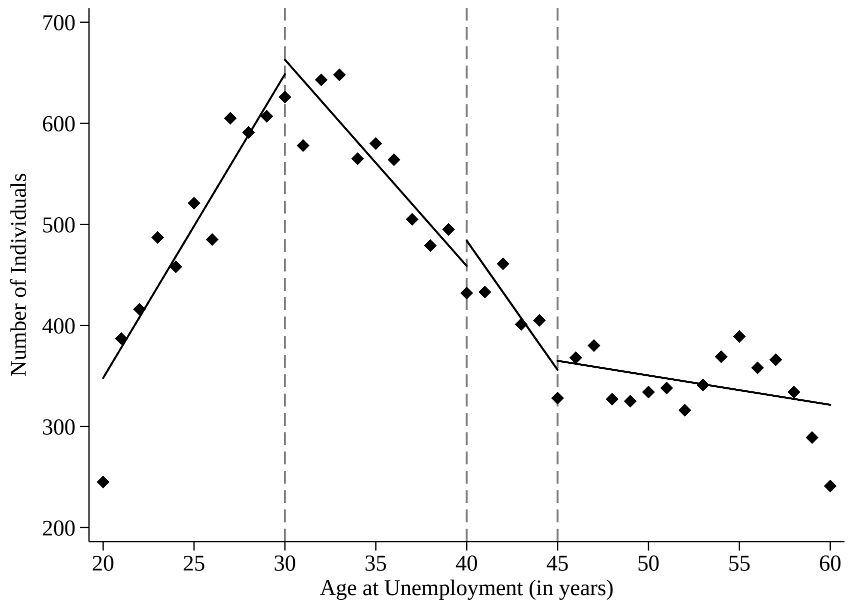
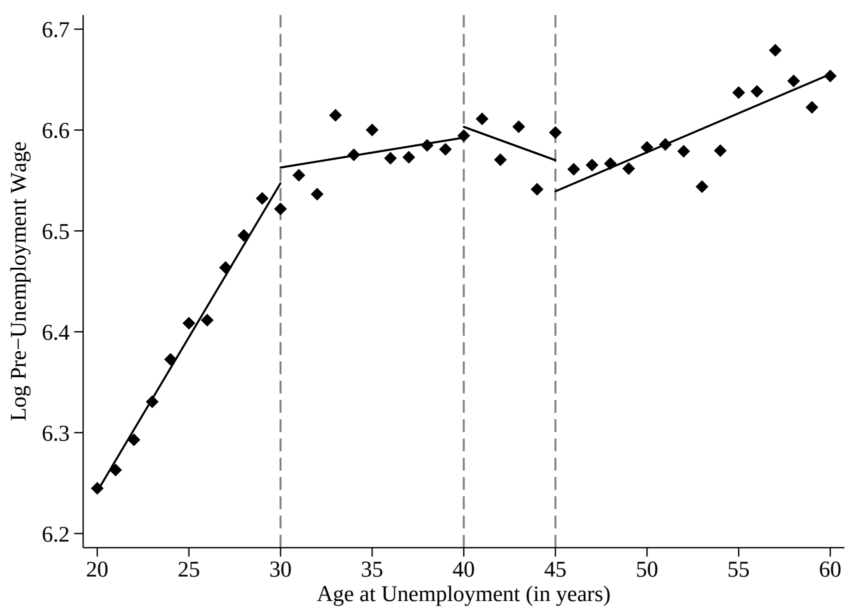


Figure 2.5: Validity of second instrumental variable - pre-unemployment wages



B2.2 First Stage Estimates

Table 2.15: Accelerated Failure Time Unemployment Duration Equations

Variable	Strategy (1)		(2)		(3)	
Ln(Potential Duration)	1.719***	(.067)				
Old			.170**	(.084)		
Difference in Potential Rules					.089***	(.0002)
Ln(Previous Wage)	-.130***	(.030)	-.019	(.032)	-.116***	(.029)
Male	-.244***	(.026)	-.229***	(.027)	-.261***	(.025)
Age (groups)						
[30,40[	-.399***	(.040)			.284***	(.004)
[40,45[	-.608***	(.063)			.592***	(.004)
≥45	-.393***	(.077)			1.449***	(.004)
Age (discontinuities)						
29 or 30			-.427***	(.052)		
39 or 40			-.100	(.063)		
44 or 45			.015	(.073)		
Reasons of unemployment						
expiration/extinction	-.126**	(.053)	-.255***	(.057)	-.155***	(.054)
employer initiative	.222***	(.055)	.198***	(.059)	.198***	(.056)
termination agreement	.425***	(.061)	.472***	(.065)	.421***	(.062)
Tenure	.047***	(.006)	.095***	(.006)	.049***	(.006)
Tenure^2	-.0006	(.000)	-.0010	(.0002)	-.0006	(.0002)
Unemployment Rate	.260***	(.009)	.327***	(.009)	.266***	(.009)
Constant	-3.127***	(.261)	.200	(.227)	.991***	(.216)
Log likelihood	-18		-19		-18	
	341.716		355.121		480.659	
N	18 543		2 088		18 543	

Notes: standard errors in parenthesis next to the estimates. The equations also include industry and region dummies

*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

B2.3 Selection Equations from Heckman Two-Step Estimation

Table 2.16: Selection Equations from Heckman Two-step Estimation

	Specification					
Variable	(1)		(2)		(3)	
ln(T)	.017***	[.001]	.012***	[.003]	.017***	[.001]
ln(Potential Duration)	-.999***	[.059]	-.594***	[.150]	-.999***	[.061]
ln(Previous Wage)	.052**	[.025]	.066	[.077]	.052**	[.024]
Male	.173***	[.024]	.126*	[.076]	.173***	[.027]
Age (groups)						
[30,40[	.088**	[.038]			.088**	[.033]
[40,45[	.136**	[.052]			.136**	[.054]
≥45	-.169**	[.070]			-.169**	[.070]
Age (discontinuities)						
30			.160	[.124]		
39			-.044	[.126]		
40			-.048	[.184]		
44			-.169	[.142]		
45			-.001	[.209]		
Reasons of unemployment						
expiration/extinction	.198***	[.053]	.102	[.162]	.198***	[.059]
employer initiative	-.051	[.051]	-.176	[.147]	-.051	[.064]
termination agreement	-.174***	[.053]	-.108	[.163]	-.174***	[.064]
Tenure	-.035***	[.005]	-.039**	[.018]	-.035***	[.005]
Tenure^2	.000**	[.000]	.002**	[.001]	.000**	[.000]
Unemployment Rate	-.247***	[.013]	-.288***	[.040]	-.247***	[.014]
Constant	4.127***	[.294]	3.643***	[.810]	4.127***	[.263]
F-Stat	455.077		67.268		1006.740	
N	18 543		2 088		18 543	

Notes: standard errors in parenthesis next to the estimates. The equations also include industry and region dummies.
*** Significant at 1% level ** Significant at 5% level * Significant at 10% level

Chapter 3

Job Security and Fertility Decisions

3.1 Introduction

Technology change and globalization have motivated the flexibility of labour markets during the last decades. New forms of employment have been created, such as casual work, task-based contracts, and fixed-term contracts. These jobs are characterised by a higher degree of insecurity as the worker is not insured against lay-off at the end of the contract. Temporary employment is more common among young workers, until they learn about the worker-firm match quality, and women, who are more likely to move to non-market employment than men (Lazear and Rosen, 1990; Booth et al., 2002b). Therefore, job security must be taken into account when young women make investment decisions, especially when they are irreversible, such as fertility choices (Becker, 1960).

The primary research question of this chapter is: How does job security affect fertility decisions? Theoretically, there are three fundamental channels through which one might expect job insecurity could reduce fertility. First, if re-employment is more difficult, working on a contract with higher probability of dismissal means higher volatility in expected future earnings, which are crucial to support childcare costs over the life-cycle. Second, in case of wage differentials that compensate temporary labour for the lack of job security, women working on temporary contracts have a higher opportunity cost when substituting work for child-rearing. Finally, in case temporary arrangements are mainly used as screening devices prior to more secure jobs, women might delay their fertility decisions while they keep working on fixed-term contracts to accumulate tenure until promotion.

The empirical challenge is to distinguish whether a woman is working on a temporary job because she does not plan to give birth or whether she does not give birth because she holds a highly insecure job. To disentangle the two effects, I build and estimate a structural life-cycle model where women choose both fertility and labour supply, conditional on job security. The degree of job security corresponds to two different types of job contracts. Fixed-term contracts have a limited duration, a maximum number of renewals, and high probability of firing. Permanent contracts are open-ended but have a low probability of dismissal. This type of duality characterises the labour market in many European countries, where young women on fixed-term contracts account for almost half of the young female employees.[1]

The use of fixed-term contracts has increased significantly in many European countries to tackle the high unemployment rates in the end of the 1970s. In a contrasting policy direction, the employment protection of permanent contracts was reinforced to avoid job destruction (Cahuc and Postel-Vinay, 2002). Both measures resulted in an excessive use of fixed-term contracts as the conversion rate into permanent contracts never exceeds 50% (Boeri et al., 2011). In this chapter I am going to focus on a particular country. Portugal is an interesting case as it has the lowest rate of contract conversion (12%) among European countries and it has been classified as the strictest

[1]According to the Eurostat statistics, on average, 53% of the young (15-24 years old) women in the Euro area held temporary contracts in 2018. Slovenia ranked highest with 81.1% and high rates were also observed outside the Euro area in countries such as Croatia, with 67.8%.

country in terms of employment protection for permanent contracts.[2] In the context of the research question in this chapter, Portugal is also a relevant example as it only provides 6 weeks of mandatory paid maternity leave.[3] Optional benefits paid by Social Security are available but, in 2017, only 9% of the mothers opted for the extended maternity leave, according to the Social Security statistics.[4] In line with this, Portugal has the second highest employment rate among mothers with at least one child between 0 and 2 years old among OECD counties (73.2%).[5]

Considering the institutional background, I abstract from maternity leave policies and focus on the link between job security and fertility decisions. Using survival analysis, I provide preliminary evidence that fixed-term contracts are an important determinant of the delay of fertility choice. This effect is especially important for the first birth decision whereas the income/time trade-off (Becker, 1965) seems to be more relevant for the second birth decision. However, one should keep in mind the simultaneity issue, between fertility decisions and the type of contract accepted, and be cautious not to interpret these results as causal. Ideally, one would need to have exogenous variation on factors that affect women working on fixed-term contracts but does not affect the fertility decision directly. The case of Portugal allows me to study such variation.

In this chapter I explore a 2003 policy reform that increased the potential accumulated duration of fixed-term contracts within the same firm from 3 to 6 years. The reform induced a higher probability of staying in the same firm on a fixed-term contract, but the conversion rate into permanent contracts did not change significantly

[2]Classification attributed according to the OECD indicator, published in various *Employment Outlook* reports. See OECD (2018) for the latest reference.

[3]Average duration of mandatory paid maternity leave among OECD countries is 19 weeks. Australia also provide only 6 weeks of paid maternity leave, but women are compensated with additional 6 weeks of pre-birth leave. Most of the states in United States have no mandatory paid maternity leave.

[4]Optional benefits are available in two-tiers. The initial maternity leave benefit is 4 months and, since 2010, the extended maternity leave benefit allows women not to work for an additional 3 months.

[5]The country leading this ranking is Netherlands, with 73.5%, and the average among OECD countries is 53.9%. Note however that a big part of the remaining percentage is out of the labour force whereas inactivity rates in Portugal are lower than the average among OECD countries.

between the pre- and post-reform periods. Therefore, this reform extended the period of job insecurity. In a reduced-form setting, I show that women on a contract under the new legislation had a lower probability of giving birth. However, this negative effect was attenuated for women working on fixed-term contracts with longer duration, as these have higher probability of conversion into permanent contracts. For validation purposes, I estimate the life-cycle model with pre-reform data and use the post-reform labour market conditions to simulate the implications of the reform. The model is able to reproduce the reduced-form results, confirming the conversion into a permanent contract to be an important channel affecting fertility decisions.

To explore the other two channels, I examine the application of two policy reforms that have been discussed in the public debate: a single contract with long probation period, and contract-specific tax rates. For the channel on future earnings expectations I implement a single contract where fixed-term workers are automatically converted into a permanent contract in case there is no dismissal during the probation period. This reform eliminates the dismissal spike at the end of the accumulated duration of fixed-term contracts and doubles the conversion rate. The model predicts an increase of the fertility rate by shifting births to periods where expected future earnings were less volatile (end of fixed-term contract and permanent contracts), and by decreasing the number of childless mothers. This result reinforces the reduced-form evidence that job security is especially important for the decision to have a first child. Note however that such reform would bring additional costs to the firms, which in turn could adjust the labour demand. For this purpose, I consider three alternative scenarios where there is a decrease of the permanent contract firing costs, a decrease in the wage seniority bonus, or a decrease on the hiring probability. In general, I conclude that the positive effect on fertility could be attenuated by each of these three adjustments, but large modifications would be required in order to bring it down to zero.

To test the wage differential channel I introduce contract-specific tax rates. Taxing fixed-term contract wages would have a nearly null effect on fertility rate, but would generate a shift from fewer children conceived during fixed-term contracts towards more children conceived during non-employment. This result sheds light on the hypothesis that women working on fixed-term contracts are less prone to give birth as the opportunity cost of child-rearing is higher when this type of contracts pay better to compensate for the lack of job security. For completeness, I also test the implementation of a credit on permanent contract wages together with the tax in fixed-term contract wages. The results show a considerable increase in fertility that is essentially generated by a compositional change of the number of children in the household. This policy reform does not change the number of mothers in the economy but rather increases the number of mothers with two children, thus indicating that more income helps to compensate for the lack of time at home in order to raise multiple children.

Literature Review This chapter contributes to four distinct parts of the literature. The major contribution of this chapter is that it adds to the scarce literature that relates job security and fertility decisions. Adsera (2004) was the first, to the best of my knowledge, to establish a relationship between unstable contracts and low fertility in Southern Europe. In a two-period theoretical model the author shows that larger uncertainty in income, from either unemployment or marginal employment arrangements, transforms the fertility decision into a risky and costly choice.[6] In a reduced-form setting De la Rica and Iza (2005) found that fixed-term contracts for women in Spain delay the motherhood decision although the same seems not to be relevant for men. Auer and Danzer (2016) found an identical result for Germany, but both studies claim the results cannot be interpreted as causal as they lack exogenous variation on fixed-term

[6] Outside the context of job security, but also accounting for idiosyncratic uninsurable earnings risk, Sommer (2016) concluded that increases in earnings risk are associated with reduction on the number of children and the delay of fertility decisions.

contracts. This chapter adds to this literature by exploiting exogenous variation in the rules of fixed-term contracts, by simulating additional policy reforms, and by using a country that has not been studied in this context before.

Second, this chapter contributes to the literature using structural life-cycle models, which estimate the impact of different job characteristics on both fertility and labour supply decisions. Moffitt (1984) was the first to estimate a dynamic model of discrete choice where both fertility and wage profiles are endogenous over the life-cycle. The results revealed that shifts in the level of the lifetime wage profile are associated with both lifetime profiles of fertility rates and female employment rates. Building on this outcome, Blackburn et al. (1993) introduced investment in human capital and showed that this investment is proportionally related to the age of first birth. Francesconi (2002) contributed to the literature by distinguishing between part-time and full-time employment. When comparing lifetime utilities between recent mothers in part-time jobs with those that interrupted their career, the author found no substantial differences.

More recently new features of the labour market have been introduced in this literature. Edwards (2014) explored the impact of *flexitime* in the reduction of career interruptions related to fertility. The author found that flexible hours schedule is more valuable to women with children and even more so to women with infant children. Adda et al. (2017) incorporated occupational choices in the model and allowed skill atrophy to be occupation specific. The authors concluded that women in abstract occupations face higher atrophy rates, thus higher opportunity costs of not working. Therefore, women in these occupations tend to have children at a later stage than women in routine occupations. In a structural model, only Guner et al. (2018) and my work have included different types of job contract in a life-cycle analysis of fertility and labour supply choices. Our papers differ in many ways, the main difference being that I account for the duration and renewals of fixed-term contracts in the model. This allows me, not only to approximate the model to reality, but also to study the impact of different policies that affected these two rules over time.

Third, and more generally, this chapter also contributes to the literature that studies the impact of family policies on fertility and labour supply. Gauthier (2007) and Olivetti and Petrongolo (2017) provided an extensive literature review on parental leave, subsidised childcare, and in-work benefits. In a cross-country comparison, Ruhm (1998) and Thévenon and Luci (2012) concluded that paid parental leave can increase female employment in 3 to 4 percent, but the result was not always positive to every country. Schönberg and Ludsteck (2014) and Lalive and Zweimüller (2009) found a negative effect in maternal labour market attachment in the short-run, but a rather small impact in the medium run. The latter study has also built a structural model to simulate the combination of parental leave and subsidised childcare and concluded that this policy could actually increase maternal labour market attachment in the medium-run. Sheran (2007) also emphasised the reduction of childcare costs as generating positive impacts on both employment and fertility. Finally, Erosa et al. (2010) and Haan and Wrohlich (2011) studied the impact of tax-credits on labour income and concluded that this instrument was also beneficial for both outcomes. This chapter provides an additional policy instrument, job security, that although not directly targeting family, it generates positive effects on both fertility and maternal employment.

Finally, this work is also related to the recent literature on the study of simulated policy changes to the dual system present in Europe. Pérez and Osuna (2014) analysed the introduction of a single-contract in Spain by eliminating the duality between fixed-term and permanent contracts and concluded that this policy would be more effective in reducing unemployment and job destruction than the decrease in severance payments. Cahuc et al. (2016) contrasted the decrease in severance payments in open-ended contracts with the taxation of temporary jobs in France and concluded that the second reform would generate opposite effects, by decreasing employment and reducing job creation. Sestito and Viviano (2016) evaluated the reforms of the Jobs Act in Italy that provided a subsidy on permanent hiring and a decrease in firing costs. The authors found that both measures shifted employment towards more permanent contracts and raised overall employment rates. Dolado et al. (2018) simulated a unified system of employment protection legislation in Spain and concluded that the welfare effects would be heterogeneous across different types of workers, being the older workers the ones with largest losses. This chapter adds to this literature by discussing both the introduction of a single-contract with and without adjustments on the labour demand, and also the taxation of fixed-term contracts, in Portugal.

The remainder of the chapter is organised as follows. Section 3.2 provides an overview of the job contracts legislation and describes the incidence of fixed-term contracts in Portugal. Section 3.3 presents the data used in both reduced-form analysis and in the model estimation. Section 3.4 provides some preliminary evidence with a survival analysis, and evaluates the policy reform that affected the rules of fixed-term contracts in Portugal, after 2003. Section 3.5 describes the life-cycle dynamic discrete choice structural model. Section 3.6 is devoted to the estimation and identification issues. Section 3.7 presents the results, which include the parameters estimates, the fit of the model and an experiment for the model validation. Section 3.8 is dedicated to the policy counterfactuals. Finally, Section 3.9 concludes.

3.2 Institutional Background

3.2.1 Legislation of Fixed-term Contracts in Portugal

The regulation of employment contracts in Portugal dates back to 1937, but it was only in 1976 that fixed-term contracts (FTCs) became legislated. As explicit regulation on dismissals was implemented in that same year, this type of contract, with low firing costs, became an attractive screening device to the firms. Back then, there was no limit on the maximum number of renewals for the FTC. However, the period for which the worker could be employed for the same firm was limited to three years. Figure 3.1 summarises the legislation of FTC in terms of renewals and cumulative duration since 1976.[7]

Figure 3.1: Rules of FTC over time: maximum number of renewals and cumulative duration

	1976–1989	1989–2003	2003–2009	2009–2012
maximum number of renewals	no limit	2	2	3
maximum cumulative duration	3 years	3 years	6 years	3 years

Back in the 1980s, when the maximum number of renewals was unlimited, the FTC could either be a succession of short-term contracts or a three years single contract. Once the probation period was exhausted, the employer was obliged to convert the contract into a permanent contract (PC) in order to keep the worker in the same firm. Both renewals of FTC and conversions of FTC onto PC were automatic unless the employer explicitly communicated otherwise within a minimum of eight days before the end of the contract.

In 1989, after the integration of Portugal in the Economic European Community, the maximum number of renewals of FTC was capped at two, without changing the

[7]The timeline finishes in 2012 as the legislation that was introduced after it is not straightforward to put in such a simple diagram, but for most of the workers in FTC, the legislation in place before 2012 still applies.

maximum duration of three years in the same firm. The period for the communication of non-renewal was extended to 15 days in cases of contracts with less than six months of duration and to 30 days in case of longer contracts. Short severance payments were also introduced in a ratio of two days for each month of FTC.

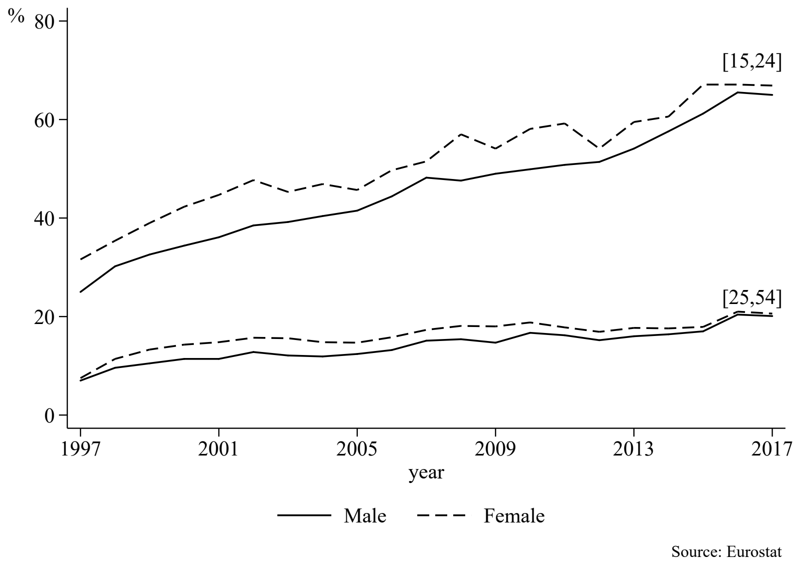
Before 2001, the replacement of a worker in a FTC job position required three months of interval between the firing date of the former worker and the hiring date of the new one. After 2001 the waiting period increased to six months; the severance payments have increased to three days for each month of the contract; and the minimum number of days for the communication of non-renewal decreased to five days only.

In 2003 the maximum number of renewals did not change, but the maximum duration of a FTC within the same firm was extended for up to six years, with the remark that the last renewal should have a duration between one and three years. More recently, in 2009, the maximum number of renewals was increased to three times while the maximum duration with renewals has reversed to three years, like in pre-2003.

3.2.2 Fixed-term Contracts Over Time, Gender, and Age

Portugal and Varejão (2009) showed that, in the context of the Portuguese dual labour market, firms look at the fixed-term contracts as an attractive device to screen workers with low firing costs. For that reason, the share of employees with FTC has always been relatively large (22% in 2017), especially when compared to other European countries (average 14.3% in 2017). This disparity is even larger for younger workers. In 2017, almost two-thirds of employees between 15 and 24 years old, in Portugal, had a fixed-term contract whereas the European average was only 44%. Figure 3.2 illustrates these trends by gender and age group during the last two decades.

Figure 3.2: Percentage of fixed-term contract employees, by year, gender and age group



According to Weiss and Gronau (1981), during the early stage of their careers, women are still deciding between market and home production. Therefore, women are more likely to defer investment in specific human capital. Under these conditions firms are "gender blind" at the hiring stage, but men are more likely to receive a promotion offer as their specific human capital grows faster (Lazear and Rosen, 1990). In other words, men with longer duration of fixed-term contract jobs may be perceived to have lower ability than women in the same position (Booth et al., 2002b).

3.3 Data

In this section I describe the two main sources of data used in this chapter. Most of the analyses rely on data from European Community Household Panel, presented in the first subsection, but in order to get more accurate estimates in terms of labour market transitions between different job contracts, I have also used the Portuguese

Labour Force Survey, presented in the second subsection. This section includes details on sample construction as well mas a brief discussion on some descriptive statistics. In Appendix C2, I provide a succinct description of two auxiliary datasets occasionally used throughout the chapter: the National Household Budget Survey and the National Fertility and Family Survey.

3.3.1 European Community Household Panel

The European Community Household Panel (ECHP) is a longitudinal harmonised survey, coordinated by Eurostat. The panel includes data from 15 countries in the European Union and, for most countries, consists of 8 waves, between 1994 and 2001.[8] For the purpose of this study, I use the panel for Portugal between 1994 and 1999.[9]

The choice of this database is related to the reasonable amount of information both at the individual and household level for such a long period. The ECHP provides information on a broad set of questions related to demographics, income, social life, housing conditions, health, education, and employment. The Portuguese Labour Force Survey also includes both information on fixed-term contracts and fertility, but only follows each household for a period of six quarters.

Even though the information contained in the ECHP is relatively broad to answer the main research question of this chapter, I complement it with the data from the Labour Force Survey conducted in Portugal during the same period. This survey provides clearer information on the duration and renewals of fixed-term contracts and includes a larger sample. This additional information should improve the modelling of labour market transitions of women along the life-cycle.

[8]The 15 countries are: Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Spain, Sweden, Portugal, and the United Kingdom

[9]I have excluded the years 2000 and 2001 as there was a change in definition of fixed-term contracts. ECHP has a follow up for more recent years, the European Union Survey on Income and Living Conditions (EU-SILC), but this has no explicit information on the duration of the fixed-term contracts.

3.3.2 Labour Force Survey

The Labour Force Survey (LFS) in Portugal (*"Inquérito ao Emprego"*) is a CPS-type household survey conducted by the Statistics Portugal (*INE - Instituto Nacional de Estatística*). Conducted on a quarterly basis, it asks information about approximately 45,000 individuals, in 15,000 households. Individuals aged 10 years old or less do not answer the survey, but are included in the database with information on demographics. Each survey respondent answers to around 150 questions about their participation in the labour market. The survey questions are divided into six broad topics including main activity, secondary activity, education and training, experience, job search, and labour market status in the previous year. In all sections, the survey follows the definitions of *Eurostat* making the labour market indicators comparable with other European countries.

In each quarter, one sixth of the sample is rotated out. Therefore, I can only compute transitions between labour market states for five sixths of the workers in the sample. However, measurement error is not a serious issue in this survey (Blanchard and Portugal, 2001). Inconsistencies in the observed labour-market transitions are negligible due to the low attrition rate and also to relatively low frequency of movements across labour market states.

To be consistent with the years selected in the ECHP data I have restricted the data in LFS to the same period. However, due to changes in the survey I cannot follow transitions that occurred between the last quarter of 1997 and the first quarter of 1998.

3.3.3 Construction of Samples and Descriptive Statistics

In both datasets I focus my analysis on women aged between 23 and 50 years old. As both of them are constructed on a rotation basis, I have two unbalanced panels.

In the sample from ECHP I have 2,283 women and in the sample from LFS I have 34,988 women. Due to the restrictions I impose relative to the period and age, in both samples I have roughly about 50% of the women being observed in all periods. For all women I collect information on age, marital status, household composition, education, employment status, employment history, type of contract, and wages, of the woman and the husband, if available.

Some definitions should be mentioned for clarity. In this chapter, a woman is employed if she holds a full-time job (either with fixed-term or permanent contract) with a certain firm that pays a given wage. I exclude part-time jobs as their proportion in Portuguese female employment is historically low, especially when compared with other countries.[10] For the purpose of this study, I also exclude self-employment, family business and contracts with undetermined ending date. All the employment-type restrictions accounted for about 24% of the employed female in the sample before restrictions.[11] Education is, throughout the chapter, classified into two categories only: non-university and university, therefore, women with unknown education were also excluded (1.4%).

Table 3.13, in the Appendix of this chapter, presents the main variables from the ECHP sample. This sample includes a relatively low percentage of women with university degree due to the period I am studying.[12] Slightly above 30% of the observations are non-workers and most of the employees hold a permanent contract. In terms of fixed-term contracts, most of them have duration of one year and zero renewals. This combination is reflected in the average tenure for all types of contracts – 20% of the

[10] Average part-time proportion of jobs held by female workers during 1994 to 1999 period was 14.3%, which is similar to the average rate over a longer period of time, whereas the average female part-time proportion in European countries is 26% and in the OECD countries is 24.5%.

[11] I follow Blundell et al. (2016a) and exclude women in all years starting from the one where the transition to one of the excluded categories is observed.

[12] According to the Portuguese Census the share of women above 15 years old with a university degree was 3.6% in 1991, 9.3% in 2001, and 16.9% in 2011.

workers have one year (or less) of tenure. In 2000's euros the average monthly wage is about 477, the minimum is 218.39 (minimum wage in 1994), and the maximum wage observed is about five times the mean wage. More than 70% of the observations are women with a partner and 34% are non-mothers. Regarding the number of children, about half of the observations are women with two children.

Table 3.14, in the Appendix of this chapter, presents the main indicators that describe the female labour market in both samples. The two samples are fairly similar when it comes to the percentage of women holding a university degree, the percentage of non-workers (including inactive and unemployed), the average unemployment duration, the percentage of fixed term contracts and average tenure. Larger differences are found in the classification of fixed-term contracts. The duration of the contract in ECHP is split into 4 classes ("less than 6 months", "6 months to 1 year", "1 year to under 2 years" and "2 years or more") whereas it is defined in number of months in the LFS database. This difference between the two datasets implies some differences in the labour market flows, presented in Table 3.15. Transitions to non-employment and short fixed-term contracts are fairly similar, but major differences occur in transitions to longer fixed-term contracts and to permanent contracts. However, the difference becomes negligible when the last two categories are summed up.

3.4 Preliminary Evidence

In this section I present preliminary evidence on the relationship between job security (fixed-term vs permanent contracts) and fertility in Portugal, during 1994-1999.

Following De la Rica and Iza (2005), I run a survival analysis on the time to first and second birth. The authors used a proportional hazards model but did not account for the increasing childlessness in the last decades, especially in Southern European countries (Frejka, 2008). In this analysis, I use the discrete version of the proportional hazards model (complementary log-log) with two mass-point unobserved heterogeneity, which accounts for the fact that some women take much longer to give birth – denoted as the *later birth* group – and eventually, because the fertile period is limited, some of them do not give birth at all – either because they are biologically or behaviourally sterile, as justified in Heckman and Walker (1990).[13]

The following hazard model is estimated for first and second births:

$$h(t|\boldsymbol{x'_{it}\beta}) = \begin{cases} 1 - \exp[-\exp(\gamma_1 \mathbb{1}FTC_{it} + \gamma_2 \mathbb{1}PC_{it} + \boldsymbol{x'_{it}\beta} + \gamma_3 \ln(t))], \\ \qquad \text{if } i \text{ is a } \textit{later birth woman} \\ 1 - \exp[-\exp(m_{ebw} + \gamma_1 \mathbb{1}FTC_{it} + \gamma_2 \mathbb{1}PC_{it} + \boldsymbol{x'_{it}\beta} + \gamma_3 \ln(t))], \\ \qquad \text{if } i \text{ is a } \textit{earlier birth woman} \end{cases} \tag{3.1}$$

where t indexes the years in which woman i is observed, $\mathbb{1}FTC_{it}$ is a dummy that takes the value 1 if the woman holds a fixed-term contract job, and 0 otherwise, $\mathbb{1}PC_{it}$ is a dummy that takes the value 1 if the woman holds a permanent-contract job, and 0 otherwise – being the base group the women not working (either inactive or unemployed) –, x_{it} is a set of controls that includes age, education, labour market experience, marital status, partner's income, and year and regional dummies, $\ln(t)$ captures the pattern of the duration dependence, and $m_{ebw} > 0$ distinguishes the two hazard rates in a discontinuous fashion. In this context likelihood contribution of each individual i becomes:

[13]See Varga (2014) for another application on fertility research with the two mass-point unobserved heterogeneity model, which is equivalent to the split population model.

$$\mathcal{L}_i = \pi \times \left[\left(\frac{h_{lbw}(t_i|\boldsymbol{x'_{it}\beta})}{1 - h_{lbw}(t_i|\boldsymbol{x'_{it}\beta})} \right)^{\delta_i} \prod_{j=1}^{t_i} (1 - h_{lbw}(t_i|\boldsymbol{x'_{it}\beta})) \right]$$
$$+ (1-\pi) \times \left[\left(\frac{h_{ebw}(t_i|\boldsymbol{x'_{it}\beta})}{1 - h_{ebw}(t_i|\boldsymbol{x'_{it}\beta})} \right)^{\delta_i} \prod_{j=1}^{t_i} (1 - h_{ebw}(t|\boldsymbol{x'_{it}\beta})) \right] \qquad (3.2)$$

where $\pi \in (0,1)$ is the proportion of women that are *later birth*, $h_k(t|\boldsymbol{x'_{it}\beta})$, k=lbw,ebw is the respective branch of (3.1), and δ_i indicates whether the spell is censored or not, i.e., whether the last month when individual i is observed, t_i is not equal to the month when the individual i gives birth.[14]

The estimation sample is restricted to women, with a partner (either formally married or not), between 23 years old, when most of the women already finished schooling, and 40 years old, by when most of the women realized their completed fertility.[15]

Table 3.1 presents the results for the complementary log-log regressions, with and without two-point unobserved heterogeneity.

Two results stand out from column (1) of Table 3.1. First, the effect of having a fixed-term contract on the probability of giving first birth is substantially large and negative. Second, it is not clear whether, for first birth decision, women prefer to have a permanent contract and a more stable source of income, or whether they prefer not to work (base group) and dedicate more time to child-rearing.

To understand which aspect drives this relationship, I use an event study. In Figure 3.3, I plot the (first) birth rate of women in two different situations: those that have transitioned from a fixed-term contract to a permanent contract, and those that transitioned from a fixed-term contract to non-work.

While the transition to a permanent contract has a positive impact on first birth

[14]See Allison (1982) and Jenkins (1995) for the construction of the complementary-log-log likelihood contribution.

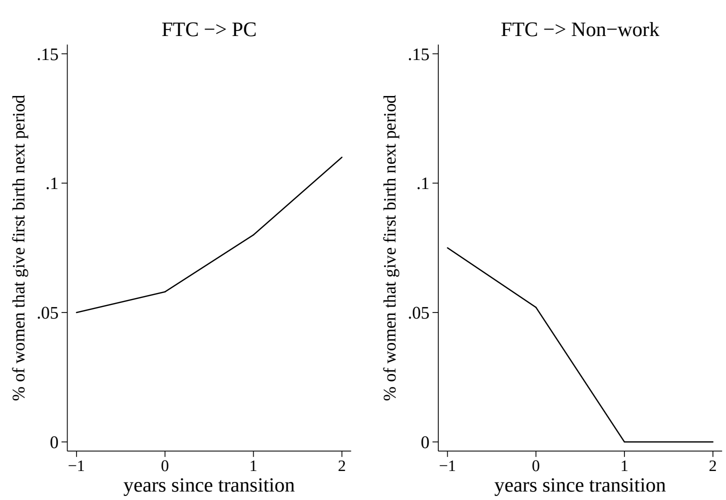
[15]Note that restricting the women to have a partner and to be older than 23 years old mechanically decreases the number of women at risk of first birth. On the other hand, the fact that I do not restrict the age of first child mechanically increases the number of women at risk of second birth.

Table 3.1: Estimates from complementary log-log regressions on first and second births, with and without two-point unobserved heterogeneity

Variable	First Birth		Second Birth	
$\mathbb{1}$ Permanent Contract (PC)	-0.105	-0.393	-0.666***	-1.895***
	(0.238)	(0.288)	(0.252)	(0.514)
$\mathbb{1}$ Fixed-term Contract (FTC)	-0.822***	-1.144***	-1.159**	-2.441***
	(0.313)	(0.348)	(0.554)	(0.719)
χ^2 PC – FTC	7.620	7.350	0.850	0.840
p-value	0.006	0.007	0.356	0.360
Constant	-18.806	-35.425	-8.648	-14.350
m_{ebw}		15.026		5.049
Prop. *later birth women*		0.195		0.588
Nr observations	972		1 925	

Notes: all regressions control for duration dependence (log form), age of the woman, university degree of the woman, husband's income, years of marriage, working experience of the woman, time effects, and regional effects. The second regression also controls for the age of the first child. Standard errors are reported in parentheses. * Significant at 10%, ** Significant at 5%, *** Significant at 1%

Figure 3.3: % of Women who give *first* birth next period, by transition from FTC



Source: author's calculations (ECHP)

decision, the loss of income in the second type of transition seems to overcome the increase in time for child-rearing.[16]

To explain what drives low fertility rates, one should also look at higher order births. I restrict my analysis to two births only as the number of mothers of three or more children is rather low in Portugal. In fact, this is also reflected in the proportion of women in *later birth* group, which is much larger for second than first births.

In columns (3) and (4) of Table 3.1 two main results stand out. For second birth, the fact that the woman is working has a negative and statistically significant impact, and this is not statistically dependent on the type of contract as the coefficient on the permanent contract is not statistically different from that for the fixed-term contract.

Two main results from this analysis should be kept in mind when modelling women's preferences. First, the different impact of job insecurity relative to the birth order. Second, the non-negligible proportion of women assigned to the *later birth* group.

As De la Rica and Iza (2005) reported, the survival analysis is important to get some evidence on the relationship between job security and fertility, but there is also a potential issue of simultaneity. Therefore, in the next subsection I explore exogenous variation in fixed-term contracts and look at its effect on fertility choices.

3.4.1 Reform on the Potential Cumulative Duration of FTC

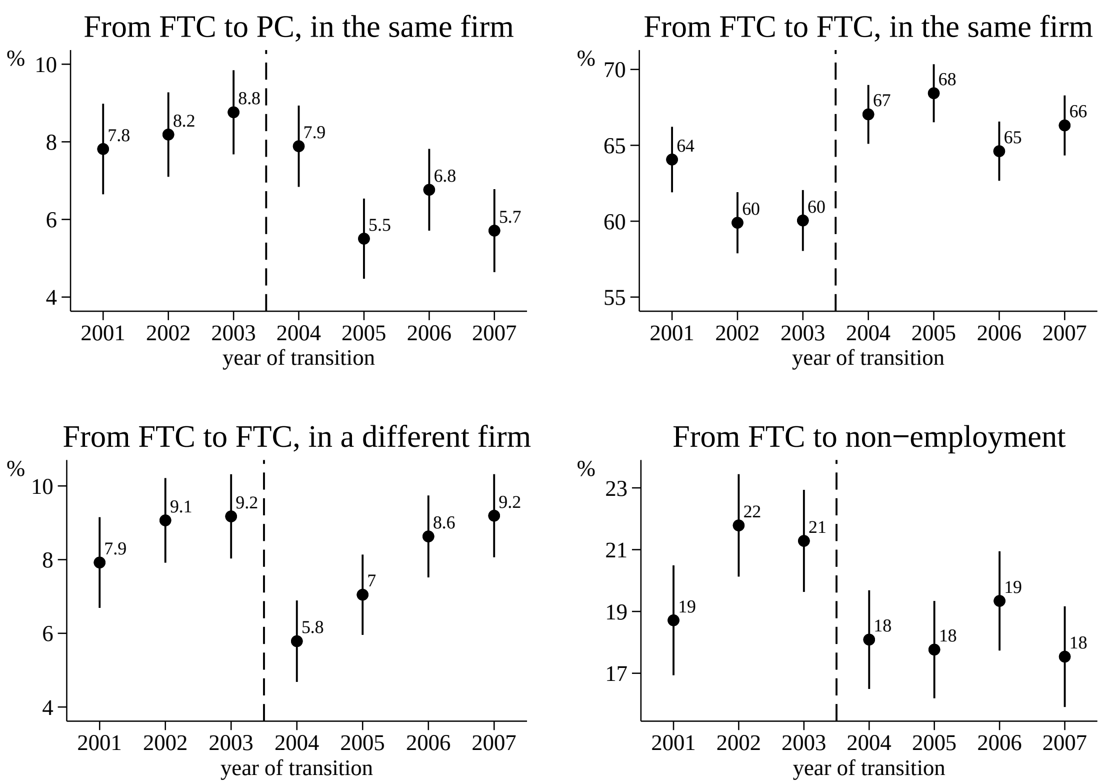
As explained in Section 3.2, in August 2003, the Portuguese government approved a law that changed the maximum duration rule of fixed-term contracts. Under the new scenario, in place since December 2003, a firm could keep an employee on a FTC for up to six years, while the cap on two renewals remained the same. In Figure 3.4, I plot the transitions from FTC to four different states: permanent contract within the same

[16]These results are in line with Del Bono et al. (2012) that show a drop on fertility rate after job loss, which comes mainly from difficulty women face in re-establishing their careers.

firm, FTC within the same firm, FTC in a different firm, and non-employment.[17]

The top left panel of Figure 3.4 illustrates a decrease in the probability of conversion into permanent contracts after the reform. However, firms did not adjust immediately on this dimension as some workers that were converted in 2004 started the fixed-term contract before the reform and, therefore, the contracts of those workers were still constrained by the pre-reform rules. The top left panel sheds light on the most significant change. The probability of staying in the same firm on a fixed-term contract increased after the reform. This effect is also reflected in the two panels at the bottom. Transitions across firms decreased in the short-run and transitions to non-employment decreased slightly. In conclusion, it seems that employment levels improved slightly, but job security decreased. Therefore, the effect on fertility is not clear.

Figure 3.4: Transitions from fixed-term contracts, by year of transition



[17]For presentation purposes I exclude transitions to permanent contracts in other firms as these represent 1% or less for all the years in this analysis.

To study the impact of this reform I use the LFS sample. As explained in Section 3.3, this dataset is more accurate in terms of the job contracts' characteristics. Moreover, this dataset provides consistent definition of fixed-term contracts during the pre- (2001-2003) and post-reform (2004-2007) periods.[18] In Table 3.2, I provide the descriptive statistics of mothers' characteristics before and after the reform.[19] In this comparison I have only considered women, with a partner, who were employed on a FTC in their first year of the contract.

Table 3.2: Descriptive statistics of labour market characteristics before and after the reform

Variable	Pre	Post	Raw difference	w. Time Trend
Age	30.917	31.517	.6 *	.145
			(.351)	(.669)
𝟙 University degree	.115	.121	.006	.047
			(.027)	(.049)
Years since 1st job	13.104	14.311	1.208 **	.809
			(.539)	(.988)
Number of previous jobs	2.787	3.441	.654 *	.092
			(.336)	(.852)
Contract duration (months)	7.99	7.519	-.471	-.249
			(.385)	(.609)
𝟙 White collar	.572	.631	.059	.055
			(.036)	(.071)
Log(wage)	5.958	6.109	.135 ***	.009
			(.032)	(.066)
Nr observations	825	1 298		
Sample size	478	747		
Population size	151 657	240 692		

Notes: The table reports descriptive statistics of married women employed with a FTC in their first year of the contract. *Pre* includes women for which the contract started before the reform and *Post* includes women for which the contract started after the reform was implemented. Column (3) reports differences in means between both groups, and column (4) the differences between pre- and post- averages ,controlling for a linear time trend. * Significant at 10%, ** Significant at 5%, *** Significant at 1%

[18]The year 2000 was excluded from the analysis due to a difference in fixed-term contracts definition. As this change could potentially affect the years of 2001, I have also performed this exercise with a shorter period around the reform, 2002-2005. Because the results did not change I kept this version to provide more information on the medium-run effects.

[19]For a similar analysis see Lalive et al. (2014)

In the third column of Table 3.2 there are some statistically significant raw differences between the pre and post sample. Naturally in the post sample I find older women as I observe some women both during pre- and post-reform periods. As a consequence, the career is also longer for the women observed in the post-reform. Not surprisingly, the wages are also higher as I am not taking into account inflation and productivity growth in the raw-differences. To account for these patterns, I test these differences controlling for a linear time trend in the last column. As no difference is statistically significant, I conclude that none of these characteristics has changed discontinuously at the same time as the reform.

In this exercise I run two regressions:

$$\text{GiveBirth}_{it+1} = \alpha_0 + \alpha_1 \mathbb{1}_{\text{after}_{it}} + \boldsymbol{x'_{it}\beta} + u_{1it}$$

$$\text{GiveBirth}_{it+1} = \alpha_0 + \alpha_1 \mathbb{1}_{\text{after}_{it}} + \alpha_2 \text{Duration}_{it} + \alpha_3 \mathbb{1}_{\text{after}_{it}} \times \text{Duration}_{it} + \boldsymbol{x'_{it}\beta} + u_{2it}$$

The first one tests if women, with FTC starting after the reform was implemented (after December 2003), were less likely to give birth in the following year, controlling for the characteristics described above. The second regression tests whether the duration of the first contract in the firm affected the impact of the reform. Results are presented in Table 3.3.

The negative coefficient on the dummy, which equals one if the FTC started after the reform, indicates the reform has negatively affected the probability of giving birth in the following year. The positive and statistically significant coefficient in the interaction term means that a longer FTC in the post-reform actually attenuates the decrease in the birth probability. This result might be explained by the fact that I have only included women in their first year of a contract in the sample as these might then give birth and still have time to come back to work and recover any career cost they might have

Table 3.3: Regressions for the effect of the 2003 reform on birth likelihood

Dep Variable: Give birth next year	(1)	(2)
𝟙 Contract after reform	-.086 *	-.143 *
	(.047)	(.076)
Contract duration (months)		-.002
		(.004)
𝟙 Contract after reform × Contract duration (months)		.008 *
		(.005)
Age	.010 **	.010 **
	(.005)	(.005)
Log(husband wage)	-.009	-.012
	(.036)	(.037)
𝟙 One child	-.019	-.037
	(.036)	(.032)
𝟙 University degree	.018	.041
	(.024)	(.033)
Years since 1st job	-.004	-.005
	(.003)	(.003)
Number of previous jobs	.003***	.012**
	(.001)	(.005)
𝟙 White collar	-.025	-.040
	(.022)	(.025)
Log(wage)	-.011	-.040
	(.018)	(.025)
Year FE	Yes	Yes
Quarter FE	Yes	Yes
F-stat	2.03***	24.64***
Observations	2123	

Notes: The sample includes all married women in the first year of a FTC between 2001 and 2007. Robust standard errors are reported in parentheses. * Significant at 10%, ** Significant at 5%, *** Significant at 1%

had. The magnitude of the effect of the reform goes in line with the result of Milligan (2005) on the impact of an increase of 1000$ in childcare benefits in Canada, which is reassuring in terms of the size of the impact of this reform on fertility decisions.

Regarding the remaining coefficients, older women (during their fertile period) with fixed term contracts are more likely to give birth as they have less time to postpone their fertility decision. The number of previous jobs also reveals a positive and statistically significant effect, which might be related to smaller losses for experienced women interrupting their career to give birth. Finally, I want to stress that the non-significance of the husband wage might be explained by the large contribution of women's wage to the household income in Portugal (40%). This importance is also revealed in the larger magnitude (even though not significant) of the coefficient on the woman's wage.

From this analysis I conclude that both the type and duration of a fixed term contract are relevant to the fertility decision. Therefore, I incorporate both features in the dynamic discrete choice structural model that is presented in the next section.

3.5 Model

I develop a structural model to describe the dynamic decisions of women on fertility and labour force participation over the life-cycle. Fertility choices are taken conditionally on the existence of a partner (including both formal marriage and cohabitation), which may change over the life-cycle with certain marriage and divorce probabilities.

Labour supply decisions are taken conditionally on job offers, which arrive with a certain probability. Jobs differ both in wages and contract duration (one year fixed-term, two year fixed-term, or permanent). Note that I assume the workers do not choose the type of contract nor the duration. The reasons for this option are twofold. Firstly, I find in the LFS data that 74% of the women working on a fixed-term contract would have liked to have a permanent contract instead, but could not find one. Secondly, Portugal and Varejão (2004) show that, for Portugal, fixed-term contract workers receive, on average, lower returns to both experience and tenure, thus meaning workers are not compensated over time for the higher risk of job loss associated.

Following Van der Klaauw (1996), Eckstein and Wolpin (1989), and Hotz and Miller (1988) I assume that women can only work full-time. The percentage of part-time workers in Portugal is small in comparison to other countries, as highlighted in the Section

3.3. According to André (1991) this phenomenon is due to the residential proximity of relatives, which are essential to support employed women and to the sizeable contribution of women's labour income to household income in Portugal. Below, I describe the main components of my model.

3.5.1 Timing and Decisions

Time is discrete, and a period lasts for a year in order to match the data frequency. The decision horizon for each woman starts at age 23, after school, and terminates at age 50, when the number of old-age pensioners increases substantially in the data. Note however that fertility decisions can only be taken until the age of 40, the age at which I assume a woman is no longer fecund.[20]

In every period, a woman has to decide both on fertility (n_{it}) and work (p_{it}). Fertility choices are only conditioned on the existence of a partner (h_{it}). I assume complete and costless control over the ability to give birth at each age, like Wolpin (1984), Moffitt (1984), Happel et al. (1984), and Cigno and Ermisch (1989). However, I impose a restriction of a maximum of two children (k_{it}) who must be born in separate years as I exclude the possibility of twins in my model.[21] Working choices are conditional on the offers received and lay-off outcomes at the beginning of the period.

[20]The starting age follows the same reasoning as Eckstein and Lifshitz (2011). Alternatively one could follow Van der Klaauw (1996), which starts the model in the first year of school-leaving. I chose 23 as starting age because according to Portuguese data there is a big drop in school attendance after 22. Moreover, according to the Fertility Survey only 5% of women in Portugal gave birth to the first child before 23 years old. Regarding the age which terminates the fertile period I follow Francesconi (2002). Also, according to Fertility Survey only 5% of the women in Portugal had the first child after 40 years old.

[21]According to *Instituto Nacional de Estatística* the number of households with more than two children was 6.5% and the less than 2% of births were twins.

3.5.2 Job Contracts and Wages

Jobs are characterised by monthly net wages and contract duration, which can be fixed or permanent. Both wages and contracts' rules follow the labour market characteristics in Portugal during the period of 1994 to 1999. To match the data frequency and the distribution of fixed-term contracts duration I allow fixed-term contracts to have the duration of one or two years.[22] Fixed-term contracts can be renewed twice, as long as the maximum duration with renewals, does not exceed three consecutive years. To control for the maximum number of renewals I also include years of tenure in both contracts rather than years of experience, as experience is not contract-specific and creates extra computational burden to the model. Once the maximum duration for fixed-term contracts is exhausted, the worker is either promoted to a permanent position or laid off. The promotion rate is lower than the renewal rate of fixed-term contracts in order to incorporate the fact the permanent contracts have higher firing costs associated, when compared to fixed-term contracts.

In order to match these labour market dynamics I allow the following combination of contracts: one-year contracts can be renewed twice if replaced with similar contracts, or once if replaced with a two-year contract; two-year contracts can be renewed once if replaced with a one-year contract; in case the worker achieves the maximum number of renewals and keeps working in the following year, the fixed-term contract is replaced with a permanent one. Note however that a worker has always the chance to receive a permanent job offer, and therefore does not necessarily have to wait for the end of the maximum fixed-term duration.[23] I also allow for hirings into positions with permanent contracts to match the cases where the probation period/fixed-term contract is less

[22] According to the law, the fixed-term contracts are even allowed to have duration of three years, but I have excluded them from the model as their existence is rather low in the market.

[23] According to Portugal and Varejão (2010), promotions to permanent positions are more likely to occur in the first two years of the fixed-term contract.

than one year.

Women's wages follow a functional form that is contract-specific (FTC, fixed-term contract, or PC, permanent contract):[24]

$$\ln(w_{it}^{FTC}) = \alpha_1^{FTC} + \alpha_2^{FTC} S_{i23} + \alpha_3^{FTC} \mathbb{1}(D_{it} = 2) + \alpha_4^{FTC} R_{it} + \eta_{it}^{FTC} \tag{3.3}$$

$$\ln(w_{it}^{PC}) = \alpha_1^{PC} + \alpha_2^{PC} S_{i23} + \alpha_3^{PC} t_i + \alpha_4^{PC} t_i^2 + \alpha_5^{z} X_{it} + \eta_{it}^{PC} \tag{3.4}$$

where S_{i23} is equal to 1 in case the woman had a university degree at 23 years old, zero otherwise; t_i is the age, which proxies for labour market experience and in this model has values between 1 and 18; $\mathbb{1}(D_{it} = 2)$ is a dummy that takes the value of one if the fixed-term contract has the duration of two years; X_{it} accounts for the number of years of tenure, which can be equal to 1, 2 or 3, where 3 actually means 3 or more years of tenure; R_{it} is the number of renewals, which can be 0, 1 or 2; finally, η_{it}^{PC} and η_{it}^{FTC} capture random variations in wages, which are independent of the decision taken. Exponentiation of the r.h.s. or log of the l.h.s. ensures non-negative wages.

Note that both specifications depend on schooling, but differ in the other determinants. It is expected that the average wage for fixed-term contracts to be higher to compensate for the absence of employment protection and for schooling returns to be larger for permanent contracts, as these should represent a better match between the worker skills and the firm needs (Dias da Silva and Turrini, 2015). To match the features of the data, fixed-term contract wages do not depend on tenure, but rather on their duration and number of renewals. Both coefficients associated to these variables should be positive as both of them can be interpreted as investments in human capital at the firm. In a sense, longer duration contracts and more renewals also capture the positive effect of tenure, which is included in permanent contracts, as a sign of the quality of the match between the worker and the firm (Booth et al., 2002b). Age also

[24] See Francesconi (2002) for a similar approach.

enters the permanent contracts' specification as a way to capture that, over time, the appeal for the outside option is decreasing (Pinheiro and Visschers, 2015).

The random components of wages are i.i.d technology shocks with zero mean, finite variance, and a nonzero contemporaneous covariance between them. These assumptions are consistent with the fact that the data used for estimation does not cover a long period (6 years), and that during this period the variation of wages was flat.[25] In subsection 3.6.1 I discuss the estimation procedure for wages.

3.5.3 Job Offer and Dismissal Probabilities

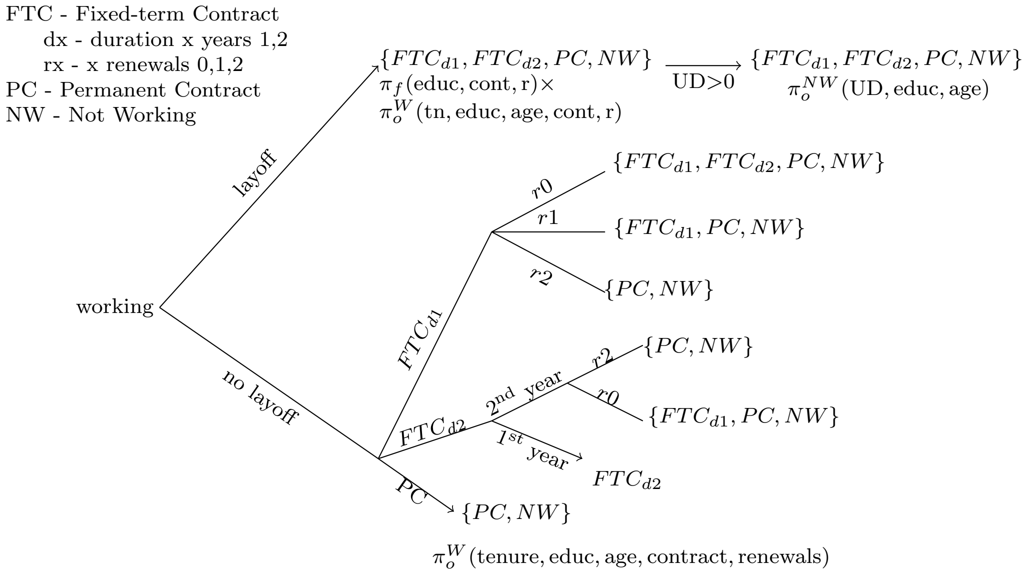
At most one job offer arrives each period with a certain probability. Even though I do not allow for search-on-the-job I do allow for job offer arrivals in the same period the woman is dismissed.[26] In case the woman was working in the previous period, the probability of getting a job offer is given by π_o^W, which depends on education, age, type of contract held before, tenure, and renewals. In case the woman was not working in previous period, the probability is defined by π_o^{NW}, which depends on unemployment duration (UD), education, and age. In Figure 3.5, I provide a diagram that represents the possible transitions in the labour market modelled here, each of them with a different probability.

Unemployed women can potentially get any type job contract offer. Women that have been dismissed at the end of last period are not penalised. However the arrival rate decreases with the duration of unemployment (1, 2, or 3, or more years). Women working on fixed-term contracts can always be promoted as long as the duration of the contract has elapsed. Transitions from a one-year contract to a permanent contract will be less frequent than transitions from a two-year contract to a permanent contract

[25]See Blundell et al. (2016b) for a similar approach.

[26]Contrary to Llull and Miller (2016) and Edwards (2014) I do not allow women to receive job offers while employed, but I do differentiate between whether they were dismissed in the same year or not.

Figure 3.5: Possible labour market transitions, by layoff, contract duration, and renewals



(Gagliarducci, 2005). Job offers of fixed-term contracts with a duration of one year may also arrive if the number of renewals has not reached the maximum. Once the woman knows in which final point of the tree she is, then she decides whether to participate in the labour market.

Conditionally on being employed, the worker is dismissed with a given probability, which depends on the schooling level at age 23 ($S_{i,23} = 1$ in case of a University degree), on tenure ($X_{i,t}$) (weighted by the number of renewals ($R_{i,t}$) in case of a fixed-term contract as for permanent contracts $R_{i,t}$=0), and also on the duration of the contract (D_{it}). All the coefficients are expected to be negative as a complement to their positive impact on wages and promotion probability.

$$P(\text{dismissal}) = \frac{\exp(\pi^d)}{1 + \exp(\pi^d)} \tag{3.5}$$

$$\pi^d = \varsigma_0^d + \varsigma_1^d S_{i,23} + \varsigma_2^d X_{i,t}(1 + R_{i,t}) + \varsigma_3^d \mathbb{1}(D_{i,t} = 2) + \varsigma_4^d \mathbb{1}(D_{i,t} = 3)$$

For simplicity, I assume that this probability is also the probability of non-renewal. Therefore, if at the end of a fixed-term contract the worker was not dismissed, this means that the firm neither offered her a renewal of the fixed-term contract nor a permanent contract.

3.5.4 Partner

In this model I do not distinguish between formal marriage and cohabitation, hence *marriage* means getting a partner with whom the woman will potentially have children.[27] I define the probability of marriage as a function of the age of the woman and the type of job contract she holds.

$$P(\text{marriage}) = \frac{\exp(\pi^m)}{1+\exp(\pi^m)} \tag{3.6}$$
$$\pi^m = \theta_0^m + \theta_1^m t_i + \theta_2^m t_i^2 + \theta_3^m \mathbb{1}(D_{i,t} \in [1,2]) + \theta_4^m \mathbb{1}(D_{i,t} = 3)$$

Contrary to other models, such as Adda et al. (2017), marriage does not depend on schooling. As this is a permanent characteristic in this model and few women have a high university degree in the data, there is not enough variation to include schooling as a determinant of marriage.

The probability of *divorce* (leaving a partner) is defined as a function of the number of children (k_{it}) only and it is assumed to be zero in the year of birth (when the youngest child reaches one year old ($a_{it} = 1$)). As for couples without children this probability is constant to reflect the lack of statistical significant evidence I found for correlation between divorce and age of the woman. This probability is thus expressed

[27]I analyse both married and cohabiting couples because, according to Portuguese census data, around 20% of the births in Portugal during the period of the analysis were conceived out of formal marriage.

in the following way:

$$P(\text{divorce}) = \begin{cases} 0 & ,\text{if } a_{it} = 1 \\ \dfrac{\exp\left\{\theta_0^d + \theta_1^d k_{i,t}\right\}}{1 + \exp\left\{\theta_0^d + \theta_1^d k_{i,t}\right\}} & ,\text{if } a_{it} \neq 1 \end{cases} \tag{3.7}$$

3.5.5 Utility Function

The decision to participate in the labour market depends on the contract duration, but the disutility from work does not, except when the woman has children. Therefore, I simplify notation hereafter by defining $p_{i,t}$ to be equal to one when the woman works and zero otherwise, regardless of the time of contract. Note that the utility will depend on the choice of labour supply, but not directly on the choice of fertility, as children are only born in the period after the decision. However, different idiosyncratic shocks arrive according to each of the four possible decisions. The utility of woman i at period t takes the following form:

$$\begin{aligned} U_{it} = {} & c_{it} + \gamma_1 p_{i,t} c_{it} + \gamma_2 k_{i,t} c_{it} + (\gamma_3 + \zeta_i^p) p_{i,t} + (\gamma_4 + \zeta_i^n) k_{it} + \gamma_5 k_{it}^2 \\ & + \gamma_6 (D1_{i,t} + D2_{it}) \times k_{it} + \gamma_7 D3_{i,t} \times k_{it} \\ & + \gamma_8 (D1_{i,t} + D2_{it}) \times \mathbb{1}(a_{i,t} = 1) + \gamma_9 D3_{i,t} \times \mathbb{1}(a_{i,t} = 1) \\ & + \gamma_{10} \mathbb{1}(k_{i,t} = 2) \times \mathbb{1}(a_{i,t} = 1) + \sum_{j=1}^{4} \mathbb{1}(\text{choice j}) \epsilon_{it}^j \end{aligned} \tag{3.8}$$

Consumption is denoted by c_{it}, which represents the level of a composite good for which the price was normalised to 1. Following the female labour supply literature (Van der Klaauw, 1996; Francesconi, 2002; Keane and Wolpin, 2010; Eckstein and Lifshitz, 2011), consumption enters linearly in the utility function and there are no assets, thus neither borrowing nor saving are allowed in the model. Following Edwards (2014), I allow for income effects through the interaction of consumption with the choice of labour supply (γ_1) and with the number of children (γ_2). The first interaction accounts for the consumption-leisure substitution effect and the intuition for the second interaction is the fact that consumption choices might be influenced by the presence of children in the household (e.g. different film choices at the cinema). The disutility from work does not depend on the type of contract, but differs with taste (ζ_i^p).[28] I also introduce heterogeneity (ζ_i^n) in the utility derived from the number of children (γ_4). See the next subsection for a description of the unobserved heterogeneity process.

In the second row, I allow for utility from children to vary with the type of job contract in order to represent the preliminary evidence I found in Section 3.4. In a fixed-term contract, it is expected the woman to derive a lower utility from children, as she is unsure about whether she will be working in the near future and be able to bear the costs of children. In the third row I allow interactions between each type of job contract and the fact that the woman has a one-year-old child (infant), which may require greater investments both in terms of time and money, than in other periods of children's lives. I also allow the utility from children to vary with this feature in order to control for the birth spacing that is observed in the data. γ_{10} is restricted to be negative in order to capture the additional difficulties that a second child may bring.[29] Finally, ϵ_{it}^j are choice-specific random preference shocks, which follow a type I extreme value distribution: $F(\epsilon_{it}^j) = \exp\left\{-\exp\left\{-\epsilon_{it}^j/\tau\right\}\right\}$ with mean $\tau\gamma$ and variance $\tau^2\pi^2/6$.

[28] There is no utility level associated with not working because I assume this as the baseline in my model.

[29] In the National Fertility Survey, 60% of the mothers of one child reported the "additional problems and complications" of subsequent children to be important for their decision to not have more children.

Permanent Unobserved Heterogeneity

In terms of unobserved heterogeneity, I follow Heckman and Singer (1984), and allow women to belong to one of 4 different types, as reported in Table 4. At age 23, when individuals start to be modelled, each woman draws a random permanent preference for work-fertility from a multinomial distribution, which depends on tenure, presence of a partner, and number of children. HW stands for high taste for children and high

Table 4: Unobserved permanent heterogeneity

	work	non-work
high fertility	HW	HU
low fertility	LW	LU

taste for work, HU for high children for fertility, but low taste for work, LW for low taste for children, but high taste for work, and LU for low taste for both children and work. In the first two cases ζ_i^k is positive and when the taste for work is high ζ_i^p is also positive; otherwise they are both zero. These parameters act as taste shifters towards working and fertility choices.

3.5.6 Budget Constraint

The household budget constraint depends on the guaranteed minimum income (GMI), the husbands' income (y), the woman's wage (w), the unemployment benefits (ub), and the children's costs (CC). It takes the following form:

$$c_{i,t}^H = GMI(h_{i,t}, p_{i,t}, k_{i,t}) + y_{i,t}h_{i,t} + w_{i,t}p_{i,t} + ub_{i,t}(1 - p_{i,t}, X_{i,t}) - CC(p_{i,t}, h_{i,t}, k_{i,t}, a_{i,t}) \quad (3.9)$$

Household consumption is defined as $c_{i,t}^H = c_{i,t} \times (1 + 0.5 \times h_{i,t} + 0.3 \times k_{i,t})$. That is, the woman's consumption ($c_{i,t}$) scaled by the "number of adult equivalent" given by the "OECD modified scale", that is, 1 plus 0.5 for a second adult plus 0.3 for each child in the household (Hagenaars et al., 1994).[30]

As the model does not allow for either borrowing or saving, the income of the household does not depend on assets. In order to ensure non-negative consumption and to take into account the welfare benefits available in Portugal, I also include the guaranteed minimum income (GMI), which depends on the presence of a partner, the labour supply choice and on the number of children. Once the woman has a partner I assume he works full-time as, in the data, 93% of the partners of women in the age range between 23 and 50 years old are working full-time. However, instead of considering husband's wage I consider husband's wage I consider husband's personal income to account for the fact that 27% of the husbands are self-employed. To avoid increasing the state space I also express husband's income in terms of the woman's observables, namely on her schooling level and age.[31]

$$\hat{y}_{it} = \hat{\alpha}_1^h + \hat{\alpha}_2^h S_{i,23} + \hat{\alpha}_3^h t_i \tag{3.10}$$

The woman is also entitled to unemployment benefits in case she was laid off in the previous year, i.e., $p_{it} = 0$ and $X_{it} = 1$. Note that in case the woman decided not to work conditional on having received a job offer she is not entitled to unemployment benefits. In terms of duration of the unemployment benefits I do not allow the woman to receive them for more than a year, which is not a strong assumption considering the rules in Portugal during the period in analysis – the maximum potential duration for individuals between 23 and 30 years old was 12 months and for individuals between 30 and 40 years old was 18 months, regardless of previous working experience. For older individuals, the potential duration of unemployment benefits could go up to 2 years

[30]Contrary to Adda et al. (2017) I will not estimate the weights of the OECD modified scale. Also, according to Burniaux et al. (1998), sensitivity analyses suggest that while the composition of income poverty is affected by the use of different equivalence scales, trends over time are much less affected.

[31]See Van der Klaauw (1996), Sheran (2007), or Adda et al. (2017) for similar specifications.

and a half, but because they have to comply with the obligations to find a job, most of the beneficiaries did not take more than one year of unemployment benefits.

Finally, I introduce a cost for childcare, CC, which depends on the labour supply of the mother, the presence of the father in the household, the number of children, and the age of the youngest child. In case the mother is not working the childcare costs are lower as she has time to look after the child. However, the costs are not assumed to be zero in order to capture the large utilization of formal childcare in Portugal, which is not free.[32] Therefore, the costs are larger if there are infants present in the household. In case the husband is not present in the household the costs are reduced to account for possible cash transfers from the father, whose income is not entering in the budget constraint directly.

3.5.7 Dynamic Decision Problem

Each year the woman makes her decisions conditional on the following state space

$$\Omega_{i,t} = \left(p_{i,t-1}, n_{i,t-1}, k_{i,t}, h_{i,t}, a_{i,t}, S_{i,23}, D_{i,t}, X_{i,t}, R_{it}; \epsilon_{i,t}, \eta^{z}_{i,t}, \zeta^{p}_{i}, \zeta^{n}_{i}\right) \tag{3.11}$$

which includes, work choice, fertility choice, children, husband, age of the youngest child, university degree, duration of the contract, tenure, renewals of the contract, shocks on preferences and wages, and unobserved heterogeneity in tastes for work and fertility.

The value function for woman i in period t is given by:

$$V^{j}_{t}(\Omega_{it}) = \max_{j \in J(\Omega_{i,t})} U^{j}_{it}(\Omega_{it}) + \beta E_t(V_{t+1}(\Omega_{t+1}) | j \in J, \Omega_t) \tag{3.12}$$

[32]See the *Education at a Glance* reports produced by OECD. In the 2018 edition, Portugal was the country with the third largest private expense on pre-schooling, where more than 90% of the children are enrolled.

where β is a discount factor, and E_t is the expectation operator conditional on information in period t.[33] Note that J, the set of possible choices, may not be equal in every period as it depends on the arrival of job offers, on the existence of a partner, on the number of children (limited to a maximum of 2), and on the age of the woman (no longer fertile from age 40 onwards). In Appendix C3, I present all the possible conditional value functions.

Terminal Condition

By the age of 40, women make their last fertility decision in the model. However, as noted by Adda et al. (2017) children have costs in the life-cycle career of the woman. Therefore, I include 10 additional years of labour supply decisions in the model in order to capture these costs before women start to retire. At the age of 50 I assume that the future value is function of children still present in the household, and on the type of the last job contract.

3.6 Estimation

The parameters of the model are recovered in two steps. First, I estimate the equations that describe the evolution of the exogenous elements in the model. These include the dynamics of marriage, divorce, lay-off, job offers, and the earnings for wives and husbands. The discount factor and the costs of childcare are externally set. Second, the remaining parameters related to the utility function are estimated by a combination of the Method of Simulated Moments (MSM) and Indirect Inference. Details of estimation for the exogenous elements in the model can be found in Appendix C4, except for wages and probabilities estimation for which the estimation is described

[33] β is set to 0.98, following Attanasio et al. (2008).

in the next subsection.

3.6.1 Wages and Offer/Lay-off Probabilities

When estimating wages one should take into account selection for work as the wages of workers might be higher than the potential wages of non-workers. In this chapter I estimate wages of women using a Heckman (1979) selection correction approach. This estimation is done in a two-step procedure where the first step is a probit for work choice and the second-stage is the wage equation, which accounts for the correlation between unobservables of work choice and unobservables of wages (Mills ratio). In the probit, I include the variables in the model that do not directly determine wages, but are important determinants for the labour supply decision. Results in Table 3.19 indicate the number of children as an important variable to take into account when women choose to work whereas being married seems to be less relevant. For both contracts, the Mills ratio turned out to be close to zero and non significant.

The results on selection should be taken with a grain of salt as the variables chosen as exclusion restrictions might not be the best determinants of work choice. For robustness purposes I have also estimated the selection equations accounting for additional variables that existed in the data, but do not enter the structural model. Personal non-labour income includes private transfers and returns from investments whereas social transfers include unemployment benefits, family-related allowances, sickness/invalidity benefits, education-related allowances, social assistance and housing allowance. Social transfers are quite important in the context of Portugal, which reduce poverty risk for around half.[34] Following the evidence in the literature showing that female participation rises when households move into home ownership, I have also included dummies for rent-free housing, rented housing, and second-home ownership (Del Boca and Lusardi,

[34]For a similar approach with United States data, see Low et al. (2018)

2003). Results corroborate the evidence in the literature and also emphasise the importance of home allowances in Portugal. Even after including all these variables, there is no evidence of selection into work.

Although there is no evidence for selection, both estimates of fixed-term contract wages and permanent wages presented in Table 3.18 are the ones resulting from the selection correction approach. For husband's wage no correction is done as the model assumes the husband always works and the average participation rate in the data is 93%. As justified in the model description, all wages refer to full-time work as part-time is not particularly relevant in Portugal, especially during the estimation period. For accuracy, the wages estimation is also restricted to wages above the minimum wage defined by law in each year.

In terms of offer probabilities, each node of the tree presented in Figure 3.5 is assigned a value that depends on the type of contract, contract duration, number of renewals, unemployment duration, schooling and age.[35] Lay-off probabilities instead follow a functional form that depends on university degree, tenure, and type of contract. Results are presented in Table 3.17. Note that job offers are not directly observed in the data (estimation uses transitions observed in the Labour Force Survey), but lay-offs can be distinguished from quits and therefore, only this type of job separation is used in the last estimation.

3.6.2 Method of Simulated Moments and Indirect Inference

The estimation of the parameters in the utility function is done by combining MSM and indirect inference.[36] Firstly, assuming an initial set of parameters in order to calcu-

[35]For presentational purposes the results on offer probabilities' estimates are omitted from this chapter, but are available upon request.

[36]See Pakes and Pollard (1989) and Duffie and Singleton (1993) for MSM and Gourieroux et al. (1993) for indirect inference

late all the possible conditional value functions, I solve the model backwards (starting at the age of 50). Secondly I simulate the choices of 2985 women (5 times the number of women I have aged 23 years old following Blundell et al. (2016a)), over the life-cycle and save them in a panel data format. Initial conditions are taken from the data. Decisions are made at the end of each period conditional on both observables and outcomes from random draws, which are known at the beginning of each period.

Using both datasets (real and simulated) I compute the moments that are described in the next subsection as well as the coefficients with respect to the indirect inference approach. Finally, using the method of Bound Optimization by Quadratic Approximation (Powell, 2009), the parameters are the solution of the following problem:

$$\min_{\theta}(M(\theta) - M_R)' W_R^{-1} (M(\theta) - M_R) \tag{3.13}$$

where θ is the vector of parameters to be estimated; $M(\theta)$ is the vector of moments computed from the simulated data; M_R is the vector of moments computed from the real data; W is the weighting matrix, which contains sample variances of M_R in the diagonal and hence the moments with greater variance will be less important. The sample variances were bootstrapped with 10000 iterations.[37]

3.6.3 Moments

The moments used for the utility function parameters are listed in Table 3.5.

The proportion of non-workers, FTC, and PC workers, by age group, identifies the disutility from work; the proportion of non-mothers, mothers of one child, mothers of two children, by age group, and the quantiles of age at first birth and second birth, identify the utility from children; the proportions of women with a one-year-old child

[37]Under the regulatory conditions stated in Pakes and Pollard (1989) and Duffie and Singleton (1993), the estimator $\hat{\theta}$ is consistent and asymptotically normally distributed.

Table 3.5: Moments

Description	#
Proportion of women with a one-year-old child in each contract	4
Transitions between each labour market status	16
Proportion of non-mothers by age group	3
Proportion of mothers of one child by age group	3
Proportion of mothers of two children by age group	3
Proportion of non-workers by age group	5
Proportion of workers in fixed-term contracts by age group	5
Proportion of workers in permanent contracts by age group	5
Proportion of married women with children by age group	5
Quantiles of age at first birth and second birth	10
OLS regression coefficients of household income net of childcare costs on age, number of children and work	3
OLS regression coefficients of children (0/1) on age, $\mathbb{1}$(husband), and contract	4
OLS regression coefficients of children (1/2) on age, $\mathbb{1}$(husband), and contract	4

Note: the age groups are [23,30],]30,35],]35,40],]40,45],]45,50]. When calculating the proportion of mothers I do not consider the age groups outside the fertile age. OLS regressions also account for year fixed effects.

who do not work, work on a FTC, or work on a PC identify the complementarity between newborns and work; the coefficients on the OLS regression of household income net of childcare costs on age, number of children, and work identify the complementarity between consumption and work and the complementarity between consumption and children; and finally the coefficients OLS regression of children (0/1) on age, and non-work, FTC, PC, and the coefficients OLS regression of children (1/2) on age, and non-work, FTC, PC, identify the complementarity between the type of contract and children. Additional moments that relate outcome variables with endogenous decisions, such as the transitions between each labour market status and the proportion of married women with children by age group are also included in the estimation as they are crucial to identify the dynamics in the model (Eisenhauer et al., 2015).

3.7 Results

In the first subsection I report and discuss the estimates of the parameters in the utility function. Afterwards, I present the fit of the model, both in aggregate statistics and in life-cycle graphs, which depict the main outcomes along the life-cycle.

3.7.1 Parameter Estimates

Parameter estimates for the utility function and respective standard errors are shown in Table 3.6. Despite the fact that I assume a utility function that is linear in consumption, there is evidence for complementarity between consumption, labour and the number of children, as both parameters (γ_1 and γ_2) give a negative value.

Table 3.6: Utility function parameters:

Parameter	Value	Standard Error
γ_1	-0.45	(0.03)
γ_2	-0.01	(0.01)
γ_3	-0.52	(0.06)
γ_4	2.63	(0.83)
γ_5	-0.21	(0.13)
γ_6	-0.50	(0.19)
γ_7	0.42	(0.05)
γ_8	-1.03	(0.21)
γ_9	-0.61	(0.18)
γ_{10}	-2.05	(0.87)
Unob.Heterogeneity		
ζ^p	0.18	(0.00)
ζ^n	0.10	(0.04)
Proportions (%)		
HW	18.40	
HU	20.17	
LW	35.59	
LU	25.84	

The contribution of children to the utility function indicates a quadratic shape (γ_4

and γ_5), which sheds light on the larger proportion of single-child mothers relative to two-child ones. Illustrating the usual spacing between the first and second births, γ_{10} is a relatively large negative value.

The interplay between labour supply and fertility decisions is reflected in γ_6, γ_7, γ_8, and γ_9. The first two parameters have opposite signs reflecting the stronger preference for children when the woman is permanently employed. As the job security is lower in a fixed-term contract the woman is likely to face an income shock, which might have (negative) implications in terms of children (monetary) investment, therefore, the coefficient associated with children during the period of fixed-term employment (γ_6) is negative. Note however that the magnitude (in absolute terms) of the parameter associated to children while the woman is permanently employed (γ_7) is smaller than that for fixed-term employment to highlight the fact that, while employed, the mothers will have less time to spend with their children.

When I look at these interactions in the presence of an infant (i.e., the youngest child is one-year-old), the coefficient associated to permanent employment (γ_9) is no longer positive, highlighting the importance of time investment in children in the first year of life. However, to compensate for the lack of time while employed, the mothers might choose formal childcare. Such costs become relatively more expensive if the mother becomes unemployed in the following years as the income necessary to keep such monetary investment might decrease (with a larger probability) in case she holds a fixed-term contract (γ_8).[38]

Finally, in terms of unobserved heterogeneity, the types that have relatively stronger preference towards working have less disutility from labour force participation (ζ^p) and the types that are more prone to have children enjoy it more (ζ^n) over the life-cycle. The proportions for each type are presented at the end of the table. From these I

[38]Remember from subsection 3.5.6 that there is a high take-up of formal childcare in Portugal.

observe that the groups with relatively more taste for children (HW and HU) are less than 40% of the sample and the groups with relatively more taste for work (HW and LW) represent more than half of it. As expected the most representative group is the one with lower taste for children and higher taste for work (LW, with 35.59%).

3.7.2 Model Fit

In Table 3.7, I list the proportions of women according to different characteristics, both in actual and simulated data.In the third column I provide the p-value resulting from the t-test that compares the means between the data and the model.

Table 3.7: Comparison of means between actual and simulated data

	Actual	**Simulated**	**t-test p-value**	**χ^2 p-value**
Non-employment Rate	.3323	.3282	.27	
Employment FTC 1 year	.0544	.0425	.05	.000
Employment FTC 2 year	.0170	.0151	.18	
Employment PC	.5964	.6141	.07	
Fertility Rate	.0370	.0354	.27	
Fert. Rate, non-employment	.2450	.3186	.07	.000
Fert. Rate, FTC	.1125	.0555	.08	
Fert. Rate, PC	.6425	.6259	.31	
% Women with no children	.3470	.2940	.02	
% Mothers of 1 child	.3127	.3825	.01	.000
% Mothers of 2 child	.3404	.3235	.07	
% Working mothers	.6555	.6486	.21	

This model achieves reasonably good fit in terms of the labour market indicators. Figure 3.6 shows the fit on labour supply along the life-cycle. Employment in permanent contracts is the one that presents a better fit, which is consistent with the the fact that this group has the closest match in terms of fertility rate (Table 3.7).

Figure 3.6: Model fit: Labour market status distribution along the life-cycle

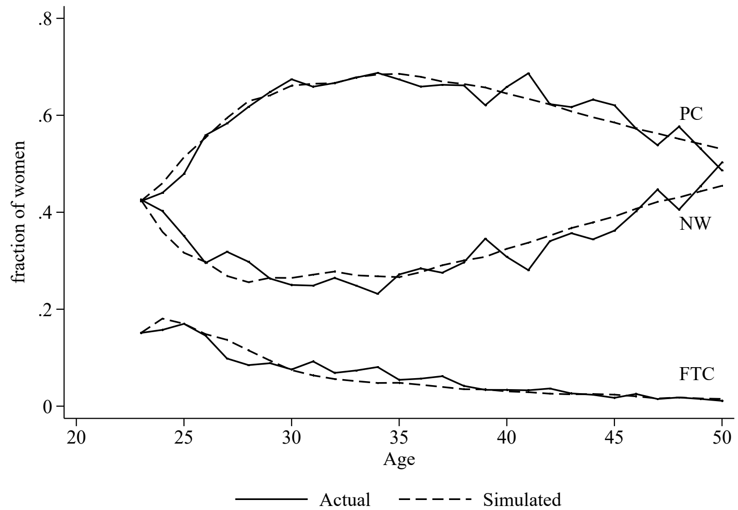
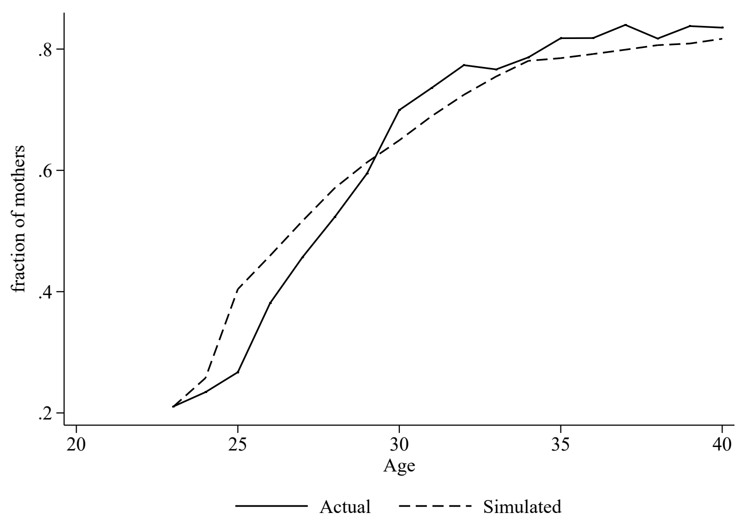


Figure 3.7: Model fit: cumulative fertility choice along the life-cycle



On average, the model predicts fertility rate reasonably well. Figure 3.7 illustrates the fit of motherhood along the life-cycle. Despite some local disparities, especially between 23 and 25 years old, the model is able to capture the general concavity of this indicator.[39]

[39] An extension of the model that could improve the fit on fertility would be to include the value for

3.7.3 Model Validation

To validate the model, estimated with data between 1994 and 1999, I extend the maximum duration of successive FTC in the same firm to 6 years, as in the reform of 2003. In this exercise, I take out the restrictions on renewals that previously limited the FTC cumulative duration to 3 years. For example, in the baseline, a 2-year contract would only be renewed into a 1-year contract or it would be converted into a permanent contract, if such an offer arrived. In this simulation, I allow the 2-year contract to be renewed twice with the same duration before it is eventually converted into a permanent contract.

As I change the structure of the labour market, but only model the labour supply I should also adapt the probabilities for each transition in the labour market. To achieve this, I re-estimate the same functional forms of the probabilities using LFS data between 2004 and 2008 in order to capture the legislation in place at the time, which was precisely the one set in this experiment. I could use data up until 2009, when legislation changed again, but I decided not to do so in order to avoid the contaminating the results by the crisis in Portugal at that time, even though I control for general trends by including dummies for both cohorts and years in all the regressions. Table 3.8 presents the results of the first experiment compared to the baseline of the model and also to the actual data from the Labour Force Survey.

As expected, and in line with Güell and Rodríguez Mora (2010), I observe a lower employment rate (a fall of almost 10%) when the prevalence of fixed-term contracts is higher. Such an effect is also observed in the significant decrease of contract conversions. The higher degree of job uncertainty should have repercussions in terms of fertility. Indeed, I observe lower fertility rate, and a higher proportion of childless women. Such

quality of the children and children investment decisions (Becker and Tomes, 1976; Chiswick, 1986; Del Boca et al., 2014; Carneiro and Ginja, 2016). Note however that the availability of data on early child investment and child performance in Portugal is very limited

Table 3.8: Model Validation

	Baseline 1994-1999	**Prediction** 2004-2008	**Actual Data** 2004-2008
Employment (%)	67.18	60.81	59.28
Permanent contracts (%)	91.42	82.14	81.77
FTC converted into PC (%)	9.33	2.82	4.21
Fertility rate	1.23	1.17	1.18
Childless at the end of fertile period (%)	17.05	19.19	21.22

predictions go in line with the channel of promotion as an important determinant of the fertility decision. To check whether this validation is also in line with the reduced-form evidence presented in subsection 3.4.1, I have run the same regression using the baseline version of the model as the pre-period and the simulated version of the model as the post-period. Results on the coefficients of interest are presented in the graph 3.11 in the Appendix of this chapter. As I have achieved a reasonable model validation, in the next section I set up two experiments for policies that have been discussed in the literature of job security.

3.8 Policy Experiments

In this section I predict the impact of two counterfactuals on labour supply and fertility decisions: (1) A single contract with a 3-year probation period; (2) A higher tax rate on fixed-term contracts compensated by a decrease on the tax on permanent contracts.[40]

[40] All policy experiments are implemented permanently from t = 1 with no prior announcement. All policies only affect women as the model does not include men working on fixed-term contracts.

3.8.1 Single Contract

To evaluate the outcome of the first policy experiment I simulate two other extreme scenarios: one without permanent contracts and another without fixed-term contracts. As the fertility mechanism under study is the level of job security, the outcomes of these two scenarios should work as bounds to the impact of the new policy experiment.

To eliminate all the permanent contracts, I force the ones in the simulated data, to be defined as 2-year fixed-term contracts instead. In terms of probabilities I keep the same estimates and sum the probability of a permanent contract to that of a 2-year fixed-term contract for consistency. As I do not adjust the rest of the labour market characteristics (restrictions on the number of renewals, lay-off probabilities and wages) one should expect the unemployment rate to increase as the duration of fixed-term contracts is exhausted and the rate of job offer arrivals is not adjusted.

To complement the previous counterfactual I eliminate all the fixed-term contracts by converting them into permanent ones since the period t=1. Once again, I do not adjust the rest of the labour market characteristics as these are assumed to be exogenous in the model. Therefore, in a situation with extremely low job insecurity one should expect the employment rate to go up.

Table 3.9: Job security bounds: average levels and deviations from baseline on fertility

	Baseline	FTC Only	(Δ)	PC Only	(Δ)
Fertility rate	1.23	0.84	(-32 %)	1.38	(12 %)
Childless women at age 41 (%)	17.25	25.03	(45 %)	14.4	(-17 %)

After computing the bounds for the extreme cases of job security, I simulate the single contract with a probation period of three years. During this period, I assume employers have the right to terminate the contract as if it was a fixed-term contract (which can be interpreted as a reduction in the severance payments for the short tenure dismissals), but following the probation period, employers are required to employ the workers permanently. Such a design is in line with what was adopted in Italy in 2015.[41] Under these conditions it is expected job security to be low in the first years of the single contract (fixed-term contract part), but also a higher conversion into the permanent position of the single contracts. Given the two counteracting forces, ex-ante, the direction of fertility change under this scenario is not clear. The results are summarised in Table 3.10, which exhibits the average effects of the counterfactual on eight main statistics.

Table 3.10: Single Contract experiment: average levels and deviations from baseline

	Baseline	Single C.	(Δ%)
Employment (%)	67.18	73.12	(8.8 %)
Permanent contracts (%)	91.42	94.6	(3.5 %)
FTC converted into PC (%)	9.33	28.58	(206 %)
Fertility rate	1.23	1.28	(4.1 %)
Childless women at age 41 (%)	17.25	15.82	(-8.3 %)
Kids conceived during FTC (%)	5.55	5.07	(-8.7 %)
Kids conceived during unemp. (%)	31.86	27.02	(-15.2 %)

This policy affects essentially the complementarity between employment and fertility choices. The percentage of births during employment is predicted to increase from 68% to 73%.[42] This result is mainly due to the increase in the percentage of births during permanent contracts (8.5%) as the number of births during fixed-term contracts has decreased. The overall effect on fertility rate is 4.1%. The magnitude of this

[41] See Lepage-Saucier et al. (2013) for a discussion of other alternatives.

[42] According to the OECD, in 1999, in Portugal, the employment rate for women between 25 and 54 years old, with at least one child between 0 and 2 years old, was 73.4% whereas that with no children (between 0 and 14 years old) was 69.4%. Only Slovenia, Croatia and Denmark presented the same difference whereas, for example, the United States presented a difference of 7.3 p.p. towards more childless working women. Eurostat presents the same evidence for more recent years as well as for women between 25 and 49 years old.

result is remarkable as I compare with the results on childcare cost reduction in the literature.Haan and Wrohlich (2011) found a 4.6% effect from a 20% increase in overall childcare benefits in Germany, and Sheran (2007) found an 8.3% effect from a decrease of 60% on childcare costs in the United States. For completeness, I have also simulated these two policies and the effects on fertility rate are respectively 4.2% and 11%. This means that improving the job security would be an effective way of increasing fertility without relying on heavy costs of childcare subsidies.

According to my simulations, the rate of conversion is predicted to double under the rules of the single contract. As a consequence, the employment rate also increases. As this is a partial equilibrium model one could be concerned about the overshooting on the employment rate, but even with a general equilibrium model Dolado et al. (2018) predicted an increase in the total employment rate in the scenario of a single contract for Spain, which is similar to Portugal in these features of the labour market. Under such scenario, the higher fertility rate together with a higher female labour supply would contrast with the policies of job protection that took place in Germany in the 80s and 90s and decreased the labour market attachment for mothers, especially in the short-run (Schönberg and Ludsteck, 2014).

For robustness, I test the sensitivity of the result to labour demand adjustments.[43] As the increase in the share of permanent contracts would bring large costs to the firms, I test the sensitivity of the result with respect to adjustments in the firing probabilities on permanent contracts, on the new hiring probabilities on fixed-term contracts, and on the seniority wage differential.

[43]This analysis is important as the feedback effects of the policy reform on labour demand could potentially reverse the conclusions drawn above by the partial equilibrium experimental evaluation (Lise et al., 2004)

Figure 3.8: Sensitivity to a Decrease on the Ratio of Firing Probabilities

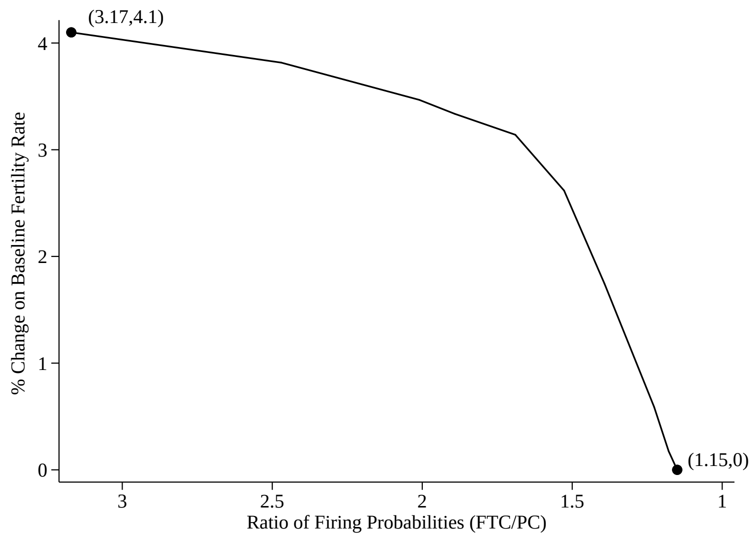


Figure 3.9: Sensitivity to a Decrease on the Probability of New FTC Hiring

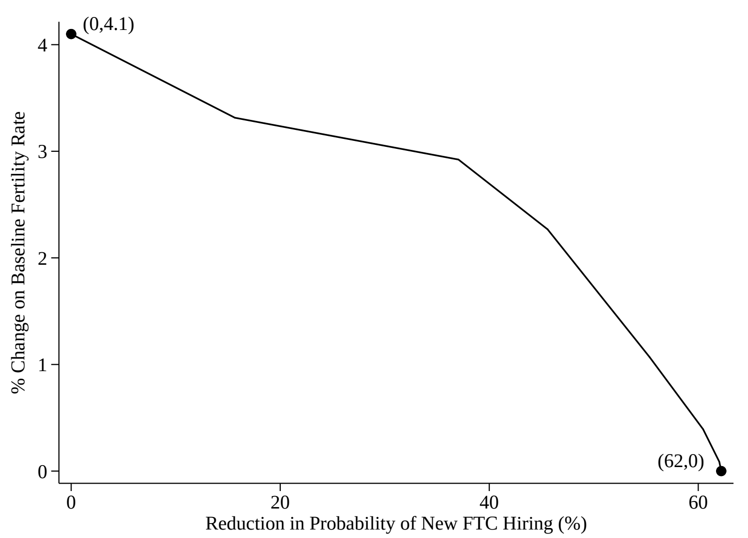
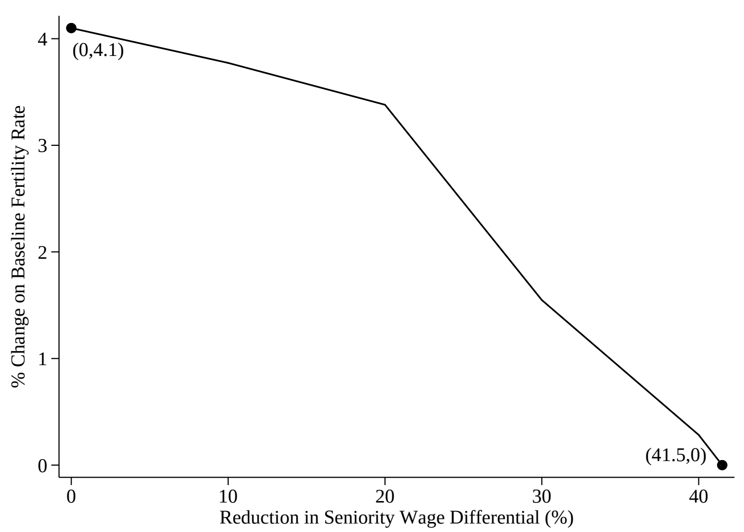


Figure 3.10: Sensitivity to a Decrease on the Seniority Wage Differential



In the first test I change the ratio of firing probability of permanent contracts relative to that of fixed-term contracts, as a proxy for the decrease in severance payment for permanent contracts. As Figure 3.8 depicts, there would have to be a reduction of around 3 times the actual ratio of firing probabilities for the positive effect on fertility to disappear. Dolado et al. (2018) predict a decrease from 3.32 to 2.66, for Spain.

Another adjustment that firms might choose is to decrease the new hirings on fixed-term contracts to compensate for the increase the labour force on permanent contracts within each firm. Figure 3.9 shows there would have to be a reduction of 62% in the rate of new FTC hiring to cancel out the positive effect on fertility. Dolado et al. (2018) actually predict an increase on hiring of 0.045%, for Spain.

Finally, firms could adjust the seniority wage differential. As the conversion rate between a fixed-term and a permanent contract increases, firms would, everything else constant, employ the permanent workers for longer. In that case, firms might have an incentive to decrease the wage seniority in the firm. According to Figure 3.10, there would have to be a reduction of 41.5% in the seniority wage differential in order to cancel the positive effect of single contract on fertility rate. According to Dolado et al. (2018), the average wage of workers aged 55 to 64 years old should decrease by 14% while the average wage of workers between 25 and 54 years old should increase 2.6%, which once again indicates that, even if there is a labour market adjustment, the impact on fertility rate should still be non-negligible.

In conclusion: even though the effect of a single contract on fertility might vary according to the different responses of the demand side of the labour market, there is still a high chance the effect is still positive under labour market adjustments.

3.8.2 Contract-specific Tax Rates

As it was observed with the sensitivity tests to the previous policy, the wages also play an important role in a dual labour market. In the data, fixed-term contracts compensate for the job insecurity with higher wages whereas permanent contracts take some time to catch up those levels as they start at a lower amount for the same tenure. As a way to decrease the share of fixed-term contracts and increase the duration of job spells, some European countries such as France, Italy, Portugal, and Spain, introduced supplements on the income taxes for fixed-term contracts. Even though the idea is the same across countries, the rules vary a lot. In France, the most penalised contracts were those shorter than a month (whereas the ones between one and three months were less penalised); in Portugal, the penalization targeted those shorter than two weeks; in Spain, those shorter than one week, and only Italy applied the same penalization to all fixed-term contracts regardless of their duration. Moreover, in both France and Italy the penalization was refunded in case the contract resulted in a conversion to a permanent one whereas Portugal and Spain did not include that option.

In this subsection I simulate the impact of two different policies, one in which I penalise the wages of fixed-term contracts by 5%, and another in which that penalization is compensated with a credit of 5% on the permanent contracts. Note that as I do not model firms decisions I have two options to simulate this policy in the model: adjust net wages accordingly, or adjust probabilities of hiring in the same direction. As it was shown with the previous experiment, the decrease in the probability of hiring workers on a fixed-term contract, compensated by the increase of hiring workers on a permanent contract would increase fertility as the share of permanent contracts would increase. Therefore, I opt to show the effects of an adjustment in net wages. Results are presented in Table 3.11.

Table 3.11: Policy experiments: average baseline and deviations from baseline

	Baseline	Tax FTC	(Δ%)	Tax FTC/ Credit PC	(Δ%)
Employment (%)	67.18	66.9	(-0.4%)	72.98	(8.6%)
PC/employment (%)	91.42	91.71	(0.3%)	92.31	(1.0%)
FTC 1 year/employment (%)	6.33	6.07	(-4.1%)	5.66	(-10.6%)
FTC 2 year/employment (%)	2.25	2.22	(-1.3%)	2.02	(-10.2%)
Fertility Rate	1.23	1.229	(-0.1%)	1.32	(7.3%)
Mothers of 1 child (%)	43.23	43.29	(0.1%)	34.4	(-20.4%)
Mothers of 2 children (%)	39.73	39.73	(0.0%)	48.56	(22.2%)
Conceived during FTC (%)	5.55	3.36	(-39.5%)	3.9	(-29.7%)
Conceived during unemp. (%)	31.86	33.9	(6.4%)	23.25	(-27.0%)

Even though I do not use a general equilibrium model, the prediction of employment decrease as a result of a higher income tax on fixed-term contracts is in line with Cahuc et al. (2016). However, as I did not differentiate the tax by duration, I observe a decrease in both types of fixed-term contracts included in the model. Because the effect of the reduction in the proportion of fixed-term contracts was counteracted with the increase in unemployment the overall impact on the fertility rate was nearly null.

On the contrary, if such a tax would be compensated by a credit on the wages for permanent contracts the effect on fertility would be much larger than that predicted by the single contract. This magnitude comes closer to the 11% impact found by

Whittington et al. (1990) on a reform of the personal exemption for dependents in income tax, in United States.

Note that the increase in fertility rate is driven by the increase in the percentage of women with two children and by the decrease in the percentage of women with one child, that is, the percentage of childless women at the end of the fertile period would be exactly the same. This result corroborates what the evidence showed in the reduced-form results: job security plays a larger role in the decision of having children, whereas income seems to be more relevant to deciding how many children to have. This result is in line with the fact that the percentage of children conceived during a permanent contract is much larger under the last policy experiment than under the single contract scenario.

3.8.3 Welfare Analysis

Despite the positive effect of both single contract and tax/credit policies on the fertility rate, it is important to check if women were better off under the counterfactual scenarios. In this subsection I evaluate the different policy experiments in terms of percentage change in welfare and lifetime consumption. Welfare is defined as lifetime utility and measured in consumption equivalent units. Lifetime income is the total income earned by the household throughout the life-cycle.

The results on Table 3.12 show that both reforms yield a similar average impact in terms of employment rate, but while the first brings a larger impact on the duration of employment spells, the second has a larger impact on welfare. The last result is driven not only by the larger impact on the fertility rate, but also by the larger effect on lifetime income, as the utility function in the model is assumed to be linear.

In terms of the impact across the different groups of unobserved heterogeneity, three results should be emphasised. First, the difference on the impact on employment

Table 3.12: Policy Experiments: Changes in Welfare and Lifetime Consumption

	Average	High F High W	High F Low W	Low F High W	Low F Low W
A. Single Contract					
Δ % in Employment Rate	8.85	4.76	3.90	10.71	13.46
Δ % in Employment Duration	19.41	2.39	10.32	23.68	33.78
Δ % in Fertility Rate	4.07	1.37	1.24	6.73	2.58
Δ % in Welfare	3.10	1.22	0.98	5.50	3.26
Δ % in Lifetime Income	6.30	3.61	2.77	8.00	9.17
B. Tax FTC / Credit PC					
Δ % in Employment Rate	8.63	0.24	8.82	0.87	35.64
Δ % in Employment Duration	10.68	1.75	12.28	2.90	38.85
Δ % in Fertility Rate	7.26	-0.12	2.17	0.38	38.79
Δ % in Welfare	5.85	1.81	3.16	2.94	18.59
Δ % in Lifetime Income	11.04	6.10	9.77	6.90	24.67

duration comes mainly from the larger impacts on the groups with high taste for work (low taste for leisure). Second, in both policy experiments, the result on fertility comes essentially from the groups with low taste for fertility. Finally, the first policy caused a larger impact on fertility rate on the group with high taste for work (low taste for leisure), whereas the second policy change in fertility comes especially from the group with low taste for work (high taste for leisure). This difference is also reflected in terms of welfare but not in terms of lifetime income.

3.9 Conclusions

In this chapter, I study the impact of labour market duality on the fertility decision. First, I provide reduced-form evidence that women working on fixed-term contracts delay maternity until they find a permanent contract that boosts their job security. This effect is particularly strong for the decision on having a first child. However, this analysis cannot provide a causal relationship due to the presence of simultaneity in the woman's decisions on labour supply and fertility. By observing a childless woman on a fixed-term contract, it is not possible to identify whether the woman is solely focussing on her career by choosing a contract type that has a higher degree of job insecurity, or whether the woman is childless because she holds a job where the likelihood of being fired is higher.

As a source of exogenous variation, I use a policy change that has modified the rules of fixed-term contracts in Portugal in 2003. Under the new rules, a firm was entitled to keep the worker in a fixed-term contract for six years, rather than three. As there was no explicit change regarding the conversion into permanent contracts, this policy change has basically extended the period for which the job security is low. Comparing women's decisions before and after the policy change I study the impact of this reform on the likelihood of giving birth. The results indicate that women working on a fixed-term contract that started after the reform, with lower prospects of permanent employment had, on average, a lower probability of giving birth, keeping everything else constant. However, the negative effect was reduced when the duration of the contract was longer, as longer contracts tend to have a higher conversion rate.

Following the reduced-form evidence, I develop a dynamic structural model in which women take their decisions, conditionally on the type of job contract they hold, if any. The main goals of the model are twofold: to study the channels through which the type of contract (fixed-term or permanent) or absence of job affect fertility decisions along the life-cycle; and to simulate counterfactual scenarios of different job security. The model is estimated with a combination of the method of simulated moments and indirect inference, using data for Portugal between 1994 and 1999, and validated with an out-of-sample prediction for the period between 2004 and 2008, after the reform.

Both policy counterfactuals come from adaptations of actual policy changes in Italy in 2015. The first experiment simulates the creation of a single contract in which fixed-term contracts are automatically converted into permanent contracts in case the woman is not fired during a probation period of three years. Results show that this policy could increase fertility in Portugal by 4.1%. As this policy change could potentially bring high costs to the firms, I perform three sensitivity analyses with respect to possible adjustments in the labour market. In case firms decrease the ratio of firing probability of permanent contract to that of fixed-term contract, as a way to incorporate a decrease in severance payment for permanent contracts, the effect on fertility rate would still be positive. In fact, a negative result would only arise in case the ratio was close to unity. Another possible adjustment one could think of would be the reduction in the seniority wage differential. According to the results, there would have to be a reduction of about 40% for the effect in fertility to disappear. Finally, firms would have to reduce new FTC hiring rate by 62% to completely offset the positive effect on fertility rate.

As evidenced by the first policy counterfactual, wage dynamics are also important determinants of the dual labour market. In an alternative policy experiment I simulate the effect of taxing fixed-term contracts and the effect of compensating this taxation by a credit on permanent contracts. According to the simulation results, this policy would bring no effect to the percentage of childless women, but rather change the composition of fertility in the households by increasing the number of mothers with two children, as opposed to one. Hence, the results of the counterfactual analysis corroborate the evidence found in the reduced-form results: job security plays a role in fertility decisions, especially for the first child, whereas income seems to be more relevant for deciding the number of children.

Appendix of Chapter 3

C1 Descriptive Statistics: ECHP and LFS

Table 3.14: Comparisson between ECHP and LFS - Labour Market Indicators

	ECHP		**LFS**	
Variable	**Mean**	**Std Dev**	**Mean**	**Std Dev**
University degree	0.105	0.307	0.106	0.308
Not working	0.325	0.469	0.343	0.475
Unemployment Duration	1.907	0.946	1.857	0.856
FTC (1 year)	0.081	0.274	0.073	0.261
FTC (2 years)	0.025	0.155	0.046	0.210
0 FTC Renewals	0.872	0.334	0.866	0.341
1 FTC Renewals	0.059	0.236	0.099	0.299
2 FTC Renewals	0.069	0.253	0.035	0.184
PC	0.894	0.308	0.880	0.325
Contract tenure 1 year	0.200	0.400	0.215	0.411
Contract tenure 2 years	0.064	0.245	0.075	0.263
Contract tenure 3+ years	0.735	0.441	0.710	0.454
N. individuals	2283		34988	
N. panel observations	10277		125484	

Note: read section 3.3.3 for explanation of main differences

Table 3.13: Descriptive statistics of ECHP

Variable	**Mean**	**Std Dev**	**Min**	**Max**
Age	35.479	8.174	23	50
University degree	0.105	0.307	0	1
Not working	0.325	0.469	0	1
Working				
FTC (1 year)	0.081	0.274	0	1
FTC (2 years)	0.025	0.155	0	1
0 FTC Renewals	0.872	0.334	0	1
1 FTC Renewals	0.059	0.236	0	1
2 FTC Renewals	0.069	0.253	0	1
PC	0.894	0.308	0	1
Contract tenure 1 year	0.200	0.400	0	1
Contract tenure 2 years	0.064	0.245	0	1
Contract tenure 3+ years	0.735	0.441	0	1
Full-time Monthly Wage in 2000's euros	476.745	269.561	246	2544
With partner	0.723	0.447	0	1
No children	0.338	0.473	0	1
Mothers				
1 child	0.473	0.499	0	1
2 children	0.527	0.499	0	1
Years in sample	4.502	1.559	2	6
N. individuals	2283			
N. panel observations	10277			

Note: the minimum value observed for wages is the minimum wage in 1994 in 2000's euros.

Table 3.15: Comparisson between ECHP and LFS - Labour Market Flows

From/To	Non-employment	FTC (1 year)	FTC (2 years)	PC
ECHP				
Non-employment	0.858	0.040	0.007	0.095
FTC (1 year)	0.145	0.501	0.063	0.290
FTC (2 years)	0.051	0.095	0.416	0.438
PC	0.049	0.015	0.005	0.932
ES				
Non-employment	0.876	0.053	0.017	0.053
FTC (1 year)	0.171	0.551	0.119	0.159
FTC (2 years)	0.071	0.050	0.656	0.223
PC	0.032	0.010	0.005	0.954

Note: read section 3.3.3 for explanation of main differences

C2 Auxiliary Data

To calibrate the childcare costs in the model, I used the National Household Budget Survey of 1994/1995, 2000, and 2005/2006. The first two years calibrated the baseline version of the model as well as the policy simulations whereas the latter was used in the validation of the model. This cross sectional survey is collected every 5 years by the National Statistics Institute, in Portugal. Each household is asked about their expenditure, net revenues, and household composition. Unfortunately the survey does not provide monetary information on assets but rather descriptive information on the house and durables.

For each survey year I have selected the households based on the same criteria as for the ECHP and LFS, presented in subsection3.3.3: households with at least one female between 23 and 50 years old that is either employed under a fixed-term or permanent contract, or not employed. To construct the measures of interest I used all the expenses related to (both formal and informal) education of children.

To grasp some insights about women's preferences I have used the National Fertility and Family Survey for the years of 1997 and 2013. This cross sectional survey has no fixed periodicity and was not collected in any other years. Even though my model is designed according to the characteristics of the Portuguese Economy during the period of 1994-1999, the second wave of this survey is particularly useful as it also contains information on the type of the job contract.

The National Fertility and Family Survey gives information on household composition, education, labour market status, respondent's parents, marital history, children history, desired children, fertility disruptions, and general opinions about family.

C3 Conditional Value Functions

According to the model presented in the paper, the number of choice possibilities in each period depends on the existence of a partner and on the arrival of job offers/renewals and not necessarily on the choice that was made in the previous period.[44] Hence, for simplicity in this section, the conditional value functions related to fertility choices will not incorporate the labour supply decision and vice-versa.[45]

In what follows $V_{i,t}^{N_k}$ denotes the value function associated to the decision of choosing to giving the k$^{\text{th}}$ birth in the next period. Not choosing to have a new child is denoted by $\bar{N}$. In case the woman gives birth, the state space will be denoted as $\Omega_{i,t}^{k}$.

In case the woman is not married, she will get a partner with probability μ, which depends on the age of the woman and on the duration of the job contract the woman holds in that period. In case she is married she will lose the partner with probability ρ, which solely depends on the number of children.[46]

Value of being single

$$V_{i,t}^{\bar{m}}(\Omega_{it}) = \; U_{it}^{\bar{m}}(\Omega_{it}) + \beta\Big\{\mu\,\mathbb{E}\max\Big[V_{t+1}^{N}(\Omega_{t+1}), V_{t+1}^{\bar{N}}(\Omega_{t+1})\Big] + (1-\mu)\,\mathbb{E}\Big[V_{t+1}^{\bar{N}}(\Omega_{t+1})\Big]\Big\}$$

Value of having a partner

$$V_{i,t}^{m}(\Omega_{it}) = \; U_{it}^{m}(\Omega_{it}) + \beta\Big\{\rho\,\mathbb{E}\max\Big[V_{t+1}^{\bar{N}}(\Omega_{t+1})\Big] + (1-\rho)\,\mathbb{E}\Big[V_{t+1}^{N}(\Omega_{t+1}), V_{t+1}^{\bar{N}}(\Omega_{t+1})\Big]\Big\}$$

Value of choosing to have the first child

$$V_{i,t}^{N_1}(\Omega_{it}) = \; U_{it}(\Omega_{it}) + \beta\Big\{\,\mathbb{E}\max\Big[V_{t+1}^{N_2}(\Omega_{t+1}^{k}), V_{t+1}^{\bar{N}}(\Omega_{t+1}^{k})\Big]\Big\}$$

[44] Note that because each period is one year and the maternity leave in Portugal is no more than 4 months then I can allow for the woman to give birth and work in the same period.

[45] Note that conditional value functions related with fertility decisions are assumed to be zero once the woman has two children and/or is older than 40 years old.

[46] I have tried different reduced form specifications for the probability of divorce and the years of marriage turned out to be non-significant in all cases.

Value of choosing to have the second child

$$V_{i,t}^{N_2}(\Omega_{it}) = \quad U_{it}(\Omega_{it}) + \beta\Big\{ \mathbb{E}\left[V_{t+1}^{\bar{N}}(\Omega_{t+1}^{k})\right] \Big\}$$

In terms of labour supply conditional values functions I denote $V_{i,t}^{P_{d,r}}$ denotes the value function associated to the decision of working on a contract of duration d and renewals r. Not working is denoted by $\bar{P}$. In case the woman is receiving unemployment benefits, the state space will be denoted as $\Omega_{i,t}^{ub}$.

In case the woman decides to work she is fired with a probability δ, which depends on the duration of the current contract, the tenure of the current contract and the number of renewals. In every period, regardless of her employment status, each woman receives at most one offer with duration D with probability λ^D, which depends on the duration of the last contract in case she was just fired (if not this duration is set to zero), the tenure of the current contract in case she was just fired or the number of years in unemployment in case she was fired in a previous period, and her age.

Value of working on a fixed-term contract with $d = 1, r = 0$

$$\begin{aligned} V_{i,t}^{P_{1,0}}(\Omega_{it}) = \; & U_{it}^{P_{1,0}}(\Omega_{it}) + \beta\Bigg\{ \left(1 - \sum_{D=1}^{3} \lambda_D\right) \mathbb{E}\left[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub})\right] \\ & +\delta \sum_{D=1}^{3} \left(\lambda^D \,\mathbb{E}\max\left[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\right) \\ & +(1-\delta) \sum_{D=1}^{3} \left(\lambda_D \,\mathbb{E}\max\left[V_{t+1}^{P_{D,1^*}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\right) \Bigg\} \end{aligned}$$

Note that when the renewal is to a 2-year contract, the firm cannot use more renewals at the end of that contract

Value of working on a fixed-term contract with $d = 1, r = 1$

$$V_{i,t}^{P_{1,1}}(\Omega_{it}) = U_{it}^{P_{1,1}}(\Omega_{it}) + \beta \Bigg\{ \Bigg[\Bigg(1 - \sum_{D=1}^{3} \lambda_D \Bigg) + (1-\delta)\lambda_2 \Bigg] \mathbb{E} \Big[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub}) \Big]$$

$$+\delta \sum_{D=1}^{3} \Big(\lambda_D \, \mathbb{E} \max \Big[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Big)$$

$$+(1-\delta)\Big(\lambda_1 \, \mathbb{E} \max \Big[V_{t+1}^{P_{1,2}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] +$$

$$\lambda_3 \, \mathbb{E} \max \Big[V_{t+1}^{P_{3,1}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Big) \Bigg\}$$

Value of working on a fixed-term contract with $d = 1, r = 2$

$$V_{i,t}^{P_{1,0}}(\Omega_{it}) = U_{it}^{P_{1,0}}(\Omega_{it}) + \beta \Bigg\{ \Bigg[\Bigg(1 - \sum_{D=1}^{3} \lambda_D \Bigg) + (1-\delta)(\lambda_1 + \lambda_2) \Bigg] \mathbb{E} \Big[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub}) \Big]$$

$$+\delta \sum_{D=1}^{3} \Big(\lambda_D \, \mathbb{E} \max \Big[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Big)$$

$$+(1-\delta)\lambda_3 \, \mathbb{E} \max \Big[V_{t+1}^{P_{3,2}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Bigg\}$$

Value of working on a fixed-term contract with $d = 2, r = 0, x = 1$

$$V_{i,t}^{P_{2,0}}(\Omega_{it}) = U_{it}^{P_{2,0}}(\Omega_{it}) + \beta\Bigg\{ \Bigg[\Bigg(1 - \sum_{D=1}^{3} \lambda_D\Bigg) + (1-\delta)\lambda_1\Bigg] \mathbb{E}\left[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub})\right]$$

$$+\delta \sum_{D=1}^{3} \left(\lambda_D \,\mathbb{E} \max\left[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\right)$$

$$+(1-\delta)\Big(\lambda_2 \,\mathbb{E} \max\left[V_{t+1}^{P_{2,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right] +$$

$$\lambda_3 \,\mathbb{E} \max\left[V_{t+1}^{P_{3,1}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\Big)\Bigg\}$$

Value of working on a fixed-term contract with $d = 2, r = 0, x = 2$

$$V_{i,t}^{P_{2,0}}(\Omega_{it}) = U_{it}^{P_{2,0}}(\Omega_{it}) + \beta\Bigg\{ \Bigg[\Bigg(1 - \sum_{D=1}^{3} \lambda_D\Bigg) + (1-\delta)\lambda_2\Bigg] \mathbb{E}\left[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub})\right]$$

$$+\delta \sum_{D=1}^{3} \left(\lambda_D \,\mathbb{E} \max\left[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\right)$$

$$+(1-\delta)\Big(\lambda_1 \,\mathbb{E} \max\left[V_{t+1}^{P_{1,2}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right] +$$

$$\lambda_3 \,\mathbb{E} \max\left[V_{t+1}^{P_{3,1}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1})\right]\Big)\Bigg\}$$

Value of working on a fixed-term contract with $d = 2, r = 1$

$$V_{i,t}^{P_{2,0}}(\Omega_{it}) = U_{it}^{P_{1,0}}(\Omega_{it}) + \beta \Bigg\{ \Bigg[\Bigg(1 - \sum_{D=1}^{3} \lambda_D \Bigg) + (1-\delta)(\lambda_1 + \lambda_2) \Bigg] \mathbb{E} \Big[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub}) \Big]$$

$$+\delta \sum_{D=1}^{3} \Big(\lambda_D \, \mathbb{E} \max \Big[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Big)$$

$$+(1-\delta)\lambda_3 \, \mathbb{E} \max \Big[V_{t+1}^{P_{3,2}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Bigg\}$$

Value of working on a permanent job contract $(d = 3)$

$$V_{i,t}^{P_3}(\Omega_{it}) = U_{it}^{P_3}(\Omega_{it}) + \beta \Bigg\{ \delta \Bigg(1 - \sum_{D=1}^{3} \lambda_D \Bigg) \mathbb{E} \Big[V_{t+1}^{\bar{P}}(\Omega_{t+1}^{ub}) \Big]$$

$$+\delta \sum_{D=1}^{3} \Big(\lambda_D \, \mathbb{E} \max \Big[V_{t+1}^{P_{D,0}}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Big)$$

$$+(1-\delta) \, \mathbb{E} \max \Big[V_{t+1}^{P_3}(\Omega_{it+1}), V_{t+1}^{\bar{P}}(\Omega_{it+1}) \Big] \Bigg\}$$

C4 Auxiliary Estimations

Externally set parameters: I set the discount factor β to 0.98 (see Blundell et al. (2016a) and Attanasio et al. (2008)). I set the mean of the extreme value distribution to 0 and the scale to 0.2. I also set the children's cost to be the averages spent by age of youngest child and number of children from the Family Expenditure Survey of 1995. The probabilities for job offers are estimated from the Labour Force Survey sample. Results are not presented here for presentational purposes, but are available upon request.

Table 3.16: Partner probabilities estimates:

	P(Marriage)	P(Divorce)
Age	0.041	
	(0.069)	
Age^2	-0.008^{*}	
	(0.005)	
PC	0.694^{***}	
	(0.185)	
FTC	0.498^{*}	
	(0.261)	
N. Kids		-0.776^{***}
		(0.164)
Constant	-2.450^{***}	-3.373^{***}
	(0.219)	(0.198)
Observations	1799	4338

Standard errors in parentheses

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

Table 3.17: Lay-off probability estimates:

	P(Laid-off)
Univ. Degree	-1.806***
	(0.273)
Tenure	-0.660***
	(0.048)
FTC (2 years)	-0.423***
	(0.162)
PC	-1.320***
	(0.104)
Constant	-0.670***
	(0.096)
Observations	23826

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

Table 3.18: Wages estimates:

	FTC Wage	PC Wage	Husband's Wage
𝟙 FTC (2 years)	0.069**		
	(0.034)		
Nr. renewals	0.022		
	(0.025)		
𝟙 Univ. Degree	0.491***	0.828***	0.548***
	(0.034)	(0.050)	(0.018)
Age		.015***	.010 ***
		(0.003)	(0.001)
Age^2		-0.002*	
		(0.000)	
Constant	5.725***	5.369***	6.045***
	(0.053)	(0.076)	(0.018)
N	415	4651	4029

All regressions refer to full time wages and control for time effects.

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

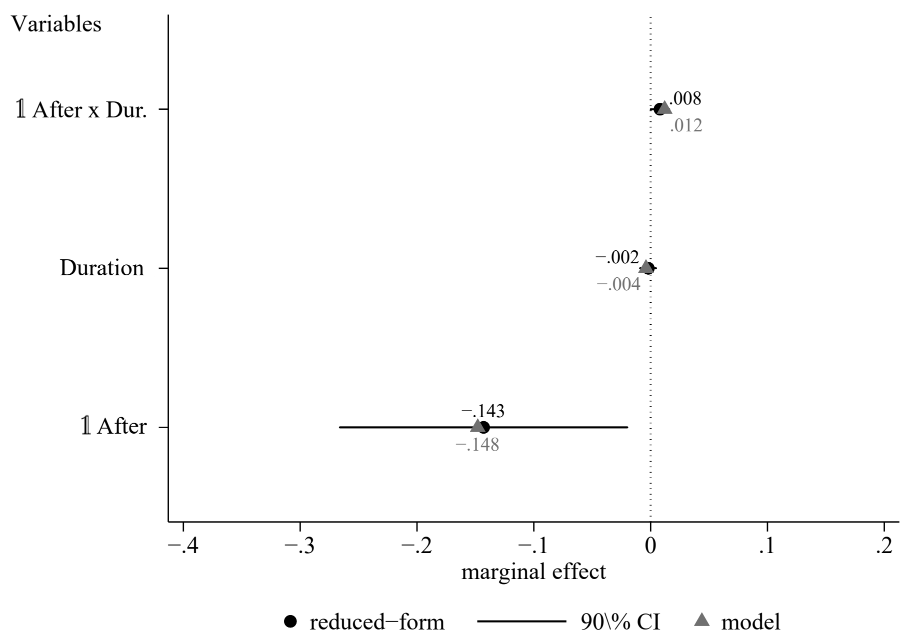
Table 3.19: Selection Equations Estimates:

	FTC		PC	
Number of children	-.193***	-.211***	-.077***	-.086***
	(.053)	(.054)	(.022)	(.022)
𝟙 Married	.161*	.164*	.037	-.018
	(.087)	(.09)	(.038)	(.039)
Personal non-labour income (000s €)		-.145		-.033
		(.167)		(.033)
Social Transfers (000s €)		-.001		-.042***
		(.012)		(.005)
𝟙 Rent-Free House		.052		-.107**
		(.106)		(.047)
𝟙 Rented House		.459***		.132***
		(.078)		(.036)
𝟙 Second Home		.073		.046
		(.125)		(.055)
Mills Ratio	-.01	.013	-.016	-.023
	(.039)	(.036)	(.095)	(.047)
N	4021		8257	

All regressions include the explanatory variables from the wage equations as well as age and time effects. The instruments included in first and third columns are the ones included in the structural model as well, whereas the instruments included in second and fourth columns exist in the data but are not part of the model. The latter serve as robustness checks for selection. The base group for house ownership is owners with no second-house. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

C5 Model Validation - Replicating the reduced-form results

Figure 3.11: Model vs Reduced-form in Itention to Treat Exercise: Coefficients of Interest



Bibliography

Adamchik, V. (1999). The Effect of Unemployment Benefits on the Probability of Re-employment in Poland. *Oxford Bulletin of Economics and Statistics*, 61(1):95–108.

Adda, J., Dustmann, C., and Stevens, K. (2017). The Career Costs of Children. *Journal of Political Economy*, 125(2):293–337.

Addison, J. T., Machado, J. A. F., and Portugal, P. (2013). The Reservation Wage Unemployment Duration Nexus. *Oxford Bulletin of Economics and Statistics*, 75(6):980–987.

Addison, J. T. and Portugal, P. (1989). Job Displacement, Relative Wage Changes, and Duration of Unemployment. *Journal of Labor Economics*, 7(3):281–302.

Addison, J. T. and Portugal, P. (2003). Unemployment Duration: Competing and Defective risks. *Journal of Human Resources*, 38(1):156–191.

Addison, J. T. and Portugal, P. (2004). How Does the Unemployment Insurance System Shape the Time Profile of Jobless Duration? *Economics Letters*, 85(2):229–234.

Addison, J. T. and Portugal, P. (2008a). How Do Different Entitlements to Unemployment Benefits Affect the Transitions from Unemployment into Employment? *Economics Letters*, 101(3):206–209.

Addison, J. T. and Portugal, P. (2008b). Six Ways to Leave Unemployment. *Scottish Journal of Political Economy*, 55(4):393–419.

Adsera, A. (2004). Changing Fertility Rates in Developed Countries. the Impact of Labor Market Institutions. *Journal of Population Economics*, 17(1):17–43.

Allison, P. D. (1982). Discrete-time methods for the analysis of event histories. *Sociological methodology*, 13:61–98.

Altonji, J. G. and Shakotko, R. A. (1987). Do Wages Rise with Job Seniority? *The Review of Economic Studies*, 54(3):437–459.

André, I. M. (1991). The Employment of Women in Portugal. *Iberian Studies*, 20(1-2):28–41.

Arranz, J. and Muro, J. (2004). An Extra Time Duration Model with Application to Unemployment Duration under Benefits in Spain. *Hacienda Pública Española*, 171(4):133–156.

Attanasio, O., Low, H., and Sánchez-Marcos, V. (2008). Explaining Changes in Female Labor Supply in a Life-Cycle Model. *The American Economic Review*, 98(4):1517–1552.

Auer, W. and Danzer, N. (2016). Fixed-Term Employment and Fertility: Evidence from German Micro Data. *CESifo Economic Studies*, 62(4):595–623.

Becker, G. S. (1960). An Economic Analysis of Fertility. In *Demographic and economic change in developed countries*, pages 209–240.

Becker, G. S. (1962). Investment in Human Capital: A Theoretical Analysis. *Journal of Political Economy*, 70(5, Part 2):9–49.

Becker, G. S. (1965). A Theory of the Allocation of Time. *The Economic Journal*, 75(299):493–517.

Becker, G. S. and Tomes, N. (1976). Child Endowments and the Quantity and Quality of Children. *Journal of Political Economy*, 84(4, Part 2):S143–S162.

Belzil, C. (1995). Unemployment Duration Stigma and Re-Employment Earnings. *Canadian Journal of Economics*, 28(3):568–585.

Belzil, C. (2001). Unemployment Insurance and Subsequent Job Duration: Job Matching Versus Unobserved Heterogeneity. *Journal of Applied Econometrics*, 16(5):619–636.

Bentolila, S. and Bertola, G. (1990). Firing Costs and Labour Demand: How Bad Is Eurosclerosis? *The Review of Economic Studies*, 57(3):381–402.

Blackburn, M. L., Bloom, D. E., and Neumark, D. (1993). Fertility Timing, Wages, and Human Capital. *Journal of Population Economics*, 6(1):1–30.

Blanchard, O. (2007). Adjustment Within the Euro. the Difficult Case of Portugal. *Portuguese Economic Journal*, 6(1):1–21.

Blanchard, O. and Portugal, P. (2001). What Hides Behind an Unemployment Rate: Comparing Portuguese and Us Labor Markets. *American Economic Review*, 91(1):187–207.

Blanchard, O. J. and Diamond, P. (1994). Ranking, Unemployment Duration, and Wages. *The Review of Economic Studies*, 61(3):417–434.

Blanchard, O. J. and Tirole, J. (2008). The Joint Design of Unemployment Insurance and Employment Protection: A First Pass. *Journal of the European Economic Association*, 6(1):45–77.

Blundell, R., Costa Dias, M., Meghir, C., and Shaw, J. (2016a). Female Labor Supply, Human Capital, and Welfare Reform. *Econometrica*, 84(5):1705–1753.

Blundell, R., Pistaferri, L., and Saporta-Eksten, I. (2016b). Consumption Inequality and Family Labor Supply. *American Economic Review*, 106(2):387–435.

Boeri, T. et al. (2011). Institutional Reforms and Dualism in European Labor Markets. *Handbook of Labor Economics*, 4:1173–1236.

Boone, J. and van Ours, J. C. (2012). Why is There a Spike in the Job Finding Rate at Benefit Exhaustion? *De Economist*, 160(4):413.

Booth, A. L., Dolado, J. J., and Frank, J. (2002a). Symposium on Temporary Work: Introduction. *The Economic Journal*, 112(480):F181–F188.

Booth, A. L., Francesconi, M., and Frank, J. (2002b). Temporary Jobs: Stepping Stones or Dead Ends? *The Economic Journal*, 112(480):F189–F213.

Bover, O., Arellano, M., and Bentolila, S. (2002). Unemployment Duration, Benefit Duration and the Business Cycle. *The Economic Journal*, 112(479):223–265.

Burda, M. C. and Mertens, A. (2001). Estimating Wage Losses of Displaced Workers in Germany. *Labour Economics*, 8(1):15–41.

Burniaux, J.-M. et al. (1998). Income Distribution and Poverty in Selected OECD Countries. OECD Economics Department Working Papers, No. 189, OECD Publishing, Paris.

Cahuc, P., Charlot, O., Malherbet, F., Benghalem, H., and Limon, E. (2016). Taxation of Temporary Jobs: Good Intentions with Bad Outcomes? IZA Discussion Paper No. 10352.

Cahuc, P. and Postel-Vinay, F. (2002). Temporary Jobs, Employment Protection and Labor Market Performance. *Labour Economics*, 9(1):63–91.

Caliendo, M., Tatsiramos, K., and Uhlendorff, A. (2013). Benefit Duration, Unemployment Duration and Job Match Quality: A Regression-Discontinuity Approach. *Journal of Applied Econometrics*, 28(4):604–627.

Card, D., Chetty, R., and Weber, A. (2007a). Cash-On-Hand and Competing Models of Intertemporal Behavior: New Evidence from the Labor Market. *The Quarterly Journal of Economics*, 122(4):1511–1560.

Card, D., Chetty, R., and Weber, A. (2007b). The Spike at Benefit Exhaustion: Leaving the Unemployment System or Starting a New Job? *American Economic Review*, 97(2):113–118.

Card, D. and Levine, P. B. (2000). Extended Benefits and the Duration of UI Spells: Evidence from the New Jersey Extended Benefit Program. *Journal of Public Economics*, 78(1):107–138.

Carling, K., Edin, P.-A., and Harkman, A. (1996). Unemployment Duration, Unemployment Benefits, and Labor Market Programs in Sweden. *Journal of Public Economics*, 59(3):313–334.

Carneiro, A., Portugal, P., and Varejão, J. (2014). Catastrophic Job Destruction During the Portuguese Economic Crisis. *Journal of Macroeconomics*, 39(PB):444–457.

Carneiro, P. and Ginja, R. (2016). Partial Insurance and Investments in Children. *The Economic Journal*, 126(596):F66–F95.

Centeno, M. and Novo, Á. A. (2009). Reemployment Wages and Ui Liquidity Effect: A Regression Discontinuity Approach. *Portuguese Economic Journal*, 8(1):45–52.

Centeno, M. and Novo, Á. A. (2014). Do Low-Wage Workers React Less to Longer Unemployment Benefits? Quasi-Experimental Evidence. *Oxford Bulletin of Economics and Statistics*, 76(2):185–207.

Chiswick, B. R. (1986). Labor Supply and Investment in Child Quality: A Study of Jewish and Non-Jewish Women. *The Review of Economics and Statistics*, 74(4):700–703.

Cigno, A. and Ermisch, J. (1989). A Microeconomic Analysis of the Timing of Births. *European Economic Review*, 33(4):737–760.

Couch, K. A., Jolly, N. A., and Placzek, D. W. (2011). Earnings Losses of Displaced Workers and the Business Cycle: An Analysis with Administrative Data. *Economics Letters*, 111(1):16–19.

Couch, K. A. and Placzek, D. W. (2010). Earnings Losses of Displaced Workers Revisited. *The American Economic Review*, 100(1):572–589.

Cox, D. R. (1972). Regression Models and Life-Tables. *Journal of the Royal Statistical Society: Series B (Methodological)*, 34(2):187–202.

De la Rica, S. and Iza, A. (2005). Career Planning in Spain: Do Fixed-Term Contracts Delay Marriage and Parenthood? *Review of Economics of the Household*, 3(1):49–73.

Del Boca, D., Flinn, C., and Wiswall, M. (2014). Household Choices and Child Development. *The Review of Economic Studies*, 81(1):137–185.

Del Boca, D. and Lusardi, A. (2003). Credit Market Constraints and Labor Market Decisions. *Labour Economics*, 10(6):681–703.

Del Bono, E., Weber, A., and Winter-Ebmer, R. (2012). Clash of Career and Family: Fertility Decisions After Job Displacement. *Journal of the European Economic Association*, 10(4):659–683.

Dias da Silva, A. and Turrini, A. (2015). Precarious and Less Well Paid? Wage Differences between Permanent and Fixed-term Contracts across the EU. *Economic Papers*, 544(105):1–44.

Dolado, J. J., Lalé, E., and Siassi, N. (2018). From Dual to Unified Employment Protection: Transition and Steady State. Available at SSRN: https://ssrn.com/abstract=2702404.

Duffie, D. and Singleton, K. J. (1993). Simulated Moments Estimation of Markov Models of Asset Prices. *Econometrica*, 61(4):929–952.

Eckstein, Z. and Lifshitz, O. (2011). Dynamic Female Labor Supply. *Econometrica*, 79(6):1675–1726.

Eckstein, Z. and Wolpin, K. I. (1989). Dynamic Labour Force Participation of Married Women and Endogenous Work Experience. *The Review of Economic Studies*, 56(3):375–390.

Edwards, R. (2014). Women's Labor Supply-Motherhood and Work Schedule Flexibility. *mimeo*.

Eisenhauer, P., Heckman, J. J., and Mosso, S. (2015). Estimation of Dynamic Discrete Choice Models by Maximum Likelihood and the Simulated Method of Moments. *International Economic Review*, 56(2):331–357.

Erosa, A., Fuster, L., and Restuccia, D. (2010). A General Equilibrium Analysis of Parental Leave Policies. *Review of Economic Dynamics*, 13(4):742–758.

Fallick, B. C. (1991). Unemployment Insurance and the Rate of Re-Employment of Displaced Workers. *The Review of Economics and Statistics*, 73(2):228–235.

Farber, H. S. and Valletta, R. G. (2015). Do Extended Unemployment Benefits Lengthen Unemployment spells? Evidence from Recent Cycles in the US Labor Market. *Journal of Human Resources*, 50(4):873–909.

Filges, T., Geerdsen, L. P., Knudsen, A.-S. D., and Jørgensen, A.-M. K. (2015). Unemployment Benefit Exhaustion: Incentive Effects on Job-Finding Rates. *Research on Social Work Practice*, 25(1):21–43.

Francesconi, M. (2002). A Joint Dynamic Model of Fertility and Work of Married Women. *Journal of Labor Economics*, 20(2):336–380.

Frejka, T. (2008). Overview Chapter 2: Parity distribution and completed family size in Europe: Incipient decline of the two-child family model. *Demographic research*, 19(4):47–72.

Gagliarducci, S. (2005). The Dynamics of Repeated Temporary Jobs. *Labour Economics*, 12(4):429–448.

Gauthier, A. H. (2007). The Impact of Family Policies on Fertility in Industrialized Countries: A Review of the Literature. *Population Research and Policy Review*, 26(3):323–346.

Geerdsen, L. P., Lyk-Jensen, S. V., and Weatherall, C. D. (2018). Accelerating the Transition to Employment at Benefit Exhaustion: Still Possible After Four Years of Unemployment? *Empirical Economics*, 54(3):1107–1135.

Giersch, H. (1985). Euroscerosis. Kiel Discussion Papers, No 112.

Gonzalo, M. T. (2002). A New Look at the Ui Effect on Transitions from Unemployment into Wage Employment in Spain: The Limited Duration of the Ui Benefits Entitlement. *Applied Economics*, 34(17):2177–2187.

Gourieroux, C., Monfort, A., and Renault, E. (1993). Indirect inference. *Journal of Applied Econometrics*, 8(S1):S85–S118.

Gritz, R. M. and MaCurdy, T. (1992). Unemployment Compensation and Episodes of Nonemployment. *Empirical Economics*, 17(1):183–204.

Güell, M. and Rodríguez Mora, J. V. (2010). Temporary Contracts, Incentives and Unemployment. CEPR Discussion Paper No. 8116.

Guner, N., Kaya, E., and Sanchéz Marcos, V. (2018). Labor Market Frictions and Lowest Low Fertility. *mimeo*.

Haan, P. and Wrohlich, K. (2011). Can Child Care Policy Encourage Employment and Fertility?: Evidence from a Structural Model. *Labour Economics*, 18(4):498–512.

Hagenaars, A., de Vos, K., Zaidi, M., and of the European Communities, S. O. (1994). *Poverty Statistics in the Late 1980s: Research Based on Micro-data*.

Ham, J. C. and Rea Jr, S. A. (1987). Unemployment Insurance and Male Unemployment Duration in Canada. *Journal of Labor Economics*, 5(3):325–353.

Han, A. and Hausman, J. A. (1990). Flexible Parametric Estimation of Duration and Competing Risk Models. *Journal of Applied Econometrics*, 5(1):1–28.

Happel, S. K., Hill, J. K., and Low, S. A. (1984). An Economic Analysis of the Timing of Childbirth. *Population Studies*, 38(2):299–311.

Hausman, J. A. (1978). Specification tests in econometrics. *Econometrica: Journal of the econometric society*, pages 1251–1271.

Heckman, J. and Singer, B. (1984). A Method for Minimizing the Impact of Distributional Assumptions in Econometric Models for Duration Data. *Econometrica*, 52(2)(2):271–320.

Heckman, J. J. (1976). The Common Structure of Statistical Models of Truncation, Sample Selection and Limited Dependent Variables and a Simple Estimator for Such Models. In *Annals of Economic and Social Measurement*, volume 5, pages 475–492.

Heckman, J. J. (1979). Sample Selection Bias as a Specification Error. *Econometrica*, 47(1):153–161.

Heckman, J. J. and Walker, J. R. (1990). Estimating fecundability from data on waiting times to first conception. *Journal of the American Statistical Association*, 85(410):283–294.

Holen, A. (1977). Effects of Unemployment Insurance Entitlement on Duration and Job Search Outcome. *ILR Review*, 30(4):445–450.

Hotz, V. J. and Miller, R. A. (1988). An Empirical Analysis of Life Cycle Fertility and Female Labor Supply. *Econometrica*, 56(1):91–118.

Hunt, J. (1995). The Effect of Unemployment Compensation on Unemployment Duration in Germany. *Journal of Labor Economics*, 13(1):88–120.

Jenkins, S. P. (1995). Easy Estimation Methods for Discrete-Time Duration Models. *Oxford Bulletin of Economics and Statistics*, 57(1):129–136.

Jenkins, S. P. (2005). Survival Analysis. *Unpublished manuscript, Institute for Social and Economic Research, University of Essex, Colchester, UK*, 42:54–56.

Jenkins, S. P. and García-Serrano, C. (2004). The Relationship between Unemployment

Benefits and Re-employment Probabilities: Evidence from Spain. *Oxford Bulletin of Economics and Statistics*, 66(2):239–260.

Jurajda, Š. and Tannery, F. J. (2003). Unemployment Durations and Extended Unemployment Benefits in Local Labor Markets. *ILR Review*, 56(2):324–348.

Kalbfleisch, J. D. and Prentice, R. L. (2011). *The Statistical Analysis of Failure Time Data*, volume 360.

Katz, L. F. and Meyer, B. D. (1990a). The Impact of the Potential Duration of Unemployment Benefits on the Duration of Unemployment. *Journal of Public Economics*, 41(1):45–72.

Katz, L. F. and Meyer, B. D. (1990b). Unemployment Insurance, Recall Expectations, and Unemployment Outcomes. *The Quarterly Journal of Economics*, 105(4):973–1002.

Keane, M. P. and Wolpin, K. I. (2010). The Role of Labor and Marriage Markets, Preference Heterogeneity, and the Welfare System in the Life Cycle Decisions of Black, Hispanic, and White Women. *International Economic Review*, 51(3):851–892.

Lalive, R. (2007). Unemployment Benefits, Unemployment Duration, and Post-Unemployment Jobs: A Regression Discontinuity Approach. *The American Economic Review*, 97(2):108–112.

Lalive, R. (2008). How do Extended Benefits Affect Unemployment Duration? A Regression Discontinuity Approach. *Journal of Econometrics*, 142(2):785–806.

Lalive, R., Schlosser, A., Steinhauer, A., and Zweimüller, J. (2014). Parental Leave and Mothers' Careers: The Relative Importance of Job Protection and Cash Benefits. *Review of Economic Studies*, 81(1):219–265.

Lalive, R., Van Ours, J., and Zweimüller, J. (2006). How Changes in Financial Incentives Affect the Duration of Unemployment. *The Review of Economic Studies*, 73(4):1009–1038.

Lalive, R. and Zweimüller, J. (2004). Benefit Entitlement and Unemployment Duration: The Role of Policy Endogeneity. *Journal of Public Economics*, 88(12):2587–2616.

Lalive, R. and Zweimüller, J. (2009). How Does Parental Leave Affect Fertility and Return to Work? Evidence from Two Natural Experiments. *The Quarterly Journal of Economics*, 124(3):1363–1402.

Lancaster, T. (1979). Econometric methods for the duration of unemployment. *Econometrica: Journal of the Econometric Society*, pages 939–956.

Layard, R., Nickell, S., and Jackman, R. (1991). *Macroeconomic Performance and the Labour Market*.

Lazear, E. P. and Rosen, S. (1990). Male-Female Wage Differentials in Job Ladders. *Journal of Labor Economics*, 8(1):S106–S123.

Le Barbanchon, T. (2016). The Effect of the Potential Duration of Unemployment Benefits on Unemployment Exits to Work and Match Quality in France. *Labour Economics*, 42:16–29.

Lepage-Saucier, N., Schleich, J., and Wasmer, E. (2013). Moving Towards a Single Labour Contract: Pros, Cons and Mixed Feelings.

Lindeboom, M. and Theeuwes, J. (1993). Search, Benefits and Entitlement. *Economica*, 60(239):327–346.

Lise, J., Seitz, S., and Smith, J. (2004). Equilibrium Policy Experiments and the

Evaluation of Social Programs. National Bureau of Economic Research Working Paper 10283.

Ljungqvist, L. and Sargent, T. J. (2008). Two Questions About European Unemployment. *Econometrica*, 76(1):1–29.

Llull, J. and Miller, R. (2016). Internal Migration in Dual Labor Markets. Society for Economic Dynamics.

Low, H., Meghir, C., Pistaferri, L., and Voena, A. (2018). Marriage, Labor Supply and the Dynamics of the Social Safety Net. National Bureau of Economic Research, Working Paper No. 24356.

Lubyova, M. and Van Ours, J. (1997). Unemployment Dynamics and the Restructuring of the Slovak Unemployment Benefit System. *European Economic Review*, 41(3-5):925–934.

Meyer, B. D. (1990). Unemployment Insurance and Unemployment Spells. *Econometrica*, 58(4):757–782.

Milligan, K. (2005). Subsidizing the Stork: New Evidence on Tax Incentives and Fertility. *Review of Economics and Statistics*, 87(3):539–555.

Moffitt, R. (1984). Profiles of Fertility, Labour Supply and Wages of Married Women: A Complete Life-Cycle Model. *The Review of Economic Studies*, 51(2):263–278.

Moffitt, R. (1985). Unemployment Insurance and the Distribution of Unemployment Spells. *Journal of Econometrics*, 28(1):85–101.

Moffitt, R. and Nicholson, W. (1982). The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits. *The Review of Economics and Statistics*, 64(1):1–11.

Mortensen, D. T. (1977). Unemployment Insurance and Job Search Decisions. *ILR Review*, 30(4):505–517.

Mortensen, D. T. (1988). Matching: Finding a Partner for Life or Otherwise. *American Journal of Sociology*, 94:S215–S240.

Narendranathan, W. and Stewart, M. B. (1993). How Does the Benefit Effect Vary as Unemployment Spells Lengthen? *Journal of Applied Econometrics*, 8(4):361–381.

Neal, D. (1995). Industry-specific human capital: Evidence from displaced workers. *Journal of labor Economics*, 13(4):653–677.

Nekoei, A. and Weber, A. (2017). Does Extending Unemployment Benefits Improve Job Quality? *The American Economic Review*, 107(2):527–561.

Nickell, S. (1979). Estimating the Probability of Leaving Unemployment. *Econometrica*, 47(5):1249–1266.

Nickell, S. (1997). Unemployment and Labor Market Rigidities: Europe Versus North America. *Journal of Economic perspectives*, 11(3):55–74.

Nickell, S., Nunziata, L., and Ochel, W. (2005). Unemployment in the OECD Since the 1960s. What Do We Know? *The Economic Journal*, 115(500):1–27.

Nickell, S. J. (1978). Fixed Costs, Employment and Labour Demand over the Cycle. *Economica*, 45(180):329–345.

OECD (2018). *OECD Employment Outlook 2018*.

Olivetti, C. and Petrongolo, B. (2017). The Economic Consequences of Family Policies: Lessons from a Century of Legislation in High-Income Countries. *Journal of Economic Perspectives*, 31(1):205–30.

Pakes, A. and Pollard, D. (1989). Simulation and the Asymptotics of Optimization Estimators. *Econometrica*, pages 1027–1057.

Pellizzari, M. (2006). Unemployment Duration and the Interactions between Unemployment Insurance and Social Assistance. *Labour Economics*, 13(6):773–798.

Pérez, J. I. G. and Osuna, V. (2014). Dual Labour Markets and the Tenure Distribution: Reducing Severance Pay or Introducing a Single Contract. *Labour Economics*, 29:1–13.

Pinheiro, R. and Visschers, L. (2015). Unemployment Risk and Wage Differentials. *Journal of Economic Theory*, 157:397–424.

Portugal, P. (2006). Wage Setting in the Portuguese Labor Market: A Microeconomic Approach. *Banco de Portugal Economic Bulletin, Autumn 2006*, 78:89–100.

Portugal, P. (2008). Unemployment Duration in the Portuguese Labour Market. *Banco de Portugal Economic Bulletin, Winter 2008*, pages 55–72.

Portugal, P. and Varejão, J. (2004). Matching Workers to Jobs in the Fast Lane: the Operation of Fixed-term Contracts.

Portugal, P. and Varejão, J. (2009). Why Do Firms Use Fixed-Term Contracts? IZA Discussion Paper No. 4380.

Portugal, P. and Varejão, J. (2010). The Hidden Side of Temporary Employment: Fixed-Term Contracts as a Screening Device. Banco de Portugal, Economics and Research Department, Working Papers, No. 2010-29.

Powell, M. J. (2009). The BOBYQA Algorithm for Bound Constrained Optimization Without Derivatives. University of Cambridge, Cambridge, Cambridge NA Report NA2009/06.

Puhani, P. A. (2000). Poland on the Dole: The Effect of Reducing the Unemployment Benefit Entitlement Period During Transition. *Journal of Population Economics*, 13(1):35–44.

Raposo, P. S., Portugal, P., and Carneiro, A. J. M. (2019). Decomposing the wage losses of displaced workers: the role of the reallocation of workers into firms and job titles. *Journal of Human Resources (Forthcoming).*

Rebollo-Sanz, Y. (2012). Unemployment Insurance and Job Turnover in Spain. *Labour Economics*, 19(3):403–426.

Robalo Marques, C., Martins, F., and Portugal, P. (2010). Price and Wage Formation in Portugal. European Central Bank, Working Paper Series No. 1225.

Røed, K. and Zhang, T. (2003). Does Unemployment Compensation Affect Unemployment Duration? *The Economic Journal*, 113(484):190–206.

Rogers, C. L. (1998). Expectations of Unemployment Insurance and Unemployment Duration. *Journal of Labor Economics*, 16(3):630–666.

Rotnitzky, A., Robins, J. M., and Scharfstein, D. O. (1998). Semiparametric Regression for Repeated Outcomes with Nonignorable Nonresponse. *Journal of the American Statistical Association*, 93(444):1321–1339.

Ruhm, C. J. (1998). The Economic Consequences of Parental Leave Mandates: Lessons from Europe. *The Quarterly Journal of Economics*, 113(1):285–317.

Schmidt, P. and Witte, A. D. (1989). Predicting Criminal Recidivism Using "Split Population" Survival Time Models. *Journal of Econometrics*, 40(1):141–159.

Schmieder, J. F., von Wachter, T., and Bender, S. (2012). The Long-Term Effects of UI Extensions on Employment. *American Economic Review*, 102(3):514–19.

Schmieder, J. F., von Wachter, T., and Bender, S. (2016). The Effect of Unemployment Benefits and Nonemployment Durations on Wages. *American Economic Review*, 106(3):739–77.

Schmitz, H. and Steiner, V. (2007). Benefit-Entitlement Effects and the Duration of Unemployment: An Ex-Ante Evaluation of Recent Labour Market Reforms in Germany. IZA Discussion Paper No. 2681.

Schönberg, U. and Ludsteck, J. (2014). Expansions in maternity leave coverage and mothers' labor market outcomes after childbirth. *Journal of Labor Economics*, 32(3):469–505.

Sestito, P. and Viviano, E. (2016). Hiring Incentives And/Or Firing Cost Reduction? Evaluating the Impact of the 2015 Policies on the Italian Labour Market. Bank of Italy Occasional Paper No. 325.

Shapiro, C. and Stiglitz, J. E. (1984). Equilibrium Unemployment as a Worker Discipline Device. *The American Economic Review*, 74(3):433–444.

Sheran, M. (2007). The Career and Family Choices of Women: A Dynamic Analysis of Labor Force Participation, Schooling, Marriage, and Fertility Decisions. *Review of Economic Dynamics*, 10(3):367–399.

Sommer, K. (2016). Fertility Choice in a Life Cycle Model with Idiosyncratic Uninsurable Earnings Risk. *Journal of Monetary Economics*, 83:27–38.

Steiner, V. (2001). Unemployment Persistence in the West German Labour Market: Negative Duration Dependence or Sorting? *Oxford Bulletin of Economics and Statistics*, 63(1):91–113.

Stigler, G. J. (1961). The Economics of Information. *Journal of Political Economy*, 69(3):213–225.

Stigler, G. J. (1962). Information in the Labor Market. *Journal of Political Economy*, 70(5, Part 2):94–105.

Stovicek, K., Turrini, A., et al. (2012). Benchmarking unemployment benefit systems. Directorate General Economic and Financial Affairs (DG ECFIN), European Commission, No. 454.

Terrell, K. and Sorm, V. (1999). Labor Market Policies and Unemployment in the Czech Republic. *Journal of Comparative Economics*, 27(1):33–60.

Thévenon, O. and Luci, A. (2012). Reconciling Work, Family and Child Outcomes: What Implications for Family Support Policies? *Population Research and Policy Review*, 31(6):855–882.

Van den Berg, G. J. (1990). Nonstationarity in Job Search Theory. *The Review of Economic Studies*, 57(2):255–277.

Van der Klaauw, W. (1996). Female Labour Supply and Marital Status Decisions: A Life-Cycle Model. *The Review of Economic Studies*, 63(2):199–235.

Van Ours, J. C. and Vodopivec, M. (2006). How Shortening the Potential Duration of Unemployment Benefits Affects the Duration of Unemployment: Evidence from a Natural Experiment. *Journal of Labor Economics*, 24(2):351–378.

Van Ours, J. C. and Vodopivec, M. (2008). Does Reducing Unemployment Insurance Generosity Reduce Job Match Quality? *Journal of Public Economics*, 92(3-4):684–695.

Varga, M. (2014). The effect of education, family size, unemployment and childcare availability on birth stopping and timing. *Portuguese Economic Journal*, 13(2):95–115.

Vishwanath, T. (1989). Job Search, Stigma Effect, and Escape Rate from Unemployment. *Journal of Labor Economics*, 7(4):487–502.

Vodopivec, M. (1999). *Unemployment Insurance and Duration of Unemployment: Evidence from Slovenia's Transition*. The World Bank.

Weiss, Y. and Gronau, R. (1981). Expected Interruptions in Labour Force Participation and Sex-Related Differences in Earnings Growth. *The Review of Economic Studies*, 48(4):607–619.

Whittington, L. A., Alm, J., and Peters, H. E. (1990). Fertility and the Personal Exemption: Implicit Pronatalist Policy in the United States. *The American Economic Review*, 80(3):545–556.

Winter-Ebmer, R. (1998). Potential Unemployment Benefit Duration and Spell Length: Lessons from a Quasi-Experiment in Austria. *Oxford Bulletin of Economics and Statistics*, 60(1):33–45.

Wolpin, K. I. (1984). An Estimable Dynamic Stochastic Model of Fertility and Child Mortality. *Journal of Political Economy*, 92(5):852–874.

Wooldridge, J. M. (2015). Control Function Methods in Applied Econometrics. *Journal of Human Resources*, 50(2):420–445.