

THE USE OF PERMANENT CONTRACTS ACROSS SPANISH REGIONS: DO REGIONAL WAGE SUBSIDIES WORK?

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This article evaluates the effectiveness of regional wage subsidies to foster permanent employment using the "Muestra Continua de Vidas Laborales". This dataset, which is used here for the first time to evaluate Spanish labour market policy, offers a complete employment history for each individual, with no aggregation bias. The policy analyzed consists of a one-time subsidy offered by some Spanish regions for new permanent contracts signed for certain fixed-term employees and unemployed workers. Because our policy variable presents individual, regional and temporal variation, we apply a triple difference estimator to identify the average treatment effect of this policy. We conclude that the outflow into permanent employment of eligible workers improves only minimally.

Keywords: Difference-in-difference-in-difference, causal evaluation analysis, wage subsidies, hiring subsidies.

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1. Introduction

During the last decade, fixed-term contracts have taken center stage in the economic debate over European labour market reform as the number of temporary contracts has steadily risen. While some analysts suggest that temporary contracts may act as stepping-stones towards permanent employment (Booth, Francesconi and Frank, 2002), there is increasing evidence that such contracts only serve to deepen the rift between insiders holding permanent contracts and outsiders who find themselves confined to the margins of the labour market, trapped between alternating spells of unemployment and temporary work (for the Spanish case, some recent studies are Güell and Petrongolo, 2007 or Rebollo, 2007). In an attempt to prevent such negative results, some governments have both restricted the conditions under which temporary employees can be hired and sought to entice employers with incentives for increasing the hiring under permanent contracts (OECD, 2002). The Spanish case offers a paradigmatic example since it is the European country with the highest rate of temporary contracts. Moreover, in order to reduce the presence of temporary contracts different levels of the Spanish administration have implemented both kinds of policies since the mid-1990s. In this paper, we aim at evaluating the effectiveness of regional wage subsidies –those designed as a tool to fostering permanent employment- by estimating their incidence over the exit rate from unemployment and from a temporary contract into a permanent one.

Although economists have long advocated employment subsidies, only a few theoretical models yield analytical results that can help to clarify the effect of employment subsidies on labour market flows. Katz (1986) provides a partial equilibrium dynamic analysis of wage subsidies for low-wage workers. He points out that, in a situation of structural unemployment in which the effective labour supply is completely elastic, a proportional wage subsidy will not affect the worker's wage, but rather will increase employment in proportion to the elasticity of labour demand for benefited workers. Yet as Calmfors (1994) has noted, the indirect effects of wage subsidies might weaken this result, at least in the short-run. He studied three indirect effects (deadweight loss, substitution effect and displacement effect) of wage subsidies. More recent theoretical work on the way subsidies affect job creation offers some interesting new insights. Mortensen and Pissarides (2001) integrate wage subsidies in a search and matching equilibrium framework

with endogenous job destruction. For economies with strong unemployment compensation packages and stringent employment protection laws, they conclude that hiring subsidies can actually decrease permanent employment by inducing a disproportionate number of firms to replace old jobs with new ones, thereby leading to a higher labour market turnover. Their main argument is that, while hiring subsidies do indeed stimulate job creation, once a job has been created, the opportunity cost of the match rises, since a firm needs only create a new position in order to receive the same subsidy again.

In Spain, the number of employees holding temporary contracts rose from 18% in 1987 to 33% in 1994 as a consequence of government-enacted flexibility measures, such as the well-known 1984 labour market reform. This rise induced the government to pass different laws in 1997 and 2001 designed to reduce the number of fixed-term contracts by lowering the firing costs of new permanent hires¹. Since 1997, the Spanish government has also subsidized the creation of new permanent contracts by applying what often amounts to a substantial discount in firm payroll taxes for new permanent hires; and this, in turn, has encouraged some regional governments to establish their own wage subsidy programs. These regional programs consist of two types of subsidies, paid to the firm only at the time of hiring: one for new permanent contracts signed for fixed-term employees *at the same firm* and other for unemployed workers. Thus, the promotion of permanent contracts has become an important labour market strategy in Spain. In fact, between 1999 and 2002, Spain devoted 0.28% of its GDP –more than any other OECD country– to subsidizing permanent employment². While a rigorous evaluation of the Spanish government's active labour market policies would help to clarify the possible benefits brought on by these efforts, we have only been able to locate two studies on the subject though none of them are devoted to the study of the regional wage subsidies mentioned above.

¹The 1997 reform designed a new permanent contract for which the main differential characteristic was its lower firing costs. However, this reform targeted certain population groups, mainly women, the long-term unemployed and younger workers. For further details see Kugler, Jimeno and Hernanz (2005).

²Between 1996 and 2002 Spain devoted 50% of its total ALMP expenditure on subsidized employment; in this regard it was only outspent by Italy and Belgium, who devoted 53% and 56% of their ALMPs, respectively, to this expense. The average for European Countries is about 30%.

One, by Kugler, Jimeno and Hernanz (2005), analyzes the effects of the 1997 national reform on the country's permanent employment rate and on the worker flows to a permanent contract. In this study the single group of workers not targeted by this reform, men aged between 30 and 45, is used as the control group and the data is extracted from the Spanish Labour Force Survey. Their results point to an increase of approximately 40% –four percentage points- in the quarterly flow of unemployed younger men into permanent positions, and of 20% –five percentage points- in the same flow of older men relative to that of middle-aged men during the reform years. On the other hand, they find that the subsidies and reduced dismissal costs have little effect on women. They also find an increase in post-1997 male conversions, as compared with the pre-1997 rate, of roughly 30% for men and 15% for women³. The other paper, by Arellano (2005), examines how the unemployment rate and the employment quality of workers in Madrid changed as a result of Spain's 2001 labour market reform⁴. The author concluded the reform does not appear to have been particularly effective at inducing firms to convert temporary contracts into permanent ones, and that during periods of clear economic expansion firms may prefer to generate new permanent contracts instead. To reach these conclusions, he used a database provided by the National Employment Agency (INEM), basing his identification strategy on the various age-related eligibility conditions defined by the 2001 subsidies.

Our paper presents the first formal evaluation of regional wage subsidies in Spain. They offer an additional source of identification (regional) which will help in the evaluation of such policies designed to enhance permanent employment. Moreover, this is also the first study using the *Muestra Continua de Vidas Laborales* (MCVL), a very large administrative dataset based on the complete labour history of more than one million workers, to evaluate the causal effect of this important labour policy. Our analysis may be said to improve on that of Kugler, Jimeno and Hernanz (2005) in two key ways. First, our database eliminates the possibility of aggregation bias since it allows us to observe all of any given worker's labour market transitions even within a given month, in contrast to the quarterly transitions studied

³Nevertheless, this last result for women should be used with caution, since the authors do not account for any age-group variability in the eligibility requirements of the reforms in question for women.

⁴This reform extended similar measures established by the 1997 reform to other worker groups.

by these authors. Secondly, there is more variability in our policy variable since the differences between eligible and non-eligible workers do not rest exclusively on age⁵ but also on regional variation. Additionally, for the pool of temporary workers, we study the conversion rate from a temporary to a permanent contract *at the same firm*, which is the specific objective of the policy we analyze.

The multidimensional variability of the eligibility rules for these subsidies allows us to identify their causal effect over the inflow to permanent employment. The decision whether or not to issue wage subsidies is entirely up to the regional authorities, as are the timing, individual eligibility requirements and the amount of any subsidies to be paid. Hence, regional wage subsidies show greater variability than do national ones, which affect all Spanish workers in a given gender-age group equally. In addition, there exists a significant regional variation in permanent employment rates. This rate is around 60% in southern Spain but come close to European standards (80-85%) in areas such as Madrid and La Rioja. At the same time, regions of both groups have implemented important wage subsidies on new permanent contracts.

Our exercise consists of a competing risk approach applied to two duration models in order to trace the impact of wage subsidies on the outflow to a permanent contract. The definition of the competing alternatives depends on the eligibility conditions established by the regional wage subsidies. First, we study the monthly conversion rate of workers who shift from temporary to permanent contracts *within the same firm*, in light of the existing alternatives (no change, becoming unemployed or changing firms, continuing with the same firm under a new temporary contract or getting into a permanent contract at the same firm). Second, we examine the monthly transition rate of workers who shift from unemployment to a permanent contract, when the alternatives are no change or temporary employment.

Given the characteristics of the policy to be analyzed and the fact that we have longitudinal data, a difference-in-difference-in-difference estimator (DDD, hereafter) would appear to be our best tool for studying the causal effect of this policy. Hence, we identify the treatment effect using three variability dimensions: 1) that between eligible and non-eligible workers who differ in their personal characteristics but belong to the same region; 2) that between eligible workers, relative to non-

⁵This is especially relevant for the pool of temporary workers, since the limitations of their data impede them from defining a proper comparison group.

eligible ones, who have the same personal characteristics but belong to different regions; and 3) that between periods when wage subsidies were in effect and periods when they were not. Our policy variable will be based on the maximum possible wage subsidy per eligible individual for any given region.

The basic result of our analysis is that regional wage subsidies affect the transition to a new permanent contract, either from unemployment or from a temporary contract at the same firm, but only minimally. In fact, the incidence we observe is even smaller than that found by Kugler, Jimeno and Hernanz (2005). Specifically, we find that wage subsidies increase the outflow from temporary to permanent employment within the same firm by about 67% for middle-aged eligible female workers. For all other workers, the effect is non-significant. The incidence of the subsidies over the entrance probability to a permanent contract from unemployment is smaller than the one found for temporary workers, and is only statistically significant for young eligible ones. This probability increases by about 10% for young female workers and by about 4% for male ones. Nevertheless, the outflows to permanent contracts are relatively low and the overall change in the intra-firm outflow from temporary to permanent employment is very small. For instance, the estimated transition probability, in quarterly terms, for renewals moving from temporary into permanent contract positions increases from 0.65% to 1.09% for middle-aged female workers.

The rest of the paper is organized as follows. First, we describe the main characteristics of wage subsidies designed and implemented at the regional level in Spain. Sections 3, 4 and 5 describe our econometric approach, our data and our main results. Finally, Section 6 discusses the main conclusions to be drawn from this study.

2. Regional wage subsidies on new permanent contracts

The promotion of permanent contracts has emerged as an important active labour market policy at both, the regional and national level in Spain. Since 1997, when the national government implemented important payroll tax discounts for new permanent employees, several Spanish regions established similar wage subsidies. In general, only temporary workers who are offered a permanent contract in the same firm and unemployed workers who are hired on a permanent basis can benefit from these regional subsidies. Since the decision of whether or not to implement such subsidies is a strictly regional one, as are the

timing, individual eligibility conditions and the amount to be paid, regional wage subsidies show strong variability in comparison with those stipulated by the national policy, which affects all Spanish workers equally. Specifically, for the period between 1997 and 2004, this variation in the individual eligibility conditions and timing of wage subsidies can be traced both between and within regions for 1) eligibility conditions and 2) subsidy amounts. Eligibility requirements are mainly based on certain individual worker characteristics, such as age and gender.⁶ For example, some regions offer subsidies only to women, whereas others restrict them to younger workers. In addition, eligibility conditions, timing and subsidy amounts can vary even within a given region. This variability allows us to examine the response of different worker groups to changes in the incentive being offered.⁷

The wage subsidies we are concerned, consists of a one-time payment to the firm once the new permanent contract is signed. Other important characteristic of this policy is that it requires that the hiring firm apply for the subsidy in order to receive it. Since our dataset comes from the Social Security archives, this condition prevents us from identifying whether or not a given worker with a new permanent contract actually benefited from the regional subsidies or merely fulfilled the corresponding eligibility criteria. In other words, given that the eligible worker in our sample can be identified as having fulfilled these criteria (but may or may not eventually have been offered permanent contracts) our analysis measures the *potential* impact of regional subsidies but cannot detect situations in which the policy fails to address certain worker groups for whom it was designed.⁸ Nevertheless, since accepting the subsidy always implies a reduction in the firm's overall

⁶The eligibility conditions of these subsidies may also depend on other individual characteristics (such as whether the worker is underrepresented in her occupation or is disabled) or certain aspects of the firm itself, such as its type of production activity. Nevertheless, since our database does not allow us to control for some of these variables, we have finally opted to consider only the age and gender based eligibility conditions of the subsidies studied here.

⁷Another dimension of variability not taken into account here is the small differences there exist in some regions among subsidies for conversions and subsidies for new permanent contracts addressed at unemployed workers.

⁸Thus, our analysis might overestimate the incidence of these subsidies over the outflow to a permanent contract. One possibility to correct this bias is to control also for the intensity of the policy in each region by using, for example, the total number of contracts that have been subsidized in any given year. We are already exploring this dimension but the data on this is not yet completely available.

labour costs, it seems reasonable to assume that firms will apply for such subsidies whenever the eligibility conditions are fulfilled.

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

Our data on wage subsidies, which was obtained from regional government sources, only includes data for subsidies issued at the regional level.⁹ The main characteristics of these subsidies are described in Tables 1 and 2. The policy shown in Table 1 was implemented beginning in 1997 in some regions, after 1997 in others, and in other regions (such as Cataluña and Navarra) it was never implemented at all.¹⁰ Table 1 also shows that the eligibility conditions for this policy vary markedly by region. In some regions, such as Andalusia, the policy applies to all workers, while in others it is geared towards certain worker groups such as women or younger adults.

⁹It may be that some local governments also offer wage subsidies or some other type of public subsidies as a means of encouraging firms to offer permanent contracts.

¹⁰While regional policies aimed to increase employment, in general, among specific worker groups did exist before 1997, we exclude them from this analysis on the grounds that they were not specifically designed to foster permanent employment.

TABLE 2
Regional incentives to the creation of permanent contracts

Regions	Gender		Age			Total	% relative to labour costs
	Males	Females	Age < 30	Age 30-45	Age > 45		
Andalusia	3,504	4,153	4,437	2,803	4,317	3,713	20.3%
Aragon	3,543	4,944	4,317	4,089	4,828	4,326	22.1%
Asturias	3,954	4,280	4,055	4,060	4,288	4,086	20.3%
Balearic Islands	0	2,912	2,907	2,906	2,956	2,912	16.3%
Canary Islands	3,256	3,600	3,498	3,331	2,991	3,422	19.7%
Cantabria	3,117	4,366	3,832	2,991	4,303	3,636	19.3%
C. Leon	3,573	4,114	4,593	3,225	2,413	3,788	22.1%
C. Mancha	3,097	3,812	3,392	3,817	3,413	3,473	18.4%
Catalunya	0	0	0	0	0	0	0.0%
Valencia	3,414	3,166	3,419	3,032	3,459	3,307	18.2%
Extremadura	9,149	8,668	8,389	9,061	10,543	8,974	55.5%
Galicia	2,962	4,411	3,886	3,327	3,211	3,639	20.9%
Madrid	8,094	8,374	8,154	8,245	8,462	8,218	33.3%
Murcia	4,591	5,224	4,981	4,538	5,258	4,826	29.3%
Navarra	0	0	0	0	0	0	0.0%
Basque Country	6,299	6,977	6,591	6,560	6,503	6,571	28.3%
Rioja	5,564	5,604	5,556	5,603	5,648	5,583	30.5%
Total	4,621	5,036	5,010	4,498	4,564	4,797	23.8%

Note:

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

Table 2 displays the average regional subsidy paid to workers from different age and gender groups, calculated in 2002 Euros. One may wonder whether the effectiveness of these wage subsidies depends on their amount. Since the company labour costs associated with hiring a new permanent worker decrease as the amount of the subsidy increases, the incidence of the policy over the entrance rate to these contracts should also rise with the level of the subsidy. Table 2 shows that the subsidy amount varies considerably by region and individual eligibility requirements. Despite such variations, the contribution represented by this particular subsidy appears to be quite generous. For the year 2000, the last column of Table 2 shows the subsidy-driven regional percentage drop in the labour costs associated with one-year contracts.¹¹ Overall, the table indicates that regional wage subsidies imply a reduction in labour costs ranging from 9% in the Balearic Islands (where

¹¹ Recall that this wage subsidy is a one-time payment that the firm receives after the new permanent contract is signed. The total discount this subsidy implies over the total labour costs assumed by the firm with the new contract depends, of course, on the duration of the contract.

the subsidy was lowest) to more than 60% in Extremadura. In average terms, the subsidy represents about 24% of labour costs.

This type of data has one important shortcoming: it refers to the maximum allowable wage subsidy for any given firm, per contract for each eligible worker characteristics. While it might initially seem more reasonable to use the minimum wage subsidy, in many cases the available information does not specify a minimum¹². Because we have chosen to use the maximum wage subsidy, our results will measure the maximum incidence of the subsidy.

3. The empirical approach: Identification and estimation method

Our goal is to measure the causal effect of regional wage subsidies on the outflow to a permanent contract. For this exercise, we estimate two duration models using a multiple risk approach. Hence, our objective is to estimate the exit or hazard rate. For each individual in our sample, we observe the duration of an episode of temporary work or unemployment: at time t the individual's situation changes as she/he moves into a competing alternative. A common strategy to estimating the hazard rate is to transform the duration model into a sequence of discrete choice equations defined on the surviving population at each spell's duration (see Jenkins, 1995). In this case, we define a binary variable y_{nt} that takes the value of one when the worker changes at time t from a given state to state n , and is otherwise assumed to be zero. This expression has exactly the same form as the likelihood function of a multinomial discrete choice model where y_{nt} is the binary endogenous dependent variable, once we have rearranged the database so that there are as many rows per individual as there are time units -in this case, months- of worker permanence in the initial situation (Allison, 1982; Jenkins, 1995). Hence, our likelihood function for each individual is:

$$\ln l = \sum_{t=1}^T \sum_{n=1}^N y_{nt} * \ln (\Pi_{nt}) + (1 - y_{nt}) * \ln (1 - \Pi_{nt}) \quad [1]$$

¹²In many cases, the corresponding regional government defines the maximum of the subsidy but nothing is said about the minimum. Thus, each regional government finally decides which the exact quantity of the subsidy is depending, as explained before, on different worker and firm characteristics.

where Π_{nt} is the conditional exit rate at time t , that is, the hazard rate to the destination state n . To estimate this transition probability, we will use a *multinomial logit* specification. Hence, the expression of the conditional exit rate from a certain initial state to the n -th destination state is:

$$\Pi_{nt} = \Pr (T_{nt=t} / T_{nt} > t) = \frac{\exp(P_t^n)}{\sum_{k=1}^K \exp(P_t^k)} \quad [2]$$

where P_t^n will be defined bellow.

3.1 The identification approach of the causal effect of the policy

To estimate the causal effect of one policy on the economic behaviour of any given individual, one needs some source of policy variation. It is usual to estimate the effect of government policies on economic outcomes using the information provided by spatial and temporal variables. Our own strategy for identifying the policy incidence takes into account not two, but rather three dimensions of policy variability (individual, regional and temporal variability). Hence, our econometric approach is based in a triple difference estimator (DDD, hereafter). More specifically, our identification strategy not only rests on pre and post-treatment variation but it also rests on the comparison of similar workers –with respect to age and gender- from treated and untreated regions and different workers within the same region.¹³ In this regard, our identification approach improves that of Kugler, Jimeno and Hernanz (2005) and also from that of Arellano (2005), since we do not rely exclusively on age-based variation in eligibility conditions.

The treatment to be analysed here is the regional wage subsidy. Hence, our treatment group includes all workers who are potentially affected by the policy at time t and region j . They will be compared to all

¹³Much of the debate over the validity of a Difference-In-Differences estimate typically revolves around the possible endogeneity of the interventions themselves. Besley and Case (2000) show that a potential source of bias is the presence of unobservable variables, which may determine both, the policy and the outcome of interest. In our case, it may be that some unobservable measure of pessimism about the region's potential for economic growth influences both the existence and the generosity of the policy and the contract type or types in a particular region. Hence, the variability of individual eligibility conditions plays an important role in this study.

workers who are not potentially affected by the wage subsidy at time t . A general specification of the DDD model is expressed as follows:

$$P_{ijt}^n = \lambda^n(t - t_0) + x'_{ijt}\alpha^n + \beta^n D_{ijt} + \eta_i^n + \mu_j^n + \delta_t^n + \xi_{it}^n + v_{jt}^n + \psi_{ij}^n + \varepsilon_{ijt} \quad [3]$$

where n stands for the corresponding destination state; the function $\lambda(t - t_0)$ controls for the duration dependence of the process and t_0 is the time when the spell began; the vector x_{ijt} contains covariates that includes time-varying personal, firm and job characteristics (age, country of origin, access to unemployment benefits, wage category, firm's size, etc).

The wage policy variable is D_{ijt} , and takes a positive value representing the maximum wage subsidy when the eligible worker i , located in region j , lives in a region with wage subsidies at time t ; otherwise, it is zero.¹⁴ From a policy point of view, the sign, size and significance of the estimated parameter β , which measures the true effect of the policy once we have controlled for all of the covariates that can simultaneously affect treatment and outcome, is the key result here. We obtain our estimation of the policy incidence on the potentially treated group from the post-treatment change in the outcome for this group, once we have controlled for the mean change in periods with and without treatment, for the mean differences among eligible and non-eligible workers' outcomes and for the mean differences among regions with and without subsidies.¹⁵

The remaining variables in equation [3] allow for a general specification of our model and also for the identification of our policy's causal effect. These controls are crucial in policy evaluation studies in order for β to be interpreted as the causal effect of the policy. They are fixed effects for individuals, regions and time and their interactions, which initially, are left unrestricted and are not specific to the treatment status. For

¹⁴Initially, our policy variable was a dummy variable that took the value of one when the worker i , located in region j at time t was eligible to receive them, and was otherwise equal to zero. However, this specification deprived us of one dimension of variability in our policy variable: the variation in the amount of the subsidy.

¹⁵One advantage of the DDD approach is that it controls for unobserved time-invariant differences between eligible and ineligible workers. Therefore, the estimation of the parameter β should be the same when estimating the model with and without control for individual unobserved heterogeneity. In any case, we also estimate the model with individual unobserved heterogeneity, assumed to be independent of all regressors, as a specification test over the policy variable. We follow the approach proposed by Heckman and Singer (1984) when specifying the heterogeneity term, assuming that each hazard rate has two support points.

instance, v_{jt} and μ_j stand for the fact that regions can have different transitions probabilities to jobs with permanent contracts and these differences can be time-varying and/or permanent. One reason for the time variation of these differences could be region-specific business cycles while sector specialization could be the source of the permanent differences among regions. The terms ξ_{it} and η_i stand for the effects that are common across regions but differ among individuals, time-varying in the first case and fixed along time in the second case. Indeed, the term ξ_{it} is crucial in our study since regional wage subsidies took place roughly at the same time as the national reduction in payroll taxes for specific groups of workers (1997 and 2001). Thus, the term ξ_{it} may be considered as a source of control for the effect of such national reforms.

Obviously, a model like the one specified in [3] is difficult to estimate in full generality. Therefore, we must put some structure on some of these terms. In this sense, the time-varying aggregate effects, δ_t , are controlled for by a set of time varying –yearly– dummy variables. The regional effects, μ_j , will be captured by regional dummies, and individual fixed effects (η_i) will be approximated by age dummies.¹⁶ Moreover, we will estimate each model separately by gender.¹⁷ With respect to the interactions terms, v_{jt} will be proxied by both a dummy variable, D_{jt} , indicating whether the worker lives in a region with subsidies in year t or not and also by the unemployment rate in each region, u_{jt} .¹⁸

Finally, the interactions of individual effects with regions (ψ_{ij}) and time (ξ_{it}) effects are more difficult to capture and some restrictions

¹⁶We introduce also *worker's age* as a continuous variable to allow for a more general effect of this variable over the hazard rates to be estimated. Nevertheless, the results concerning the policy effect on these flows remain almost the same when we restrict the effect of this variable just to age dummies.

¹⁷As will be discussed in the following section, the impact of these wage subsidies varies considerably across age groups, making it likely that the overall characteristics and behaviour of non-eligible workers will match that of the eligible ones. Hence, we will estimate the treatment effect allowing for different effects for different age categories.

¹⁸Alternatively, we could proxy v_{jt} by interactions of regional and year dummies or by region-specific time trends. However both alternatives have proved to be very difficult to be identified in our sample. In any case, we have also included in our specification an interaction between regional dummies and two variables measuring the pre-1997 and the 1997-2001 periods in our sample. This is a more parsimonious way of allowing for more general regional time-varying effects, taking into account the two labour market reforms that took place in our sample period.

are needed to make the model more parsimonious. Given that the individual fixed effect is captured by age dummies, we represent ψ_{ij} by interactions between age and region dummies. Finally, the time varying effects that are common across regions but differ among individuals (ξ_{it}) are modelled by interactions between age and time dummies. This last group of variables may be capturing any potential asymmetric age-related effects resulting from the 1997 national reform.

One identification assumption is that the composition of the eligible and non-eligible worker groups must remain stable over time.¹⁹ This assumption might be violated if a person eligible for subsidy funding were to learn of and react to this policy before it became effective. Nevertheless, we feel that the strong regional and temporal variability in the eligibility conditions and the use of regional and individual time-varying covariates provides us with variables to fulfil these requirements. Admittedly, individual anticipation of the policy might affect the results when the period under analysis is short and when it immediately follows the effective start date of the new policy. In this context, one could argue that firms act strategically in an effort to gain as much subsidized funding as possible. But this is not the case here, since our study includes regions in which the policy remains unchanged over the course of several years.

4. Data description: muestra continua de vidas laborales

The data for our analysis has been obtained from a new dataset that has recently become available in Spain, called the *Muestra Continua de Vidas Laborales* (MCVL, hereafter). This is an administrative dataset based on a random draw from the Spanish Social Security records, which contains a sample of 4% of all affiliated employed and unemployed workers and pensioners for the year 2004. Because it covers each worker's entire labour history, it provides us with a sample in which individual work experiences are not left-censored for periods prior to that under analysis.²⁰

This dataset contains information for approximately 1.1 million people and covers their entire labour history. Because the amount of information available for each individual is quite large, we have imposed some restrictions. The first of these responds to the fact that the same

¹⁹These assumptions are discussed in detail in Blundell and MaCurdy (1999).

²⁰The previous labour history of workers is known to have a strong impact on the outflow to a permanent contract (Rebollo, 2007).

employment spell may be covered by more than one register. For example, it is quite common in Spain for firms to optimize their labour costs by firing workers for short periods of time and re-hiring them later.²¹ When this occurs, we treat the employment spell as a continuous period, despite the short interruption that appears mid-spell. More specifically, we unify successive registers when they correspond to the same worker in the same firm with the same type of contract, and when the interruption lasts for less than 15 days.²² Second, we disregard simultaneous employment spells and, instead, use the information corresponding to the longer-lasting of these. Third, we unify any two registers that present overlapping contracts, i.e., when one of the contracts begins before the previous one has ended. Fourth, we eliminate any register that is incomplete or inaccurate, such as those in which key information is missing or is clearly incorrect (incompatible starting and ending dates, etc.). Fifth, we only consider the labour histories of workers within the so-called *Regimen General*, that is, the pool of regular paid employees for any given firm. Our sample excludes self-employed workers, those working under the *Fishing and Agriculture* rubric, and other minor special cases. Finally, given that there are significant lacunae in the available data on contract type for pre-1994 employment spells, we only study post-1994 employment (and unemployment) spells.

This database stores the entire labour history for each worker, for whom it provides information relating to the worker's age, gender, occupation and the exact duration of each unemployment/employment spell. The latter is especially relevant here. Since the information gathered by other databases is either annual or quarterly, they exclude any data relating to individual labour market transitions taking place over the course of an entire year or even within a given quarter. As a result, they tend to under-represent the real number of short-term temporary contracts issued for any given period. Moreover, for periods of unemployment, the MCVL database allows us to distinguish between those that coincide with periods of payroll tax payment (during which the worker receives unemployment benefits) and those

²¹ A high percentage of workers in our database sign a new contract each week, mainly at the same firm.

²² In order to avoid studying extremely short spells due to reallocation or strong turnover within the firm, we disregard employment spells that last less than 30 days. Since one can assume that they really are of a temporary nature, they fall outside the scope of policies that aim to promote permanent employment.

that do not. The latter can represent either periods of unemployment without benefits or periods of inactivity. Nevertheless, we use the term *unemployment* to denote all periods of inactivity with respect to work.²³ Another interesting characteristic of this database is that it assigns a unique authentication code to each firm. This is fundamental to our analysis, since it helps us to identify situations in which the worker remains with the same firm after her contract has changed, or after an unemployment spell has ended.

TABLE 3
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
TC 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
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-Low Qualification	28.4%	31.1%	25.8%	37.3%	40.7%	34.2%	40.3%	36.6%
Low Qualification	37.4%	46.6%	51.7%	41.3%	41.8%	55.7%	52.1%	50.1%
Immigrant	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%
Private Firm	74.9%	88.9%	80.6%	95.9%	90.5%	92.0%	94.3%	98.3%
Sample Size	57,570	221,985	17,352	2,934	80,363	324,253	29,449	3,197

Note:

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

For our established selection criteria, we obtain 472.406 and 585.602 unemployment spells and 299.841 and 437.262 spells of temporary contracts for female and male workers, respectively. The main characteristics for unemployed and temporary workers in our sample are displayed in Tables 3 and 4. The mean duration for temporary contracts is between 5 and 9 months. The same figure for unemployment spells is 5-6 months.

²³We will not estimate the exit from unemployment for durations larger than 12 months. These spells will be considered in the analysis but they will be censored at duration 12.

TABLE 4
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-Low Qualification	31.3%	33.4%	37.0%	34.1%	32.7%	31.4%
Low Qualification	49.1%	44.7%	41.9%	52.1%	57.3%	51.0%
Immigrant	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=1)	35.4%	27.0%	32.8%	36.5%	31.7%	33.9%
Unemployment Benefits (t=T)	20.0%	18.1%	22.1%	24.0%	23.9%	24.5%
Sample size	61,961	350,018	60,697	63,826	458,307	63,472

Note:

PC=Permanent Contract; TC=Temporary Contract

In Tables 5 and 6 we show the distribution of destinations, conditional on exits, from temporary and unemployed workers, respectively. From Table 5 we learn that the outflow into a permanent contract is relatively low independently of the worker's age, gender or region. The main destination state after unemployment is a temporary contract. At the same time, the data shows that after a temporary contract a worker is most likely to exit to unemployment or into a job at a different firm. Thus, the probability that a female or male worker will obtain a permanent contract, conditional on exiting from unemployment, is only about 13% and 11%, respectively. The renewal probability from a temporary to a permanent contract is much lower and interestingly, again, larger for women (1.2%) than it is for men (0.9%). Since regional wage subsidies tend to be larger for women, we might ask ourselves whether these gender-specific differences are the result of these regional policies. Some age-based differences can also be found in these tables. For instance, younger workers are more likely than workers from other age groups to exit from unemployment into a permanent contract; at the same time, the conversion probability of temporary to permanent contracts within any given firm is larger for older workers than it is

for younger ones. The renewal probability from a temporary to a permanent contract is 1.5% and 1.1% for older female and male workers, respectively.

TABLE 5
Distribution of destinations, conditional on exit, for unemployed and temporary workers by age and gender

	Unemployed			
	U. - P.C.		U. - T.C.	
	Women	Men	Women	Men
<30	14.14%	12.39%	85.86%	87.61%
30-45	12.00%	9.66%	88.00%	90.34%
>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
<30	1.13%	0.85%	6.21%	7.06%
30-45	1.22%	0.82%	8.97%	10.13%
>45	1.53%	1.14%	9.36%	10.51%
Average	1.19%	0.88%	7.32%	8.45%

Note:

PC=Permanent Contract; TC=Temporary Contract

TABLE 6
Distribution of destinations, conditional on exit, for unemployed and temporary workers by region and gender

	Unemployed		Temporary Contract			
	U-P.C.		TC- T.C. Same Firm		TC-P.C. Same Firm	
	Women	Men	Women	Men	Women	Men
Andalucía	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%

Note:

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

Table 6 shows also the distribution of destinations, conditional on exits, for our sample. There exists also a significant regional variability with regard to outflows. Thus, the observable rates for Madrid, la Rioja or Catalonia are almost double those of Andalusia and Extremadura. If we compare this information to the data on regional subsidy amounts, it becomes evident that there is no clear relationship between these two variables.

FIGURE 1
Empirical Hazard rate from unemployment to a permanent contract by gender

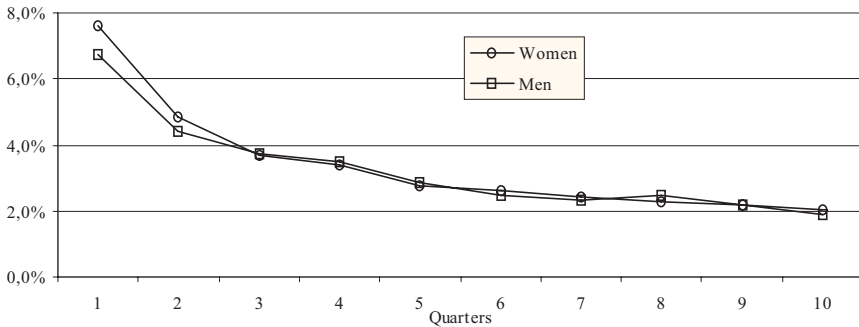
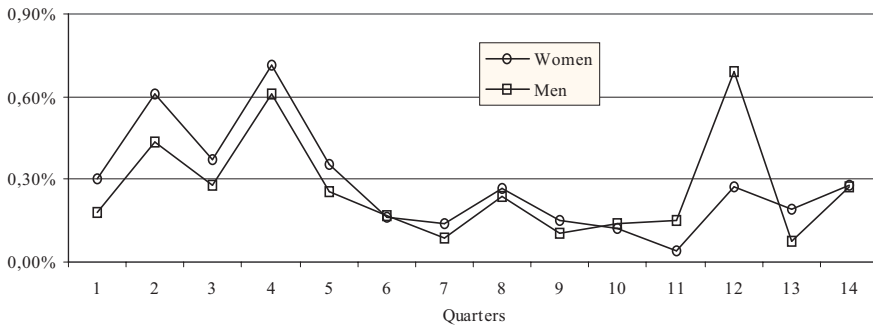


FIGURE 2
Empirical Hazard rate from a temporary to a permanent contract at the same firm by gender



In order to get a general picture of how transition rates change over the course of any given spell, Figures 1 and 2 show the quarterly exit rate to a permanent contract from unemployment and from a temporary contract at the same firm, respectively, for male versus female workers.²⁴ As expected, the exit rate from unemployment to a permanent contract decreases notably as the duration of the spell increases.

²⁴Though we have monthly information on spells duration we show the quarterly transition rate in order to make this information comparable to the one displayed in Kugler, Jimeno and Hernanz (2005).

Thus, during the first quarter this figure is close to 7%, falling one year later to approximately 3%. Minor gender-based differences can be observed at the beginning of the spell; women also face a slightly higher exit rate to a permanent contract than do men. By contrast, the pattern of change in the conversion rate from a temporary to a permanent contract tends to decrease as the duration of the temporary contract increases, with the exception of specific key moments, particularly the six-month, twelve-months and three-year marks, when it tends to spike. This rate is greater for women at the beginning of the contract period, while for men it is notably higher at the three-year peak.

FIGURE 3
Unconditional entrance probability to a permanent contract from a temporary contract at the same firm (1995-2004)

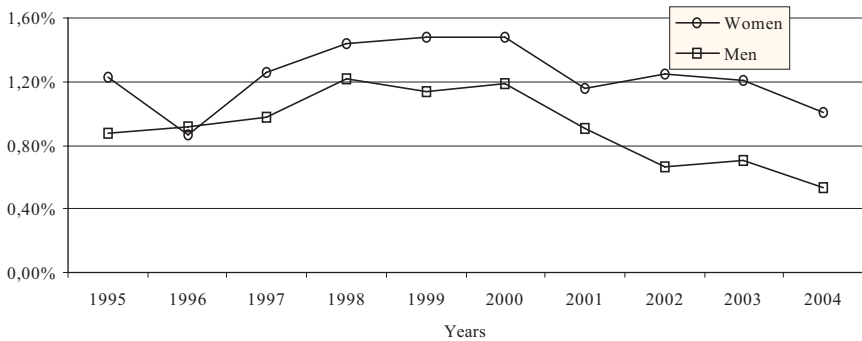


FIGURE 4
Unconditional entrance probability to a permanent contract from unemployment (1995-2004)



Figures 3 and 4 represent the unconditional transition probability from unemployment to a permanent contract and from a temporary to a permanent contract at the same firm, respectively between 1995 and 2004 among the workers in our sample. Remarkably, the outflow to a permanent contract increases between 1998 and 2001. Not surprisingly, this trend coincides with the post-1995 period of economic expansion; but, and this is particularly relevant here, it also coincides with the major labour market reform of 1997 and with the beginning of the national and regional policies to foster permanent employment. In sum, both figures highlight the relevance of controlling for cyclical factors as well as for periods of national reform.

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

This section addresses the causal incidence of regional wage subsidies over the outflow to a permanent contract for eligible workers, both from a temporary contract and from unemployment. In both cases, we follow a competing risk approach because it allows for a better understanding of the effects of the policy variable on all the labour market transitions the worker may have. For instance, for the pool of temporary workers, we can evaluate whether wage subsidies also tend to favour the transition from one temporary contract to a renewed version of that contract, instead of to unemployment. It might be that the effect on the conversion rate to a permanent contract is low, since firms might find it profitable to exhaust their legal limits before using the subsidy to convert this contract into a permanent one (Güell and Petrongolo, 2007). This could imply that these wage subsidies are carrying deadweight costs since they are subsidizing new permanent contract which would be anyway signed.

We have obtained our estimation using a sample of workers between the ages of 18 and 64, which we then divided up by gender in order to gain homogeneity between eligible and non-eligible individuals. We also allowed for age-based heterogeneous treatment effects and modelled the policy variable as an interaction term between the policy variable and the worker's age group. We studied workers from three age groups: 30 and younger, 31 to 45, and 46 and older. Since wage subsidies tend to vary by age, this distinction is relevant. We define the policy variable as the maximum permissible amount of subsidized funding per eligible individual. This allows us to use the individual,

regional and temporal variations in this quantity for eligible groups, to estimate the response of eligible individuals to a change in the incentives. We introduce in the model the exact amount of the subsidy and its square in order to capture any nonlinear pattern in the treatment effect. Using this specification, we can check whether the wage subsidy does in fact generate a rise in permanent contracts and whether (and to what extent) the amount of that subsidy plays a role in that effect.²⁵

We also specify a number of personal, firm and job characteristics in order to control for certain differences between eligible and non-eligible individuals, which might affect our outcomes. These variables include *worker's age, worker's country of origin, qualification category of the job, firm size, firm age, full-time job, activity sector, firm's ownership and temporary employment agency*. These variables refer to the worker's current job situation when employed and to the previous job when the worker is unemployed. As it was already mentioned in the previous section, we also consider variables that handle for national and region-specific time effects and regional fixed effects, thereby ensuring that our coefficients do not reflect smoothly trending omitted variables or permanent regional differences that are potentially correlated with the outcome under study. Because we follow a competing risk approach for each duration model, we obtain a specific vector of parameters for each alternative. Moreover, each duration model is estimated by gender, both with and without controls 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 to present only the most relevant of our results here.²⁶

Before we analyze the specific results of each duration model, a few general observations are in order. The duration dependence of the exit probability is specified as a polynomial in the log of the spell duration. In addition, we build some dummy variables that control for specific contract durations. For instance, for temporary workers these variables are: less than 4 months and 6, 12, 24, and 36 months.

²⁵Nevertheless, we have also estimated the model omitting the square term of the policy variable and we did not reach to different conclusions from the ones presented in this paper. Moreover, the statistical significance of the policy variable improves with the specification shown in the paper.

²⁶In the Appendix, we present the results regarding the policy variables. For the rest of the results see García-Pérez and Rebollo (2008).

For unemployed workers these variables are: less than 4 months and 6 and 12 months. The form of the empirical hazard rates justifies their consideration.

It is interesting to note that all interaction terms defined in order to control for the three dimensions of our policy variable are jointly significant²⁷ and that they differ among different alternatives (see García Pérez and Rebollo, 2008)). Moreover, these results support our view that a competing risk model is the best way to correctly identify the effect of the wage subsidy over the outflow to a permanent contract, in light of each of the competing alternatives. Furthermore, the results for these variables suggest that there are permanent and time-varying differences between eligible and non-eligible worker groups and between regions that would affect the outcome under study had it not been included in the analysis. The remaining variables, which control for permanent and time-varying differential effects by age group, are also statistically significant, in general. For instance, our results confirm the appropriateness of considering age-based differences in the effects of the business cycle and in the effects of the national labour market policies over the outflow to a permanent contract, as well as age-based differences.

For ease of exposition, we will summarize the main results for each of our duration models. Each summary includes the odd ratio, the total change in the exit probability for each alternative, both of which are measured at the average wage subsidy of 4.800 Euros and the Wald test of the joint statistical significance of the level and quadratic effect of the policy variable.²⁸ The odd ratio offers an easy way to summarize the main effects of the subsidies over the ratio of the outflow to a permanent contract relative to each competing alternative, comparing for each case the outflow for eligible versus non-eligible workers.²⁹ For instance, this ratio allows to test whether wage subsidies benefit more the accumulation of temporary contracts than the conversion of this temporary contract into a permanent one. By contrast, the total

²⁷We have tested the joint significance (Wald test) of the different groups of interactions for each competing alternative and they are all highly significant.

²⁸Standard errors are corrected to take into account the group structure of our policy variable.

²⁹Thus, the odd ratio could be expressed as follows: $[(P_i(e)/P_i(ne))/(P_j(e)/P_j(ne))] - 1$ where " $P(\cdot)_j$ " is the probability of a certain event, the indexes "i" and "j" describe competing alternatives, "e" refers to eligible workers and "ne" refers to non-eligible workers.

change in the exit probability shows whether wage subsidies have a substantial causal impact for eligible workers on the transition probability to each of the competing alternatives in relation to remaining at the same position.

5.1 The effects of regional wage subsidies on the conversion rate from a temporary to a permanent contract at the same firm

For temporary workers, the only existing wage subsidies are those that support the conversion of a temporary contract into a permanent one within the same firm. This restriction affects our analysis in two ways. First, it affects the way we define the competing alternatives. These are: 1) remaining at the same temporary contract;³⁰ 2) accepting another temporary contract with the same employer; 3) accepting a permanent contract with the same employer; 4) becoming unemployed or accepting a contract with a different employer. We matched the unemployment alternative with that of a direct transition to a different firm, since it seems reasonable to assume that workers who change firms will be unemployed for at least one day. For ease of exposition, this latter alternative will henceforth be denoted as *exit to unemployment*. Second, workers directly hired by a Temporary Employment Agency will not benefit from any subsidies unless they become unemployed. In order to sharpen our definition of a potentially eligible worker, we have therefore excluded all temporary workers who were hired through such agencies from the estimation of this model. However, these workers will figure among the sample of unemployed workers we will analyze below.

Table 7 gives a summary of our main results. We can evaluate the impact of the subsidies by looking at the total change in the transition probability for eligible workers versus non-eligible ones brought about as a direct result of these funds. The results presented in Table 7 show that the impact of wage subsidies over the employee outflow to a permanent contract at the same firm varies considerably by worker type. It is positive for young and middle-aged female workers and for old male ones, but it is only statistically significant for middle-aged women. In this last case, this probability increases by 66.91% for eligible workers relative to non-eligible ones.³¹ For the rest of workers

³⁰Temporary contracts are censored at duration 36.

³¹It is also interesting to note that, for middle age female workers, the change in the intra-firm transition probability to a permanent contract increases as the wage subsidy rises to the observed maximum value (12.200 Euros). For instance, the odd

the effect of wage subsidies over this outflow is negative but it is not statistically significant, as well.

TABLE 7
Main results of the effects of the wage subsidy by gender and age (exit from a temporary contract)

			To U	To T. C. same firm	To P. C. same firm	Staying at TC
Women	<30	Odd Ratio	10.32%	15.08%	-	10.94%
		Total Change in Probability	0.41%	-3.74%	10.77%	-0.16%
		Significance Test	(3.94)	(7.13) **	(0.19)	
	30-45	Odd Ratio	66.05%	82.47%	-	67.24%
		Total Change in Probability	0.52%	-8.53%	66.91%	-0.20%
		Significance Test	(11.87) **	(8.15) **	(5.93) **	
>45	Odd Ratio	-26.32%	-13.50%	-	-25.92%	
	Total Change in Probability	0.44%	-13.82%	-25.81%	0.14%	
	Significance Test	(0.86)	(1.28)	(2.06)		
Men	<30	Odd Ratio	-26.12%	-33.64%	-	-27.13%
		Total Change in Probability	-1.06%	10.14%	-26.91%	0.30%
		Significance Test	(0.49)	(3.58)	(1.59)	
	30-45	Odd Ratio	-15.75%	-18.36%	-	-13.26%
		Total Change in Probability	2.24%	5.50%	-13.86%	-0.69%
		Significance Test	(2.35)	(1.97)	(1.29)	
>45	Odd Ratio	34.89%	30.65%	-	43.37%	
	Total Change in Probability	4.72%	8.12%	41.26%	-1.47%	
	Significance Test	(2.47)	(1.04)	(0.86)		

Note:

U= Unemployment or Different Firm; TC= Temporary Contract, Same Firm; PC = Permanent Contract and Same Firm.

The Odd Ratio measures the change in the ratio of the probability of exiting to a permanent contract for eligible workers versus non-eligible ones, relative to the same ratio -eligible versus non-eligible-, for the available alternatives, when the wage subsidy takes the average sample value (4,800 Euros).

The total change in probability is measured at the average wage subsidy and at t=6.

The significance test shown is the result of the Wald Test of the joint statistical significance of the level and quadratic effect of the policy variable: statistically significant at 90% (*) and at 95% (**).

Looking at the odd ratios we can observe that the probability of obtaining a permanent contract at the same firm, for eligible workers versus non-eligible ones, relative to the same ratio among eligible and non-eligible for the exiting to unemployment is positive in cases involving young and middle-aged female workers and old male ones but the effect of the policy is statistically significant only in the case of middle-aged women. For them, the odd ratio is 66.05%. That is, the

ratio of exiting to a permanent contract is 101% when the wage subsidy is 6.800 Euros and 210% when the wage subsidy is 12.200 Euros. Therefore, increasing the wage subsidy would seem to offer some leverage for improving the transition rates to permanent contracts among women.

increase in the probability of exiting to a permanent contract for eligible workers is 66.91% while the growth in the probability of exiting to unemployment for eligible workers relative to non-eligible ones is only 0.52%. We obtain also that the odd ratio of obtaining a permanent contract instead of exiting to another temporary contract at the same firm is also positive for female workers and for old male ones but the effects are only statistically significant for young and middle-aged female ones. These odd ratios are 15.08% and 82.47%, respectively. In both cases, these odd ratios are larger than the ones relative to the alternative of unemployment because the probability of renewal with a temporary contract at the same firm drops for eligible workers (-3.74% for young female workers and -8.53% for middle-aged female ones). Interestingly, this result shows that regional wage subsidies do not seem to favour the accumulation of temporary contracts at the same firm for these eligible workers.

TABLE 8
Estimated conversion rate from a temporary to a permanent contract at the same firm (Quarterly transition rates)

		Estimated Transition Rate to PC (2 ^o Quarter)*		
		Without the RWS	With the RWS	Growth Rate
Women	<30	0.83%	0.92%	10.84%
	30-45	0.65%	1.09%	66.97%
	>45	1.49%	1.11%	-25.58%
Men	<30	0.56%	0.41%	-27.00%
	30-45	0.48%	0.42%	-12.26%
	>45	0.57%	0.80%	41.57%

Note:

We display all the estimated transition rates but to correctly interpret these results, one must to take into account the statistical significance of these estimations showed in Table 7.

Table 8 shows the estimated intra-firm transition rate to a permanent contract, in quarterly terms, with and without the wage subsidy. This quarterly transition rate is computed at the mean of the explanatory variables and is measured during the second quarter of the temporary contract. This table shows that since the transition rates to a permanent contract at the same firm are fairly low, the increase in the employee outflow to a permanent contract at the same firm generated by these changes is quite small. For instance, the quarterly estimated transition probability rises from 0.65% to 1.09% for middle-aged women, at the second quarter of the temporary contract.

Since this low sensibility of the flow to a new permanent contract 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 duration longer than three months. Though, there are not relevant differences, we find interesting to point out that the effects of wage subsidies over the outflow to a permanent contract at the same firm are slightly smaller. Specifically, the total change in the exit probability to a permanent contract caused by regional wage subsidies is around 46% for middle-aged women. Interestingly, for the case of young workers the incidence of regional wage subsidies slightly varies from the previous estimation. In the case of young female workers we obtain that regional wage subsidies increase the exit probability to unemployment and decrease the probability of renewal the temporary contract, for eligible workers. We obtain that the first probability increases by 6% while the second one decreases by 12%, for young female eligible workers. For young male workers we obtain that the probability of renewal the temporary contract increases for the eligible ones by around 9%. The incidence of regional wage subsidies for the rest of workers is not statistically significant. Overall, thus, we cannot conclude that the low sensibility of the flow to a new permanent contract to regional wage subsidies is due to the temporary nature of the jobs covered by these contracts.

Finally, we have 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, hence, we do not find evidence of relevant differences between eligible workers in private or public firms. For instance, the total change in the exit probability to a permanent contract caused by regional wage subsidies is around 45% for middle-aged women and as before, the effect for the rest of workers is not statistically significant.

5.2 Effects of the regional wage subsidies over the transition rate from unemployment to a permanent contract

In the case of exiting from unemployment, some of the explanatory variables used for the estimation differ slightly from those explained before. Specifically, we add two variables that identify whether the worker (1) is eligible for unemployment benefits and (2) receives them in each month analyzed. We also consider here whether she was pre-

viously hired by a Temporary Employment Agency and whether she returns to a firm for which she has previously worked or begins with a new one.

TABLE 9
Main results by gender and age of the policy variable
(exit from unemployment)

			To T. C.	To P. C.	Stayed Unemployed	
Women	<30	Odd Ratio	14.69%	-	9.05%	
		Total Change in Probability	-4.46%	10.00%	0.48%	
		Significance Test	(5.28) **	(10.09) **		
	30-45	Odd Ratio	7.98%	-	6.51%	
		Total Change in Probability	-1.31%	6.25%	0.05%	
		Significance Test	(1.73)	(1.24)		
>45	Odd Ratio	-8.72%	-	-8.14%		
	Total Change in Probability	0.79%	-9.80%	0.15%		
	Significance Test	(5.51) *	(1.71)			
Men	<30	Odd Ratio	5.01%	-	3.29%	
		Total Change in Probability	-1.43%	3.51%	0.22%	
		Significance Test	(8.30) **	(5.34) *		
	30-45	Odd Ratio	-1.01%	-	4.05%	
		Total Change in Probability	4.08%	3.02%	-0.99%	
		Significance Test	(4.29)	(1.93)		
>45	Odd Ratio	1.04%	-	-5.31%		
	Total Change in Probability	-5.21%	-4.23%	1.15%		
	Significance Test	(3.70)	(0.46)			

Note:

TC = Temporary Contract; PC = Permanent Contract

The Odd Ratio measures the change in the ratio of the probability of exiting to a permanent contract for eligible workers versus non-eligible ones, relative to the same ratio -eligible versus non eligible-, for the available alternatives, when the wage subsidy takes the average sample value (4,800 Euros).

The total change in probability is measured at the average wage subsidy and at $t=3$.

The significance test shown is the result of the Wald Test of the joint statistical significance of the level and quadratic effect of the policy variable: statistically significant at 90% (*) and at 95% (**).

In Table 9, we display the predicted treatment effects by age and gender. The first result emerging from this table is that the impact of wage subsidies over the transition to a permanent contract is lower for unemployed workers than it is for temporary ones. In addition, we find that the largest incidence of wage subsidies over the outflow to a permanent contract occurs for younger workers: the odd ratios of exiting to a permanent contract relative to a temporary one are positive for younger eligible workers, both males and females, while for the rest of cases the odd ratios are quantitatively irrelevant or the coefficients are not statistically significant. This odd ratio for young

workers is 14.69% for females and 5.01% for males. In fact, for these two worker groups the exit to a temporary contract for eligible workers decreases, by about 4.46% for women and 1.43% for men, while the exit to a permanent contract increases by 10.00% for women and 3.51% for men. Hence, as it is shown in Table 10, regional wage subsidies are causing the exit from unemployment to a permanent contract to increase from 7.90% to 8.64% for young female workers and from 7.32% to 7.57% for the young male ones.

In sum, the impact of regional wage subsidies on the outflow to a permanent contract from unemployment is low. While younger workers benefit the most from such funds, the policy has no effect for male workers over the age of 30 and just marginally favours the exit from unemployment to a temporary contract for older women.

TABLE 10
 Expected growth rate of the entrance probability to a permanent contract from unemployment as a result of the policy
 (Quarterly transition rates)

		Estimated Transition Rate to PC (1 st Quarter)		
		With RWS	Without RWS	% Change
Women	<30	8.64%	7.90%	9.40%
	30-45	8.07%	7.59%	6.39%
	>45	7.44%	8.07%	-7.82%
Men	<30	7.57%	7.32%	3.44%
	30-45	7.88%	7.67%	2.73%
	>45	8.06%	8.38%	-3.73%

Note:

We display all the estimated transition rates but to correctly interpret these results, one must take into account the statistical significance of these estimations showed in Table 9.

To check the robustness of our results, we perform two additional estimations, one restricted to workers who had previously held temporary contracts and other restricted to workers who had previously held permanent contracts. In the first case, the results hardly vary from the ones presented. For instance, for eligible young female women, the probability of entering into a permanent contract increases by 9.79% due to regional wage subsidies. On the contrary, when we focus on the second case, some differences emerge that are worth to mention. Now the incidence of regional wage subsidies are stronger for young male workers since the probability of entering into a permanent contract for eligible workers increases by 29.32%. Meanwhile, the probability of existing from unemployment into temporary contract decreases by

25.62% and the outflow to a permanent contract increases by 6.77% for eligible young women.³²

6. Conclusions

Spain has one of the highest rates of temporary contracts in all of Europe, a situation that may redound negatively on the national economy by decreasing its efficiency and undermining equity conditions. As a preventative measure, the national and some regional governments have designed different policies geared towards the creation of permanent jobs. These measures include the national initiative to cut payroll taxes on new permanent contracts and the many regional wage subsidies for new permanent hires that have sprung up around the country since 1997, and which target a number of different worker groups.

In this paper, we have estimated the incidence of regional wage subsidies over the transition from a temporary to a permanent job at the same firm and from unemployment to a permanent job. In both cases, we have used a competing risk model and we followed a Difference-in-Difference-in-Difference (DDD) approach. Our analysis makes use of the regional and temporal variability in regional wage subsidies (similarly to what is usually done in the US), and the differences in individual eligibility criteria, as tools that can help identify the treatment effect. In order to gain homogeneity between both groups in our sample, we also divide it into gender groups and measure the effect of the policy on workers from different age groups.

Our main results suggest that subsidizing new permanent contracts has a positive but small effect on the flow to a permanent job, and only for certain eligible worker groups. At any rate, a number of differences in our results associated with the initial job status, gender and age are worth mentioning here. Middle-aged female workers, for example, appear to derive the greatest benefit from wage subsidy schemes, followed by young ones. Nevertheless, the quantitative effects of these wage subsidies over the entrance probability to a permanent contract are fairly low. The intra-firm conversion rate for eligible workers moving from a temporary to a permanent contract increases by about 67%

³²These results could be interpreted as evidence of some churning effect, in the sense of some permanent workers being fired and re-hired after being unemployed in order to get the subsidy again. This might be tested by studying in more detail the transitions of permanent workers and to the effects of the policy over such transitions.

for middle-aged female workers (from 0.65% to 1.09% in the second quarter of the temporary contract). The probability of exiting from unemployment to permanent employment increases only by 10% for younger female workers and by about 4% for younger male ones.

Wage subsidies to new permanent contracts work basically in two ways. First, they provide firms that offer permanent contracts to eligible individuals with funding incentives that may enhance the worker's probability of accessing to that position. Second, they decrease the likelihood that certain non-eligible individuals will obtain such a position. Therefore, one could claim that substitution effects between eligible and non-eligible workers might explain the low change in the permanent employment rate observed during the period analyzed. Nevertheless, given the results presented, substitution effects between eligible and non-eligible temporary or unemployed workers seem not to be the reason since we do not find strong effects on the inflow to a permanent contract for eligible workers and on the outflow to unemployment or to a temporary contract for the non-eligible ones. Our results seem to match the predictions of recent theoretical models (see Dolado, Garcia-Serrano and Jimeno, 2002): wage subsidies only serve to temporarily reduce hiring costs and they do not help close the significant existing gap between the overall labour costs for employees holding temporary and permanent contracts.

Appendix

TABLE A.1
Main results for the exit rate from a temporary contract

	Men						Women					
	Unemployment		T.C Same Firm		P. C. Same Firm		Unemployment		T.C Same Firm		P. C. Same Firm	
	Coef.	t-S	Coef.	t-S	Coef.	t-S	Coef.	t-S	Coef.	t-S	Coef.	t-S
Dijt*(Age<30)	-0.005	-0.68	0.024	1.39	-0.086	-0.90	0.006	0.58	0.013	0.41	0.029	0.39
Dijt*(Age<30)^2	0.000	0.70	-0.001	-0.77	0.004	0.59	-0.001	-1.21	-0.004	-1.55	-0.002	-0.29
Dijt*(Age 30-45)	0.010	1.26	0.012	0.54	-0.025	-0.27	0.008	0.87	-0.011	-0.37	0.116	1.20
Dijt*(Age 30-45)^2	-0.001	-1.45	0.000	0.02	-0.001	-0.14	-0.001	-1.94	-0.001	-0.66	-0.002	-0.27
Dijt*(Age >45)	0.018	1.56	0.024	0.86	0.107	0.90	0.003	0.27	-0.039	-0.73	-0.120	-1.03
Dijt*(Age >45)^2	-0.001	-1.33	-0.001	-0.51	-0.007	-0.77	-0.001	-0.58	0.002	0.39	0.012	1.32

Notes:

The estimation includes all the variables described in Section 5 concerning the sample of temporary workers.

Standard Errors are adjusted for the cluster groups defined in terms of our policy variable.

To estimate the model, the policy variable is expressed in thousand Euros.

TABLE A.2
Main results for the exit rate from unemployment

	Men				Women			
	T.C		P.C		T.C		P.C	
	Coef.	t-s	Coef.	t-s	Coef.	t-s	Coef.	t-s
Dijt*(Age<30)	0.000	-0.03	0.010	1.56	-0.015	-2.06	0.033	1.96
Dijt*(Age<30)^2	-0.001	-1.18	-0.001	-0.49	0.001	1.85	-0.003	-2.56
Dijt*(Age 30-45)	0.014	1.08	0.017	0.66	-0.002	-0.30	0.019	1.05
Dijt*(Age 30-45)^2	-0.001	-0.44	-0.002	-1.01	0.000	-0.23	-0.001	-0.85
Dijt*(Age >45)	-0.020	-1.92	-0.019	-0.57	-0.004	-0.29	-0.022	-0.73
Dijt*(Age >45)^2	0.001	1.85	0.002	0.66	0.001	1.06	0.001	0.39

Note:

The estimation includes all the variables described in Section 5 concerning the sample of unemployed Standard Errors are adjusted for the cluster groups defined in terms of our policy variable.

To estimate the model, the policy variable is expressed in thousand Euros.

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Resumen

Este artículo evalúa la efectividad de los subsidios regionales a la creación de empleo indefinido y para ello hace uso de la información contenida en la Muestra Continua de Vidas Laborales. Esta base de datos, usada aquí por primera vez para hacer evaluación de políticas sociolaborales, ofrece el historial laboral completo de más de un millón de trabajadores españoles, si problemas de agregación ni de censuras. La política que analizamos consiste en un subsidio pagado al inicio del contrato indefinido y que se ofrece por algunas comunidades autónomas para los nuevos contratos indefinidos firmados con trabajadores temporales y con los desempleados. Debido a que la política tiene variabilidad temporal, individual y regional, podemos aplicar una técnica de triples diferencias que identifica el efecto medio del tratamiento de dicha política. Nuestra principal conclusión es que el flujo a empleos indefinidos de los trabajadores elegibles por esta política cambia solo mínimamente. El efecto, no obstante, es mayor para trabajadores temporales que para desempleados y es mayor para jóvenes y mujeres de entre 30 y 45 años.

Palabras clave: Análisis de causalidad, subvenciones salariales, subvenciones a la contratación, modelos de duración.

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