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Abstract

This article studies how job creation subsidies designed for several Spanish regional governments to foster the creation of new permanent contracts during the period 1997-2004 might affect the subsequent employment stability of the eligible workers. We use a triple difference approach that focuses on regional and temporal variability in individual eligibility conditions of these subsidies to obtain the causal effect of the policy. Our data comes from the Muestra Continua de Vidas Laborales (MCVL) and from a database that contains information on the policy analyzed. Our main result is that workers who are eligible for these subsidies face a higher probability of exiting from their current permanent contract than those who do not. These effects vary by age and gender, as well as by contract duration and contract type. This result is particularly relevant for male workers whose contracts also benefited with nationally subsidized payroll deductions and for women with such deductions but only during their first year of employment. During that initial first-year period, the exit rate among eligible workers in our sample increased by 9%, 21% and 16% for younger, middle-aged and older female workers, respectively, and by about 13% and 25% for younger and older male workers, respectively.

Keywords: labour market rotation, permanent contracts, wage subsidies, triple difference, causal inference, average treatment effects, duration model.

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1 Introduction

Since the early 1990s, rising temporary employment rates in Spain have induced the national and some regional governments to implement a number of active labour market policies (ALMPs) designed to bolster the number of permanent hires and thus to forestall the perceived threat of temporary contracts over the country's economic efficiency and equity. Indeed, Spain invests more public funding in this type of ALMP than does any other OECD country. Between 1999 and 2002, for example, it dedicated roughly 0.28% of national output to this end (0.4% in 2006). Yet between 1996 and 2006, the proportion of permanently employed Spanish workers rose by a mere 0.3 percentage points, from 66.4% in 1996 to 66.7% in 2006.

This paper focuses on labour market policies that use targeted subsidies to increase employment stability. Since 1997, when the national government issued an important labour market reform (see Kugler, Jimeno and Hernanz, 2003 or Mendez, 2008 for a description) many Spanish regional governments have offered one-time payments to firms issuing new permanent contracts to certain groups of workers.[2] In our initial evaluation of this policy (García-Pérez and Rebollo, 2009), we concluded that the causal incidence of such subsidies over the entrance probability to a permanent contract was very low;[3] specifically, our results indicated that while such subsidies increased by 67% the conversion of temporary to permanent contracts re-hired by the same firm among eligible female workers aged between 30 and 45, it had no effect on other groups of temporary workers. However, the rise in this conversion rate (from 0.65% to 1.09%) was so small as to be economically irrelevant in terms of its final effect over permanent employment. We also obtained that the incidence of the subsidies over the pool of unemployed workers was only statistically significant among workers younger than 30, for whom the increase in the transition probability to a permanent contract for eligible workers ranged from 4% for female workers to 10% for male ones.

Despite this evidence, the available data on regional expenditure rates shows that such subsidies have been used intensively in some regions,[4] where they represent a significant reduction in labour costs. In fact, cross-regionally, the joint availability of both national and regional subsidies can reduce the total labour costs of the average worker's first two years of permanent contract by 15.4% for women aged 30 to 45, and by almost 22.8% for older female workers. It

[2] These subsidies, as we discuss later in this paper, target unemployed workers and workers with temporary contracts who obtain permanent ones under the same employer.

[3] The results of this paper accord with those obtained in other studies that evaluate the causal incidence of the different national labour market reforms implemented in Spain since 1994. That is, employer hiring policies seem to experience no significant change in response to the 1997 and 2001 reforms (Kugler, Jimeno and Hernanz, 2003; Arellano, 2005). In Mendez (2008), the author concludes that the reforms of 1994 and 1997 only increased the probability of transitioning from unemployment to permanent employment. Cebrian, Moreno and Toharia (2005) show that firing costs do not appear to be the main element in the determination of the proportion of employees with a temporary contract in Spain.

[4] In García Pérez and Rebollo (2007) we show that regional wage subsidies are used most intensively in Murcia, Baleares, Valencia and Galicia.

seems that regional subsidies affect total labour costs to a greater degree than do national ones. In fact, only between 3.6% and 10.4% of this cost reduction can be attributed to national payroll tax deductions.

The available literature indicates no overall positive effect of these ALMPs on the permanent employment rate. Katz (1994) shows that in a world marked by wage rigidities, the cost of labour drops when a firm receives a subsidy. If this cost reduction occurs during the worker's term of employment, the subsidy can, in fact, increase job's length. However, if the subsidy consists of a one-time payment at the beginning of the contract its effects on employment duration are more uncertain. In particular, in a situation where labour costs increase with the duration of the contract, the relevance of any subsidy-induced drop in labour costs diminishes as contract tenure increases. Hence, subsidized workers may in fact have shorter employment durations than other worker groups, particularly when the worker hired under a subsidized contract would not otherwise have been offered a position. The idea is that job creation subsidies counterbalance the lower labour productivity of eligible workers as compared against ineligible ones. In addition, the literature on causal evaluation points to a number of other unforeseen consequences of these policies. For instance, Calmfors (1994) argues that subsidizing permanent hires carries deadweight costs and substitution effects, which make it hard to evaluate the net effect of that strategy.[5] Martin and Grub (2001) argue that most evaluations focusing on firm behaviour have pointed to large deadweight and substitution effects when private-sector employment is subsidized. As a result, such schemes yield small net employment gains.[6] In a more recent paper, Mortensen and Pissarides (2001) show that job creation subsidies[7] might increase labour market rotation. Following this lead, the aim of this paper is to assess whether regional subsidies may favour the labour market rotation of eligible workers, by reducing the average duration of their subsidized permanent contracts. In broader terms, we wish to contribute to current knowledge regarding the effect of job creation subsidies on employer hiring and firing practices.

This evaluation exercise draws on sample data taken from the "*Muestra Continua de Vidas Laborales* (MCVL-2005)", a database compiled in 2005 by the Spanish Social Security administration. For the purposes of this paper, we have also compiled a database that provides detailed cross-regional information on the eligibility conditions for the regional subsidies we have been discussing—those that aim to bolster permanent employment designed by regional

[5] Deadweight costs refer to the hiring activities that benefited from the policy, but that would have taken place even in its absence. Substitution effects mean that some subsidized contracts are used to substitute other contracts held by ineligible workers, since one of the effect of the policy is to create a gap between the labour costs associated with hiring eligible workers versus ineligible ones. In his work, Calmfors shows that these effects reduce the proportion of regular employment (unsubsidized jobs) and increase that of irregular employment, although he does not discuss how subsidies affect overall employment.

[6] For instance, evaluations of wage subsidies in Australia, Belgium, Ireland and the Netherlands have suggested combined deadweight and substitution effects amounting to around 90 per cent, implying that for every 100 jobs subsidised by these schemes only ten were net gains in employment.

[7] They distinguish between employment subsidies, wage subsidies and hiring subsidies. An employment subsidy is an employment-contingent flow paid throughout the duration of a job-worker match. A wage subsidy can be regarded as a reduction in the payroll tax. A hiring subsidy is a payment that can be either a direct transfer to the employer or a tax rebate made to offset job creation costs.

governments —for the years 1997 through 2004, the only period for which relevant data is available.[8] One outstanding characteristic of the MCVL is that it allows us to observe contract modifications occurring within a single employment spell. Without this information, we would have risked biasing the estimated effect of the policy's causal incidence by excluding those who had worked first as temporary and then as permanent workers under the same employer, with no employment gap between the two contracts. Other valuable information provided by this dataset refers to the identification of those contracts that are benefited from discounts in payroll taxes at the national level. This information also enables us to measure where the causal effect of the policy differs by type of indefinite contract. This kind of information can not be found in other labour market databases.

We estimate a single risk duration model for a sample of workers with permanent contracts. Since our policy variable –the hiring subsidy-, varies by region, year and individual eligibility conditions, we use a triple difference approach to identify its causal effect over the exit rate from the permanent contract. When determining causal inference, it is crucial to properly define the control and treatment groups. In the interests of obtaining a homogeneous database, our reference sample only includes individuals whose prior employment records made them eligible for regional subsidies. That is, since these subsidies specifically target workers with unstable employment histories (unemployed and temporary workers), our estimation sample is comprised of only those workers whose employment history rendered them eligible for subsidized hire at the time of sampling. Within this overall pool, workers whose year of employment, region, age and/or gender rendered them eligible for such funding are included in the treated group; workers rendered ineligible for subsidy on the basis of these same criteria are placed in the control group.

Our main results show that both the worker's eligibility status and her contract length and type – with or without discounts in payroll taxes-, must be considered when measuring the influence of regional subsidization on the exit rate from a permanent contract. We find that such subsidies indeed increase the exit rate from permanent employment among eligible workers, particularly those whose permanent contracts also provides for national payroll tax deductions. Thus, during the first year of employment, the exit rate among eligible workers increases by 9% to 21% for female workers and by 13% to 25% for male workers. Since quarterly exit rates increase from 4.1% to 5.1% for eligible women and from 2.9% to 3.2% for eligible men, these estimated effects would appear to be relevant from an economic point of view. Nevertheless for certain cases we obtain the opposite result, i.e., a decreased exit rate among eligible workers. This result particularly arises for workers aged 45 and over, after their first year of employment. However, looking at the net effect of theses subsidies on predicted employment duration we observe that average duration decreases for eligible workers, versus illegible ones, by 20.52% for women aged 45 and over and by around 9-10% for older male workers and female middle-age ones.

[8] This explains why we do not use more recent versions of the MCVL.

Hence, the increase in the hazard for the first year of the contract seems to dominate to the opposite result found for some workers after one year has elapsed.

The rest of the paper is organized as follows. Sections two and three describe the data and the econometric model used for our analysis. Our main results are discussed in section four, and section five presents our main conclusions.

2 Data

The data for this evaluation exercise derives from two sources. Comprehensive employment histories for a sample of workers were drawn from the Muestra Continua de Vidas Laborales (MCVL), a database compiled and published by the Spanish Social Security administration.[9] In addition, regional eligibility requirements and subsidy amounts were taken from a dataset compiled by ourselves for the purpose of evaluating the policies discussed here (See García Pérez and Rebollo, 2007 for further details).

2.1 Regional Subsidies

Regional subsidies are entirely independent of those administered through the national government, but both offer essentially two kinds of aid: that intended to help unemployed workers find permanent work, and that which aims to shift temporary workers into permanent positions within the same firm. Thus, an employee from our sample of permanent workers is assumed to have benefited from the subsidy only if she had been previously unemployed or had held a temporary contract with the same employer. Many regional governments further narrow the pool of eligible workers by targeting those for whom it is most difficult to obtain permanent work. In such cases, if an unemployed worker had held a temporary contract in her previous job, she was eligible for subsidization regardless of how much time had elapsed between the end of that job and the beginning of the one sampled. However, if her previous contract had been a permanent one, she was only eligible if she had been unemployed for longer than three months at the time of hiring (if the new job was with a new employer) or if she has been two years out of the firm (if the new employer is the same than in the previous job). All of these eligibility requirements, which relate to the worker’s previous job spell, are considered in our analysis.

The main eligibility requirements –additionally to those referred to the characteristics of the worker´s prior labour market position explained above-, for the regional subsidies analyzed here are given in Table 1 (for male workers) and Table 2 (for female ones). They show the policy years, region of application, and eligibility rules regarding age and gender for workers in each of the two target groups: (1) temporary employees who move into permanent positions under

[9] See García Pérez (2008) for a complete description of this dataset.

the same employer, and (2) unemployed workers who obtain permanent positions. Table 2 shows the average wage subsidy (in 2002 euros) by age and gender for each region.[10]

Table 1: Regional Incentives: Eligibility Conditions by Age and Gender across Spanish Regions (Males, 1997-2004)

	Unemployed		Temporary Contract	
Andalusia	all ages	1997-2002	18-30	1997-2002
			30 or more	1997
Aragon	18-30	1999-2004	18-40	2002-2004
	30-40	1999-2003		
	40 and over	1998-2004	40 and over	1998-2004
Asturias	all ages	1997, 2001	all ages	1997-1998, 2000-2003
	18-30 and >45	1998, 2000, 2002		
Balearic Islands	NO		NO	
Canary Islands	18-25	1.998	18-25	1.998
	all ages	1.999	NO	NO
Cantabria	all ages	1998, 2000-2004	18-30	1998, 2001-2004
			45 and over	1998, 2000-2004
C. Leon	all ages	1998-2004	all ages	1998-2004
C. Mancha	16-30	1.998	16-30	1.998
	16-29 & 45 and over	1999-2003	45 and over	1999-2003
Catalunya	NO		NO	
Valencia	all ages	1998-2001, 2003-2004	18-30	1998-2001, 2003-2004
			45 and over	1998-2000
Extremadura	all ages	1997-2004	all ages	1997-2004
Galicia	18-30 & 45 and over	1998	18-30 & 45 and over	1998
	all ages	1999-2004	all ages	1999-2004
Madrid	all ages	1998-2004	all ages	1999-2004
Murcia	all ages	1998-2004	all ages	1998-2004
Navarra	NO		NO	
Basque Country	all ages	1998-2004	all ages	1998-2004
Rioja	all ages	1998-2004	all ages	1998-2003

Source: García-Perez and Rebollo (2007)

A close look at Tables 1 and 2 shows that regional recourse to this kind of policy varies widely; hence, some regions only began to implement the policy in 1997 while others, such as Catalonia and Navarra, have never offered such subsidies. Individual eligibility rules also show significant

[10] Although in Table 2 we give the average wage subsidy, in the estimation presented below we have matched the wage subsidy to each eligible worker in accordance with the year of hiring, her age, gender, and of course with her type of prior labour market position since in some regions wage subsidies differ between unemployed an temporary workers.

regional and temporal variations for workers of both genders, and especially for men; in several regions, these eligibility conditions also varied by the worker's prior labour state at the time of hire. For instance, regions such as Extremadura and the Basque Country offered subsidies targeted at all worker groups, while in the Balearic Islands they were reserved for women and in Valencia and the Canary Islands they mainly targeted younger workers. In some regions, including Aragón, Asturias, Castile-La Mancha and Valencia, eligibility conditions also varied according to the worker's most recent job spell and state of employment at the time of hire.

Table 2: Regional Incentives: Eligibility Conditions by Age and Gender across Spanish Regions (Females, 1997-2004)

	Unemployed		Temporary Contract	
Andalusia	all ages	1997-2002	all ages	1997-2002
Aragon	all ages	1998-2004	all ages	1998-2004
Asturias	all ages	1997-1998, 2000-2003	all ages	1997-1998, 2000-2003
Balearic Islands	all ages	2000-2004	all ages	2000-2004
Canary Islands	all ages	1.998	all ages	1.998
	all ages	1.999	all ages	1.999
Cantabria	all ages	1998, 2000-2004	18-30	1998, 2001-2004
			30 and over	1998, 2000-2004
C. Leon	all ages	1998-2004	all ages	1998-2004
C. Mancha	all ages	1.998	all ages	1.998
	all ages	1999-2003	all ages	1999-2003
Catalunya	NO		NO	
Valencia	all ages	1998-2004	18-45	1998-2004
			45 and over	1998-2000
Extremadura	all ages	1997-2004	all ages	1997-2004
Galicia	all ages	1.998	all ages	1.998
	all ages	1999-2004	all ages	1999-2004
Madrid	all ages	1998-2004	all ages	1998-2004
Murcia	all ages	1998-2004	all ages	1998-2004
Navarra	NO		NO	
Basque Country	all ages	1998-2004	all ages	1998-2004
Rioja	all ages	1998-2004	all ages	1998-2003

Source: García-Perez and Rebollo (2007)

Table 3 also indicates several regional variations in terms of subsidy amount. First, the high subsidies offered in regions such as Madrid or Extremadura (6.674 and 7.818 Euros per contract, in average terms) contrasts strikingly with the much lower ones offered in areas like Valencia or Galicia (where they fall to 1.807 and 2.639 Euros, respectively, on average). Second, the subsidized funds also vary in accordance with the worker's gender and age; thus, regional subsidies seem to be higher for women and older workers than they are for men and younger ones. It should be stressed that these regional subsidies represent a significant discount in labour costs. In fact, the joint availability of both national and regional subsidies may reduce the total labour costs associated with the average worker during his or her first two years of permanent contract from 13.5% for men aged 30 to 45 to almost 21.5% for young female workers, across regions.

Table 3: Regional Incentives: Average Subsidies by age and gender

	Males			Females		
	Age < 30	Age 30-45	Age > 45	Age < 30	Age 30-45	Age > 45
Andalusia	3,202	2,402	2,402	3,202	3,304	3,304
Aragon	2,850	2,888	4,317	3,137	2,870	3,030
Asturias	2,650	2,250	2,854	3,187	3,100	3,350
Balearic Islands	0	0	3,005	2,854	1,464	2,854
Canary Islands	3,000	3,600	3,400	3,000	3,000	3,000
Cantabria	2,423	2,400	3,040	3,239	3,077	3,478
C. Leon	3,456	2,401	2,401	3,456	2,651	2,651
C. Mancha	3,000	0	2,760	3,440	3,440	3,440
Catalunya	0	0	0	0	0	0
Valencia	1,424	1,400	1,400	2,584	2,584	2,854
Extremadura	5,379	6,158	8,944	5,896	5,896	8,454
Galicia	2,300	1,900	2,100	3,200	3,200	3,200
Madrid	7,200	7,200	7500	8,100	8,100	8,100
Murcia	3,540	2,850	3,214	3,540	3,514	3,514
Navarra	0	0		0	0	0
Basque Country	4,440	4,301	4,443	4,666	4,533	5,525
Rioja	3,844	3,006	3,757	4,700	4,700	4,700
Total	4,621	5,036	5,010	4,498	4,564	4,797

Source: García-Pérez and Rebollo (2007)

To compare these variations in the average cost reduction brought about by regional subsidies for different types of workers in different regions, Table 4 presents the cost reduction associated with both nationally-subsidized payroll tax deductions and regional subsidies for the year 2002, for each type of worker. Total labour costs in this table are computed for the first two years of employment under a permanent contract, during which it is assumed that the worker does not leave her job, i.e., that there is no firing cost involved. The resulting data show that, on average, regional subsidies cover 16% of total labour costs. Only in regions where there are no wage subsidies (in cursive) does the percentage of labour costs saved drop below 10%. For all other regions these savings are quite substantial –they are greatest in Madrid—especially for female and older workers. For a firm that fires a worker with a severance payment of 45 days per year worked, such payments represent about 10% of the wage and payroll taxes paid by the firm in the previous two years. One indication of the importance of the subsidization approach is the fact that employers are more than fully compensated for such firing costs in nearly every region offering such subsidies. Notice that these firing costs will be even lower for the new indefinite contract created in the labour market reform of 1997.

Table 4: Average discount in labour costs due to both National and Regional subsidies in the year 2002

	Males			Females		
	Age < 30	Age 30-45	Age > 45	Age < 30	Age 30-45	Age > 45
Andalusia	18,43%	14,38%	18,35%	21,37%	18,75%	24,29%
Aragon	15,28%	13,35%	17,28%	18,90%	16,91%	23,41%
Asturias	9,75%	13,97%	16,97%	19,91%	17,45%	22,06%
Balearic Islands	12,90%	8,35%	15,99%	14,90%	10,72%	19,31%
Canary Islands	*3,63%*	*3,63%*	*8,61%*	*4,53%*	*4,53%*	*10,43%*
Cantabria	12,77%	14,02%	19,21%	22,45%	17,56%	24,49%
C. Leon	13,91%	11,93%	16,49%	18,28%	15,79%	23,06%
C. Mancha	19,29%	11,96%	16,57%	21,84%	14,55%	20,30%
Catalunya	*3,63%*	*3,63%*	*8,61%*	*4,53%*	*4,53%*	*10,43%*
Valencia	9,11%	8,58%	11,94%	13,18%	12,85%	16,45%
Extremadura	10,88%	10,88%	25,84%	13,60%	13,60%	31,28%
Galicia	9,76%	8,86%	13,52%	12,99%	11,90%	17,98%
Madrid	**41,26%**	**33,20%**	**35,93%**	**45,73%**	**37,73%**	**44,20%**
Murcia	21,00%	18,54%	22,99%	25,37%	22,42%	25,62%
Navarra	*3,63%*	*3,63%*	*8,61%*	*4,53%*	*4,53%*	*10,43%*
Basque Country	12,10%	10,48%	24,12%	31,18%	18,80%	38,18%
Rioja	13,11%	11,36%	15,53%	21,61%	20,56%	25,01%

Source: García-Pérez and Rebollo (2007) and own calculations based on the MCVL.

2.2 Labour Market Data for Individual Workers

Our data on workers' individual employment histories was taken from the 2005 edition of the *Muestra Continua de Vidas Laborales* (MCVL hereafter). The MCVL is a sample of more than one million worker case-histories which provides very detailed information about their current and previous labour activities including the worker's wage category, type of contract and reasons for its termination, as well as the hiring firm's size, age, ownership, location and activity sector, among other job and firm characteristics. Since the database assigns each worker the corresponding identification code for the firm where she works, it allows us detecting whether or not a specific worker changed firms when moving from one employment spell to the next. Obviously, this is a critical factor in our research, since the eligibility requirements for the regional subsidies studied here[11] take into account whether a would-be candidate for subsidized employment has previously worked with the same hiring firm.

In terms of contract type, the MCVL provides two other groups of relevant data that are available in no other database. First, it tells us whether or not the worker's contract was a permanent one and, if so, whether or not he benefited from national payroll tax deductions. Consequently, in this analysis we distinguish between permanent contracts that also included national subsidies, and those that did not.[12] Second, the database allows us to observe contract

[11] This information is also relevant for the national payroll tax deduction policy.

[12] The 1997 national reform gave rise to a new type of permanent contract with lower firing costs (named "Contrato de Fomento de la Contratación Indefinida"). However, our dataset does not allow us to identify whether each permanent contract is subject to lower firing costs or not. Nevertheless, a high proportion of permanent contracts with

modifications taking place during a single employment spell; for example, it indicates whether the worker began her current job as a temporary worker and then obtained a permanent position with the same firm, or whether she originally held a permanent contract subject to national payroll tax deductions before moving into a new permanent contract with no such benefits. Disregarding this critical information might have led to a bias in the estimated causal effect, since it would have caused us to exclude from our eligible group those who had worked for as temporary workers with a given employer before becoming permanent ones. It would also have potentially biased the effect of the causal incidence of the national policy over the entrance probability into a permanent contract, by falsely lowered the observed number of permanent contracts that included national payroll tax deductions.

We measure the duration of each contract in quarters and on the basis of the specified starting and ending dates. Since the database also gives the ending dates for each contract modification, we also compute contract durations that take into account any variations that may have occurred within the same employment spell.[13] Likewise, we compute the duration of each unemployment spell by measuring the time lapse between the end date of the worker's previous contract and the start date of her new one.

In order to obtain a more homogeneous and comprehensive estimation sample, we have applied the following rules when selecting the employment spells for our sample. First, when two employment spells overlapped such that one of the spells encompassed the other, we used only the longer one.[14] Second, when two employment spells were simultaneous at a given point in time (but not all of the time)[15] we kept only the most recent one; however, when the simultaneity lasted for less than 15 days, we treated both spells as part of a job-to-job transition.[16] Third, we only consider employment spells from the so called "*Régimen General*", censoring any spells that lead to a position outside of this category.[17] Fourth, the final sample is composed by workers aged between 18 and 60 years. Finally, we omit all job spells for which any information is missing.[18] As we pointed out in the introduction, we are only interested in the duration of the employment spells when the job is a permanent one. Hence, once we had finished selecting our data according to the above rules, we eliminated spells not associated with permanent contracts.

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national subsidies (66%, according to the information offered by the Public Employment Agency on subsidized contracts in 2006) also specify low firing costs.

[13] That is, we use two variables from the MCVL-2005 called the "first contract modification" and "second contract modification".

[14] For instance, when we observe an employment spell for the period 2000-2001 and another one for the period 1999-2002 we omit the first one and we keep only the second one for our sample.

[15] That is, when the first contract ends after the second contract has begun.

[16] In this sense, we assume that the unemployment duration in this case is zero and we consider both spells as a unique employment spell.

[17] This definition includes the pool of regular paid employees for any given firm.

[18] For instance, lack of information regarding the contract beginning or ending dates and, more importantly, regarding contract type. This last restriction is the reason why we begin to collect data from 1995 onwards. Before 1995 the information on contract types suffers from a high percentage of non-observation. On the contrary, from 1995, this is not a real problem.

In causal analysis a proper definition of the treated and control group is crucial to obtaining an unbiased estimate of the policy. In order to obtain a homogeneous sample of workers,[19] and in light of the restrictions imposed by the policy in terms of the worker's employment status and her most recent job spell, we keep a sample of permanent workers whose previous experience rendered them eligible for subsidization. Hence, all workers whose most recent contract had been a permanent one and who had been unemployed for less than three months at the time of hiring (or 24 months in the case of re-employment in the same firm) were deemed ineligible and eliminated from our estimation sample, regardless of their age or gender.[20]

3 Descriptive Evidence

Let us now take a closer look at our final sample of workers with permanent contracts. Table 5 gives its main characteristics by age and gender. Here we observe, first of all, that contract duration as well as the percentage of censored observations are increasing with age.[21] Thus, average tenure ranges from 7 quarters (for older women) to 11.7 (for older men).

We can also see in this table, that more than 59% of the observed spells were associated with a permanent contract from the outset. Interestingly, this number tends to be greater among female workers, rising to 81% among older female ones. The remaining workers began their current spell under a temporary contract before moving to a permanent position within the same firm. Hence, these statistics highlight the importance of taking into account contract modifications, since we have found that between 10% and 30% of our spells began as temporary ones. In the absence of this information, we would have classified as ineligible a significant number of employees holding permanent contracts who may have benefited from the policy analyzed here.

The portion of permanent contracts carrying national payroll tax deductions is significant for all age groups. Obviously, the national eligibility rules for this type of contract (see Mendez, 2008) means that only about 26% of men aged between 30 and 45 held it from the time they began their spell of permanent employment. For certain group of workers, however, the majority of these contracts benefited from national payroll tax deductions. For instance, 48% of female workers under 30 and over 45 years of age with permanent contracts also had national deductions. This rate rises to 59% among older male workers.

[19] As it is well known (see, for example, Meyer, 1995), this is a basic requirement of any well defined difference-in-difference estimation.

[20] The number of employment spells that become ineligible given their previous unemployment spell's duration is quite large. They are around 23% for young workers, 48% for those aged 30-45 and almost 59% of all the observed employment spells of workers aged more than 45. Thus, it seems that the rotation across permanent contracts is also quite standard, especially among not very young workers. As explained in the text, these transitions are not considered in our estimation sample given they are not fulfilling one of the basic requisites of the policy analyzed.

[21] To avoid bias stemming from lack of data, we censored each spell at the 20th quarter (that is, after five years).

Table 5: Main characteristics of the estimation sample, by age and gender

	< 30		30-45		> 45	
	Women	Men	Women	Men	Women	Men
Current Spell						
Contract Duration (completed spells)	8.4	9.9	7.9	10.7	7.0	11.7
% of Censored Observations	46.29%	47.18%	54.44%	53.09%	56.33%	56.42%
Perm. contract since the beginning of the spell	66.98%	63.99%	65.07%	59.99%	81.55%	81.22%
Perm. Contract with national subsidies	48.04%	38.30%	33.71%	26.17%	48.17%	59.88%
Part-Time	22.59%	9.80%	29.00%	6.90%	32.38%	5.86%
Layoff	56.89%	47.49%	77.08%	64.13%	87.77%	74.71%
New Firm	26.81%	24.93%	26.33%	30.98%	24.99%	34.50%
Private Firm	95.66%	97.40%	89.15%	94.67%	89.71%	94.80%
<u>Sector of Activity</u>						
Industry	14.56%	28.36%	17.31%	29.88%	18.12%	38.00%
Construction	2.03%	8.40%	1.61%	8.33%	1.31%	11.09%
Services	85.41%	63.24%	81.07%	61.78%	80.57%	50.91%
<u>Firm Size</u>						
< 5 employees	41.42%	36.15%	38.28%	34.48%	38.80%	36.08%
5-20 employees	16.60%	18.39%	14.22%	16.41%	13.33%	14.61%
20-100 employees	18.05%	21.41%	19.01%	21.37%	18.27%	19.03%
> 100 employees	23.93%	24.04%	28.49%	27.73%	29.60%	30.28%
Age	24.98	25.11	36.63	36.66	50.29	50.57
Immigrant	3.07%	2.95%	3.15%	3.41%	1.42%	1.40%
<u>Qualification</u>						
High	12.08%	13.76%	14.63%	20.28%	8.50%	18.50%
Medium-High	24.66%	20.29%	24.14%	22.75%	18.51%	21.20%
Medium-Low	37.09%	33.62%	27.25%	36.09%	22.91%	39.84%
Low	26.17%	32.33%	33.98%	20.88%	50.07%	20.46%
Previous Trajectory						
Previous Unemployment Spell (months)	5.87	6.41	6.58	5.70	6.39	6.65
Nº of Temp. Contracts	3.7	3.6	3.8	3.6	3.3	2.9
Nº of Unemployment Spells	2.9	3.2	5.2	5.9	5.2	6.5
Number of Spells	90,967	111,291	42,172	57,089	15,433	23,790

The rest of employment characteristics of the workers in our estimation sample differ markedly by age and especially by gender, reinforcing the importance of carrying out gender-specific estimations of the model. Part-time jobs were more common among female workers than among male ones, with the percentage of workers holding such jobs ranging from 6% among middle-aged men to 29% among women over the age of 45. The main reason for leaving a job was involuntary separation. The percentage of workers for whom this was the case increases by age and varies from 47% (young male workers) to 87% (older female ones). The proportion of sampled workers holding jobs in the service sector was greater for women than for men, while the opposite was true of industry-sector jobs. While gender-based differences regarding firm size were less important, women did tend to work for smaller firms more frequently than men did. Finally, high-skill jobs were more common among men and older workers than they were among women and younger ones.

The final rows of Table 5 provide data on the worker's job experience just prior to taking the permanent position analyzed. For workers who had been previously unemployed, this period of unemployment lasted an average of five to seven months. Employees generally experienced quite a few temporary and/or unemployment spells prior to the permanent contract under study. Specifically, the number of temporary contracts held during this period ranged from 2.9 (for male workers over the age of 45) to 3.8 (for female workers under the age of 30). The number of unemployment spells varied from 2.9 among younger male workers to 5.9 among middle-aged female ones.

Workers with permanent contract can experience different type of jobs transitions depending on the type of contract in the new job and whether they experience an unemployment spell in between or not (Table 6). Here we focus on those transitions scenarios that represent a risk to the worker's labour market stability. Hence, we classify the observed spells into two different job-transition destination states whenever one is observed (that is, when the unemployment spell is finished by entering into a new job): a worker in transition either found a new permanent job or got a job as a temporary worker after a certain unemployment spell.[22] Two other transition scenarios were censored from our duration analysis: when the employee returns to the same firm after a spell of unemployment lasting less than one month, and when she begins working with a different firm after having been unemployed for less than one week.[23]

In general terms, a high proportion of the observed transitions tend to lead to temporary contracts. This suggests that, for the workers in our sample, holding a permanent contract did not guarantee that the next contract would be a permanent one. Nevertheless, a number of interesting gender-based differences in this regard can be observed. For male workers, the more likely exit was to a temporary contract. Over 60% of the observed transitions show this type of transition, with this rate decreasing slightly by age. Among female workers, the results vary widely by age group. The probability of obtaining a temporary contract after ending the observed permanent contract decreases sharply as the worker's age increases, falling from 60% for younger workers to 53% and 43% for middle-aged and older female workers, respectively.

[22] A third option is lapsed into unemployment from which there is no observed exit. Employment transitions that end into unemployment with no observed exit correspond with censored unemployment spells. These transitions are not relevant from the analysis since they vary from 0.56% -older male workers-, to 1.16% -older female ones.

[23] We have observed that a significant number of job-to-job transitions take place during the first week of unemployment and that more than 50% of them lead to a new permanent contract. As explained in the text, we are not considering these transitions as an exit from permanent employment and, hence, they are treated as censored spells.

Table 6: Employment Transitions by Age and Gender (1995-2004)

	<30		30-45		> 45	
	Women	Men	Women	Men	Women	Men
Nº of Censored spells	42,317	52,933	23,058	30,347	8,657	13,559
Nº of Completed Spells	48,650	58,358	19,114	26,742	6,776	10,231
Exit to a Temp. Contract	64.12%	64.05%	53.00%	61.18%	42.96%	60.35%
Exit to a new Perm. Contract	34.90%	35.01%	45.84%	37.87%	56.36%	39.10%

Table 7: Main sample characteristics: eligible versus ineligible workers (1995-2004)

	Women		Men	
	Ineligible	Eligible	Ineligible	Eligible
Age	32.6	33.0	33.9	35.0
Current Spell				
Exit from the current perm. contract	31.75%	34.16%	27.99%	31.96%
Contract Duration (Uncensored)	9.0	8.3	9.4	8.9
Perm. Contract since the beginning of the spell	66.48%	69.11%	61.99%	67.86%
Perm. Contract with national subsidies	22.82%	48.98%	31.93%	44.05%
Part-Time Job	21.56%	30.34%	5.87%	8.84%
Layoff	79.25%	84.32%	49.52%	65.25%
Private Firm	89.25%	90.25%	92.26%	93.54%
Activity Sector				
Construction	2.52%	2.13%	11.99%	12.09%
Services	84.47%	85.66%	64.50%	62.90%
Industry	14.01%	12.21%	23.50%	25.01%
Firm Size				
< 5 Employees	40.40%	41.26%	36.27%	38.80%
5-20 Employees	18.74%	17.78%	20.66%	21.11%
20-100 Employees	18.60%	17.64%	22.80%	21.65%
> 100 Employees	24.27%	23.32%	21.27%	18.45%
Job Qualification				
Highly skilled	9.89%	10.37%	13.19%	13.29%
Medium skilled to highly skilled	23.26%	22.02%	19.45%	17.92%
Medium to low skilled	27.95%	35.28%	38.12%	37.60%
Low skilled	40.89%	32.32%	30.23%	31.19%
Previous Spell				
Same firm	61.04%	61.33%	60.12%	51.40%
Previous Temp. Contract	87.08%	90.39%	87.99%	93.02%
Nº of Temp. Contracts	3.5	3.4	3.5	3.7
Nº of Spells of Unemployment	3.7	3.8	4.5	4.3
% of Spells	43.47%	56.53%	54.59%	45.41%
Nº of Spells	64584	83987	104905	87264

Table 7 gives the main sample characteristics for eligible and non-eligible workers by gender. It shows that there are no important differences between them in terms of basic job characteristics, which suggest that our control and treated groups are quite similar. Important differences do arise, however, when we consider the probability of exiting from the current contract, average contract length and type of contract held. Thus, ineligible workers tend to hold onto their current permanent positions longer and are less likely than eligible ones to exit from them. For example, about 28% of ineligible male workers exited from their current permanent contract, while this ratio rose to almost 32% among eligible ones. In the case of women, we observe that about 32% of ineligible workers exited from the current contract while this ratio increased to 34% among eligible ones. Interestingly, the share of workers whose permanent position at the time of sampling started out as a temporary one was greater among eligible workers (67% to 69%) than among ineligible ones (62% to 66%). The same can be observed with respect to national payroll tax deductions: eligible workers (44% to 48%) were more likely than ineligible ones (22% to 31%) to benefit from such deductions. The latter result suggests that firms often ask for both regional subsidies and national ones (in the form of payroll tax deductions) when hiring a new worker.

Table 8 shows the type of labour market transitions for eligible and non-eligible workers of each gender group[24]. With regard to the spell following the current one, eligible workers appear to behave somewhat differently than non-eligible ones. Of particular interest is the fact that eligible workers are more likely (27% for women and 26% for men) to obtain a new permanent contract with national deductions in payroll taxes than are ineligible ones (23% for women and 19% for men).

Table 8: Labour Market Transitions by Eligibility and Gender (1995-2004)

	Women		Men	
	Ineligible	Eligible	Ineligible	Eligible
Temporary contract	52.66%	52.91%	54.44%	56.31%
Perm. Contract	46.41%	46.23%	44.66%	43.00%
With national subsidies	23%	27%	19%	26%

Finally, Figures 1 and 2 give the empirical exit rate from a permanent contract by gender for eligible versus ineligible workers. Here we find, first, that the exit rate tends to rise during the first year of the contract and decreases monotonically afterward. Second, when we compare the exit rate between eligible and ineligible workers we find that this rate tends to be higher among eligible workers regardless of gender, although the observed differences seem to be greater among female workers, particularly during the first year of hire. Third, the differences between eligible and ineligible workers with regard to the exit rate from a permanent contract vary in accordance with the duration of the contract, with the observed difference being a bit larger during the first year of employment.

[24] As in Table 6, a third option is not exiting from unemployment after having being separated from the previous job. As before this transitions is highly insignificant as it varies from 0.69% (older male workers) to 0.93% for young female workers.

Figure 1: Exit rate from a permanent contract (eligible versus non eligible, Women)

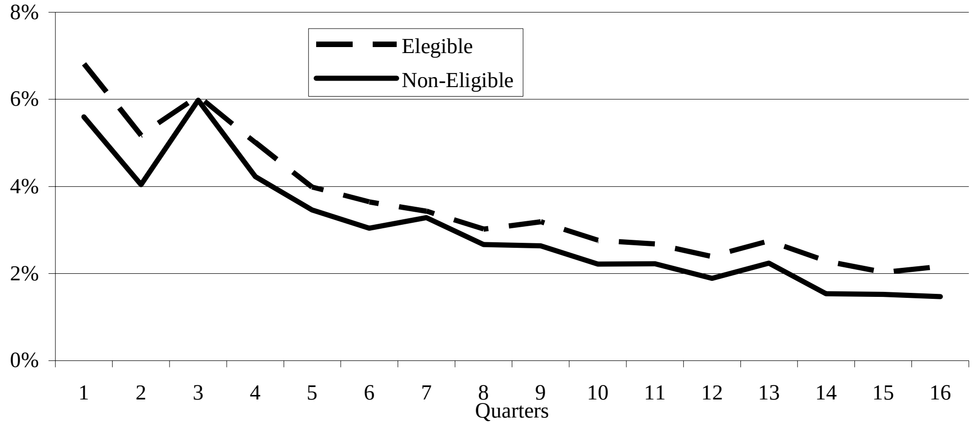
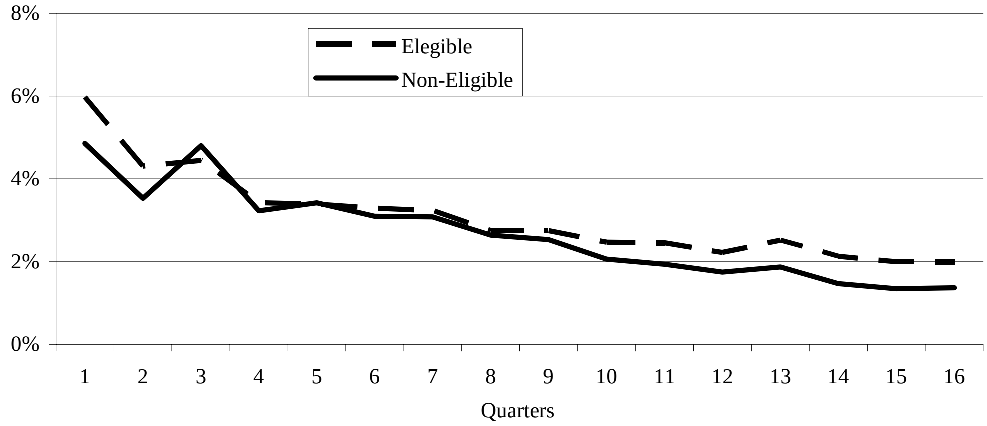


Figure 2: Exit rate from a permanent contract (eligible versus non eligible, Men)



Further differences arise when different types of permanent contracts are considered. Figures 3 and 4 distinguish between workers holding permanent contracts with and without national payroll tax deductions. Here, the exit rate from permanent contract for those without national benefits is higher at the beginning of the employment spell and decreases substantially during the first two years of employment. By contrast, the same exit rate for those benefiting from the tax deductions is basically flat, with almost no duration dependence. Interestingly, during the first year the exit rate of this type of nationally-subsidized contract remains lower than that from permanent contracts without national benefits. This fact may have something to do with the qualification rules and benefits associated with such nationally-supported contracts (national deductions in payroll taxes tend to last for two years).

Figure 3: Exit rate from a permanent contract by contract type (women)

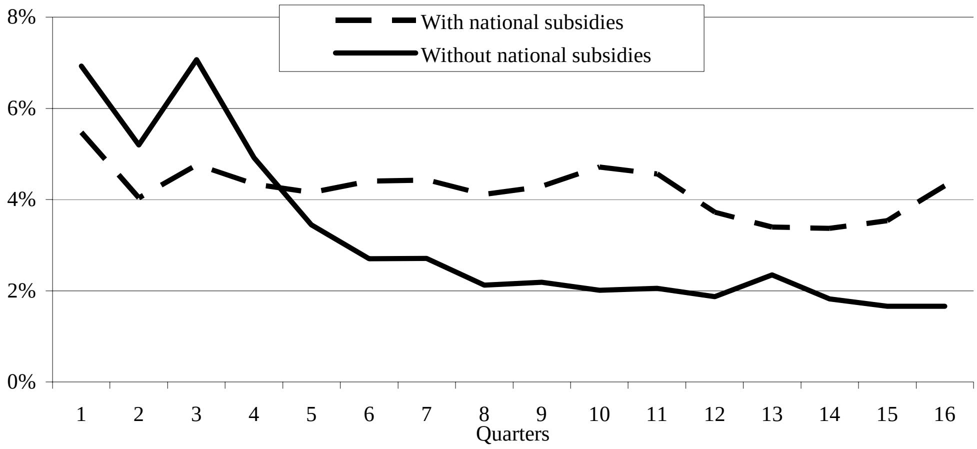
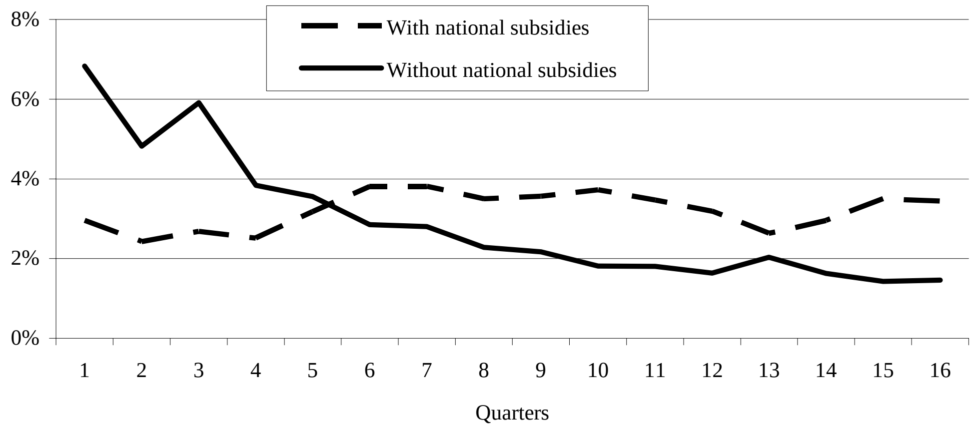


Figure 4: Exit rate from a permanent contract by contract type (men)



In sum, our dataset indicates that the risk of being separated from the job seems to be greater among eligible workers. We have also shown that the pool of eligible workers does not greatly differ from that of ineligible ones in terms of observed characteristics. On the basis of this evidence, one may be tempted to conclude that regional subsidies underlie the shorter employment spells observed for eligible workers. We attempt to disentangle the regional policy's causal effect by estimating a duration model that uses a triple difference estimator. We have also provided evidence that suggests the combined use of regional subsidies and national payroll tax deductions. Nevertheless, since some eligibility requirements are common to both policies, our conditional approach must confirm any differential effect between the regional subsidies for workers holding permanent contracts with deductions in payroll taxes, in comparison with those without.

4 Econometric procedure: Identification and Estimation

The aim of this paper is to measure whether regional subsidies cause or partly contribute to the observed differences in permanent contract duration among regional subsidies' eligible versus ineligible workers. To this end, we estimate a duration model that establishes the determinants of the exit rate from the current permanent contract. We identify the average treatment effect of the policy under study for its three dimensions of variability (region, time period and individual eligibility rules). Thus, time variation across regions, regional variation across time and eligibility variations across regions and time allow us to identify the causal effect of regional subsidies over the duration of permanent contracts.

4.1 Identification of the causal effect

In this analysis, we use the term "eligible" rather than "treated" because our database lacks information on real treatment. Thus, while we are able to observe the worker's individual characteristics and recent employment transitions, we do not know whether she finally benefited from the regional policy or not, when she was hired under a permanent contract. Likewise, we cannot observe whether the firms actually applied for the subsidy when hiring an eligible worker under a permanent contract. Consequently, the treatment effect we identify should be described as a "potential" effect, since we can only measure the policy benefits for workers who were *potentially* treatable, but who may or may not actually receive treatment. Nevertheless, given that such subsidies represent an important discount in hiring costs, it is reasonable to assume that most of eligible workers finally benefited from the policy.

In this context, our model must be carefully and appropriately specified in order to capture all observed and unobserved differences between the treatment and the control group of workers. As discussed earlier, in order to maximize the similarity between workers in the treated and control groups, we have restricted our sample to all workers whose job histories just prior to sampling rendered them eligible for subsidized hire. The treatment group is comprised of workers eligible for subsidized hire on the basis of their age, gender and prior labour market position, who were living in the region offering the subsidy at the time of its implementation. Similarly, the control group is comprised of workers deemed ineligible on the basis of age or gender or prior labour market position, and who lived in a region -or time period- for which no such funding was available.

Our triple difference model can be represented as follows:

$$P_{ijt} = \lambda(t - t_0) + x'_{ijt}\,\alpha + \beta D_{ijt} + \eta_i + \mu_j + \delta_t + \xi_{it} + \nu_{jt} + \psi_{ij} + \varepsilon_{ijt} \qquad (1)$$

where i refers to individuals, j to regions and t to time (quarters); the function $\lambda(t\text{-}t_0)$ comprises the duration dependence of the exit rate from the permanent contract, specified as a polynomial of degree two.[25] The index P_{ijt} is the argument of the probability statement being estimated as explained bellow. The variable that identifies the causal effect of the policy is D_{ijt}, which takes the value of the maximum wage subsidy for each eligible worker with individual characteristics i, in region j and period t, and zero otherwise. The aim of our econometric exercise is to obtain an unbiased estimate of the effect of this variable on the exit rate from permanent employment. To do so, we must control for all the covariates that can simultaneously affect the treatment and outcome and that present individual, regional and temporal variations. García-Pérez and Rebollo (2009) present a detailed description of the identification approach used also here to assess the causal effect of the policy. Specifically, we control for temporal variation with annual dummy variables, δ_t, regional variation through regional dummy variables, μ_j, and individual variation in eligibility conditions, η_i, which are proxied by dummy variables that control for age groups and previous employment history.[26] Finally, we must also consider how these three dimensions interact with one another. Thus, ν_{jt} represents the interactions of regional dummy variables and temporal dummy variables which are grouped in three periods (1995-1996, 1997-2000 y 2001-2004) for the purposes of identification; ψ_{ij} represents the interaction of age group dummy variables with the regional ones; and ξ_{it} is represented by the interaction of age group and year dummy variables for the three periods specified above (1995-1996, 1997-2000 y 2001-2004). Note that the variables in this last group, as well as those contained in ν_{jt}, play a crucial role in the identification of the causal effect of the regional policy, since an important national labour market reform which brought a new permanent contract and subsidies for new permanent contracts, was implemented during the same period and eligibility rules were also related to the age of the worker.

Finally, the vector x_{ijt} comprises variables (contract type, job skill level, activity sector, firm size, firm ownership, and so on) that may differ by individual, region and time period and that allow us to control for observable differences between eligible and ineligible workers which could bias our results. Since the eligibility conditions also address the worker's prior employment history, we also include variables that describe certain aspects of the worker's job experience. These variables include the number of prior unemployment spells experienced by the worker and a set of binary variables that indicate whether her job position at the time of sampling corresponded to her first employment spell (denoted here as *first spell*), whether she previously had held a temporary contract and then been unemployed (*temp. contract*), and whether she had previously held a temporary contract with the same firm (*conversion from temp. contract*).

Since we have also found interesting differences in the average duration of permanent contracts depending on whether they have national subsidies or not, we also perform a second exercise designed to assess whether the causal effect of the subsidies varies by contract type. While our

[25] Additionally we add some binary variables to control for specific contract durations at 4th, 8th and 12th quarters. Their inclusion here is justified by the behaviour of the empirical exit rate shown in the statistical section.
[26] We also estimate each duration model separately for each of the two gender groups, given the previously shown differences among both groups.

first estimation covers the period 1995-2004, the second one is restricted to the years between 1997 and 2004, given that data on national subsidies is only available from 1997 onward. It should be stressed that when estimating the model for the period from 1997 to 2004, we partially lose one dimension of our identification; since we drop all of the data corresponding to the period before the subsidy took effect. Yet we can still identify the average treatment effect of the policy in this case by looking at regional and temporal variability, as well as that regarding individual eligibility conditions.[27]

4.2 The estimation method

For both exercises, we estimate a duration model using the single risk approach. Hence, our objective is to estimate the exit or hazard rate. A common strategy to estimating the hazard rate is to transform the duration model into a sequence of discrete choice equations defined on the surviving population at each spell's duration. In this case, we define a binary variable y_t that takes the value of one when the worker exit from the current state, and is otherwise assumed to be zero. This expression has exactly the same form as the likelihood function of a discrete choice model where y_t is the binary endogenous dependent variable, once we have rearranged the database so that there are as many rows per individual as there are time units -in this case, quarters- of worker permanence in the initial situation (Allison, 1982; Jenkins, 1995). Hence, our likelihood function for each individual is:

$$\ln L = \sum_{t=1}^{T} \left[\ln(\Pi_t) * y_t + \ln(1-\Pi_t) * (1-y_t) \right] \tag{2}$$

where Π_t is the conditional exit rate at time t . To estimate this transition probability, we use a *logit* specification. Hence, the expression of the conditional exit rate from the permanent contract:

$$\Pi_t = \Pr(T = t / T > t) = \frac{\exp(P_t)}{1+\exp(P_t)} \tag{3}$$

where P_t has been defined above (expression 1).

5 Do regional subsidies influence the exit rate from a permanent contract?

In this section, we discuss the main results of our analysis. As stated earlier, the latter centres on the estimation of a duration model that uses a triple difference approach to identify the causal

[27] From Tables 1 and 2 we learn that the policy was only used in certain regions for several years. In addition, some regions upheld the policy for the duration of the period in question, while other only began to implement it a few years after the period had begun and/or ended it before the period ended.

effect of regional subsidies on the exit rate from a permanent contract. In order to obtain a better understanding of the causal incidence of these subsidies, we proceed in several steps.

First, we estimate the average treatment effect of the subsidies on the exit rate from a permanent contract for potentially eligible workers and we also check to see whether these effects vary by type of firm or employment position. Second, we wonder whether the incidence of the subsidies can vary by job tenure. Finally, since a number of national policies designed to boost stable employment were also effective during the period we analyzed, our sample includes workers who stood to benefit from both national and regional subsidies. In order to test whether or not the incidence of regional subsidies varied by the presence or absence of the national ones for any given contract, we estimate another duration model in which the policy variable interacts with a dummy variable introduced to control for the existence of national payroll tax deductions benefits.

Our statistics have shown that eligible workers face shorter employment durations. These observed differences could be attributed to the regional subsidy or to other individual and/or job characteristics. If, from our estimation, we find that eligible workers tend to leave their permanent contract sooner than ineligible ones, then we should conclude that regional subsidies negatively affect the labour market stability of subsidized permanent workers. This result would imply the existence of some unexpected side effects produced by the regional subsidies. These side effects should be considered when evaluating the benefits of this policy since they decrement its potential positive effects on the permanent employment rate. In this vein, Mortensen and Pissarides (2001) state that a possible indirect effect of job creation subsidies is to raise the hiring and firing rate of permanent workers without altering the permanent employment rate. In particular, they argue that the hiring subsidy could have positive effects on the permanent employment rate only when it is equal or lower than the firing costs. Hence, this hypothesis is even more interesting when applied to the Spanish case, where regional and national policies coexist simultaneously.

Our study allows for heterogeneous effects of the policy by age group. We cross the policy variable "D_{ijt}" with the worker's age group, dividing the latter into the three main age categories established by both regional and national eligibility rules: under 30, between 30 to 45, and over 45. Gender is another important policy determinant and is considered in our analysis by estimating separate duration models for men and women. Finally, we allow for the fact that the policy variable may have a non-linear effect on the exit from permanent employment, by modelled it as a polynomial of degree two.

For ease of exposition, the main results of each of our estimated duration models are summarized below (Table 9) and further detailed in the appendix (Table A.1).[28] Table 9 first presents the estimated exit probability from a permanent contract, computed at the mean of the explanatory variables for each group of workers -defined by age and gender. Then, in the right hand side of the table, it gives the estimated causal effect of the regional subsidies on the exit

[28] The remaining results are available upon request.

from the permanent contract. We obtained this measurement from the total change in the estimated exit probability of eligible versus ineligible workers that can be directly attributed to the presence of subsidy funds. When both of the coefficients associated with the policy are statistically significant, we interpret this as evidence of the existence of a causal effect.[29]

Table 9: Causal effect of the policy on the exit rate from permanent employment (1995-2004)

	Exit Prob. (Without Policy)		**Change in Prob. due to the policy**	
	Women	**Men**	**Women**	**Men**
< 30	3.77%	3.05%	6.43%	10.98%**
30-45	3.30%	2.93%	5.23%	9.62%*
>45	3.69%	3.06%	-6.72%	6.05%*

Notes: The estimated probability is obtained at the average of the observed characteristics and at the eighth quarter of the contract. The average change caused by the policy is computed at the average of the wage subsidy for each group of workers. The symbol "*" means that the coefficients associated with the policy variable were both statistically significant at 90% and "**" means that they were both statistically significant at 95%.

From the results shown in Table 9, we learn that regional subsidies seem not to influence the exit rate from a permanent contract among eligible female workers, whereas they have a low positive effect among eligible male ones.[30] The results obtained for women are not statistically significant for any age group, while for men the effect is positive and statistically significant. The exit rate from a permanent contract among male eligible workers increases by 10.98% for younger workers, by 9.62% for middle-aged ones and by 6.05% for older ones.[31]

Nevertheless, these results refer to the average treatment effect, which may vary by job tenure and/or contract type. In particular, our statistical analysis has shown that the exit rate from a permanent contract during the first year behaves differently than it does thereafter, a tendency which may be related to the effect of job creation subsidies on hiring costs relative to firing costs (which in Spain depend on job tenure). That is, since the job creation subsidies reduce hiring costs but have no effect on firing costs, such funds may simply encourage the job turnover rate among eligible workers. At the point which firing costs go beyond the drop in

[29] Standard errors are corrected to take into account that we may have different spells of the same individual. That is, an individual might fulfil more than once the criteria to be in the sample and this is taken into account in the estimation.

[30] It is worth to mention that when the control group includes all types of workers with permanent contracts, regardless of whether or not their previous work situations made them eligible for RWS, the results differ notably from those presented in Table 9. The effect of RWS appears to be positive and statistically significant for all workers. For instance, in this estimation RWS appear to raise the female exit rate from a permanent contract by 14% for middle-aged women and 17% for women over the age of 45. The effect is even more marked for men, rising by 21%, 24% and 5% for younger, middle-aged and older male workers, respectively. However, as we have presented in Table 9, the results totally change when considering a proper control group, what confirms the need for a properly definition of such control group in causal analysis.

[31] We have also estimated the model by including the data on job transitions lasting less than 7 days. The estimated causal effects are slightly lower in this case. Thus, for young male workers the effect is 8.8%, while for middle-aged men it is 7.0%. The rest of results do not change.

hiring costs, the incentives of firing the subsidized worker increase.[32] In order to determine whether the treatment effects differ among workers with different job tenures, we also estimate a model in which the policy variable is interacted with a dummy variable that takes the value one during the first year of the contract, and is otherwise fixed at zero.

Table 10 gives the main results of this new estimation. The most striking result here is that, in all cases, the estimated effect of the policy is statistically significant and positive during the employee's first year of contract. That is, this exercise suggests that regional subsidies may actually increase the exit rate from a permanent contract among eligible workers of both genders during the first year of hire. This effect seems to increase with age, mainly among men (for whom the causal effect ranges from about 13%, for workers aged 45 and under, to 25.74% for older ones). The effect vanishes as contract tenure lengthens, even becoming negative among relatively older workers (-21.15%). That is, once an eligible worker has held onto a contract for more than one year, her probability of exiting from that contract is lower than it is for ineligible female workers over the age of 30 and male workers over the age of 45. Thus, the exit probability drops by 20.37% for middle-aged female workers and by 21.15% and 54.16% for male and female older workers, respectively. This last effect is more relevant for men than for women in the sense that more than 40% of exits in the sample take place during the first year of the contract in the case of women, while in the case of men this portion drops to around 28%. For younger female workers and middle-aged male ones, the effects are not statistically significant. Only for young male workers, the sign of the effect remains the same as those obtained for shorter contract periods, although at a lower scale (about 8%). When we look at the net effect of theses subsidies on employment duration by computing the expected duration we obtain that the negative incidence of the subsidies over eligible workers predominates in all cases, although it is large for females than for males. The effect is more important for older workers than for younger ones. For instance, average expected duration decreases for eligible older women by 20.52% and for older workers by around 10%, versus illegible ones. For the rest of workers average expected duration for eligible workers versus illegible ones drops around 2%, except for the case of middle age female ones who also experiment drops in expected duration in permanent employment of around 9.6%. Thus, although we estimate a reduction in the hazard rate for older workers once one year has elapsed, the effect over the first year of the contract is more important what makes the complete effect over expected duration to be negative in all cases, and specially for older workers.

These results are highly interesting since it might indicate that some firms effectively use regional subsidies to reduce the cost of hiring workers on a permanent basis. Meanwhile, also the opposite result emerges indicating that other firms might be using these subsidies to reduce the hiring costs of new employees who finally experience short-term jobs. That is the incidence of the regional subsidies partially depends on the hiring practices of Spanish firms. Notice that previous empirical studies have already pointed out that some firms do use temporary contracts

[32] This is relevant since in many regions subsidized workers should remain, at least, one year with the current permanent contract.

as screening devices while some Spanish firms seem to be almost exclusively reliant on temporary contracts as a means of maintaining external and internal flexibility (Guell and Petrongolo, 2007 and Rebollo 2008). It seems that subsidized permanent contracts may be also having this screening role documented before for temporary contracts.

Table 10: Estimated Causal effect of the policy on the exit rate from permanent employment: the influence of job tenure (1995-2004)

		Exit Prob.(Without Policy)		**Change in Prob. due to the policy**	
		Women	**Men**	**Women**	**Men**
Contract Length<=12 months	< 30	5.11%	3.22%	9.26%**	12.90%**
	30-45	4.13%	2.83%	21.55%**	13.09%**
	>45	4.15%	2.94%	16.67%**	25.74%**
Contract Length>12 months	< 30	3.22%	2.60%	4.61%	8.82%*
	30-45	2.59%	2.28%	-20.37%**	5.76%
	>45	2.60%	2.37%	-54.16%**	-21.15%**

Notes: The estimated probability is obtained at the average of the observed characteristics and at the 12th and the 24th quarter of the contract. The average change caused by the policy is computed at the average of the wage subsidy for each group of workers. The symbol "*" means that the coefficients associated with the policy variable were both statistically significant at 90% and "**" means that they were both statistically significant at 95%.

We are also interested in studying whether these average treatment effects differ by job type. For this purpose, we have estimated the initial specification –Table 9-, allowing for a heterogeneous effect by firm size, activity sector and job skill level. Improved knowledge of how treatment effect responds to different employer characteristics may help the design of the policy we are discussing. The results displayed in Table 11 indicate that the causal effect for men under the age of 30 tends to arise in the context of smaller and medium-size firms, in those firms associated with the service sector and when the job skill level is low or medium. In the case of middle-aged men, the effects are only statistically significant in the industry sector and for low qualified jobs. Among female workers, the effects on younger workers appears to be similar to those obtained for young male workers; that is, they are positive and statistically significant when the job is with a smaller firm, one associated with the service sector, and/or when it is low qualified. In the case of women over the age of 45, all of the statistically significant effects are negative and generally occur in firms with more than 20 employees and in low qualified jobs.

In sum, regional subsidies may favour the exit from a permanent contract among eligible workers, especially during the first year of the contract. These effects seem to concentrate around low qualified jobs and those in the service sector.[33] By contrast, the effects are not statistically significant or are even negative for longer contract durations. The latter effect is particularly relevant for women over the age of 30 and older male workers.

Table 11: Estimated Causal effect of the policy on the exit rate from permanent employment: the interaction with other firm and job characteristics (1995-2004)

		Women			Men		
		<30	30-45	>45	<30	30-45	>45
Firm Size	< 20	14.61%**	4.57%	4.81%	9.20%**	4.26%	7.10%
	20-100	11.77%	-5.41%	-34.92%**	8.44%**	-0.26%**	14.63%
	>100	3.18%	3.98%	-35.99%**	-0.63%	9.86%	-4.70%
Activity Sector	Industry	0.48%	-15.60%	-14.60%**	2.00%	30.61**	29.37**
	Construction	25.72%	-40.49%	-82.15%**	-2.18%**	7.63%	11.48%
	Services	10.74%**	1.94%	-11.53%**	8.33%**	-1.40%	-1.67%**
Job Skill Level	Low	-1.78%**	-7.75%	-22.75%**	6.73%**	12.36%**	19.01%
	Medium	8.88%	7.14%**	-14.89%*	10.41%**	11.37%**	2.71%
	High	0.61%**	-11.22%	16.84%	0.68%	-7.68%**	6.43%

Notes: See Table 9.

5.1 Does the causal effect of regional subsidies vary with the availability of national subsidies?

In the previous section, we argued that an unintended effect of the policy under discussion is to reduce the duration of permanent contracts for some eligible workers, thereby increasing their job turnover rate. Since national payroll tax deductions were available for some permanent hires during the same analytical period, we wonder whether these two policies may have had complementary effects or not. We are particularly interested in analysing whether the causal effect of regional subsidies on an eligible worker who also benefits from payroll tax deductions is greater or lower than that for an eligible worker with no such deductions. In the statistical section, we have shown that the empirical exit rate from a permanent contract differs markedly by the availability of national subsidies, particularly during the first year of employment.

As we noted at the beginning of this paper, employers who open new permanent positions for eligible workers can obtain important deductions in total labour costs by combining both type of subsidies. Mortensen and Pissarides (2001) show that for economies with strong unemployment compensation packages and stringent employment protection laws, hiring subsidies can actually decrease permanent employment by inducing a disproportionate number of firms to replace old

[33] We have also estimated the model by assuming heterogeneous treatment effects for these job characteristics and, at the same time, allowing the effect of the policy being different over the course of the employment spell. Again, we obtain that the strongest effect occurs during the first year of employment. Thus, among low qualified workers the exit rate from permanent contract is maximum (20% for women aged more than 45 and 26% among the young ones). In the service sector, these effects ranged from 11% to 16% for women and from 15 to 18% for men.

jobs with new ones, thereby leading to a higher level of turnover. Their main argument is that, while hiring subsidies do indeed stimulate job creation, once a job has been created the opportunity cost of keeping the match rises, since a firm need only create a new position to receive the same subsidy again. This opportunity cost is even higher when the job creation subsidy is combined with a discount in payroll taxes, since the latter normally last for two years (excepting those for older workers). In this regard, both policies may be favouring the labour market rotation of the workers who benefit from them.[34]

In this estimation we allow the policy effect to vary with the contract duration as we did in the previous section. Table A.2 (Appendix) gives detailed results of these estimations. The main result obtained here (shown in Table 12) is that the causal incidence found previously mainly applies to eligible workers—particularly female ones—who are in their first year of contract *and* were hired under a contract carrying national payroll tax deductions (CIF contract type).

These results differ notably by gender and age group. With respect to female workers under the age of 45, we obtain that regional subsidies increase the exit rate from permanent employment during the first year of employment among eligible workers -versus illegible ones-, holding contracts with national subsidies. In general, the exit rate increases by 23.98% among eligible young women and by 19.00% among middle-aged ones. Among these eligible female workers whose contract does not include payroll tax deductions, the exit rate decreases during the first year of employment by 19.05% for young women and by 7.34% for middle age ones. After one year of employment the incidence of regional subsidies on the exit rate from the permanent contract seem not be as much as relevant as before and the exit probability only increases 3.57% for eligible workers in the case of young workers holding a permanent contract without discounts in payroll taxes. For the rest of female workers groups aged below 45 the regional subsidies decrease the exit rate by between 1.5% and 1.9%. The results are slightly different for women over the age of 45.[35] For them, the effects of regional subsidies are not statistically during the first year of the contract. For contracts longer than a year, the exit rate among eligible workers appears to have decreased by around 24% for contracts without national subsidies while the effects are not statistically significant for the other types.

Table 12: Causal effect of the policy on the exit probability from a permanent contract: the effect of national subsidies availability (1997-2004)

		Estimated exit Prob. (no policy)				Change in Prob. due to the policy			
		Women		**Men**		**Women**		**Men**	
		CIO	**CIF**	**CIO**	**CIF**	**CIO**	**CIF**	**CIO**	**CIF**
Contract	< 30	10.39%	6.58%	8.56%	4.42%	-19.05%*	23.98%**	5.08%**	7.46%**

[34] As we pointed out earlier, the MCVL dataset does not allow us to identify the types of firing costs associated with each type of permanent contract. Data provided by the Spanish Employment Agency indicates that all contracts without national subsidies have high firing costs (45 days of compensation per year worked). Firing costs for the nationally subsidized contracts can be either high (45 days) or low (33 days) depending on the firm. However, a close look at all of contracts signed in 2006 that received a national subsidy shows that nearly 66% of them carried low firing costs (33 days of compensation per year worked).

[35] For this group of workers, this estimation does not capture the fact that regional subsidies might reduce the exit probability from a permanent contract after the first year of employment as we have shown previously.

	30-45	7.89%	5.37%	6.41%	3.39%	-7.34%**	19.00%**	6.70%**	-10.47%**
	>45	8.70%	4.80%	5.17%	2.97%	3.61%	-0.37%	14.01%**	28.33%**
Contract Length >12	< 30	5.54%	5.36%	5.76%	4.93%	3.57%	-1.96%**	3.04%**	17.76%**
	30-45	4.15%	4.35%	4.27%	3.79%	-1.74%**	-1.59%**	3.89%**	45.58%**
	>45	4.58%	3.88%	3.44%	3.32%	-24.62%**	-0.51%	-1.95%**	-16.54%**

Notes: See Table 10.
CIO=Permanent contract without national subsidies. CIF=Permanent contract with national subsidies.

Turning to male employees, we can observe that, as in the general model, regional subsidies increased the exit rate from the permanent contract for almost all eligible workers in our sample. For young workers, we find that these subsidies always increased their exit rate, with the effects being clearly more pronounced when the he held a nationally subsidized contract (7.46% versus 5.08% for the first year contract and 17.76% versus 3.04% afterwards). In contrast to our result for female workers, we find that for males, the estimated effects increased after the first year of the contract, when they rose from 7.46% to 17.76%.

The effect of regional subsidies on middle-aged male workers holding different types of contracts differed from that found for younger workers. For this group, we find that these subsidies increased the exit rate among those holding contracts without national tax deductions, which was not the case among those with such deductions. However, for contracts lasting longer than one year, the results obtained for middle-aged male workers resemble those obtained for younger workers, although the effect over the exit rate is even greater (45.58% among those with national subsidies). Nevertheless, these results must be treated with caution, since the proportion of workers with national subsidies is lower for this group than it is for other ones.[36]

Finally, for male workers over the age of 45 the main difference in the causal effect of the policy depends on the duration of the contract. During the first year of employment, eligible workers with national subsidies faced a greater causal effect (28.33%) than did those without national subsidies (14.01%). The economic relevance of this heightened risk of job loss among older eligible male workers should be underlined here, since it affected such a high proportion of the sample population. By contrast, for contracts lasting longer than one year the subsidies had just the opposite effect, lowering the exit probabilities among eligible workers with (-16.54%) and without (-1.95%) national discounts in payroll taxes.

Hence, we have found that the incidence of regional subsidies on the labour market stability of the workers in our sample differs markedly by contract length and type. With regard to the former, the most relevant effect of these subsidies on the exit rate from a permanent contract centres on those contracts that also benefit from national payroll tax deductions. That is, the unintended side effects of the regional policy might be reinforced when the worker also benefit from national policies. The results are less homogeneous with regard to contract duration. Here, regional subsidies have significant causal effects on eligible female workers during the first year

[36] These differences mainly result from the eligibility restrictions for permanent contract with national payroll tax deductions.

of the contract, increasing the exit rate from the permanent contract with national subsidies. The latter effect also arises among older male workers and, to a lesser extent, among male workers under the age of 30.

The interpretation of these results is not straightforward. On the one hand, one could affirm that regional subsidies applied in combination with national payroll tax discounts help firms "try out" different workers for permanent positions by financing the rotation of workers during the early stages of their contracts. Hence, it seems that the joint availability of both national and regional subsidies may be cancelling out the possible negative effect of firing costs on both the hiring and firing of specific kinds of workers (essentially younger and female ones).

An alternative view is that firms are using these new subsidized permanent contracts as a substitute for temporary contracts. That is, Spanish labour policies that target permanent employment can be seen as one way to reduce the differences, mainly in terms of labour costs, between permanent and temporary contracts. Recall that job creations subsidies seem to compensate for the firing costs associated to permanent contracts. In this regard, Cebrian, Moreno and Toharia (2005) also point to the greater instability of the new permanent contracts –benefited with lower firing costs-, introduced with the 1997 reform, as compared to the ordinary ones –with high firing costs-, and conclude that encouraging employers to reduce the number of temporary employees by subsidizing new permanent contracts which carry lower firing costs might be leading them simply to redefine these contracts without enhancing employment stability.

6 Conclusions

Policies that aim to foster stable employment by subsidising new permanent contracts currently stand out as one of the main tools to active labour market policies, not only in Spain but also across Europe. Despite this undeniable political relevance, the available empirical literature stresses the limited benefits of such policies and points to the sometimes unexpected side effects that they may produce. These conclusions have been voiced in a number of studies, including those by Calforms (1994), Martin and Grub (2001) and Mortensen and Pissarides (2001).

Since 1997, several Spanish regional governments as well as the Spanish national government have implemented a number of different policies designed to reduce the high rate of temporary employment in the Spanish labour market by targeting specific worker groups.

In this paper we study the causal influence of job creation subsidies designed by regional governments on the duration of the subsidized permanent contract. Our main goal has been to evaluate whether such subsidies might cause certain side effects (such as increased job turnover among eligible workers with a permanent contract) that would limit their potential benefits on the permanent employment rate. We work with a longitudinal database, using a triple difference approach to identify the causal effect of the policies analyzed.

From the results presented above, we conclude that regional subsidies influence the labour market stability of some workers, but that the effects of these funds differ markedly by contract duration and type. In particular, we obtain that eligible workers are more likely to exit from the permanent contract during the first year of the contract and that this probability increases when the contract also benefited from national payroll tax deductions. Moreover, these effects seem to be clustered around low-qualified male workers in smaller firms and around jobs in the service sector (for all younger workers) and the industry sector (for younger male workers).

These estimated causal effects of regional subsidies on the exit probability from a permanent contract seem to be more significant than those we obtained elsewhere for temporary and unemployed workers on the entrance probability to a permanent contract (see García-Pérez and Rebollo, 2009). Hence, it seems that regional subsidies that favour the creation of new permanent positions may be encouraging firms not only to hire more permanent workers, but also to fire these workers more frequently. As a result, such subsidies ultimately do very little to increase the prevalence of permanent employment in the labour market to which they are applied. This raises the question as to whether or not such subsidies ultimately serve to increase labour market rotation rather than labour market stability.

Our result accords with the main conclusions of Mortensen and Pissarides (2001). They argue that although regional subsidies are designed to support employment, once a job is created the opportunity cost of keeping it increases, giving rise to a higher firing rate. In the case of Spain, this effect seems also to be reinforced by the existence of national subsidies.

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APPENDIX

Table A.1: Main Results for the complete model (1995-2004): Exit from the Permanent contract

		Women		Men	
		Coef.	T-S	Coef.	T-S
Contract Duration	ln(t)	0.188	7.16	0.094	3.89
	Ln(t)^2	-0.208	-18.22	-0.150	-14.44
	T=4 quarter	0.143	7.93	-0.055	-3.05
	T=8 quarter	-0.058	-2.08	-0.038	-1.55
	T=12 quarter	-0.086	-2.05	-0.146	-3.99
Policy Variables	Dijt*(Age< 30)	0.011	0.93	0.031	3.3
	Dijt^2*(Age< 30)	4E-04	0.45	-0.001	-1.7
	Dijt*(Age30-45)	0.016	1.15	0.023	1.79
	Dijt^2*(Age30-45)	-0.001	-0.84	-4E-04	-0.34
	Dijt*(Age>45)	-0.026	-1.18	-8E-04	-0.04
	Dijt^2*(Age>45)	0.002	1.33	0.003	1.92
Personal Characteristics	Age	-0.158	-27.22	-0.135	-25.09
	Age^2	0.002	20.09	0.001	17.76
	Immigrant	0.742	25.85	0.811	36.26
Age groups	Age <30	-0.581	-0.9	-1.4	-2.63
	Age 30-45	0.499	0.74	-0.129	-0.23
Job Characteristics	Part-time	0.19	14.42	0.505	28.87
	Layoff	1.068	85.83	0.957	83.32
Job Qualification	High	-0.851	-33.62	-0.645	-30.38
	Medium-High	-0.567	-34.6	-0.325	-20.19
	Medium-Low	-0.292	-22.03	-0.215	-17.5
Firm Size	20-100 Employee	0.182	9.87	0.203	10.98
	5-20 Employee	0.203	11.01	0.276	15.01
	< 5 Employee	0.251	15.23	0.382	21.91
	New Firm	0.123	9.82	0.183	15.42
	Private	0.376	9.73	0.393	7.43
	Temporary Help Agency	0.357	4.56	0.378	3.87
	Contract with national subsidies	-0.232	-19.35	-0.373	-30.68
Sector of Activity	Construction	0.019	0.41	0.148	7.4
	Services	0.312	16.2	0.396	26.81
Previous Labour Path	Temp. Contract (Previous)	-0.198	-8.01	-0.256	-9.28
	Conversion From Temp. Contract	-1.236	-46.55	-1.243	-43.29
	First Spell	0.169	4.81	0.012	0.33
	Nº of Unemployment Spells	0.086	56.41	0.093	64.07
	Constant term	-5.566	-6.81	-4.604	-7.55

Note: In the estimation we also include dummy variables by years, quarters and regions, as well as the interactions of years and regions, years and age groups and regions and age groups.

Table A.2: Main Results of the model for the period 1997-2004: Exit from the Permanent contract

		Women		Men	
		Coef.	T-S	Coef.	T-S
Contract Length	ln(t)	0.425	12.53	0.166	5.78
	Ln(t)^2	-0.348	-24.28	-0.233	-18.64
	Ln(t)* "Fomento" type	-0.658	-13.63	-0.218	-4.42
	Ln(t)^2	0.396	18.96	0.282	13.68
	T=2 quarter	0.145	7.92	-0.033	-1.79
	T=4 quarter	-0.1	-3.5	-0.055	-2.2
	T=12 quarter	0.02	0.47	-0.124	-3.29
Policy Variables	Dijt*(Age < 30)*(T<=4)*"Fomento" type	0.046	3.49	0.017	2.96
	Dijt*(Age < 30)*(T>4)*"Fomento" type	-0.004	-3.76	0.039	5.89
	Dijt*(Age< 30)*(T<=4)	-0.046	-3.88	0.012	2.75
	Dijt*(Age<30)*(T>4)	0.007	6.55	0.007	1.43
	Dijt*(Age30-45)*(T<=4)*"Fomento" type	0.039	2.5	-0.026	-2.15
	Dijt*(Age30-45)*(T>4)*"Fomento" type	-0.003	-2.49	0.089	8.74
	Dijt*(Age30-45)*(T<=4)	-0.017	-0.88	0.016	2.65
	Dijt*(Age30-45)*(T>4)	0.004	2.02	0.009	1.33
	Dijt*(Age>45)*(T<=4)*"Fomento" type	-8E-04	-0.03	0.057	5.17
	Dijt*(Age >45)*(T>4)*"Fomento" type	-0.001	-0.51	-0.041	-2.61
	Dijt*(Age >45)*(T<=4)	0.007	2.75	0.03	2.77
	Dijt*(Age>45)*(T>4)	-0.056	-1.89	-0.004	-0.31
Personal Characteristics	Age	-0.161	-27.02	-0.136	-24.94
	Age^2	0.002	19.96	0.001	17.61
Age groups	Age <30	-0.73	-2.87	-0.201	-0.83
	Age <30*"Fomento" type	0.144	2.49	-0.127	-2.91
	Age 30-45	-0.425	-1.58	-0.004	-0.01
	Age 30-45*"Fomento" type	0.230	3.66	-0.089	-1.83
	Immigrant	0.731	25.4	0.799	35.76
Job Characteristics	Part-time	0.184	13.73	0.505	28.61
Job Qualification	High (Job Qualification)	-0.847	-32.9	-0.638	-29.63
	Medium-High (Job Qualification)	-0.557	-33.4	-0.321	-19.63
	Medium-Low (Job Qualification)	-0.285	-21.08	-0.215	-17.21
	Layoff	1.063	84.01	0.95	81.52
Firm Size	20-100 Employee	0.174	9.35	0.197	10.51
	5-20 Employee	0.203	10.86	0.278	14.94
	< 5 Employee	0.247	14.73	0.377	21.32
	New Firm	0.124	9.65	0.184	15.24
	Private Firm	0.422	10.35	0.425	7.65
	"Fomento" type	-0.397	-6.91	-0.681	-14.4
Sector of Activity	Construction	0.021	0.45	0.146	7.21
	Services	0.299	15.92	0.399	26.56
Previous Labour Path	Temp. Contract (Previous Spell)	-0.196	-7.87	-0.257	-9.29
	Conversion From Temp. Contract	-1.217	-45.43	-1.189	-41.22
	First Spell	0.194	5.4	0.044	1.14
	Nº of Unemployment Spells	0.084	54.37	0.092	62.56
	Constant term	-0.682	-2.46	-1.481	-5.5

Note: "Fomento" type means that the permanent contract has national subsidies.