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The use of permanent contracts across Spanish regions: do regional wage subsidies work?

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RESUMEN

En este trabajo se evalúa el impacto de las políticas regionales que subvencionan los nuevos contratos indefinidos tomando una muestra de trabajadores con contratos temporales y desempleados para la economía española. En concreto, se estudia la probabilidad de entrada a un contrato indefinido a partir de una nueva base de datos procedente de los registros de la Seguridad Social denominada “La Muestra Continua de Vidas Laborales”, que se utiliza por primera vez para realizar ejercicios de evaluación de políticas públicas en el mercado de trabajo español. Esta base de datos ofrece importantes ventajas respecto a la Encuesta de Población activa dado que tiene toda la historia laboral del trabajador. Al disponer de variabilidad individual, regional y temporal en nuestro indicador de política aplicamos el método de triples diferencias para identificar el efecto medio de la política. Si bien las políticas regionales que subvencionan la contratación indefinida se llevan aplicando en España casi 10 años, este es el primer estudio centrado en su análisis. Nuestros principales resultados son que, en media, esta política tiene efectos positivos pero pequeños, en la probabilidad de transición a un contrato permanente, tanto desde del desempleo como desde un contrato temporal. No obstante, la incidencia de esta política es mayor cuando el trabajador parte de un contrato temporal. Es también mayor para las mujeres menores de 30 años mientras que para los trabajadores mayores de 45 años esta política no tiene impacto alguno. El cambio total en la probabilidad de entrar en un contrato indefinido a partir de uno temporal, cuando se recibe la subvención media de 5100 Euros, es aproximadamente del 26% para las mujeres jóvenes, y 24% y 22% para las mujeres y los hombres de entre 30-45 años. Sin embargo, dado que la probabilidad de conversión de un contrato temporal en indefinido es muy baja, este cambio en la probabilidad deja prácticamente inalterada la tasa de conversión de un contrato temporal a uno fijo. Por ejemplo, en el caso de las mujeres jóvenes, dicha tasa de conversión se estima que crece desde el 0.064% al 0.075% mientras que para los hombres de entre 30-45 esta crece de 0.039% a 0.041%.

Palabras clave: Diferencias-en-diferencias, Evaluación de Políticas, Subvenciones sobre los salarios, Modelos de Riesgo en Competencia.

ABSTRACT

This paper evaluates the effect of regional wage subsidies to foster permanent employment for a sample of temporary and unemployed Spanish workers. We study the transition into permanent employment using a new dataset based on administrative Social Security registers named “La Muestra Continua de Vidas Laborales”, which is used for the first time to carry out policy evaluation exercises in the Spanish labor market. This dataset offers important advantages with respect to the Labour Force Survey, given it offers the complete labor history of the worker. Moreover, since we have individual, regional and time variation in our policy measure, we can apply a difference-in-differences estimator to identify the average treatment effect of this policy. Though, these regional policies have been implemented for almost ten years, as far as we know, this is the first attempt to evaluate them. Our main results are that, in average terms, this policy has positive but small effects on the transition rate into permanent employment either from a temporary contract or from unemployment. The incidence of these subsidies, however, is larger when the worker is in a temporary contract. It is also larger for young females while for old male workers do not have any effect. Measured at the average wage subsidy (5100 Euros) the total change in the entrance probability from a temporary to a permanent contract is around 26% for young women, 24% for middle age ones and 22% for middle age men, the most benefited workers. Nevertheless, since the transition rates to permanent contracts at the same firm are pretty low these relative changes hardly generate a change in the transition probability from a temporary to a permanent contract at the same firm. For instance, in the case of young women the estimated transition probability growths from 0.064% to 0.075% while for middle age men it changes from 0.039% to 0.041%.

Keywords: Difference-in-differences, Evaluation Analysis, Wage Subsidies, Competing risk

JEL Classification: J38, J68

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1 Introduction[1]

In the last decade, fixed-term contracts have taken center stage in the economic debate on labor market reforms in Europe. The debate has mostly focused on the magnitude and characteristics of the phenomenon: very rapidly, temporary jobs have become a major novelty in the European labor markets. In this context, Spain is a paradigmatic example since it is the European country with the highest rate of temporary contracts, and this position has been maintained since the first half of the 1990s. This rapid expansion of temporary contracts has fuelled researches´ effort to understand their effects on labor market outcomes. On one hand, it is suggested that, they provide workers with a stepping-stone towards permanent employment (Booth, Francesconi and Frank, 2002). On the other hand, there is an increasing evidence that they might represent a "dead-end", in that they further segment the labor market between insiders holding permanent contracts and outsiders who find themselves confined at the margins, trapped between spells of unemployment and temporary contracts (for the Spanish evidence see Guell and Petrongolo, 2007 or Rebollo, 2006). This kind of concern has led policy-makers to intervene on fixed-term contracts in the recent years in an attempt to reduce their negative effects. According to OECD (2002), governments have intervened both by setting restrictions on the adoption of temporary contracts and by providing employers with incentives to either hire workers on permanent positions or convert temporary jobs into permanent ones. Spain is also a paradigmatic example in this context as both kinds of policies have been implemented since the second half of the nineties.

In the eighties, Spain was one of the tightest labor markets and its rate of unemployment was one of the highest in the OECD. This led Spanish policy makers to implement flexibility measures like the well-known 1984 labor market reform. The flexibilization strategy mainly consisted of introducing the possibility of hiring workers on flexible, fixed-duration contracts. As a result of this reform, temporary contracts rose from 18% in 1987 to 33% of all employees in 1994. Given this high rate, in 1997 and 2001 there were new regulations aimed at contracting the scope for using fixed-term contracts by reducing the firing costs for new permanent employees[2]. Additionally, since 1997, the Spanish government has subsidized the creation of permanent contracts by, in some cases, large discounts in firm's payroll taxes on new permanent

[1] We wish to thank Rosario Martagón, Lola Morales and Esther Ruiz for their help in collecting all the information about regional wage subsidies. We also wish to thank the participants at the Bank of Spain and CEMFI seminars and at the Workshop on Policy Evaluation celebrated in Seville in December 2005. Special thanks are due to Alberto Abadie, Raquel Carrasco, Michael Lechner, Sergi Jimenez, Juan F. Jimeno, Luis Toharia and Ernesto Villanueva for all their helpful comments. The financial support of the Ministry of Education (SEJ2006-04803) is also gratefully acknowledged. Of course, any remaining errors are our responsibility.

[2] In the reform of 1997 a new permanent contract was designed. The main differential characteristic of this new contract was its lower firing costs. Nevertheless, it was aimed at certain population groups, mainly long term unemployed and young workers. For further details see Kugler, Jimeno, Hernanz (2003).

contracts. During the same period, some regions also implemented wage subsidies for new permanent contracts. As a result, Spain has become the OECD country with the highest percentage of GDP, 0.28% in the period 1999-2002, devoted to subsidy permanent, employment. Hence, this is an important policy that, at least potentially, has affected many Spanish workers. As such, a rigorous evaluation of the program may lead to insights regarding the benefits of this policy. Yet, so far we can only find one research devoted to the evaluation of the effects of the 1997 reform into the permanent employment rate Kugler, Jimeno and Hernanz (2003).

The aim of this paper is to evaluate the effectiveness of regional wage subsidies to foster permanent employment. It is important to mention that though, these regional policies have been implemented for almost ten years, as far as we know, this is the first attempt to evaluate them. Our evaluation exercise focuses on the effects of wage subsidies over the flow to a permanent contract. Firstly, we study the conversion rate from a temporary to a permanent contract at the same firm considering also the alternatives of unemployment, moving to a different firm or having another temporary contract at the same firm. Secondly, we examine the transition rate from unemployment to a permanent contract also taking into account the alternative of getting a temporary contract. Thirdly, we check whether there are any substitution effects by examining the incidence of regional wage subsidies over the exit probability from a permanent contract considering as competing alternatives unemployment, a temporary contract and other permanent contract.

We take advantage from the multidimensional variability of the eligibility rules of these subsidies, since they vary across individuals, regions and time to identify the policy effects. Besides, the variation of the permanent employment rate among Spanish regions is somehow large and it differs from the variability found in regional wage subsidies: meanwhile the Southern regions show rates of permanent employment lower than 60%, in regions as Madrid or La Rioja these rates are basically similar to the European standards. Hence, as it happens with unemployment, the South of Spain concentrates a larger fraction of temporary contracts than in the rest of the country. At the same time, however, regions with a low rate of temporary employment, such as Madrid, have also implemented large wage subsidies on new permanent contracts.

Since we have longitudinal data, and given the characteristics of the policy to be analyzed, we consider that the best approach to study the effect of this policy is the difference-in-differences estimator (DID, hereafter). Our policy variable is based on the maximum amount of the wage subsidy offered for the eligible individual in each region. We identify the treatment effect using three dimensions of variability: i) between workers of the same region but with different individual characteristics; ii) between similar workers of different regions; iii) between periods with wage subsidies and periods without them. Thus, we extend the familiar DID approach to a Difference-in-differences-in-differences estimator (DDD).

We use a new administrative database obtained from the Social Security records (*Muestra Continua de Vidas Laborales)* where we have available the whole labor history of the worker. One contribution of this paper is the use of this dataset for the first time. It is highly interesting since it offers information about more than one million workers to study labor market transitions in the Spanish labor market and the effectiveness of these regional wage subsidies. Such transitions may be much accurately studied with this dataset, which overcomes the typical problems caused by the pseudo-panel structure of the Spanish Labor Force Survey.

We obtain that the impact of regional wage subsidies differs by gender and age and depends on the initial labor state of the worker. Firstly, the incidence of wage subsidies is larger from a temporary contract than from unemployment. Secondly, the impact of regional wage subsidies is the largest for young female workers while is zero for old workers. When the initial state is a temporary contract the benefited workers are young and middle age women and middle age men. For the rest of workers with temporary contracts the impact is not statistically significant. For the sample of unemployed, young workers are the most benefited ones followed by middle age ones. Thirdly, wage subsidies on new permanent contract might favor the permanence of employees at the same firm but by accumulating temporary contracts. Fourthly, we do not find evidence of substitution effects except for young female workers. Interestingly, we find that wage subsidies foster the transition between permanent contracts and between permanent and temporary contracts for treated workers.

Therefore, we conclude that the labor demand by type of contract in Spain hardly varies with wage subsidies on new permanent contracts. The average wage subsidy represents a discount of around 25% of the monthly labor cost while the conversion rate from a temporary to a permanent contract increases by 26% for young women, 24% for middle age ones and 22% for middle age men. The incidence of wage subsidies over the entrance probability to a permanent contract from unemployment is even smaller and it reaches the maximum for young workers. The total variation in the exit probability to a permanent contract is 14.97% and 14.4%, respectively. The impact for middle age workers is 6% for women and 8% for men. Nevertheless, since the transition rates to permanent contracts are pretty low these relative changes hardly generate a change in the transition probability to a permanent contract. For instance, in the case of young women the estimated transition probability from a temporary to a permanent contract at the same firm growths from 0.064% to 0.075% while for middle age men it changes from 0.039% to 0.041%.

The rest of the paper is organized as follows. First, we will briefly review previous research on the use of temporary contracts, emphasizing those focused on the Spanish case. In Section 3, we describe the main characteristics of the regional wage subsidies. Sections 4, 5 and 6 describe the econometric approach, the data used and the main results. Finally, Section 7 presents our main conclusions.

2 Background

Wage subsidy programs have been implemented in many OECD countries, as an important measure to foster labor market integration of disadvantaged workers, not only in countries using the most interventionist approach to labor market policy such as France, Germany or Sweden, but also in countries firmly attached to the Anglo-Saxon model of the labor market, like USA, UK, and Canada. Nevertheless, two recent surveys of labor market policies in a number of countries have reached rather conflicting conclusions about the relative efficiency of different active labor market policies (ALMPs, from now on). Heckmam (1999) states, after surveying a large number of studies about European ALMPs, that they can not conclude any ALMP consistently yields greater employment impact than any other. On the contrary, Martin and Grubb (2001) conclude that in several OECD countries, evaluations have found that wage subsidies programs have a greater impact than public training programs on direct job creation measures.

There are various forms of wage subsidy schemes in different countries. Wage subsidies can be paid either to employers or to employees. The subsidy itself can be paid as a tax credit, as a reduction in Social Security contributions or as a direct payment. Although, wage subsidy programmes have much variance in their designs, their ultimate goal is to increase employment through lowered labour costs or increased labour supply. In the case of targeted wage subsidy, this policy aims to provide increased employment opportunities to the targeted worker group by changing labour demands to be favourable towards the targeted worker. However, wage subsidy programmes have been criticised due to their high levels of deadweight losses and substitution effects. Also, some subsidy programmes have encountered low participation of employers.

Despite the fact that economists have advocated employment subsidies for a long time, there are only a few theoretical models which provide analytical results concerning the effect of employment subsidies on labor market flows. Katz (1986) provides a partial equilibrium dynamics analysis of wage subsidies for low-wage workers. The employer-side wage subsidy is expected to stimulate the demand for workers, and shifting out the labor demand curve by reducing the cost of labor relative to other inputs. Katz points out that, in a situation of structural unemployment in which the effective labor supply is completely elastic, a proportional wage subsidy will not affect the wage of the worker, but will increase employment in proportion to the elasticity of labor demand for low-wage workers. However, these impacts could be reduced by indirect effects of wage subsidies, at least in the short-run. The indirect effects that have received special attention in relation to wage subsidies are the deadweight loss, the substitution effect and the displacement effect (Calmfors, 1994). The deadweight loss is defined as the hiring from the target group that would have occurred also in the absence of the programme. Calmfors also define the substitution effect as the extent to which jobs created for a certain group of workers simply replace jobs for other categories, because relative wage costs are changed. This effect might occur when the subsidy is targeted on a subgroup. The displacement effect is defined as the possible reduction of jobs elsewhere in the economy because of competition in goods market. This effect can occurs when a firm with subsidized workers increases output, but displaces output among firms who do not have subsidized workers and as a result, the subsidy crowds out employment.

There is a more recent theoretical work that offers new insights about the effects of subsidies on employment creation. Mortensen and Pissarides (2001) deliver insight into the effect of different types of subsidies on job creation and job destruction by integrating subsidies in search and matching equilibrium framework with endogenous job-destruction. Other interesting fact of their analysis is that they distinguish two possible scenarios. One in which the unemployment benefit replacement ratio is moderate and unemployment protection legislation is weak, as in the US, and another in which both the level of unemployment compensation and the extent of employment protection are relatively high, as in Europe. The outcome variables studied are the steady state unemployment rate, the average wage and the aggregate income. Hence, their results mainly could help us to understand the effects of wage subsidies over the transition rate from unemployment to a permanent contract. In their analysis, they distinguish two types of subsidies: hiring subsidies paid at once, at the start of an employment spell, and employment subsidies which provide a flow of subsidies during the job duration. They also present the results for two levels of worker skill (high and low productivity worker). The values of the subsidy considered in their simulations range from 0 to 20% of the average wage. These values are similar to the average wage subsidy we are analyzing. Their main findings for economies with high unemployment compensation and stringent employment protection can be summarized as follows: hiring subsidies lead to shorter unemployment durations but increase unemployment incidence for high skill workers while it reduces for low skill workers. The reason for this difference is the asymmetric effect of firing costs on job creation by skills levels: the hiring subsidy may offset the negative effect of the firing cost on job creation for low skill workers. Therefore, a hiring subsidy can decrease employment by encouraging too much replacement of old jobs with new too frequently and consequently lead to a higher turnover on the labor market. This last situation might take place mainly when the hiring subsidy offset the effects of employment protection, mainly firing costs. The idea is that a hiring subsidy stimulates job creation, however once the job is created, the opportunity cost of maintaining the match rises since the hiring subsidy can again be obtained when creating a new job.

Since in this model there are not differences on firing costs by type of contracts the theoretical conclusions can not be generally applied to the situation we analyzed in this research. Nevertheless, it offers some interesting insights. The 1997 reform created a new permanent contract with lower firing costs for certain group of workers. Additionally, it introduced discounts in payroll taxes for new permanent contracts for certain type of workers[3]. All these facts might have introduced a significant discount in labor costs for certain type of workers

[3] The 1997 reform initially reduced firing costs by around 25% and payroll taxes by between 40% and 60% for newly signed permanent contracts –disabled workers had larger discounts-, and 25% and 60% for all conversions of temporary into permanent contracts, but after 1999 these reductions applied only for conversions of young and older workers. These discounts in payroll taxes were mainly applied during the first year of the contract though in some cases they could be extended to two years. The groups of workers who received the larger discounts in labour costs for new permanent contracts were younger and older workers, long term unemployed, women under-represented in their occupations and disabled workers.

(mainly young workers)[4]. Therefore, it could happen that the introduction of wage subsidies could have given rise to a situation in which hiring subsidies compensate the dismissal costs and as a result they have encouraged transitions between new permanent contracts.

There are two additional papers that consider explicitly a labour market characterized by two types of labor contracts: one without firing costs and other with large firing costs. Dolado, Garcia-Serrano and Jimeno (2002) consider the existence of dismissal costs and argue that the equilibrium ratio of temporary to permanent employees is determined by the ratio among unit labor costs[5] under each of those contracts, the elasticity of substitution between temporary and permanent workers, the volatility of labor demand along the business cycle and the average economic growth rate. They distinguish three components within the concept of labour costs: the wage, the firing cost and the hiring costs. If we interpret the wage subsidy as a drop in the hiring cost for new permanent employees relative to temporary ones, the effect of this policy should be to increase the transition rate to permanent contracts and consequently the permanent employment rate.

Kugler, Jimeno and Hernanz (2003) analyze the effectiveness of a labor market reform consisting in a drop of firing costs and an introduction of discounts in payroll taxes of new permanent contracts to reduce the share of temporary workers. The final aim was to analyze the impact of the Spanish labor market reform of 1997 into the permanent employment rate. Moreover, given the characteristics of the reform, it might provide an interesting estimate of the elasticity of permanent employment with respect to non-wage labor costs. They develop a simple dynamic matching model similar to Blanchard and Landier (2002) but they endogenize dismissals and introduce payroll taxes. In their model the demand for permanent and temporary employment depends on two productivity thresholds that depend, among other things, on dismissal costs and payroll taxes. Given the values of the two productivity thresholds they can derive the steady-state values of temporary and permanent employment. Their theoretical discussion shows that the 1997 reform should, more likely than previous reforms, have increased employment levels. Their model suggests that a reduction in dismissal costs for permanent workers increases hiring and firing and therefore has an ambiguous effect on unemployment. On the contrary, a reduction in payroll taxes for permanent contracts increases conversions from temporary to permanent contracts but leaves dismissals unchanged so the net effect is to increase permanent employment. Their empirical results suggest that reducing the costs of permanent employment may be of special value for younger workers. They obtain that the transition probability to permanent employment increased between 30% and 46% for younger men and between 14% and 24% for younger women during the reforms years. The results for older workers were insignificant. They estimate the elasticity of permanent

[4] For instance, Kugler, Jimeno and Hernanz (2003) estimate that dismissal costs and payroll taxes account for about 1.9% and 21.6% of total labour costs for young men workers, respectively. Given the corresponding percentage change in payroll taxes and dismissal costs as a result of the reform, the percentage reductions in total labour costs implied by the reform for young men they estimate it was between 7%-9.2%

[5]In this line, García-Pérez and Rebollo (2006) show how the behaviour of unit labour costs is an important determinant of the divergences in the permanent employment rate across Spanish regions. They show that regions with larger unit labour costs will tend to have lower rate of permanent employment, once structural characteristics such as regional specialization are taken into account.

employment to non-wage labor costs and they find a fairly elastic response of permanent employment to non-wage labor costs for younger workers, for whom the payroll tax reduction was relatively more important.

Though the characteristics of the reform they analyzed were different from our policy analysis, main theoretical implications can be applied to the policy evaluated in this paper from the point of view that wage subsidies imply a drop in hiring costs of new permanent contracts keeping these costs constant for temporary contracts.

2.1 Some Previous Empirical Results

The recent microeconometric evaluation literature on wage subsidies has mainly focused on two different outcomes. A first branch of research looks at the effects of wage subsidies over the probability of being unemployment (Gerfin et al, 2005). Other studies focus on the transition rate out of unemployment (e.g. Lubjova and Van Ours, 1999; Fredriksson and Johansson, 2004, Forslund, Johansson and Lindqvist, 2004; Gobel and Cokx, 2005) and the transition rate to permanent employment (Cipollone and Guelfi, 2003). Lubjova and Van Ours (1999) have evaluated the effect on the transition rate to regular employment of two different subsidies employment program for the Slovak Republic. They report a positive effect of the short-term subsidies in the public sector and a negative effect for the two year subsidized jobs in the private sector. Forslund, Johansson and Lindqvist (2004) evaluate an employment subsidy program in Swedish and conclude that employment subsidies increased the flows into jobs of participants compared to non-participants[6]. Gobel and Cokx study the impact of subsidized employment for long-term unemployed workers in Belgium on the transition rate from employment to non-employment[7] and find that the policy decreases this transition rate only in the first year of participation whereas for men it increases after participation. Cipollone and Guelfi (2003) examine the effects of a tax credit to firms choosing to hire workers under permanent contracts for the Italian economy. They conclude that firms use this tax credit provision to hire under permanent contracts workers who, on average, turn out to have the highest probability of being permanently hired even without the subsidy.

It is difficult to draw general conclusions from these recent empirical studies. The programme implementation[8], labor market conditions and the institutional setting in the different countries seem to play a role for the effect of subsidized employment. Our paper contributes to the literature by evaluating the effect of wage subsidies on new permanent contracts over the transition rate to permanent employment from unemployment and from a temporary contract.

[6] They follow a matching and instrumental variables in their econometric approach.
[7] They account for selective participation on the basis of a multivariate duration model with correlated unobserved heterogeneity.
[8] Differences in the eligibility criteria for the treated group, the duration and amount of the subsidies might play a role for the effectiveness of a program.

3 Regional Wage Subsidies on new permanent contracts

The promotion of permanent contracts has emerged as an important labor market policy from both the regional and national level in a lot of countries. Besides, this importance has been increasing since the second half of the 1990s, so nowadays, Spain devotes an important amount of resources to the promotion of permanent employment.[9] In the case of wage subsidies on new permanent contracts, this is an initiative that several Spanish regions established since 1997 when the national government implemented important discounts in firm's payroll taxes on new permanent contracts. While this last policy is common for all Spanish workers, regional wage subsidies have varied between and within regions, by individual characteristics and along time in two dimensions during the period 1997-2004: i) the eligibility conditions; ii) the amount of the subsidies. The eligibility rules are mainly based on some individual characteristics of the worker such as age and gender[10]. For example, some regions offer these subsidies only for women whereas others restrict their use just for young workers. Besides these eligibility conditions might vary within a region and across time. On the other hand, the amount of the wage subsidy has also varied. These facts allow us to examine the response of different group of individuals to the change in incentives.

One of the main characteristics of the policy we are analyzing is that the firm that hires the worker must apply for the subsidy to receive it. We can not identify which worker has effectively been benefited from this regional policy given the origin of our dataset, Social Security registers. Therefore, our treatment group will be composed by workers who fulfill the eligibility criteria and consequently, our analysis will measure the potential impact of regional wage subsidies over the creation of new permanent contracts. That is, it is possible that workers considered as potentially treated are not effectively treated by this policy and therefore, our analysis could overestimate the incidence of wage subsidies over the entrance rate to new permanent contracts. Nevertheless, since this subsidy implies a drop in total labor costs for the firm, it is plausible to assume that the firm will apply for it whenever the eligibility conditions are fulfilled.

Our data of regional wage subsidies is taken from each regional government and it only covers wage subsidies at the regional level[11]. The main characteristics of these wage subsidies are described in Tables 1 and 2. The average subsidy for each region and by age and gender, in 2002 Euros, is shown in Table 1, with their minimum and maximum amount. As it is shown in Table 1, this policy was implemented since 1997 in some regions, whereas in others it was implemented afterwards or never, as in Cataluña or Navarra[12]. Table 1 also shows that the

[9] For instance, Spain devoted 50% of total ALMPs expenditure on subsidized employment during the period 1996-2002. Only Italy and Belgium devoted more resources (53% and 56% respectively). The average for the European Countries is around 30%.

[10] The eligibility conditions of the regional wage subsidies may also depend on other variables such as the labour state of the worker or certain characteristics of the firm such as the type of activity. Nevertheless, given we can not control for all the eligibility conditions in our database, we have finally opted to consider only eligibility conditions based on age and gender.

[11] It could be that some local governments also offer wage subsides or any other kind of public subsidy to foster permanent contracts.

[12] Before 1997 we can find regional policies to foster employment for specific group of workers. Those regional policies are not specifically designed to foster permanent employment and therefore they are not considered in this analysis.

eligibility conditions vary notably among regions and time. We find regions as Andalusia where the policy applies to all workers while other focus this policy on certain group of workers such as women or young workers.

We also wonder whether the effectiveness of these wage subsidies depends on the amount of the subsidy. The idea is that the larger the amount of the subsidy the larger is the reduction in labor costs when the firm hires a permanent contract and the larger should be the incidence of the policy over the entrance rate to new permanent contracts. That is, it could be that the policy does not foster permanent contracts because the wage subsidy does not significantly cover the larger labor costs related to the new permanent contract. In Table 2 we display a summary of the regional wage subsidies by region and individual characteristics. Again, we can observe there is a strong variability in the intensity of the policy among regions and by personal characteristics. Moreover, the contribution provided by this subsidy looks quite generous. Last column of Table 2 shows the percentage reduction in per-cápita labor costs due to the subsidy (using data for the year 2000) by regions. The evidences show a labor cost reduction which ranges from 9% in Baleares, the region with the lowest subsidy, to more than 60% in Extremadura. Therefore, we will use this additional source of variability to identify the average effect of the policy over the transition rate to a new permanent contract.

We face one shortcoming with this type of information that it is important to mention. The data available refers to the maximum wage subsidy the firm can receive per contract and year. Initially it seems more reasonable to use the minimum wage subsidy but in many cases the information available fixes this minimum at zero. The use of the maximum wage subsidy implies that our results measure the maximum incidence of the subsidy.

4 The Empirical Approach: Identification and Estimation Method

Our goal in this paper is to measure the causal effect of regional wage subsidies over the flow to a permanent contract either from the conversion of a temporary contract or from unemployment. For this exercise we will estimate three duration models following a multiple risk approach to avoid the biases related to single risk models. Firstly, we estimate the transition from a temporary contract considering that the worker can move to unemployment or other firm, other temporary contract at the same firm or to a permanent one also at the same firm. Secondly, we estimate the transition from unemployment considering that the worker can exit to a temporary contract or to a permanent one. Thirdly, we estimate the exit rate from a permanent contract considering as competing alternatives unemployment, other temporary contract and other permanent contract.

As it is traditional in the literature of duration models the objective here is to estimate the exit or hazard rate. For each individual we observe the duration in a determined state -duration of an episode of temporary contract or unemployment-, from $t=1$ up to the month k in which the individual changes of situation to any of the competing alternatives. A common alternative to estimate the hazard rate consists of transforming the duration model in a sequence of discrete choice equations defined on the surviving population at each duration (Jenkins, 1995). In this case, we define a binary variable y_{lkt}, that takes value one when the worker changes state at time t from l to state k, and zero otherwise. This expression has exactly the same form that the likelihood function of a discreet choice model where y_{lkt} is the binary endogenous dependent variable, once we have rearranged the database so there are so many rows by individual as time intervals -months in this case-, in which the worker has remained in the initial situation (Allison, 1982; Jenkins, 1995). Hence our likelihood function for each individual is:

$$\ln l_i = \sum_{t=1}^{T}\sum_{l=1}^{L}\sum_{k=1}^{J} y_{lkt} * \ln\left(\Pi_{lkt}\right) + \left(1 - y_{lkt}\right) * \ln\left(\Pi_{lkt}\right) \qquad (1)$$

where Π_{lkt} is the conditional exit rate, that is, the hazard rate from state l to the destination state k:

$$\Pi_{lkt} = \Pr\left(T_{lkt} = t \,/\, T_{lkt} > t\right) \qquad (2)$$

4.1 The identification approach

Our econometric approach to identify the treatment effect of these regional wage subsidies is directly linked to the standard causal effects analysis. In particular, we follow a Difference-In-Differences (DID, hereafter) approach.

To estimate the effect of policies on economic behaviour one needs a source of policy variation. Empirical studies tend to use spatial and temporal variation for estimating the effect of government policies on economic outcomes (Besley and Case, 2000). Our identification strategy of the policy incidence is interesting because we use several dimensions of the policy variability. The individual, regional and time variability of the wage subsidies provides us with many sources of identification of the unbiased estimator of the policy effect. That is, we use similar workers in different regions and different workers within the same region as control groups. The treatment to be analysed is the wage subsidy. Therefore, the treatment group is composed of those potentially affected by the policy at time t . Obviously, the control group[13] is those workers who are not potentially affected by the wage subsidy at time t.

The general specification of the competing risk duration model we estimate is the following for each origin state:

$$\Pi_{ijt} = F(\delta(t) + \alpha_t^{'} x_{ijt} + \gamma_t^{'} z_{jt} + \beta_0 D_{ijt} + \beta_1 D_{ij} + \beta_2 D_j + \beta_3 D_t + \beta_4 D_{jt} + \beta_5 D_i + \beta_6 D_{it} + \varepsilon_{ijt}) \qquad (3)$$

[13] Good control groups will be those whose behaviour has evolved similarly to those of the group experiencing the policy change and who respond similarly to changes in the variables that derive policies to change. The appeal of the DID estimation comes from its simplicity as well as its potentiality to circumvent many of the endogeneity problems that typically arise when making comparisons between heterogeneous individuals (See Meyer 1995, for an overview).

where i stands for individuals, j for regions and t for time; the function δ(t) control for the duration dependence of the process; the matrix x_{ijt} contains covariates that vary among individuals and are mainly related to time varying personal and labor characteristics; z_{jt} stands for time varying covariates specific to the region where the individual works[14]. These variables help to identify an unbiased estimate of the policy's effects as they adequately reflect the incidence of changes in other variables that are simultaneously influencing outcomes of the control and treated group under study. This idea is relevant since using individual and regional time varying covariates, we extend identification to those instances in which observed compositional differences between treated and controls cause non-parallel dynamics in the outcome variable.

The wage policy variable is D_{ijt} and takes a positive value that represents the wage subsidy when the worker i, located in region j at time t is living in a region with wage subsidies and she is eligible, and zero in other case[15]. The rest of "*D*" variables help to identify an unbiased estimate of the average treatment effect of the policy:

D_j=1	⇔ Worker is located in a treated region
D_i=1	⇔ Individual characteristics of the worker are eligible independently of being in a treated region at the time the policy is implemented.
D_{ij}=1	⇔ Worker belongs to the eligible group in a treated region
D_{jt}=1	⇔ Worker is in a region that applies the wage subsidy at time t, independently of being eligible within a region
D_j=1	⇔ Worker is located in a treated region
D_{it}=1	⇔ Worker has eligible characteristics at time t, independently of being in a treated region.
D_t=1	⇔ Policy is implemented at time "t"

Summarizing, the variables D_i, D_j and D_{ij} control for permanent differences between eligible and non eligible workers, treated and non-treated regions and eligible and non-eligibles individuals

[14] Besley and Case (2000) shows that fixed effects models might mislead the effect of the policy. A potential source of bias is due to the presence of unobservable variables that may determine both the policy and the outcome of interest. In our case, it could be possible that some unobservable measure of pessimism about the region's potential for economic growth may influence both the existence and generosity of the policy and the type of contracts in a particular region. Therefore, the individual variability in the eligibility conditions within regions plays an important role in this study.

[15] Initially, our policy variable was a dummy variable that took value one when the worker *i*, located in region *j* at time *t* was living in a region with wage subsidies and she was eligible, and zero in other case. But with this specification we lost one dimension of variability in the policy variable. Since, in DID approach policy variation is fundamental we have opted to analyze the incidence of the policy considering the amount of the wage subsidy.

within a region, respectively. The variables D_{it} and D_{jt} controls for time-varying individual and regional effects. Finally, D_t controls from common aggregate effects to treated and non-treated workers that could influence the outcome variable. Formally, the analysis should include all the variables pointed out above but given the characteristics of the policy we face with several restrictions. For instance, the effect of the covariates D_i and D_{it} can not be identified directly but by including in x_{ijt} the covariates that determine the eligibility conditions, basically age and gender.

Therefore, the key issue from a policy point of view concerns the sign, size and significance of the estimated parameter β_0 that measures the true effect of the policy once we have controlled for all the covariates that could simultaneously affect the treatment and the outcome. The estimation of the policy incidence on the treatment group, β_0, is estimated as the post-treatment change in the outcome for the treatment group, after controlling for the mean change in outcomes observed pre and post-treatment and for the mean differences in outcomes between the treatment and the control group. The parameters β_1 and β_2 are the treatment group and region specific effects and they account for average permanent differences between the treatment and control group in the first case and between treated regions and non-treated region in the second case[16]. The parameter β_3 and β_4 show the existence of national and regional aggregate effects from the policy. Finally ε_{ijt} is the error term whose composition is the following:

$$\varepsilon_{ijt} = \eta_i + \upsilon_{ijt} \qquad (4)$$

where η_i describes unobserved time-invariant differences and υ_{ijt} the random error term of the model. We will assume that the random component υ_{ijt} is independent of both, the individual and the region effects. Recall that one advantage of the DID approach is that it controls for unobserved time-invariant differences. Therefore the estimation of the parameter β_0 should be the same when estimating the model with and without unobserved heterogeneity. Nevertheless we will also estimate the model with a control for unobserved heterogeneity in order to test whether the policy parameter β_0 changes. Nevertheless, this is really relevant in the analysis of the duration dependence of the exit rate[17].

There are two main identification assumptions maintained in this DID estimation. The first one is that, apart from the control variables, there are no other forces affecting treatment and control groups. In addition, the composition of the treatment and control group must remain stable

[16] The inclusion of these variables comes from the fact that we are not working with a truly randomised experiment. Note that in a randomised experiment, where subjects are randomly selected into treatment and control groups, β_1 should be zero, as both groups should be nearly identical.

[17] It is well known that duration analysis produce incorrect results if unobserved heterogeneity is ignored. On average, subjects with relatively hazard high hazard rates for unobserved reasons leave the state of interest first, so that samples of survivors are selected. Differences between such samples at different times reflect behavioural differences as well as this selection effects. Nevertheless, with a Difference in Differences approach the concern of the effects of unobserved heterogeneity on the policy parameter should not be relevant. Therefore, we will estimate the model with unobserved heterogeneity as a specification test over the policy variable. We follow the approach proposed by Heckman and Singer (1984) to specify the heterogeneity term. We assume that each hazard rate has two support points. Besides, we allowed two types of individuals, so that each type is characterized by a unique set of points of support and the corresponding probability.

overtime[18]. Therefore, to provide an unbiased estimate of the treatment effect, it must be the case that either –unobserved- time varying regional and individual variables did not change between the pre and post-treatment period or that they changed in an identical manner in the control and treatment group. One reason for these assumptions to be violated is the fact that individuals eligible for the wage subsidy could react to it in anticipation of the policy.[19] Nevertheless, we consider that the strong variability in the eligibility conditions across regions and time and the use of regional and individual time varying covariates provides us with a control group that matches these two requirements. For instance, the anticipation of the policy could affect the results when the analysis cover a short period after the policy is implemented. In this context, one can argue that firms act strategically to get the highest reward from the policy. But this is not the case in this analysis since there are regions where the policy remains unchanged for several years. Therefore we consider that in this case the incidence of the anticipation effect on our estimations should not be relevant.

Moreover, much of the debate around the validity of a DID estimate typically evolves around the possible endogeneity of the interventions themselves (See Besley and Case, 2000).[20] In this paper, as it is common in many panel data studies[21], we will include time varying regional variables as well as regional effects to control for permanent and transitory differences across regions in policies and outcomes[22].

Other issue related to the definition of the comparison group is whether we can assume that the two groups of workers –treated and non treated-, are subject to the same aggregate labor market trends and react in the same way. Evidently, this assumption is more plausible when we split the effect of the wage subsidy by age and gender to the extent that the human capital of each groups is similar and also preferences for work should be the same. For instance, preferences for work between the eligible group in their early twenties and the eligible group in their middle thirties may, however, not be the same as this is the age that many people have children. This might generate differential aggregate trends across groups. So we estimate the average treatment effect by gender and age groups. Moreover, as we will show in the statistical section, there are significant differences in the impact of these wage subsidies across gender and age groups. This makes most likely that the overall characteristics and behavior of the control group match that of the treatment group. Such an approach is similar to the discontinuity design (Hahn, Todd and Van der Klauss, 1999). Nevertheless, the substitution effects are likely to be much more severe the closer are the productivity characteristics of the two groups. In the event of substitution, the impact of the program for the eligible group is biased upwards by the fact that the outcome for

[18] These assumptions are discussed in detail in Blundell and MaCurdy (1999).
[19] In our case this could apply when firms anticipate the policy.
[20] Besley and Case (2000) show that the inclusion or exclusion of variables that determine both policy and behavioural outcomes dramatically alters the estimated impact of the policy when the identification strategy relies exclusively on regional variability. Their findings are a reminder that inadequate controls for time-varying regional level variables may bias estimates of the policy incidence identified from regional-level policy variation. They suggest that one way of dealing with these concerns is the DID approach. That is to try to identify the policy effect by selecting a control group of workers in the same industry or occupation in regions where the policy variable did not change, among regions thought to be similar to that whose policy has changed.
[21] See Anderson and Meyer (1997), Gruber and Madrian (1997).
[22] If the systematic determinants of state policies are additive, time invariant regional characteristics, then will indeed remove concerns about endogeneity.

the control group is decreasing. Nevertheless we avoid this bias by including in the model the variable D_{jt} which was described above.

5 Data Description: Muestra Continua de Vidas Laborales

We use a new dataset recently available in Spain which is named "*Muestra Continua de Vidas Laborales" (MCVL,* hereafter). This is an administrative dataset based on a random draw from the Social Security archives. It contains a sample of 4% among all the affiliated workers, working or not, and pensioners in the year 2004. Since it is an administrative data set it is free of sample selection bias due to informative dropout. Besides, as it covers the whole labor market history of the worker it does not suffer from the typical problem of left censored information of previous labor market experiences to the one analyzed[23].

This dataset has information about 1,1 million people and covers their entire labor history. The amount of information for each individual in our database is quite large and we have imposed some restrictions. Firstly, there may exist one different register for each contract held. In the Spanish labor market it is quite usual to find firms that optimize their labor costs by the mean of firing the workers in short periods of time and hiring them again after that[24]. In these situations we consider that the employment spell is actually continuing although it has a short interruption in the middle. Hence, we unified successive registers when they correspond to the same worker in the same firm and with the same type of contract and when the interruption is lower than 15 days. Secondly, we have eliminated simultaneous employment spells and we kept the information about the longer one. Thirdly, we have unified each two registers when they correspond to one contract that begins before the previous one has finished. Fourthly, we have eliminated each incomplete or incorrect register. This may happen because some important information is missing or because it is clearly incorrect (dates of beginning and finishing incompatible, etc.). Finally, we are only considering labor histories of workers within the called "Regimen General", that is, regular workers being paid by a firm. We are not using information for self-employed and either for workers in Agriculture, Fish and other minor special cases. Given the problem of too much missing information regarding the type of contract for the spells before 1995, we will only study the employment spells (and non-employment spells), beginning after the end of 1994. Nevertheless, we will use the information previous to 1995 to control for the labor market history of the worker in terms of their previous employment and unemployment experiences.

This database offers information about the personal characteristics of the worker and also about all the employment spells throughout her labor history. We have information about age, gender, occupation, unemployment and employment spells and their respective exact durations. This last information is especially relevant given the aim of the paper. Other databases only gather

[23] Preceding labour market histories are known to have a strong impact on the probabilities of accessing into a permanent contract.
[24] In our database, there is a high percentage of workers who have a new contract each week, mainly at the same firm.

annual or quarterly information and therefore individual labor market transitions within a year are not available giving rise to an under-representation of short-term temporary contracts. The duration of each employment spell is built from the dates of beginning and ending the contract and it is measured in months.[25] Moreover, for the periods of non-employment, we can distinguish among the ones when the payroll taxes are being paid, that is, when the worker is receiving Unemployment Benefits and those when the worker contributions to Social Security are not paid, which can be both periods of unemployment without benefits or periods of inactivity. Hence, we use the terminology non-employment to name all these spells of not working within a firm. Other interesting characteristic of this database is that each firm has an authentication code. This last aspect is fundamental since it allows identifying whether the worker remains in the same firm after changing the contract or after the unemployment spell.

The impact of these regional wage subsidies are evaluated by looking at transition probabilities to new permanent contracts either from unemployment or from a temporary contract at the same firm. We define three types of transitions from a temporary contract: exit to unemployment or to a different firm, other temporary contract at the same firm and a conversion from a temporary to a permanent contract at the same firm. The definition of a temporary contract is a relevant issue. Guell and Petrongolo (2007) and Booth, Francesconni and Frank (2002) use a broad definition of a temporary contract[26]. We follow the same approach by considering temporary contracts the following categories: fixed term, specific task, training, and contract for circumstances of production, internship contract and replacement. With respect to the unemployment state we define two types of transitions depending on the type of contract at the new job: exit to a temporary contract and exit to a permanent contract.

Given our sample selection criteria we end up with 472.406 and 585.602 unemployment spells and with 299.841 and 437.262 spells of temporary contracts for women and men, respectively. Sample size and main sample characteristics by gender for unemployed and temporary workers are displayed in Tables 4-0. Firstly, we can observe that the entrance probability into a permanent contract is pretty low in all cases. For instance, in Tables 3 and 4 we display the transition rates to permanent contracts by regions, age and gender. These transition rates ratify the high turnover level that characterises the behaviour of the labour market for the Spanish economy independently of the age, gender and region of the worker. We can observe that the main destination state after unemployment is a temporary contract. Meanwhile, after a temporary contract the most probable option is to exit to unemployment or to other firm. Thus, the probability of accessing into a permanent contract is only around 13% and 11% for unemployed women and men, respectively. The conversion probability from a temporary to a permanent contract is much lower: 1.2% for women and 0.9% for men. Interestingly, the conversion rate to a permanent contract is slightly higher for women. Since regional wage

[25] We are not considering employment spell durations lower than 30 days in order not to study just very short spells due to reallocation or strong turnover within the firm. In this case the majority of the contracts are in fact signed through a temporary help agency. Given the interest of this research, we have opted to omit this type of spells. One can assume that they are really of temporary nature and therefore they should be independent of the policies committed to promote permanent employment.

[26] For instance, in Casquel and Cunyat (2005) some type of contract such as internship contract are not considered.

subsidies tend to be larger for women –as we will show later on-, we can wonder whether these differences by gender can be related to these regional policies. Moreover, there are also some differences by age groups. For instance, young workers have more chances to enter into a permanent contract from unemployment while the conversion rate of temporary contracts is larger for older workers. Thus, the transition rates to permanent contract from unemployment are 14% and 12% for female and male young workers, respectively. The conversion rates from a temporary contract are 1.5% and 1.1% for older workers.

Regional transition rates are displayed in Table 4. They show the existence of an important variability of transition rates across Spanish regions. For instance, we have regions such as Madrid, la Rioja or Catalunya whose transition rates almost double the ones found for Andalusia and Extremadura. When we relate this information with the intensity of regional wage subsidies we can observe that there is not a clear relationship between these two variables.

In Tables 5 and 6 we display the main sample characteristics for unemployed and temporary workers. They indicate that there are differences in the entrance probability into a permanent contract by personal characteristics. Among unemployed workers those low qualified, immigrants, and workers in the construction sector show lower probabilities of entering into a permanent contract. When a worker is in a temporary contract the probability of getting a permanent one in the same firm is also lower for low qualified and part-time workers, immigrants, and workers in the construction and services sector.

In Tables 7 and 0 we show the relation between the type of transition and the duration of the unemployment spell, in the first case and the duration of the temporary contract, in the second one. The information shown in both tables put forward again the existence of a high turnover rate in the Spanish labor market during the analyzed period. With respect to unemployment duration we can observe that more than 70% of the sample consists of unemployment spells shorter than six months. The spells that end up into a temporary contract tend to have shorter durations than those spells ending up into a permanent one. In relation to the duration of the temporary contract we again observe that the sample is mainly composed by short-term temporary contracts. More than 60% of them lengthen less than six months, independently of the following state. We also observe that the probability of having another temporary contract at the same firm is larger for short-term contracts than for the long-term ones. On the contrary, the conversion probability from a temporary to a permanent contract is more common for lengthier ones.

We display in Figures 1 and **¡Error! No se encuentra el origen de la referencia.**2 the entrance probability to a permanent contract from unemployment and from a temporary contract at the same firm, respectively. In Figures 3, 4 and 5 we represent the exit probability from unemployment to a temporary contract, from a temporary contract to unemployment and from a temporary contract to other temporary contract at the same firm, respectively. All figures are represented by gender and refer to the sample period 1994-2004. These figures illustrate our identification strategy. The first remarkably fact is that the entrance probability to a permanent contract from unemployment (Figure 1) and the conversion rate from a temporary to a permanent contract (Figure **¡Error! No se encuentra el origen de la referencia.**2) increase during the period 1998-2001. Moreover, in both cases this tendency seems to be negatively correlated with the entrance probability into a new temporary contract, either from unemployment (Figure 3) or from other temporary contract at the same firm (Figure 4). From Figure 3 we can observe that the exit probability from unemployment to a temporary contract shows a lower growth rate during this same period. In Figures 4 and 5 we show that the probability of accumulating temporary contracts at the same firm and the exit probability from a temporary contract to unemployment or to a different firm also decrease during the period 1998-2001. Evidently, this increasing trend in the entrance probability to a permanent contract should be related with the economic expansion that began in 1995 but it is also relevant to note that in 1997 there was an important labour market reform and national and regional active labour market policies to foster permanent contracts began.

As complementary information of previous figures, we display in Table 9 the main sample entrance probabilities to new permanent contracts for treated and non-treated workers before and after 1997. Given the characteristics of the policy we are analysing, we can not define properly the periods before and after the treatment as it is common in policy evaluation. Nevertheless, we find interesting to highlight the differences in entrance probability to new permanent contracts by labour state before and after 1997 between treated and non-treated workers. We can observe that treated workers show an increase in the entrance probability to new permanent contracts except for old women and young men unemployed workers and male workers with a temporary contract. Meanwhile, all non-treated workers show a decrease in this probability. In the last column of Table 9 we display the empirical counterpart of the difference-in-differences estimator. From these numbers we can say that the entrance probability to a permanent contract of treated workers relative to the one of non-treated workers increased after 1997. The percentage change in the transition probability to a permanent contract is clearly larger when workers are in a temporary contract. It is also notably larger for women, independently of the initial labour state.

When we look at the conversion rate from temporary to permanent contract we observe that young and older women experience the largest increase. When we look at unemployed workers, young and middle age women seem to be more benefited, especially the young ones. We consider that the results displayed in this table are highly interesting since they could be capturing the effect of the regional wage subsidies. The labour market reform of 1997 established general discounts in payroll taxes for any new permanent contracts and introduced a new permanent contract for certain group of workers. As it has been stated by Kugler *et al* (2003) this reform mainly affected to male young workers[27]. Therefore, the differences found in the diff-and-diff statistic shown in Table 9 can not be exclusively related to national labour

[27] They state that the effects found for young women are not always statistically significant.

market reforms and they might be related to the regional wage subsidies. Specially, the differences found between female and male young workers.

6 Results: The incidence of wage subsidies over the entrance probability to a permanent contract

In this section we deal with the incidence of regional wage subsidies over the transition rate to a permanent contract from unemployment and the conversion rate from a temporary to a permanent contract at the same firm. Moreover, since theses regional policies might give rise to substitution effects between different types of workers we also study whether these regional policies affect the exit rate from permanent employment.

Our policy variable measures the maximum amount of the wage subsidy for each eligible individual[28]. Furthermore, we introduce the exact amount of the subsidy and its square term in order to capture any nonlinear pattern in the treatment effect. With this specification we can check whether the wage subsidy fosters permanent contracts and whether the amount of the wage subsidy is relevant to explain its effect. We have shown that wage subsidies are larger for women and young workers and therefore, we could expect that the impact onto the entrance probability to a permanent contract will be larger for them. Moreover, we argue that with this specification we can take advantage of the additional variation in the quantities, even within eligible individual groups, along time and across regions to estimate the response of eligible individuals to the change in incentives.

The sample used in the estimation includes workers aged 18-64 years. In all cases, we also split the whole sample by gender in order to gain homogeneity between treated and non-treated individuals. Besides, we allow for heterogeneous treatment effects by age and we model the policy variable as an interaction term between the maximum of the wage subsidy and the age group the worker belongs. We build three age groups: younger than 30 years old, aged between 30 and 45, older than 45 years old. This distinction is also relevant since wage subsidies tend to vary by age. The duration dependence of the exit probability is specified as a polynomial in the log of the corresponding duration. Our specification also includes many personal and labor characteristics to control for differences between treated and non treated individuals that could affect the outcome under study. Hence, *qualification category*, *firm size*, *firm age*, *full time job, sector of activity, private ownership, temporary help agency and whether the new contract is in the same firm as in the following one*. We also consider variables that control for national and region-specific time trends. In this way we provide reassurance that our coefficients are not reflecting smoothly trending omitted variables that are potentially correlated with the adoption

[28] Initially, we have estimated other version of equation (1) where the variable D_{ijt} was a binary variable being equal to one for those workers potentially treated in the corresponding year and region. We got that the policy variable was statistically significant but that the incidence was pretty low. Nevertheless, this model can offer incomplete information. This could arise for two main reasons. Firstly, one can argue that the labor demand by type of contract is highly inelastic to wage subsidies. This situation can arise when jobs covered by fixed-term contracts are in fact of temporary nature. Secondly, it could be that the amount of the wage subsidy does not outweigh the gap of unit labour costs between permanent and temporary workers.

of regional wage subsidies. The variables considered are the regional unemployment rate and the national production growth rate. We also include a set of dummy variables that reflects the age group of the worker described above and we interact them with the national production growth rate and with the regional unemployment rate to control for age-specific different time varying effects. In this way we introduce a control for asymmetric cyclical effects by age groups[29].

Since we are estimating a competing risk model for each duration model, we obtain a specific vector of parameters for each alternative. Moreover, each duration model is estimated by gender and with and without control for unobserved heterogeneity. We also estimate different specifications of each model in order to check for the robustness of the results. For ease of exposition, we have opted not to present all the results but just the most relevant ones. Since we did not find relevant differences on the policy parameter between the model with and without unobserved heterogeneity we show the results of the first case. [30]

6.1 The Conversion Rate from a Temporary to a Permanent Contract

For the sample of workers with a temporary contract, wage subsidies only exist for the conversion of this contract into a permanent one at the same firm. This restriction has three main implications in the analysis. Firstly, the definition of the competing alternatives in our duration model differs slightly from the one of unemployed workers. In this context, the competing alternatives are, first, a new temporary contract with the same employer, second, a permanent contract with the same employer and, third, unemployment or other contract with a different employer. We have matched the alternative of unemployment with that of a direct transition to a different firm since we could assume that workers who change firm stay at least one day unemployed. For ease of exposition, from now we will name this last alternative as exit to unemployment. Secondly, workers directly hired by a Temporary Help Agency can not be benefited from this policy unless they become unemployed. Therefore, to get a closer definition of a potentially treated worker we have estimated the model without considering those temporary workers hired through a Temporary Help Agency[31]. Thirdly, we drop those employment spells shorter than a month. The idea is that very short-term temporary contracts will probably be jobs of temporary nature and therefore they will be highly inelastic to any subsidy to convert it into a permanent one. This restriction makes sense in this analysis given

[29] A better way to control for age-specific cyclical effects in our evaluation exercise is to compare the transition probabilities of treated and control workers during the analysed period with these transition probabilities during an earlier period with similar cyclical behaviour. Nevertheless, we do not apply this approach since in our database we can not identify properly the type of the contract before the year 1995. Moreover, given that our analysis is based on regions affected by common macroeconomic shocks, we can assume that age-specific cyclical effects can be controlled by taking into account regional variability.

[30] Anyway those results not presented will be provided upon request.

[31] This restriction does not mean that these workers are kept out of the analysis. They will be included in the sample of unemployed workers.

that the final aim of these regional wage subsidies is to reduce the use of temporary contracts for permanent tasks[32].

The duration dependence of the exit probability from a temporary contract is specified as a polynomial in the log of the duration of the temporary contract. Besides we build some dummy variables that control for specific contract durations: less than 4, 6, 12, 24, and 36 months[33].

We present the general results in Table 10. The first idea to highlight is that on average, the policy variable and the rest of "*D*" covariates are statistically significant and differ between alternatives. This result supports the interest to specify a competing risk model to correctly identify the effect of the wage subsidy over the conversion rate to a permanent contract relative to each of the competing alternatives. For instance, given the competing alternatives defined, we can check whether wage subsidies also affect the transition between temporary contracts at the same firm. It could be possible that the effect over the conversion rate to a permanent contract is low since the firms might find profitable to exhaust legal limits before using the subsidy to convert this contract into a permanent one. Besides, since the rest of "D" variables are statistically significant, our results put forward the fact that there are permanent and time varying differences among eligible workers and regions that must be controlled for in order to get an unbiased estimate of the causal effect of the policy. The variable D_{ij} is not separately identified from D_j in the case of women since regional wage subsidies tend to be generally addressed to this group. The rest of variables that control for permanent and time varying differential effects by age groups are also statistically significant. For instance, we can observe it is convenient to control for differences in the effects of the economic cycle by age group.

We obtain that there is a non linear relationship between the exit rate from a temporary contract and the wage subsidy but this relation differs among alternatives and by worker characteristics. For ease of exposition we show in Table 11 a summary of the main results. This summary includes the odd ratio, the total marginal effect[34] of wage subsidies and the total change in the exit probability when the wage subsidy moves up from zero to 5100 Euros (which is the average wage subsidy). The main problem of computing marginal effects in *multinomial logit* models is that they can change sign depending on where the independent variables are being evaluated. Thus, we also display the odd ratios because they offer an easy summary of the main effects of wage subsidies over the entrance probability to a permanent contract for treated workers relative to the non-treated ones.

The estimated marginal effects of wage subsidies show that the exit rate from a temporary contract to unemployment tends to decrease for treated workers. The effect is statistically significant for all workers. The fall is slightly larger for women and it ranges from –0.32% for old men to –1.51% for young women when measured at the average subsidy. The total change

[32] There are a few number of empirical studies that show that many Spanish firms use temporary contracts to fill permanent jobs (Guell, 2000; Rebollo 2006).

[33] The empirical hazard shows a spike at these specifics durations.

[34] The total marginal effect implies that we are taking into account any non-linear relationship between wage subsidies and the corresponding exit probability.

in this exit probability is negative for treated workers except for old workers and middle age men. Female and male old treated workers might experience a small increase in their exit probability to unemployment of around 1.21% and 1.17% respectively. This increase is lower for middle age men workers[35], 0.48%. On the contrary, the exit probability to unemployment experiences a larger drop for young workers (8.07% and 3.11% for female and male workers, respectively) and middle age women (6.34%).

The marginal effect of wage subsidies over the exit probability to other temporary contract at the same firm is larger –in absolute terms- that the marginal effect on the exit rate to unemployment. Moreover, it is statistically significant in all cases. As before, the marginal effect is slightly larger for women. It varies from 1.22% for young men to 5.10% for old women. Consequently, the total change in the transition probability to new temporary contracts at the same firm is always positive and much larger than the one found for the exit rate to unemployment. It is also interesting to note that this growth in the transition probability of accumulating temporary contracts at the same firm tends to increase with age. For instance, the estimated exit probability to a temporary contract as the wage subsidy growths to 5100 Euros, increases by 16.08%, 16.18% and 23.53% for young, middle age and old women, respectively. In the case of men, the total change in this exit probability is lower except for middle age men. These changes are 2.42%, 17.42% and 14.79% for young, middle age and old men, respectively.

The marginal effects of wage subsidies over the transition rate to a permanent contract show important differences among workers. They are only statistically significant at 90% of confidence for women and middle age male workers. Nevertheless, when it is statistically significant, the magnitude of the effect is larger than the one found for the other exits. The marginal change in the conversion probability to a permanent contract is the largest for young women and middle age male workers (6.68% and 4.45%, respectively). Consequently, the total change in the exit probability to a permanent contract is larger for these two groups of workers, 26.26% for young women and 22.40% for middle age men. On the contrary, the marginal effect is low for middle age women, 1.99% and it is even negative for old women, -0.69%. Nevertheless, the total variation in the conversion probability to a permanent contract is not low for them, around 23.59% and 11.48% for middle age and old women, respectively. This is due to the fact that the marginal effect of wage subsidies for these workers is larger for lower levels of the subsidy[36]. Nevertheless, since the transition rates to permanent contracts at the same firm are pretty low these relative changes hardly generate a change in the transition probability from a temporary to a permanent contract at the same firm. For instance, in the case of young women the estimated transition probability growths from 0.064% to 0.075% while for middle age men it changes from 0.039% to 0.041%.

It is also interesting to note that the marginal change for women keeps increasing with the level of the wage subsidy. For instance, the marginal effect is 7.68% for a wage subsidy of 10.000

[35] Though the marginal effect at the average wage subsidy is negative for middle age workers, the total change in probability is positive since the marginal effect is negative for lower amounts of the wage subsidy.
[36] For instance, the marginal effect of the wage subsidy measured at 1 euro is 4% for middle age women.

euros for young women and around 4% for middle age and old women. On the contrary, the marginal effect states at around 4.4% for middle men at 5.100 euros and around 4% at 10.000 euros. Therefore, there seem to be room for fostering the transition rates to permanent contract for women by increasing the wage subsidy.

Finally, the effects of wage subsidies over the mean duration at the temporary contract are also low though statistically significant in all cases. They are also larger for women. The estimated mean duration of the temporary contract increases from 13 to 16 months for young and middle age women and from 13 to 15 for old women. The estimated mean duration of the temporary contract for men states around 13 months.

The information provided by the odd ratios leads us to the same ideas. Their interest rest on that they offer a summary of the net effects of wage subsidies over the entrance probability to a permanent contract for treated workers. The probability of getting a permanent contract relative to exiting to unemployment for treated workers increases in all cases except for old male workers. These odd ratios are larger for young and middle age women, 27.4% and 22.4% respectively, and middle age men, 21.8%. On the contrary, the odd ratio of getting a permanent contract relative to exiting to a new temporary contract at the same firm is positive but small for young women and middle age men or negative for the rest of workers. Therefore, wage subsidies seem to reduce the incidence of unemployment for treated workers while the incidence of new temporary contracts at the same firm is not clearly lower and in fact it might increase for certain type of workers as it seems to be the case for old male workers. The odd ratios relative to staying at the same temporary contract are also positive in all cases except for old middle age men. Again the magnitude of the odd ratios is larger for young and middle age women and for middle age men. Therefore, the analysis of the odd ratio also points that wage subsidies foster the permanence of workers at the same firm partially by accumulating temporary contracts while it hardly increases the duration of the temporary contract itself.

In order to value the importance of the effect of wage subsidies over the entrance probability to a permanent contract we also must consider which is the discount received by firms in terms of their labor costs. A wage subsidy of 5100 euros implies a drop of the monthly average wage[37] of 25.3% if we consider that the worker stays one year with the permanent contract and around 12.6% if we consider that the workers remains two years with a permanent contract at the firm[38]. Evidently, as the duration of the contract increases the discount these subsidies imply over monthly labor costs decreases. At the same time, total labor costs increases with tenure –dismissal costs mainly-. Hence, the computation of an elasticity of labor demand for permanent contracts relative to the wage subsidy is not straightforward. If we take the one year discount obtained by the firm we would conclude that the demand for new permanent contract is not sensitive to wage subsidies except for young women and middle age workers. Only for this

[37] The monthly average labour cost is 1.681 Euros and is taken from the Labour Costs Survey published by INE for the year 2000.
[38] Since wage subsidies are designed to foster stability on employment we consider that the desired permanence on employment should be at least two years. Besides, in many cases, wage subsidies are conditioned to keep the worker for a period of non shorter than two years.

group of workers, the discount guaranteed by the wage subsidy makes firms to increase their demand of permanent contracts.

Since this low elasticity of the demand to new permanent contracts could be related with the temporary nature of the job we have also estimated the model restricting the sample to workers with temporary contracts of more than three months. Tough, there are not relevant differences; we find interesting to point out that the marginal effects are slightly smaller. Consequently, we can not conclude that the low elasticity of labor demand for new temporary contracts is due to the temporary nature of the jobs covered by temporary contracts. Finally, we also estimated the model for workers hired only by private firms. Previous empirical research has shown that the increasing trend in the rate of temporary contracts is due to the increasing use of these contracts from the public sector, mainly local agents (Dolado, García-Serrano, and Jimeno, 2002). Again the results obtained for the policy variables are pretty similar and therefore we do not find evidence of differences between treated workers in private or public firms.

6.2 The Transition Rate from Unemployment to a Permanent Contract

For the case of unemployed workers we have also imposed one sample restriction and we focus on workers with unemployment spells no lengthier than one year. This sample restriction is made in order to gain homogeneity for treated workers. The policy we are evaluating is not designed to foster employment for disadvantaged unemployed workers as the long-term unemployed ones. Besides, since 1994 there are many national and regional ALMPs addressed basically to foster the employment of long-term unemployed workers and we want to analyze the effect of wage subsidies without any other ALMP influence, we exclude long-term unemployed.

As before, the duration dependence of the exit probability is specified as a polynomial in the log of the spell duration. The control variables differ slightly. Specifically, we omit some variables related to previous job characteristics and we add two variables that identify whether the worker is eligible for unemployment benefits and whether she receives them when she exits from unemployment. We also use as a regressor whether the worker was previously hired by a temporary help agency and whether he returns to the same firm.

In Table 12 we present the general results of estimating the competing risk model for unemployed by gender when the policy variable is the maximum amount of the wage subsidy for each eligible worker. From this table we can observe that the policy variable and the rest of "D" variables are statistically significant in most of the cases and differ between alternatives. This leads us to conclude that it is important to control for permanent differences between eligible workers and treated regions to get an unbiased estimate of the policy. Besides, we also can observe that aggregate and regional cyclical effects impact differently into the exit rate from unemployment by age group.

In Table 13 we have the predicted treatment effects by age and gender in terms of the odd ratios of exiting to a permanent contract relative to the other two competing alternatives. With this analysis we can get a clearer idea of the net effects of wage subsidies over the entrance probability to a permanent contract from unemployment. Besides, we also display the marginal effect of the policy computed at the average wage subsidy and the total change in the probability. The first result to highlight is that the effect of wage subsidies over the entrance probability to a permanent contract for unemployed workers is notable lower than the one found for workers with a temporary contract.

The marginal effects of wage subsidies over the exit rate from unemployment to a temporary contract are negative for young workers and positive for the rest. But in any case, the values are pretty small and only for old men it reaches the magnitude of 1%. Thus, when we compute the total change in the exit probability from unemployment to a temporary contract as a result of the policy it gives small changes as they range from 0.16% for young women to 6.27% for old women[39]. It is interesting to note that old women it seems that wage subsidies could benefit them indirectly. That is, wage subsidies seem to increase their exit rate from unemployment to a temporary contract and, as we have shown in previous section, they favor their accumulation of temporary contracts at the same firm. The question that remains is whether they finally enter into a permanent contract or not.

To measure the direct effect of wage subsidies for unemployed workers we must look at the effects of wage subsidies over the exit rate from unemployment to a permanent contract. We can observe in Table 13 that the marginal effect measured at the average wage subsidy is statistically significant for all workers except for old men. They all are positive and larger for women, but in all cases they are quite small. For women, they range from 2.26% -old workers-, to 2.55% -young ones-. Meanwhile, this marginal effect for men varies from 0.22% –old workers-, to 2.38% -young workers-. In terms of the total change in the exit probability due to wage subsidies, the biggest impact is among young workers. Female and male young workers experience an increase in the entrance probability to a permanent contract of 14.97% and 14.40%. For middle age workers this change is 8.36% for men and 6.46% for women.

Finally, the marginal effects of wage subsidies over the probability of staying unemployed are statistically significant but irrelevant since they all are smaller than 0.5%. Equivalently, the total change in the probability of staying unemployed is pretty small. Therefore, we do not find evidence that wage subsidies accelerate the exit from unemployment for treated workers. The estimated mean duration of unemployment states around 7 months for men and 6 months for women.

The information provided by the odd ratios leads us to the same ideas pointed out so far. Firstly, the impact of wage subsidies is lower for unemployed workers than for temporary ones since the odd ratios are smaller than the ones found for temporary workers. Secondly, the larger incidence of wage subsidies is for young unemployed workers and middle age unemployed males.

[39] The fact that the marginal effects are negative for young workers though the total change in the exit probability to a temporary contract is positive is due to the non-linear relationship between the exit rate from unemployment and the wage subsidy.

To check the robustness of the results we have made two additional estimations. Firstly, we have included all unemployed workers. Secondly, we have restricted the analysis to workers with temporary contracts in the previous job. In both cases the results do not differ strongly. In the first case, as expected, the odd ratios are even smaller. In the second case, they are slightly larger.

6.3 Testing for Substitution Effects: The Exist Rate from Permanent Employment

We are also concerned with the possible existence of substitution effects. Substitution effects occur if participants take some of the jobs that non-participants would have got in absence of the treatment. The wage subsidy can affect the individual probability of having a permanent contract in two ways. First, the firm that hires an eligible individual receives a wage subsidy when signing a new permanent contract that may enhance the worker's probability of having a permanent contract. Second, some of the individuals who are not eligible might faced a drop in this probability. In this situation we would say that in average terms the policy has no effects when in reality it has. The extent to which this may happen will depend on some different factors. If the wage subsidy just covers the deficit in productivity of unemployed or temporary workers we would not expect any substitution effect. The eligible workers are no cheaper than anyone else. Second, it will depend on the extent that these workers are substitutable in production for the existing ones and on the extent that it is easy to churn workers, that is, to replace a worker finishing a permanent contract with a new subsidized worker. This latter point is important when the wage subsidy does not require keeping the worker for several years. Of course, if new permanent contracts are generally short, firms will be able to use subsidized workers instead of non-subsidized ones, without extra effort. For instance, Cebrian and Toharia (2005) offer some evidence that this might have happened.

To estimate the exit rate from permanent employment we follow the same approach applied in previous section. We estimate a competing risk model for workers with a permanent contract and considering they face the following alternatives: staying at the same contract, exiting to unemployment, to a temporary contract and to other permanent contract. In terms of our duration model, if wage subsidies generate important substitution effects we should observe an increase of the exit rate to unemployment or to a temporary contract, for non-treated workers. In Table 14 we display the results for these transitions. We can observe that the policy variable is statistically significant for young and middle-age workers while for old workers is only marginally significant for one of the alternatives considered. This last result is expected as we previously have shown that the policy has not significant effects for this group of workers, either when they are in a temporary contract or unemployed.

In Table 15 we display a summary of our main results by computing the odd ratio of staying at the permanent contract relative to the rest of alternatives and the marginal and total effect of the wage subsidies. In this model, and given the way we have defined the odd ratios, the information provided by the total marginal effect is equivalent to the one provided by the odd ratios.

The first idea to highlight is that we find some evidence of substitution effect. The marginal effects of wage subsidies over the exit probability to unemployment are statistically significant in all cases, nevertheless, they are relevant only in the case of women since the marginal effects for men are just close to 1%. The marginal effects measured at the average wage subsidy decrease with age and they are 3.26%, -5.36% and –14.24%, for young, middle age and old workers, respectively. However, the total change in the probability is negative, in any case, and takes the following values: –20.45%, -7.71% and –2.58 for young, middle age and old workers, respectively. The result found for young workers is due to the fact that the incidence of wage subsidies over the exit rate to unemployment is negative for lower levels of the wage subsidy[40]. Hence, we obtain that if wage subsidies foster substitution effects they concentrate on young female workers. Given that young female workers are also the most benefit ones, this result imply that wage subsidies favor the substitution of non-eligible young women with permanent contracts for eligible ones.

The marginal effect of wage subsidies over the exit probability to a temporary contract is significantly lower than the one found for the unemployment exit rate. As before, there are interesting differences by gender. Though the marginal effects are all positive, the total change in the probability is negative for men while positive for middle age and old women. The total change in probability for young women is negative but very small. Therefore, it seems that treated women face an increase in their probability to move from a permanent to a temporary contract while the opposite is found for treated men since they experience a drop in this probability. However, the magnitude of the effects is not large in any case. They are 9.63% and 14.97% for middle age and old women, respectively and –8.15%, -19.67% and –14.43% for young, middle age and old male workers.

The most interesting results rise when we look at the exit probability to other permanent contract. The marginal effects and the total change in the probability are especially relevant for women. The marginal effects are large and positive for this group and consequently, the total change in the exit probability to a permanent contract is also large. These relative changes are: 57.57%, 95.19% and 67.53% for young, middle age and old women. For men they are notably lower for middle age and old men (3.48% and 3.33%, respectively), and even negative for young workers. Hence, these results point out that wage subsidies favor the transition between permanent contracts for treated women.

Finally the effects of wage subsidies over the mean duration at the permanent contract are all statistically significant and negative. As the consequence of wage subsidies, the estimated average duration at the permanent contract increases for young workers from 96 to 113 months for women and from 116 to 120 for men. For middle age workers it also increases but in a lower extend since it varies from 96 to 97 months for women and 116 to 119 for men. Finally, the

[40] For instamce, the marginal effect is –7.80% when the wage subsidy is 1 Euro.

estimated average duration decreases for old workers from 96 to 94 for women and from 116 to 112 for men.

7 Conclusions

Spain is one of the countries with the highest rate of temporary contracts in Europe and this might bring negative costs for the economy both in terms of efficiency and equity conditions. Thus, national and regional governments have designed different policies to foster the creation of permanent employment. Since the 1997 labor reform, the national government offered discounts in payroll taxes for new permanent contracts. Simultaneously, since 1997, different regional governments have begun to encourage permanent employment by offering wage subsidies to new permanent contracts for certain type of workers, in some cases, and for all workers in others.

In this paper we use the information of these regional policies to measure the impact of wage subsidies on the creation of permanent employment. We take advantage from all the variability derived from these regional policies, that is, regional, time and individual eligibility criteria, to identify the treatment effect. We apply a Difference-in-Differences-in-Differences (DDD) approach to estimate the effects of regional wage subsidies over the probability of getting a permanent contract. With this approach we control for observed and unobserved heterogeneity between control and treatment groups. To check the robustness of the results and in order to gain homogeneity between treated and non-treated individuals, we also estimate the model splitting the whole sample by gender.

We estimate the incidence of regional wage subsidies over the transition rate to a permanent employment from a temporary contract and from unemployment using as a policy variable the maximum amount of the wage subsidy for an eligible individual. Besides, in order to check for substitutions effects we estimate the exit probability from a permanent contract. In all cases we follow a competing risk approach.

From our main results, we can conclude that the policy of subsidizing new permanent contracts shows a positive effect over the transition rate to a permanent job only for certain groups of treated workers. Though, there are several differences by the initial labor state, by gender and age we can conclude that, in general, the effect is small. Old workers are clearly not affected by wage subsidies. On the contrary, young workers, especially women, are the most benefited ones followed by middle-age workers. Young male workers seem to be benefited only when they are unemployed. Nevertheless, the quantitative effects of these wage subsidies are pretty low and it seems that the elasticity of the demand for new permanent contracts relative to wage subsidies tends to be lower than one, even for young female workers. For instance, the average wage subsidy implies a drop of around 25% in total labor costs for a one year contract. The conversion rate from temporary to permanent contracts for treated workers increases between 22-26% for young females and middle age workers. The transition rate from unemployment to a permanent contract increases only by 14% for young workers and between 6%-8% for middle age workers. Given that the empirical transition probability to a permanent contract is pretty low, this small rate of growth leads these transition probabilities practically constant.

Wage subsidies might increase the total duration of the employment spell at the same firm for treated workers and mainly for women through the accumulation of new temporary contracts. This statement comes from the fact that treated workers with a temporary contract get a drop in the exit probability to unemployment and more importantly, an increase in the exit probability to a temporary contract in the same firm. Specially relevant seem to be this result for the case of old women since we obtain that the transition probability to a temporary contract increases by 23,5%. Given these results, we could argue that firms might use wage subsidies to convert temporary contracts into permanent ones after the worker has spent a certain amount of time at the firm through several temporary contracts[41].

The results match the theoretical predictions since wage subsidies only imply a temporary drop in hiring costs and it maintains the large existing differences in total labor costs between temporary and permanent contracts. The results found are also similar to the ones presented in Kugler *et al* (2003) who analyses the effect of discounts on payroll taxes implemented in the 1997 reform on to the transition rate to new permanent contracts. They find that the elasticity of labor demand relative to non-wage labor costs is lower than one and larger for young workers.

Finally we have estimated whether wage subsidies generate any substitution effect between eligible and non-eligible workers. The results obtained do not support the idea that this policy increases the exit rate from permanent contract in the case of non-treated men and women, except for young women. In this last case, we obtain that wage subsidies increases the transition to unemployment of non-treated female workers. Interestingly, we obtain that wage subsidies increase the transition rate between permanent contracts for women, especially for middle age ones. This last result evidences the fact already put forward in Mortensen and Pissarides (2001) model commented in section 2. The idea is that a hiring subsidy stimulates job creation, however once the job is created, the opportunity cost of maintaining the match rises since the hiring subsidy can again be obtained when creating a new job. Recall that the labor market reform of 1997 created a new permanent contract with lower dismissal costs and this contract was addressed to certain group of workers included women. Besides, this group was also elected to receive larger discounts in payroll taxes for new permanent contracts.

[41] Rebollo (2006) shows that for certain type of workers temporary contracts might be a stepping-stone to a permanent one but once the worker has accumulated several temporary contracts. Guell (2000) shows that Spanish firms tend to exhaust legal limits before converting a temporary contract into a permanent one.

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List of Tables

Table 1: Regional Wage Incentives to the creation of Permanent Contracts: Eligibility conditions by age and gender across Spanish Regions (1997-2004)

	MALES		FEMALES	
Andalusia	all ages	1997-2002	all ages	1997-2002
Aragon	40 or more	1998-2004	all ages	1998-2004
Asturias	all ages	1997-1998, 2000-2003	all ages	1997-1998, 2000-2003
Balearic Islands	NO		all ages	2000-2004
Canary Islands	18-25	1.998	all ages	1.998
	all ages	1.999	all ages	1.999
Cantabria	all ages	1998, 2000-2004	all ages	1998, 2000-2004
C. Leon	all ages	1998-2004	all ages	1998-2004
C. Mancha	16-30	1.998	all ages	1.998
	16-29 & 45 or more	1999-2003	all ages	1999-2003
Catalunya	NO		NO	
Valencia	all ages	1998-2001, 2003-2004	all ages	1998-2004
Extremadura	all ages	1997-2004	all ages	1997-2004
Galicia	18-30 & 45 or more	1998	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	1998-2004

Table 2: Regional Incentives to the creation of Permanent Contracts

		Minimum	Mean	Maximum	Mean/Total Costs*
Regions	Andalusia	2.400	3.844	6.012	21.0%
	Aragon	1.200	3.684	5.160	18.9%
	Asturias	3.600	4.100	4.500	20.3%
	Balearic Islands	1.653	1.726	1.800	9.6%
	Canary Islands	3.000	3.300	3.600	19.0%
	Cantabria	1.803	3.604	4.808	19.1%
	C. Leon	1.800	3.605	5.115	21.0%
	C. Mancha	3.000	3.300	3.600	17.4%
	Catalunya	0	0	0	0.0%
	Valencia	1.800	4.513	7.466	24.8%
	Extremadura	4.166	10.076	15.177	62.3%
	Galicia	3.000	3.600	4.200	20.6%
	Madrid	3.600	7.971	12.000	32.3%
	Murcia	3.000	4.838	7.200	29.4%
	Navarra	0	0	0	0.0%
	Basque Country	3.273	4.459	7.512	19.2%
	Rioja	4.491	5.001	6.011	27.3%
Gender	Males	1.200	4.764	15.177	-
	Females	1.200	5.147	15.177	-
Age	Aged 18-30	1.200	5.159	15.177	-
	Aged 31-44	1.200	4.078	15.177	-
	Aged 45-64	1.653	4.766	15.177	-

To compute the last column we have taken total labor costs per year for each region from the survey of quarterly labor costs, INE 2000.

Table 3: Exit probability by Age and Gender (Uncensored Observations)

	Unemployed			
	U. - P.C.		**U. - T.C.**	
	Women	Men	Women	Men
Age <30	14.14%	12.39%	85.86%	87.61%
Age 30-45	12.00%	9.66%	88.00%	90.34%
Age >45	12.24%	9.30%	87.76%	90.70%
Average	13.43%	11.33%	86.57%	88.67%
	Temporary Contract			
	T.C -P.C. Same Firm		**T.C -T.C. Same Firm**	
	Women	Men	Women	Men
Age <30	1.13%	0.85%	6.21%	7.06%
Age 30-45	1.22%	0.82%	8.97%	10.13%
Age >45	1.53%	1.14%	9.36%	10.51%
Average	1.19%	0.88%	7.32%	8.45%

PC=Permanent Contract; TC=Temporary Contract

Table 4: Exit Probability by Region and Gender (Uncensored Observations)

	Unemployed		**Temporary Contract**			
	U-P.C.		**TC- T.C. Same Firm**		**TC-P.C. Same Firm**	
	Women	Men	Women	Men	Women	Men
Andalucia	9.87%	6.90%	8.68%	12.00%	1.20%	0.65%
Aragon	12.75%	12.75%	8.33%	7.83%	1.21%	1.07%
Asturias	11.18%	9.43%	7.49%	8.57%	1.17%	0.76%
Baleares	15.45%	11.89%	2.67%	5.18%	0.73%	0.62%
Canarias	14.17%	11.16%	5.35%	5.39%	1.13%	0.86%
Cantabria	10.91%	9.85%	8.34%	9.19%	1.31%	1.16%
Castilla-Mancha	10.72%	9.16%	8.52%	11.51%	1.07%	0.83%
Castilla-León	12.02%	11.06%	7.47%	8.01%	1.06%	0.76%
Cataluña	16.60%	15.13%	5.63%	5.78%	1.29%	1.07%
C. Valenciana	12.92%	11.92%	7.44%	7.71%	1.31%	1.12%
Extremadura	9.61%	8.13%	6.32%	10.41%	0.95%	0.67%
Galicia	10.51%	10.76%	11.72%	10.51%	1.38%	0.81%
Madrid	17.18%	15.63%	5.67%	7.41%	1.05%	0.78%
Murcia	12.12%	11.23%	6.82%	6.7	1.26%	1.19%
Navarra	13.63%	12.50%	9.62%	9.93%	1.58%	1.27%
País Vasco	11.11%	11.30%	11.22%	11.08%	1.30%	1.24%
Rioja	13.15%	13.26%	10.19%	9.95%	1.60%	1.81%

PC=Permanent Contract ; TC=Temporary Contract ; U=Unemployment

Table 5: Main Sample Characteristics for Unemployed Workers by Gender

	Women			Men		
	Censored	**T.C**	**P.C**	**Censored**	**T.C**	**P.C**
U. Duration	13.4	5.3	5.8	11.9	4.8	5.4
Age	31	29	29	32	30	29
Temporary Employment Agency	6.1%	8.3%	7.6%	6.4%	10.0%	8.4%
Equal Employer	-	37.7%	23.9	-	26.9%	17.5%
High Qualification	6.3%	7.5%	6.1%	4.9%	2.8%	5.4%
Medium-High Qualification	13.1%	14.1%	14.9%	9.5%	7.0%	11.7%
Medium Qualification	31.3%	33.4%	37.0%	34.1%	32.7%	31.4%
Medium-Low Qualification	49.1%	44.7%	41.9%	52.1%	57.3%	51.0%
Inmigrant	3.5%	2.5%	3.2%	5.4%	5.7%	5.2%
Part-time	40.4%	37.2%	48.7%	18.9%	15.8%	19.5%
Industry	9.8%	9.8%	10.0%	12.1%	12.7%	17.6%
Construction	2.9%	1.7%	1.7%	29.6%	34.6%	17.7%
Services	86.4%	88.7%	87.5%	56.8%	52.6%	64.6%
Unemployment Benefits (t=T)	20.0%	18.1%	22.1%	24.0%	23.9%	24.5%
Unemployment Benefits	35.4%	27.0%	32.8%	36.5%	31.7%	33.9%
Simple Size	61,961	350,018	60,697	63,826	458,307	63,472

PC=Permanent Contract; TC=Temporary Contract

Table 6: Main Sample Characteristics for Workers with a Temporary Contract by Gender

	Women				Men			
	Censored	**U or Diff. Firm**	**T.C same firm**	**P.C same firm**	**Censored**	**U or Diff. Firm**	**T.C same firm**	**P.C same firm**
T.C Duration	11.4	5.9	5.2	7.5	11.7	6.2	6.3	8.7
Age	33	30	32	31	34	31	33	32
Temporary Employment Agency	1.6%	3.3%	6.7%	0.0%	1.3%	3.1%	5.5%	0.0%
Equal Employer	4.1%	32.0%	-	-	2.7%	21.3%	-	-
High Qualification	18.7%	8.2%	10.3%	6.0%	8.9%	2.9%	2.9%	3.6%
Medium-High Qualification	14.5%	13.9%	12.6%	15.0%	8.9%	7.0%	4.6%	9.1%
Medium Qualification	28.4%	31.1%	25.8%	37.3%	40.7%	34.2%	40.3%	36.6%
Medium-Low Qualification	37.4%	46.6%	51.7%	41.3%	41.8%	55.7%	52.1%	50.1%
Inmigrant	5.2%	3.0%	1.6%	3.4%	9.0%	6.4%	4.4%	4.5%
Part-time	35.6%	37.5%	28.2%	35.0%	11.0%	15.1%	13.1%	14.7%
Industry	8.9%	9.4%	11.7%	19.4%	12.9%	12.2%	12.0%	32.3%
Construction	3.6%	1.7%	0.7%	1.6%	40.1%	34.5%	35.5%	18.3%
Services	87.0%	88.0%	87.4%	78.6%	46.0%	51.3%	51.5%	48.6%
N	57,570	221,985	17,352	2,934	80,363	324,253	29,449	3,197

U=Unemployment; PC=Permanent Contract; TC=Temporary Contract

Table 7: Temporary Contract Duration by Competing Alternatives

	Women		Men	
	T.C	P.C	T.C	P.C
< 6 Months	75.4%	70.0%	78.3%	71.6%
6-12 Months	14.9%	16.4%	14.1%	17.1%
12-18 Months	5.2%	7.0%	4.4%	6.2%
18-24 Months	2.8%	4.1%	2.1%	3.3%
24-30 Months	1.8%	2.5%	1.2%	1.8%

PC=Permanent Contract; TC=Temporary Contract

Table 0: Temporary Contract Duration by Competing Alternatives

	Women			Men		
	U. or Diff. Firm	T.C Same Firm	P.C. Same Firm	U. or Diff. Firm	T.C Same Firm	P.C. Same Firm
< 6 Months	72.3%	79.1%	64.9%	71.6%	81.5%	75.5%
6-12 Months	21.7%	16.9%	25.0%	19.8%	12.9%	15.0%
12-18 Months	3.1%	2.0%	5.4%	4.5%	2.1%	4.4%
18-24 Months	1.5%	0.9%	2.4%	2.0%	1.4%	2.2%
24-30 Months	0.5%	0.4%	1.2%	0.8%	0.7%	0.8%

U=Unemployment; PC=Permanent Contract; TC=Temporary Contract

Figure 1:Entrance Probability to a Temporary Contract from Unemployment (1995-2004)

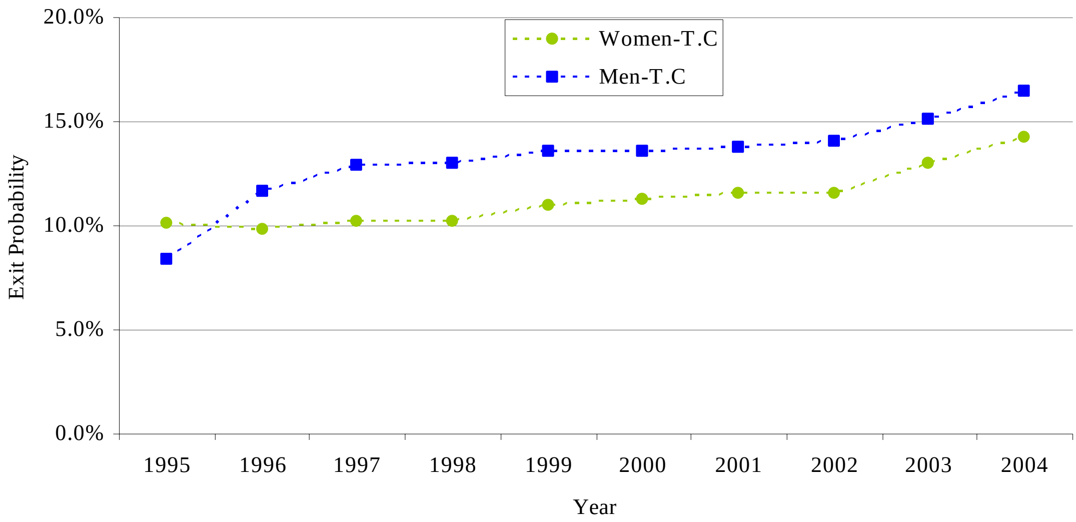

Figure 2:Transition Probability from a Temporary Contract to a Permanent Contract at the same firm (1995-2004)

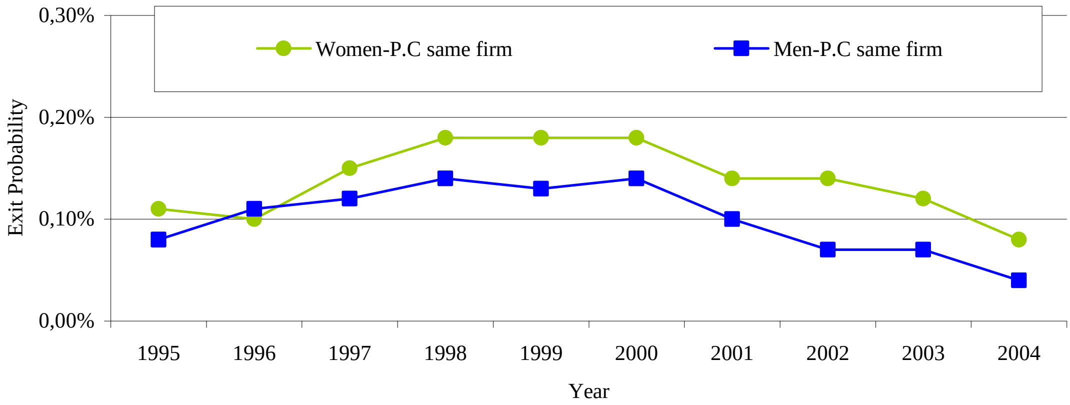

Figure 3:Entrance Probability to a Temporary Contract from Unemployment (1995-2004)

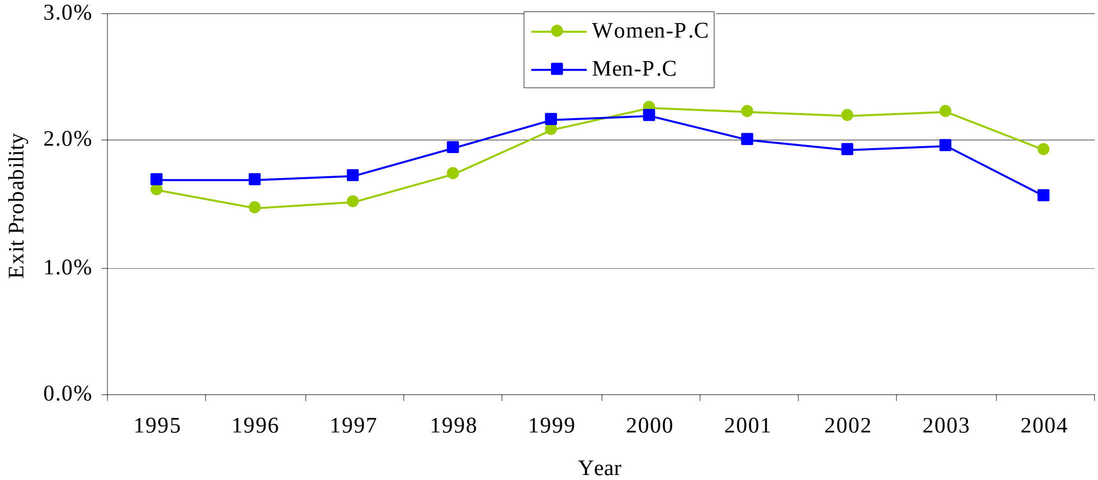

Figure 4:Transition Probability from a Temporary Contract to other Temporary Contract at the same firm (1995-2004)

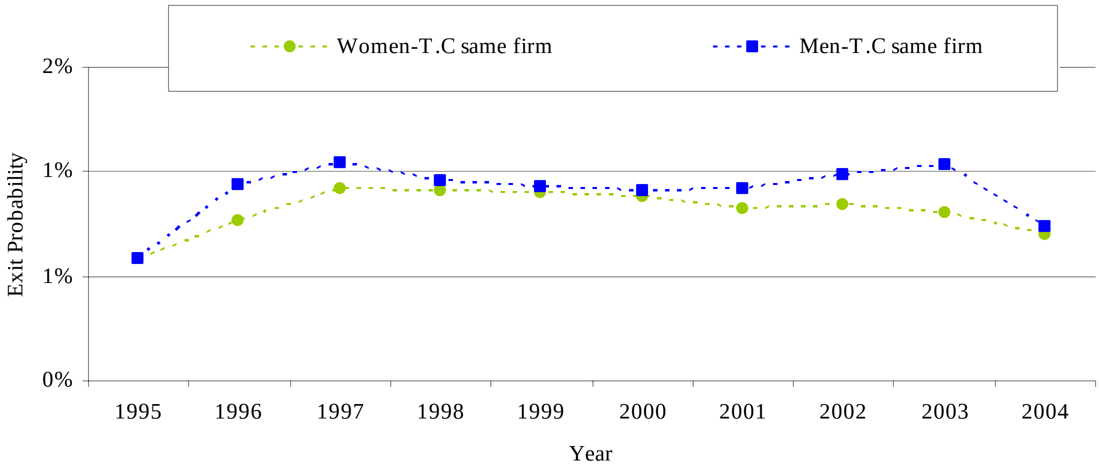

Figure 5:Transition Probability from a Temporary Contract to Unemployment or Different Firm (1995-2004)

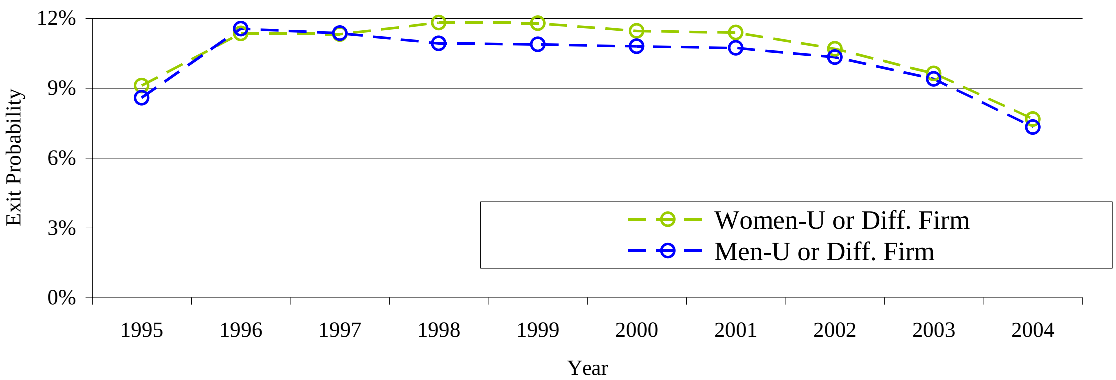

Table 9:Empirical Entrance Probability to a Permanent Contract by labour states before and after the 1997 reform

	Before 1997		After 1997		Diff before-after	Diff before-after	Diff-and Diff
	Treated	**Non-Treated**	**Treated**	**Non-Treated**	**Treated**	**Non-Treated**	
Unemployed							
Women							
Age <30	12.3	14.5	14.3	11.6	16%	-20%	181%
Age 30-45	10.0	15.0	12.3	11.1	23%	-26%	188%
Age >45	13.7	16.2	12.6	11.3	-8%	-30%	73%
Men							
Age <30	12.4	15.3	12.1	9.8	-2%	-36%	93%
Age 30-45	8.3	13.5	9.5	8.7	14%	-35%	141%
Age >45	7.6	15.3	8.4	8.8	10%	-42%	125%
Temporary Contract							
Women							
Age <30	1.00	0.83	1.12	0.65	12%	-22%	155%
Age 30-45	1.05	1.60	1.10	0.63	5%	-61%	108%
Age >45	0.88	1.43	0.98	0.84	115%	-41%	128%
Men							
Age <30	0.90	0.90	0.82	0.47	-9%	-48%	81%
Age 30-45	0.82	0.94	0.77	0.46	-6%	-51%	88%
Age >45	0.80	1.10	0.79	0.70	-1%	-36%	97%

Table 10: Competing Risk Duration Model by Gender (Initial State=Temporary Contract)

	Women						Men					
	U.		**T.C**		**P.C**		**U.**		**T.C**		**P.C**	
	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd
TC. Dur. (ln)	3.793	0.01	3.678	0.04	3.800	0.13	3.410	0.01	3.540	0.03	3.598	0.13
TC. Dur. (ln)^2	-0.959	0.00	-0.988	0.02	-0.822	0.03	-0.867	0.00	-0.890	0.01	-0.731	0.03
TC. Dur < 3 months	1.417	0.01	1.416	0.03	1.495	0.08	1.325	0.01	1.261	0.02	1.407	0.08
Month 6 (T.C. Dur)	0.992	0.01	0.761	0.03	1.369	0.05	0.797	0.01	0.532	0.02	1.362	0.05
Month 12 (TC. Dur)	1.249	0.01	1.031	0.05	1.531	0.07	0.842	0.01	0.377	0.04	1.253	0.07
Month 24 (TC. Dur)	1.624	0.03	1.381	0.16	1.410	0.20	1.160	0.03	0.591	0.11	0.714	0.20
Month 36 (TC. Dur)	2.629	0.05	3.051	0.20	2.620	0.26	2.845	0.03	3.532	0.09	2.983	0.17
D_{ijt} *(Age <30)	-0.020	0.00	0.060	0.02	0.073	0.04	0.006	0.00	0.024	0.01	0.018	0.03
D_{ijt} ^ 2*(Age <30)	0.000	0.00	-0.006	0.00	-0.009	0.00	-0.000	0.00	-0.002	0.00	-0.002	0.00
D_{ijt} *(Age:30-45)	-0.012	0.00	0.062	0.02	0.039	0.04	0.008	0.00	0.064	0.01	0.081	0.03
D_{ijt} ^2 *(Age:30-45)	0.000	0.00	-0.006	0.00	-0.003	0.00	-0.001	0.00	-0.006	0.00	-0.007	0.00
D_{ijt} *(Age >45)	0.011	0.00	0.080	0.02	0.033	0.06	0.009	0.00	0.055	0.00	-0.034	0.04
D_{ijt} ^2 *(Age >45)	-0.002	0.00	-0.007	0.00	-0.004	0.00	-0.001	0.00	-0.005	0.00	0.003	0.00
Age <30	0.540	0.04	0.417	0.12	0.403	0.27	0.635	0.03	0.659	0.08	0.037	0.23
Age 30-45	0.132	0.03	-0.024	0.11	-0.069	0.25	0.256	0.03	0.269	0.07	-0.312	0.23
D_{ij}	-0.052	0.01	0.253	0.04	-0.141	0.067	-0.101	0.01	-0.040	0.04	0.081	0.11
D_{jt}	0.061	0.01	-0.058	0.05	-0.039	0.11	-0.008	0.01	-0.053	0.03	-0.004	0.09
D_j	-	-	-	-	-	-	0.014	0.01	0.236	0.05	-0.278	0.11
Age	-0.002	0.00	0.109	0.01	0.100	0.02	0.024	0.00	0.122	0.01	0.156	0.01
Age^2	0.000	0.00	-0.001	0.00	-0.001	0.00	0.000	0.00	-0.001	0.00	-0.002	0.00
Dempleo	0.050	0.14	-2.950	0.47	-0.047	1.09	1.037	0.11	0.078	0.35	-0.712	1.00
Unemp. Rage	-0.011	0.00	0.006	0.01	-0.018	0.01	0.006	0.00	0.036	0.00	-0.044	0.01
T.c. GDP	-0.037	0.01	0.023	0.03	-0.016	0.06	-0.011	0.01	0.051	0.02	0.000	0.05
Unemp. Rage*(Age <30)	0.020	0.01	0.001	0.02	0.041	0.05	0.028	0.01	0.026	0.02	0.098	0.04
T.c. GDP*(Age <30)	0.026	0.01	0.024	0.03	-0.061	0.06	0.006	0.01	-0.003	0.02	-0.005	0.05
Unemp. Rage*(Age 30-45)	-0.005	0.00	0.008	0.01	0.007	0.01	-0.014	0.00	-0.005	0.00	0.027	0.01
T.c. GDP*(Age 30-45)	0.004	0.00	0.011	0.01	0.029	0.01	-0.006	0.00	-0.002	0.00	0.021	0.01
Part-time Job	0.133	0.01	0.092	0.02	0.065	0.04	0.311	0.01	0.209	0.03	0.324	0.054
High Wage Category	-0.415	0.01	-0.077	0.03	-0.189	0.08	-0.684	0.01	-0.422	0.04	-0.534	0.10
Medium-High Wage Cat.	-0.090	0.01	-0.067	0.03	0.246	0.06	-0.250	0.01	-0.260	0.03	-0.101	0.07
Medium-Low Wage Cat.	-0.035	0.01	-0.123	0.02	0.339	0.04	-0.118	0.00	-0.066	0.02	-0.036	0.04
Inmigrant	0.123	0.01	-0.173	0.06	0.274	0.10	0.135	0.01	-0.120	0.03	-0.088	0.09
Layoff	1.055	0.01	2.122	0.05	1.664	0.07	0.925	0.01	2.010	0.03	1.788	0.07
Big Firm	-0.188	0.01	0.583	0.02	0.281	0.05	-0.173	0.01	0.287	0.03	0.347	0.05
New Firm	-0.068	0.01	-0.092	0.02	0.007	0.04	0.017	0.00	-0.052	0.02	-0.152	0.04
Private	0.053	0.01	0.024	0.03	1.661	0.10	-0.038	0.01	0.078	0.03	1.908	0.14
Cte	-7.091	0.06	-13.44	0.22	-15.96	0.45	-6.684	0.05	-13.57	0.15	-16.25	0.44

U= Unemployment or Different Firm; TC= Temporary Contract and Same Firm; PC = Permanent Contract and Same Firm
We also control for time and quarterly dummies and for sector of activity.
The Unemployment Rate refers to Regional Unemployment Rates while the GDP is measured at the national level

Table 11: Main Results of the Effects of the Wage Subsidy by Gender and Age (Exit from a Temporary Contract)

			To U	**To T. C.**	**To P. C.**	**Staying at TC**
Women	Age <30	Odd Ratio	27.43%	0.91%	-	16.57%
		Marginal Effect	-1.51%*	2.86%*	6.68%**	0.08%*
		Total Change in Probability	**-8.07%**	**16.08%**	**26.26%**	**1.09%**
	Age 30-45	Odd Ratio	22.43%	-1.31%	-	14.23%
		Marginal Effect	-0.60%*	2.95%*	1.99%*	0.09%**
		Total Change in Probability	**-6.34%***	**16.18%***	**23.59%**	**0.97%****
	Age >45	Odd Ratio	2.60%	-16.27%	-	3.56%
		Marginal Effect	-0.58%**	5.10%*	-0.65%*	0.03%*
		Total Change in Probability	**1.21%**	**23.53%**	**11.48%**	**0.47%**
Men	Age <30	Odd Ratio	5.72%	-3.39%	-	2.21%
		Marginal Effect	-0.70%**	1.22%**	-0.50%	0.05%**
		Total Change in Probability	**-3.11%**	**6.02%**	**2.42%**	**0.21%**
	Age 30-45	Odd Ratio	21.81%	4.24%	-	22.56%
		Marginal Effect	-0.48%*	2.82%*	4.45%*	0.02%*
		Total Change in Probability	**0.48%**	**17.42%**	**22.40%**	**-0.13%**
	Age >45	Odd Ratio	-10.90%	-21.47%*	-	-9.71%
		Marginal Effect	-0.32%**	2.08%	0.27%	0.01%*
		Total Change in Probability	**1.17%**	**14.79%**	**-9.86%**	**-0.16%**

U= Unemployment or Different Firm; TC= Temporary Contract and Same Firm; PC = Permanent Contract and Same Firm
* Marginal Effect statistically significant at 95%
** Marginal Effect statistically significant at 90%
The rest of coefficients are not statistically significant
The Odd Ratio is measuring the behavior of the probability of having a permanent contract relative to the rest of alternatives for treated workers when the wage subsidy is in its sample average value (5100 Euros).
The marginal effect is also measured at the average wage subsidy.

Table 12: Competing Risk Duration Model by Gender (Initial State=Unemployed)

	Woman				Men			
	TC		PC		TC		PC	
	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd
U. Dur. (ln)	-0.422	0.01	-0.283	0.02	-0.410	0.01	-0.491	0.01
U. Dur. (ln)^2	0.030	0.00	-0.041	0.01	-0.019	0.00	0.036	0.00
D_{ijt} *(Age <30)	0.001	0.00	0.031	0.01	0.073	0.00	0.032	0.01
D_{ijt} ^ 2*(Age <30)	-0.000	0.00	-0.001	0.00	-0.001	0.00	-0.001	0.00
D_{ijt} *(Age:30-45)	0.011	0.00	0.004	0.01	0.009	0.00	0.020	0.01
D_{ijt} ^2 *(Age:30-45)	-0.000	0.00	0.002	0.00	-0.000	0.00	-0.000	0.00
D_{ijt} *(Age >45)	0.014	0.01	-0.015	0.01	0.003	0.01	-0.022	0.01
D_{ijt} ^2 *(Age >45)	-0.000	0.00	0.004	0.00	0.000	0.00	0.002	0.00
Age <30	1.043	0.04	1.129	0.08	0.967	0.03	0.592	0.06
Age 30-45	0.340	0.03	0.376	0.07	0.335	0.03	0.052	0.06
D_{ij}	-0.045	0.01	-0.275	0.02	-0.051	0.01	-0.083	0.02
D_{jt}	-0.023	0.01	-0.094	0.03	-0.024	0.01	-0.040	0.02
D_j	-	-	-	-	0.059	0.01	-0.180	0.03
Age	0.131	0.00	0.214	0.00	0.149	0.00	0.232	0.00
Age^2	-0.001	0.00	-0.002	0.00	-0.002	0.00	-0.003	0.00
Received. U.B	-0.693	0.01	-0.371	0.02	-0.555	0.01	-0.309	0.02
Dur. U.B.	0.083	0.01	0.142	0.01	0.027	0.01	0.096	0.01
Dempleo	0.127	0.12	1.250	0.26	0.591	0.10	0.730	0.23
Unemp. Rage	-0.007	0.00	-0.041	0.00	0.013	0.00	-0.055	0.00
T.c. GDP	0.016	0.01	0.034	0.02	0.007	0.01	0.034	0.01
Unemp. Rage*(Age <30)	-0.018	0.01	-0.038	0.01	-0.014	0.01	-0.037	0.01
T.c. GDP*(Age <30)	-0.031	0.01	-0.081	0.02	-0.006	0.01	-0.068	0.01
Unemp. Rage*(Age 30-45)	-0.006	0.00	0.002	0.00	-0.019	0.00	0.017	0.00
T.c. GDP*(Age 30-45)	0.001	0.00	-0.005	0.00	-0.010	0.00	0.007	0.00
Part-time Job	-0.116	0.00	-0.061	0.01	-0.195	0.00	-0.163	0.01
Temp. Help Agency	0.252	0.01	-0.018	0.02	0.255	0.01	-0.137	0.02
High Wage Category	0.202	0.01	0.056	0.02	-0.180	0.01	0.328	0.02
Medium-High Wage Cat.	0.035	0.01	0.128	0.01	-0.135	0.01	0.200	0.01
Medium-Low Wage Cat.	0.061	0.01	0.187	0.01	0.045	0.00	0.164	0.01
Inmigrant	0.164	0.01	0.238	0.03	0.179	0.01	0.153	0.02
Layoff	0.224	0.01	-0.221	0.01	0.101	0.00	-0.169	0.01
Big Firm	0.090	0.01	-0.027	0.01	0.014	0.01	-0.037	0.01
New Firm	0.010	0.00	0.040	0.01	0.023	0.00	0.016	0.01
Same Firm	0.847	0.00	0.279	0.01	0.622	0.00	0.111	0.01
Constant Term	-5.080	0.05	-7.412	0.10	-5.465	0.04	-6.781	0.09

TC= Temporary Contract; PC = Permanent Contract
* Marginal Effect statistically significant at 95%
** Marginal Effect statistically significant at 90%
The rest of coefficients are not statistically significant
The Odd Ratio is measuring the behavior of the probability of having a permanent contract relative to the rest of alternatives for treated workers when the wage subsidy is in its sample average value (5100 Euros).
The marginal effect is also measured at the average wage subsidy.

Table 13: Main Results by Gender and Age of the Policy Variable (Exit from Unemployment)

			To T. C.	To P. C.	Stayed Unemployed
Women	Age <30	Odd Ratio	14.56%	-	15.36%
		Marginal Effect	-0.01%*	2.55%**	-0.06%*
		Total Change in Probability	**0.16%**	**14.97%**	**-0.34%**
	Age 30-45	Odd Ratio	-0.16%	-	7.39%
		Marginal Effect	0.55%*	2.26%*	-0.13%**
		Total Change in Probability	**3.96%***	**6.46%**	**-0.68%****
	Age >45	Odd Ratio	-5.80%	-	1.41%
		Marginal Effect	1.07%*	2.41%**	-0.21%*
		Total Change in Probability	**6.27%**	**0.66%**	**-0.86%**
Men	Age <30	Odd Ratio	7.59%	-	14.66%
		Marginal Effect	-0.26%**	2.38%	-0.01%**
		Total Change in Probability	**0.75%**	**14.40%**	**-0.41%****
	Age 30-45	Odd Ratio	3.18%	-	8.61%
		Marginal Effect	0.43%*	1.33%*	-0.10%
		Total Change in Probability	**2.96%**	**8.36%**	**-0.66%**
	Age >45	Odd Ratio	-6.32%*	-	-5.72%
		Marginal Effect	0.52%	0.20%	-0.09%
		Total Change in Probability	**2.03%**	**-5.79%**	**-0.22%**

TC= Temporary Contract; PC = Permanent Contract
* Marginal Effect statistically significant at 95%
** Marginal Effect statistically significant at 90%
The rest of coefficients are not statistically significant
The Odd Ratio is measuring the behavior of the probability of having a permanent contract relative to the rest of alternatives for treated workers when the wage subsidy is in its sample average value (5100 Euros).
The marginal effect is also measured at the average wage subsidy.

Table 14: Competing Risk Duration Model (Initial State=Permanent Contact)

	Woman						Men					
	U		TC		PC		U		TC		PC	
	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd	Coef.	sd
PC. Dur. (ln)	-1.307	0.04	-1.538	0.10	-1.558	0.11	-1.328	0.04	-1.412	0.08	-1.540	0.11
PC. Dur. (ln)^2	0.247	0.01	0.268	0.02	0.311	0.02	0.250	0.01	0.26	0.02	0.311	0.02
Month 6 (P.C Dur)	0.412	0.03	0.200	0.10	0.188	0.12	0.582	0.03	0.456	0.08	0.283	0.11
Month 12 (PC. Dur)	0.229	0.04	0.221	0.12	0.270	0.15	0.288	0.04	0.114	0.10	0.296	0.13
Month 24 (PC. Dur)	0.071	0.06	0.018	0.17	0.087	0.20	0.022	0.06	-0.179	0.15	-0.486	0.23
Month 36 (PC. Dur)	-0.091	0.08	-0.254	0.24	0.175	0.22	-0.297	0.08	-0.025	0.16	-0.057	0.21
D_{ijt} *(Age <30)	-0.079	0.01	-0.017	0.04	0.108	0.04	-0.009	0.01	-0.065	0.02	-0.034	0.03
D_{ijt} ^ 2*(Age <30)	0.007	0.00	0.003	0.00	-0.004	0.00	0.002	0.00	0.004	0.00	0.004	0.00
D_{ijt} *(Age:30-45)	-0.036	0.01	0.035	0.05	0.186	0.06	-0.018	0.01	-0.060	0.03	-0.002	0.00
D_{ijt} ^2 *(Age:30-45)	0.004	0.00	-0.003	0.00	-0.011	0.00	0.003	0.00	0.006	0.00	-0.002	0.00
D_{ijt} *(Age >45)	-0.008	0.02	0.031	0.07	0.138	0.08	-0.004	0.02	0.012	0.05	0.025	0.06
D_{ijt} ^2 *(Age >45)	0.001	0.00	-0.001	0.00	-0.007	0.01	0.002	0.00	-0.001	0.00	-0.004	0.01
(Age <30)	1.535	0.09	1.757	0.34	1.870	0.33	1.616	0.1	1.639	0.23	1.963	0.31
(Age 30-45)	0.416	0.09	0.633	0.34	0.653	0.33	0.692	0.09	0.733	0.23	0.948	0.3
D_{ij}	-0.011	0.02	-0.099	0.07	-0.271	0.09	-0.189	0.04	0.005	0.09	0.227	0.12
D_{jt}	0.095	0.04	-0.123	0.13	-0.366	0.15	-0.055	0.03	0.121	0.08	-0.004	0.10
D_{j}	-	-	-	-	-	-	0.155	0.04	0.046	0.1	-0.407	0.13
Age	0.083	0.01	0.062	0.02	0.147	0.02	0.11	0.01	0.127	0.01	0.184	0.02
Age^2	-0.001	0.00	0.000	0.00	-0.001	0.00	-0.001	0.00	-0.001	0.00	-0.002	0.00
Unemp. Rate	-0.054	0.02	-0.033	0.02	-0.005	0.02	-0.009	0.00	0.015	0.01	-0.015	0.02
T.c. GDP	-0.026	0.02	-0.036	0.07	0.004	0.07	-0.016	0.02	-0.107	0.05	-0.052	0.07
Unemp. Rage*(Age <30)	0.015	0.00	0.055	0.07	-0.015	0.07	0.007	0.02	0.086	0.05	0.043	0.07
T.c. GDP*(Age <30)	0.022	0.01	-0.016	0.08	-0.170	0.08	-0.013	0.02	0.001	0.06	-0.046	0.07
Unemp. Rage*(Age 30-45)	-0.005	0.00	0.011	0.02	-0.023	0.02	-0.004	0.00	-0.021	0.01	-0.011	0.02
T.c. GDP*(Age 30-45)	0.004	0.00	0.030	0.02	0.022	0.02	-0.002	0.00	-0.011	0.01	-0.003	0.02
Part-time Job	0.285	0.02	0.248	0.05	0.132	0.06	0.789	0.02	0.579	0.05	0.341	0.07
High Wage Category	-0.641	0.03	-1.093	0.10	-0.741	0.10	-0.706	0.03	-1.562	0.09	-0.404	0.08
Medium-High Wage Cat.	-0.426	0.02	-0.607	0.06	-0.481	0.07	-0.369	0.02	-0.629	0.05	-0.088	0.06
Medium-Low Wage Cat.	-0.180	0.02	-0.307	0.05	-0.185	0.06	-0.305	0.02	-0.353	0.04	-0.204	0.05
Inmigrant	0.378	0.04	0.432	0.10	0.274	0.14	0.342	0.03	0.259	0.07	0.030	0.11
Layoff	1.624	0.02	0.924	0.05	0.756	0.05	1.773	0.02	1.146	0.04	0.809	0.05
Big Firm	-0.197	0.02	-0.278	0.06	-0.431	0.07	-0.333	0.02	-0.441	0.06	-0.530	0.07
New Firm	0.147	0.01	0.183	0.04	0.177	0.05	0.176	0.01	0.223	0.03	0.069	0.04
Private	0.114	0.04	-0.040	0.12	0.534	0.15	0.329	0.05	0.354	0.17	0.326	0.16
Same Firm	1.290	0.02	0.040	0.08	2.073	0.06	1.300	0.02	0.279	0.07	2.162	0.06
Temp. Help Agency	0.534	0.11	0.881	0.22	0.947	0.26	0.714	0.13	1.268	0.21	1.715	0.24
Constant Term	-6.110	0.18	-7.550	0.58	-10.05	0.64	-6.824	0.17	-8.592	0.39	-9.746	0.53

U= Unemployment ; TC= Temporary Contract; PC = Permanent Contract
We also control for time and quarterly dummies and for sector of activity.
The Unemployment Rate refers to Regional Unemployment Rates while the GDP is measured at the national level

Table 15: Main Results of the Policy Variable by Age and Gender (Exit from a Permanent Contract)

			To U.	**To T. C.**	**To P. C.**	**Staying at PC**
Women	Age <30	Odd Ratio	-20.58%	-0.98%	57.33%	-
		Marginal Effect	3.26%**	1.69%*	10.19%**	-0.03%*
		Total Change in Probability	**-20.45%**	**-0.82%**	**57.57%**	**0.15%**
	Age 30-45	Odd Ratio	-7.72%	9.62%	95.17%	-
		Marginal Effect	-5.36%**	0.95%*	21.26%**	-0.03%**
		Total Change in Probability	**-7.71%**	**9.63%**	**95.19%**	**0.01%**
	Age >45	Odd Ratio	-2.56%	15.00%	67.58%	-
		Marginal Effect	-14.24%**	2.51%*	13.61%**	-0.01%
		Total Change in Probability	**-2.58%**	**14.97%**	**67.53%**	**-0.59%**
Men	Age <30	Odd Ratio	0.57%	-19.70%	-5.92%	-
		Marginal Effect	1.28%**	0.30%**	2.04%	-0.01%**
		Total Change in Probability	**0.60%**	**-19.67%**	**-5.89%**	**0.03%**
	Age 30-45	Odd Ratio	-1.49%	-14.45%	3.48%	-
		Marginal Effect	1.58%*	2.37%*	1.46%*	-0.01%*
		Total Change in Probability	**-1.46%**	**-14.43%**	**3.48%**	**0.03%**
	Age >45	Odd Ratio	2.52%	3.76%*	3.36%	-
		Marginal Effect	1.43%	0.40%	-0.46%	-0.01%*
		Total Change in Probability	**2.50%**	**3.73%**	**3.33%**	**-0.02%**

* Marginal Effect statistically significant at 95%
** Marginal Effect statistically significant at 90%
The rest of coefficients are not statistically significant
The Odd Ratio measures the behavior of the probability of staying at the permanent contract relative to the rest of alternatives for treated workers when the wage subsidy is 5100 Euros
The marginal effect is measured at the average wage subsidy of 5100 Euros.