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

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Abstract

We present an analysis of the use of permanent employment contracts across Spanish regions during the period 1995-2002. This paper concerns the evaluation of regional wage subsidies designed to foster the creation of permanent jobs in the Spanish economy since 1997. Since we have longitudinal data we apply a difference-in-differences estimator to identify de potential average effect of the treatment for this policy. We estimate the incidence of this policy over the permanent employment rate and over the transition rate to a permanent employment using the panel data dimension of the Spanish Labor Force Survey and Social Security Records, respectively. Our main results are that this policy shows a small positive effect over the rate of permanent contracts, being this due to a highly decreasing pattern showed by the treatment effect over the duration of the policy implementation. The effect is larger for females than for males and is the largest for young workers. With respect to the effect of these policies over the transition rates, we find that the direct transition from temporary to permanent employment is almost unaffected but the one from unemployment to permanent employment is considerably larger, specially for females and also for young workers.

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1 Introduction

During the 1990s, Spain has faced several labor market reforms aimed at reducing the presence of fixed-term contracts among employees. However, the use of these contracts has only experienced a small decrease (from 33.6% in 1996 to 32.1% in 2002) and Spain continues being the European country with the highest rate of temporary or fixed-term contracts in Europe. When the labor market is characterized by excess of supply, the use of temporary employment must be related to the demand side of the market. Hence, the current extensive use of fixed-term contracts is basically due to the firm's desire of reducing total labor costs. Given a permanent contract is associated with larger firing costs, it seems Spanish firms use temporary contracts to both reduce labor costs and to face economic uncertainty. Furthermore, García-Pérez and Rebollo (2005) show how the behavior of unit labor costs is an important determinant in the use of temporary contracts across Spanish regions, especially in those where the level of education of the labor force is lower and the share of small firms is larger. During the period analyzed, Spanish unit labor costs have permanently increased at an annual average rate of 1.3%. This certainly means that labor productivity grew at a lower rate than total labor compensation per employee harming the chances to increase permanent employment during the upturn of the 1995-2002 period.

During the second half of the 1970s and the first years of the 1980s, Spain had one of the tightest labor markets in all Europe 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 implemented at that time is a paradigmatic example of what has been called two-tier selective labor market policies. Broadly speaking, the reform of 1984 consisted of introducing the possibility of hiring workers on flexible, fixed-duration contracts. The objective was to foster job creation in order to reduce the already high rate of unemployment. As a result of this reform, temporary contracts rose from 18% in 1987 to 33% in 1994 and this rapid increase positioned Spain as the European country with the highest temporality rate. In 1994, 1997 and 2001 there were new changes in the regulation of the labor market aimed at reducing the scope for using fixed-term contracts by reducing the firing costs for new permanent employees[1]. However these institutional reforms have hardly decreased the use of fixed-term contracts as the temporary employment rate was still almost three times higher than the European average.

Together with these institutional reforms, since 1997, the Spanish government has subsidized the creation of permanent contracts by, in some cases, large discounts of firm's payroll taxes. Hence, each month the firm obtains a reduction in total labor costs for these new permanent employees. Additionally, during the same period, some regions decided to implement wage subsidies for new permanent contracts. These subsidies are paid only at the beginning of each new permanent contract and, in some regions, they also depend on some eligibility conditions

[1] In 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.

based on individual characteristics, basically age and gender, while in others they affect to any unemployed or temporary worker. In this context, Spain is the OECD country with the highest percentage of GDP, 0.28% in the period 1999-2002, used to subsidy regular 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.

The main objective of this paper is to evaluate the effectiveness of these regional wage subsidies to foster permanent employment. One of the main features of this research is that we take advantage from the eligibility rules of these subsidies that vary across individuals, regions and time in order to identify the policy effects. Besides, the variation of the permanent employment rate among Spanish regions is somehow large: 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.

Two main empirical approaches are applied. Firstly, we examine the incidence of regional wages subsidies over the permanent employment rate. Secondly, we study the effect of such regional wage subsidies over the transition rate towards permanent employment. For the first exercise, we will use the Spanish Labor Force Survey for the period 1995-2002, complemented with Regional Accounting data for computing some aggregate regional information. For the second exercise we use Social Security records coming from the new “Muestra Continua de Vidas Laborales” for the period 1995-2004. 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 the policy is the difference-in-differences estimator (DID, hereafter). Furthermore, we will control for the possible incidence of unobserved heterogeneity in our estimations.

Our main result is that these regional wage subsidies increase the probability of having a permanent contract only during the first three years the policy is implemented: the maximum effect is in the first year when this probability increases by 1.9 percentage points (more for females and young workers). Therefore we find evidence of a program introduction effect. Furthermore, this effect is increasing for subsidies lower than 8.000€. The predicted elasticity (of the permanent employment rate to the subsidy in thousands Euros) is, however, very low, +0.018. The effect of the policy is larger for women and especially for young workers, an effect already found in a previous paper that analyses the 1997 labor market reform in Spain (Kugler, Jimeno and Hernanz, 2003).

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 the third section we describe the data used and the main characteristics of the regional wage subsidies. Sections fourth and fifth describe the econometric approaches used and the main results. Finally, section six presents our preliminary conclusions.

2 Background

The introduction of fixed-term contracts in Spain was based on the argument that by reducing firing costs, total labor costs could decline and subsequently they would encourage job creation. However, the experience has shown that it cannot be taken for granted that the liberalization of the labor market through the use of temporary contracts improves the functioning of the labor market, so far. Some previous works have point out (Bentolila and Dolado, 1994; Saint Paul, 1996) that the role of fixed-term contracts seems to be unclear in terms of fostering employment creation and promoting regular employment. It is generally concluded that the introduction of this type of contracts is equivalent to a reduction in firing costs[2] and its impact on unemployment is therefore ambiguous. As Bentolila and Saint Paul (1992) show, using a model specially designed to analyze the macroeconomic implications of fixed-term employment, this type of contract has mainly contributed to increase the cyclical elasticity of employment. Among more recent theoretical models we find the work of Cahuc (2001), who presents a model that analyses this issue but with the novelty that firms can create both permanent and temporary jobs and they may convert a certain share of the latter to permanent contracts at their expiration. He concludes “a more flexible regulation on fixed-term contracts may actually destroy jobs, increase unemployment and reduce aggregate welfare, especially when firing costs are high”. The intuition of this result is that the higher the firing costs of permanent contracts, the lower the share of temporary jobs transformed into permanent jobs, because large firing costs are an incentive for the employer to use temporary jobs in sequence rather than converting them to long-term contracts, which are subject to firing costs. As a consequence, the use of temporary contracts is more likely to raise unemployment and labor turnover when it comes on a labor market already regulated by stringent permanent job security provision.

Other branch of the literature has focused on the study of the determinants of the equilibrium rate of permanent to temporary employment. This point is important to evaluate if the current rate of temporary employment is mainly driven by short terms factors or long terms ones and evidently must be considered for any policy aimed at reducing the use of temporary contracts. One important result of this literature is the correlation of the permanent employment rate with labor productivity and/or total labor costs. To illustrate the relation between labor productivity and the rate of permanent employment, dynamic models of labor demand seems to be a good choice. For instance, Dolado, Garcia-Serrano and Jimeno (2002) use a basic dynamic model to argue that the equilibrium ratio of temporary to permanent employees is determined by the ratio among unit labor costs[3] 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 growth rate. Wasmer (1999) extends the matching framework of Pissarides’ Equilibrium Unemployment Theory and proves that macroeconomic factors such as productivity growth and labor force growth have an impact on the relative demand for

[2] After the reform of 1997, fixed term contracts terminate at a lower costs as the severance payments is 12 days for year worked while for indefinite contracts is 45 or 20 days for year worked, depending on the type of layoff.
[3] They distinguish three components within the concept of labour costs: the wage, the firing cost and the hiring costs.

temporary contracts, beyond the need for flexibility from workers at the microeconomic level. He shows that firms will face a trade-off between paying high turnover costs and having stable workers, or paying low turnover costs but being more frequently engaged in a search process. In this context, higher productivity means higher expected profits, which induces further hiring today, since hiring costs indexed on productivity growth are lower if paid today. This is called the *capitalization effect* of growth and implies that when productivity growth is high, firms want to retain workers by offering them long-term contracts. This author evaluates the relative importance of productivity growth and shows that when this growth rate declines from 4% to 0% the share of short-term jobs increases from 0.5% to 10%. Following a similar argument, Holmlund and Storrie (2002) conclude that the rapid growth rate of temporary employment in Sweden is more related to adverse macroeconomic shocks than to institutional labor reforms.

In this framework we also want to point out the studies of Blanchard and Landier (2002) and Kugler, Jimeno and Hernanz (2003) since they analyze the effectiveness of a labor market reform consisting in a reduction of firing costs to reduce the share of temporary workers. Interestingly, Blanchard and Landier (2002) show that in countries with little employment protection, such as United States and United Kingdom, the proportion of the workforce on fixed term contracts has been relatively low and fairly stable while in countries characterized by high levels of employment protection, such as Spain, France and Italy, the proportion of temporary workers has doubled during the 1990s. Kugler, Jimeno and Hernanz (2003) develop a simple dynamic matching model similar to Blanchard and Landier (2002) but they endogenize dismissals and introduce payroll taxes, in order to analyze the Spanish labor market reform of 1997. 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 and given the values of the two productivity thresholds they can derive the steady-state values of temporary and permanent employment[4]. 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 firing costs for permanent contracts increases the rate of conversions from temporary to permanent contracts and reduces wage differentials. Their empirical results suggest that the reform increased permanent employment probabilities for young relative to middle-aged workers. They also show increases in the relative transitions from non-employment to permanent employment for young and older men and from temporary to permanent employment for young men and women after the reform. The reason why this reform mainly affected to young workers is that the reduction in dismissal costs and payroll taxes increased both hiring and dismissals for older men, though had a positive effect on the hiring margin of young workers with little effect on dismissals. They estimate the elasticity of permanent 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.

[4] As far as we know only the paper of Kugler, Jimeno and Hernanz (2003) -theoretically and empirically- directly relates the rate of permanent employment to labour costs when they examine the effects of the Spanish reform of 1997 on permanent employment.

Summarizing, the achievement of labor market flexibility through the expansion of fixed-term contracts has not been as successful as was initially expected, since it has also brought about efficiency and equity costs. Theoretical models show that labor productivity and total labor costs influence the equilibrium rate of permanent employment. Hence, from this theoretical model one would conclude that short-term labor market policies will not be as effective as long term ones, specially when the high rate of temporary employment is considered to be related to equilibrium variables. From this perspective, temporary labor market policies such as wages subsidies on new permanent contracts will only be partially effective in lowering the high rate of temporary employment.

3 Data Description

3.1 Permanent Employment Rate

We will use two different data sources in our first empirical exercise. One offers information on individual data and the second one on aggregate data at the regional and sector level. They are the Spanish Labor Force Survey (*Encuesta de Población Activa, EPA* hereafter) and the Regional Accounting Dataset (*Contabilidad Regional de España, CRE* hereafter*)*. These databases offer us the two levels of information we need in our study, individual and region-sector data. The *EPA* has information on personal and labor characteristics of the individuals while from the *CRE* we take information on production characteristics of the sector and location where the individual works.

Our sample includes individuals that were surveyed between 1995 and 2002, covering more than a full cycle of the Spanish economy. We select non-farm employed workers between 16 and 65 years old and we exclude the self-employed. Given that the *EPA* has a rotating structure that follows individuals for a maximum of six quarters, replacing one-sixth of the sample every quarter, our sample has a panel structure.

The *CRE* has annual information on production, employment and total labor costs by sector and region. From this database we take the value added at constant prices and we construct indicators for the regional sector specialization based on eight non-farm productive sectors: *Energy, Manufacturing Industry, Construction, Commerce and Hotels, Transport and Communications, Financial Services, Professional Services and Other Services*[5].

We start the descriptive analysis by looking at the pattern followed by the percentage of permanent contracts for Spain during the period 1995-2002 using our estimation sample. Table 1 displays the rate of permanent employment for each region in that period.[6] We observe that on average terms, the period analyzed is characterized by an increase in the rate of permanent

[5] Within this label non-market services are included.
[6] All figures in the following tables are based on sample rates. Thus, they are not representing population rates, although the difference between them is quite insignificant.

employment[7] of around 4 percentage points and, despite the institutional labor reforms of the 1990s and the economic upturn in 2001, we still had a low rate of permanent employment compared to the European average in that year. To illustrate the regional dispersion, Figure 1 shows the percentage of permanent contracts for each region in the years 1995 and 2002. Those regions located over the black line experienced a positive growth in permanent employment while those located under the black line experienced a negative change. We observe that except Andalusia, Extremadura, Galicia and Asturias, the rest of regions have experienced this slight growth in the rate of permanent employment. We can see also that southern and eastern regions have the lowest rate of permanent employment. Within this group of regions we can distinguish those that have an increasing trend in this rate (Castilla-La Mancha, Valencia, Murcia and Canarias) from those whose rate has even declined (Andalucía, Extremadura and Galicia). On the opposite side we find that northern regions and Madrid have the highest rate of permanent employment. The exceptions are the Basque Country and Galicia since in both regions the rate of permanent employment is slightly lower than the national average.

Though the institutional labor reforms of the 1980s and 1990s have been the same for all the regions, the response in terms of the rate of permanent employment has been different, and the national rate of permanent employment hides regions with low and high rates of permanent employment. These differences may be due to divergences in the steady state composition of employment by type of contract across regions. For this reason, we also display the probability of having a permanent contract by different personal and labor characteristics, since differences in the composition of the workforce may be behind the observed dispersion in the rate of permanent employment across regions. In Table 2 we show the average of the permanent employment rate by some personal characteristics for the period 1995-2002, and in Table 3 by firms' characteristics for each region. In general terms, we observe that the distribution of temporary employment varies among types of workers, but we do not identify clear differences across regions. It appears that permanent employment is less prevalent among women, less educated workers and youths. Especially relevant is the difference in the rate of permanent employment by age: the highest rate of temporality is found for young workers; their percentage of permanent contracts is well below 40% for all regions except Madrid. This type of contract is also less intensively used in private and small size firms and in low skill occupations. These descriptive statistics, as stated in previous empirical studies (Bentolila and Dolado, 1994), show that the increasing employment flexibility has been achieved at the margin, that is, flexibility especially affects certain groups of workers, new entry workers with low educational attainments, in small private firms, but not the core of *permanent employees*.

[7] During this period the *EPA* has suffered several methodological changes, mainly in 1995, 1999 and 2000, which may partially affect the data. Among other things, these changes implied that permanent employment is now better described and therefore, the rate of temporality could artificially decrease. Toharia (2002) points that when one uses a homogenous series the rate of temporality remains almost unchanged during this period. Nevertheless, these methodological changes should not affect interregional differences and consequently, neither the results we present in this chapter.

3.2 The Transition Rate to Permanent Employment

We use the Social Security records coming from the recently available "Muestra Continua de Vidas Laborales" (MCVL) for the period 1995-2004 to study the transition rates towards permanent employment. This data base has not been used previously to analyze the temporality issue in the Spanish economy although the Social Security records offer detailed information on the labour market history of the individual. Other databases only gather annual information and therefore individual labour market transitions along a year are not available giving rise to an under-representation of short-term temporary contracts.

This database gathers much information about the worker labour market trajectory and individual characteristics such as age, gender, occupation, unemployment and employment spells and their durations. We have available the reasons to the end of the contract, the location, the sector of activity, type of contract and whether the contract was signed with a temporary help agency for each spell of employment. The duration of the unemployment and employment spells are built from the dates of the hiring and the end of the contract and it is measured in months. In the analysis we also consider several aggregate variables at the regional and national level to control for the labour market situation and the business cycle. Basically we use the growth rate of the domestic product and the regional unemployment rate.

The definition of a temporary contract is a relevant issue. Guell and Petrolongo (2004) and Booth, Francesconni and Frank (2001, 2002) use a broad definition of a temporary contract[8]. We follow the same approach and within the concept of temporary contract we include the following categories: fixed term, specific task, training, contract for circumstances of production, internship contract and replacement.

We start off displaying in Table 4 the distribution of new contracts at different points in time. The first remarkably fact is the high share of new temporary contracts, around 80%, along the analysed period though in 1997 and 2001 important labour market reforms aimed at the reduction of temporary contracts, among other things, took place in Spain.

We display in Table 5 the different options a worker faces after the temporary contract. The worker moves to unemployment in approximately 50% of the cases, whereas only in approximately 7% of the cases he improves its position and obtains a permanent contract. Workers move again to a temporary contract, either in the same or in another firm in the 41% of the transitions. As before, this result seems enough stable along the period 1995-2004 and they show that the probability of getting a permanent contract from a temporary position is pretty low in Spain. We observe greater variability in the probability of changing firm, either with a temporary or a permanent contract. The data show that the probability of connecting temporary contracts and the transition from temporary to permanent contracts within the same firm have tended to increase. Thus, whereas in 1996, around 15% and 24% of new contracts were signed

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

in the same company, in the year 2003 these percentages ascended to 30% and 39% respectively.

We show in Table 6 the relation between the duration of the temporary contract and the type of transition. We obtain that short term contracts tend to end into unemployment while longer term ones end into permanent contracts. Moreover, we can observe that the average duration of temporary contracts is pretty short since 53% of those contracts lasted three months and 22% between three and six months.

We display in Figure 2 the exit rates from a temporary contract to each of the destination states considered. Firstly, we observe negative duration dependence when the destinations are the unemployment and other temporal contract. Interestingly, the exit rate to a permanent contract hardly varies except at certain durations. Interestingly it reaches a local maximum at duration 36. In Figures 3 and 4 we display the exit rate from unemployment both for males and females. We can also observe the clear distinction among workers with and without Unemployment Benefits.

3.3 Regional Wage Subsidies

Wage subsidies on new permanent contracts are 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 and it has remained constant, wage subsidies have varied between regions and along time during the period 1997-2002.

These wage subsidies show regional and time variation in eligibility rules mainly based on the individual characteristics of the worker such as age and gender. This fact allows us to examine the response of different group of individuals to the change in incentives. For example, some regions offer these subsidies only for women whereas others restrict their use just for young workers. In certain cases they also depend on the characteristics of the firm such as the type of activity or the labor state of the worker. For instance, in some regions the wage subsidies are addressed for workers in temporary contracts while in others they are also extended to unemployed workers.

The data is taken from each regional government and it only covers wage subsidies at the regional level[9]. The main characteristics of these wage subsidies are described in Tables 0 and 12. As it is shown in Table 0, this policy was implemented since 1997 in some regions, whereas in others it was implemented afterwards or never, as in Cataluña. Table 12 also shows that the 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.

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

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. Given the way the data is collected, we can not observe who has effectively been benefited from the policy. Therefore, our treatment groups are those workers who fulfill the eligibility criteria. 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 when the eligibility conditions are fulfilled.

Besides we are also interested in measuring whether the effectiveness of these wage subsidies depends on the amount of the subsidy. Again, we can observe that this strongly varies between personal characteristics and regions. Therefore we will use this additional source of variability to identify the average effect of the policy. 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. 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.

The average subsidy for each region, in 2002 Euros, is shown in Table 14, with their minimum and maximum amount. This ratio represents around 9% of the regional gross annual wage in Baleares, the region with the lowest subsidy, and more than 60% in Extremadura, one of the regions with less permanent contracts. In Table 14 we display the subsidy by age and gender. The subsidy is clearly larger for women and for older workers.

4 The Empirical Approach: Identification and Estimation Methods

Our econometric approach is directly linked to the standard causal effects analysis. In particular, we follow a Difference-In-Differences (DID, hereafter) approach[10]. 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). In fact much of the debate around the validity of a DID estimate typically revolves around the possible endogeneity of the interventions themselves (See Besley and Case, 2000)[11]. The individual, regional and time variability of the policy discussed in the previous section 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 in the same region as control groups. Moreover, in this paper, as it is common

[10] Recall that the aim of the DID approach is to compare outcomes between groups before and after the treatment. The treatment group is composed of those potentially affected by the policy. We must also select a control group, either from other regions or a different group of individuals. Good control groups will be those whose behavior 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.

[11] 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.

in many panel data studies[12], we will include a regional fixed effect to control for permanent differences across regions in policies and outcomes[13].

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 overtime[14]. Therefore, to provide an unbiased estimate of the treatment effect, it must be the case that either time varying regional 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 Furthermore, there must be no time-varying regional level unobserved variables determining outcomes that vary, pre- and post-treatment, between regions. 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. Other factors that could induce differential time trends relate to the slight differences either in region or individual characteristics of the groups to be compared. 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.

4.1 The Rate of Permanent Employment: Econometric Approach

We will analyze the individual probability of being employed under a permanent contract over the period 1995-2002. We estimate a binary choice panel data model with the individual data taken from the *EPA* sample previously described and controlling for a large set of regressors dealing with individual, firm and regional heterogeneity. Moreover, we will study the treatment effect of our policy by controlling for the three dimensions of variability in our policy measure: eligibility conditions, based on individual characteristics, time and region where the worker lives.

The main issue of this section is to measure the incidence of regional wage subsidies over the permanent employment rate. Nevertheless, we are also concerned with the 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 its probability of having a permanent contract. Second, some of the individuals who are not eligible might faced a drop in this probability. The extent to which this may happen will depend on a number of factors. If the wage subsidy just covers the deficit in productivity we would not expect any substitution effect. The eligible workers are no cheaper than anyone else. Second, it will depend on the extend that these workers are substitutable in production for

[12] See Anderson and Meyer (1997), Gruber and Madrian (1997).
[13] If the systematic determinants of state policies are additive, time invariant regional characteristics, then will indeed remove concerns about endogeneity.
[14] These assumptions are discussed in detail in Blundell and MaCurdy (1999).

existing workers 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 is an important point if 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 the non-subsidized ones, without extra effort. For instance, Cebrian and Toharia (2005) offer some evidence that this might be happened. Finally, we also take care to judge how permanent the effects are likely to be and we measure if the impact of the policy depends on its duration.

There are a number of additional issues that we need to address. First, the definition of the comparison group is, of course, central to the evaluation. There is the basic issue of whether we can assume that the two groups are subject to the same aggregate labor market trends. Evidently, this assumption is more plausible when we split the sample by age to the extent that the human capital of the two groups is similar and also preferences for work should be the same. 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 also consider estimating the average treatment effect splitting the whole sample into subsamples by gender or age. 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 the control group is decreasing. To avoid this type of bias, we will include in the model a regional time varying dummy variable that takes value one when the worker lives in a region at the time the policy is implemented.

The next important issue is whether the impact of the policy is heterogeneous with respect to observable characteristics. If this is the case, we should interpret the estimate we obtain as an average impact across different effects. This procedure can also be restrictive since we do not allow for the coefficients in our model to depend on observed heterogeneity. Nevertheless, this parametric assumption can be easily relaxed under the parametric setting. The effect of treatment could be allowed to depend on the observable characteristics of the agents. Though we are working on these issues, the results presented are obtained assuming homogeneous treatment effects between individuals. Again, this assumption is too restrictive for the whole sample but more suitable when we split it by age and/or gender.

The individual probability of having a permanent contract should respond to changes in regional economic and individual characteristics. We assume that there is an underlying response variable defined by the following linear regression relationship:

$$\mathrm{T}^*_{ijt} = \alpha' x_{ijt} + \gamma' z_{jt} + \beta_0 D_{ijt} + \beta_1 D_{jt} + \beta_2 D_{ij} + \beta_3 D_j + \beta_4 D_{ijt} * Dur_{ijt} + \varepsilon_{ijt} \qquad (1)$$

where i stands for individuals, j for regions and t for years; 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[15]. 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. Using covariates we extend identification to those instances in which observed compositional differences between treated and controls cause non-parallel dynamics in the outcome variable[16]. We also consider regional fixed effects to control for permanent differences across regions in outcomes.

The policy variable is D_{ijt} and takes value one 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. The rest of "*D*" variables help to identify an unbiased estimate of the average treatment effect of the policy. The variable D_j takes value one when the worker is located in a treated region[17]. The variable D_{ij} takes value one when the individual belongs to the eligible group. The variable D_{jt} takes value one when the worker is in a region that applies the wage subsidy at time t. Summarizing, the variables D_j and D_{ij} control for permanent differences between treated regions and eligible individuals respectively whereas, the variable D_{jt} controls for time varying regional effects and it helps to identify the existence of general equilibrium effects. This, together with the fixed effects and the regional variability in the eligibility conditions, might remove concerns above the biases related to the endogeneity of the policy.

Finally, we also care about how permanent the effects are likely to be and we multiply the policy variable D_{ijt} with its duration. The idea behind this specification is to test whether there is evidence of an important "program introduction effect" in the sense that the impact of the program is much larger in the first year than in subsequent ones. That is, wage subsidies might encourage the conversion of temporary contracts into permanent ones the first year the policy is implemented and have negligible effects afterwards.

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. The estimate of policy incidence on the treatment group, β_0, is estimated as the post-treatment change in 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

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[15] Besley and Case (2000) shows that fixed effects models might also mislead the effect of the policy. In this context 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 play an important role in this study.

[16] Under the *Conditional Independence Assumption* the selection of individuals is supposed to be made on observable characteristics and thus, conditioning on those variables, the potential outcomes and the treatment status are independent. Besides the DID also allows selection on time invariant unobservables.

[17] Formally, the analysis should also consider the variable "*never treated*", that is D_i but, given the characteristics of the policy we are analysing, this variable matches with D_{ij}. Nevertheless, we include in x_{ijt} controls for the variables determining the eligibility conditions, basically age and gender).

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[18]. The parameter β_3 shows the existence of aggregate effects from the policy. If there is no general equilibrium effect we expect it to be non significant.

Finally ε_{ijt} is the error term whose composition is the following:

$$\varepsilon_{ijt} = \eta_i + \eta_j + \eta_t + \upsilon_{ijt} \qquad (2)$$

where η_j and η_j describes the individual and regional time-invariant effect, η_t the time effect 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 region effects. Recall that one advantage of the DID approach is that it controls for unobserved time-invariant differences.

The sample used in the estimation includes all individuals aged 18-64 years who were employed in the non-farm sector, excluding those always unemployed and out of the labor force. The dependent variable is the type of contract for each worker and takes value one if the worker has a job with a permanent contract and zero otherwise. Our specification also includes several personal and labor characteristics such as *age*, *gender*, *educational attainment*, and dummies for *occupation*, *firm size*, *seasonal job*, *full time job, sector of activity, size of the sector in terms of employment* and *private firm's ownership*. Besides, we include time and regional dummy variables such as *share of small firms* and *relative sector specialization*. The year effects capture the impact of macro shocks affecting the probability of getting a permanent job.

One crucial aspect in our analysis is that we have no information regarding a specific individual in our sample being treated or not. What we have is whether she is covered by this policy or not, that is, we will estimate the effect of being potentially treated by the policy.

Given we have not only the information about whether the policy has been implemented or not but also the exact maximum amount of each subsidy in each region and for different personal characteristics, we will estimate two versions of equation (1). In the first one, the variable D_{ijt} will be a binary variable being equal to one for those workers potentially treated in the corresponding year and region. The second version will 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. Furthermore, we will introduce the exact amount of the subsidy and its square term in order to capture any nonlinear pattern in the treatment effect.

One final but important remark is the usual worry about the presence of additional individual unobserved heterogeneity provoking serial correlation in our dependent variable and also within treatment units (Bertrand, Duflo and Mullainathan, 2003). We will also use two different approaches to deal with this issue. Firstly, following Chamberlain (1980), we allow the

[18] 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.

unobservable variables to be correlated with some elements of the covariates. A Mundlak version (1978) of Chamberlain´s assumption is to parameterize the aggregate unobserved heterogeneity as a linear projection on the mean of the time varying variables. In our case, this implies to parameterize the aggregate region-sector specific effects as a linear projection on the mean of time varying variables at the individual and at the region-sector level[19]. We specify these covariates:

$$\omega_{ij} = \pi Z_{ij} + \nu_{ij} \tag{3}$$

$$\omega_{kij} = \theta X_{kij} + \nu_{kij} \tag{4}$$

where Z_{ij} and X_{kij} represent average covariates over time at the aggregate and individual level respectively. A second approach to deal with unobserved heterogeneity is the usual random effect one. Another possibility would be to apply a conditional maximum likelihood estimator to our equation but, in that case, β_0 would not measure the average treatment effect, given that no distribution assumption had been made for the unobserved component so β_0 is not measuring the effect for the mean worker (See Wooldridge, 2004).

4.2 The Transition Rate to Permanent Employment: The Econometric Approach

Given that the Social Security records offer the duration of the spells of employment and unemployment on a monthly base the appropriate approach is to study the exit rate from a temporary contract as a discreet duration model. Moreover discreet time duration models allow specifying with enough flexibility the time dependence characteristics of the exit rate, as well as to incorporate in the analysis explanatory variables with temporary variability (Alison 1982). In addition, discreet duration models put in evidence the narrow existing correspondence between duration and discreet choice models.

As it is traditional in the literature of duration models the objective function is to estimate the exit rate. For each individual we observe the duration in a determined state -duration of an episode of temporary contracting-, from *t=1* up to *k-month* in which the individual changes of situation to any of the competing alternatives or remains in the same state. A common alternative to estimate the hazard rate consists of transforming the duration model in a sequence of discreet choice equations defined on the surviving population at each duration (Jenkins, 1995). In this case, we define a binary variable y_{ikj}, that takes value one when the worker changes state at time k to state j, and zero otherwise.

This expression has exactly the same form that the likelihood function of a discreet choice model where y_{jk} 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

[19] The statistical result at work is that the group mean is a sufficient statistic for estimating the unobserved effect specific to the corresponding level. The resulting slope estimators are not a function of these unobserved effects, which implies that they are consistent. By using this formulation, Mundlak (1978) maintains that the difference between the fixed and random effects approach is based on incorrect specification.

which the worker has remained in the initial situation (Allison, 1982; Jenkins, 1995). The corresponding probability term of the likelihood function when we estimate the duration model as a binary discreet choice model is:

$$P_j\left(t\right)=\frac{\exp(\varphi_{jt}+\alpha' x_{ijt}+\gamma' z_{jt}+\beta_0 D_{ijt}+\beta_1 D_{jt}+\beta_2 D_{ij}+\beta_3 D_j)}{\sum_{j=0}^{J}\exp(\varphi_{jt}+\alpha' x_{ijt}+\gamma' z_{jt}+\beta_0 D_{ijt}+\beta_1 D_{jt}+\beta_2 D_{ij}+\beta_3 D_j)} \quad (5)$$

where J=0,1,2,3 for the competing risk model. Given the above specification the competing risk model is estimated as a multinomial logit.

In the previous specification we have assumed that all the heterogeneity is due to observable characteristics. Nevertheless, it is also possible that unobserved characteristics such as the individual's preference for leisure or individual's ability constitute a factor of heterogeneity in the hazard rate among observations. As it has been extensively shown in the literature, the existence of unobserved heterogeneity in duration models might give rise to spurious duration dependence and bias the estimated coefficients[20]. In relation to the temporary hazard rate, the unobserved heterogeneity might miss-specify the duration dependence term. For instance, if more motivated workers find a permanent employment more quickly, the share of workers who remain in the temporary situation might grow along time. Moreover, one important characteristics of the competing risk model with unobserved heterogeneity is that the hypothesis of independence of irrelevant alternatives is not imposed anymore.

The unobserved heterogeneity term usually is specified as a random variable with a given distribution and it can be specific to each labour state as in Katz and Meyer, (1990) or common to all labour states as in Flinn and Heckman, 1982. The most common procedure to take into account this term is to add to the hazard function as a multiplicative stochastic error. The corresponding probability term of the likelihood function when we estimate the duration model as a binary discreet choice model is:

$$P_j\left(t\right)=\frac{\exp(\varphi_{jt}+\alpha' x_{ijt}+\gamma' z_{jt}+\beta_0 D_{ijt}+\beta_1 D_{jt}+\beta_2 D_{ij}+\beta_3 D_j+v_j)}{\sum_{j=0}^{J}\exp(\varphi_{jt}+\alpha' x_{ijt}+\gamma' z_{jt}+\beta_0 D_{ijt}+\beta_1 D_{jt}+\beta_2 D_{ij}+\beta_3 D_j+v_j)} \quad (6)$$

where v describes the unobserved heterogeneity term. We follow the approach proposed by Heckman and Singer to specify the unobserved 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, π_m. The points of support and the associated probabilities are estimated jointly. Following McFadden and Train (2000) the likelihood function of a mixed multinomial logit is:

[20] In general terms, the unobserved heterogeneity component allows controlling for specification problems as the omission of relevant variables or measurement errors in the endogenous variable (see Lancaster, 1979, 1990).

$$\ln L = \sum_{i=1}^{N} \sum_{j=1}^{J} y_{ij} * \ln\left(\Pi_{ij}\right) \tag{7}$$

where $y_{ij}=1$ if the worker changes to state j and zero otherwise. We also define the term Π_{ij}:

$$\Pi_j = \sum_{m=1}^{M} \pi_m P_j\left(t\right) \tag{8}$$

where m represents the points of support and π_m the corresponding probabilities.

5 Results

5.1 The Effects of Wage Subsidies over the Permanent Employment Rate

This section presents the results of estimating the effect of wage subsidies on the individual probability of having a permanent contract. Firstly we estimate the impact of the wage subsidy when the policy variable is represented by a dummy variable that takes value one when the individual is eligible for the wage subsidy and zero otherwise. In order to check the robustness of the results we estimate different version of this model. Moreover, given that the assumptions implicit in the DID approach may not be completely fulfilled in our sample, we also estimate the treatment effect of the policy once we split the initial by gender or age. Again theses estimators help us to evaluate the robustness of the results.

In Table 15 we display the main results relative to the policy variable for each model estimation and in Tables 16 and 17 we display all the estimated parameters for the Mundlack's correction and the random effects model.

The first column of Table 8 shows the *logit* estimates without adding additional covariates while in the second column we add regional and time dummies. We can see that the coefficient is not significant or shows a negative sign in some cases. But, once we introduce region and year dummies, we obtain a clearly positive and significant effect of this policy over the probability of having a permanent contract. The third column shows the results for our final specification but without any additional control for the presence of unobserved individual effects. We obtain that the effect of this policy, for the full sample, is positive and significantly different from zero. We can also see that the effect is larger for females than for males and also larger for young workers, being non-significant for old workers. Other relevant variable of our model is the "duration of the policy". In this case the coefficient is significantly different from zero and negative. This is telling us that the effect of these subsidies is decreasing with the number of years the policy is implemented. That is, there is a "programme introduction effect". We have computed the period in which the net effect of the policy turns to be negative, which is around four years for the full sample in this model.

The last two columns show the results when we control for the presence of unobserved individual heterogeneity. The Mundlack correction leaves our results almost the same as in the *logit* model[21]. When we assume a normal distribution for this heterogeneity the results change in the sense that the estimated effect of the policy is larger but highly decreasing. In fact, under this specification the wage subsidy has only a positive effect for the first two years of implementation. With this last specification, the evaluation of the treatment effect depends substantially on the shape of the distribution function for the unobserved component, which could be not similar to the normal one. We are now trying to estimate using a semi-parametric assumption about this distribution given the more sensible distribution of unobserved components should be one with basically two mass points: a very negative one which makes people do not work under permanent contracts and a very positive one which provokes just the opposite.

In order to quantify the effects found in our previous regressions, we have computed in Table 18 the marginal effect of being subject to the policy, varying with the number of years the policy may have been implemented. As can be seen, the effect is the largest one in the first year for the full sample and decreasing afterward. These marginal effects with the random effect specification are basically the same but depends substantially on the specific value of the unobserved component.

Other interesting result obtained is relative to the D_{jt} variable. Recall that this variable may capture equilibrium effects since it takes value one when the individual is located in a region where and at the time when the policy is implemented, independently of the eligibility conditions. This variable is statistically significant and shows a negative effect over the individual probability of getting a permanent contract except for the group of older workers. This implies that the probability of having a permanent contract is lower for non-eligible workers when the policy is implemented.

To check the robustness of the result we have estimated the random effects model for a subsample of workers with tenure lower than five years[22]. This sample selection rests on the idea that these are the workers that potentially might be strongly affected by these regional subsidies. Evidently, in this subsample the share of young workers is larger whereas the share of permanent workers is much lower. Nevertheless the estimated parameters are close to those found in previous estimations. Concretely, the results are almost the same as for the subsample of young workers.

Secondly we estimate the effect of the policy taking into account the maximum amount of the subsidy. This second exercise is interesting since it allows to compute the elasticity of the probability of having a permanent contract in relation to a marginal increase in the subsidy. Besides, we test if there is a non-linear relation between the probability of having a permanent contract and the amount of the subsidy. The idea is to measure whether the policy is efficient in

[21] This result means that the time-invariant heterogeneity is already capture by the DID specification problems.

[22] We do not present the results but they are available upon request.

the sense that an extra euro of subsidy increases this probability. As it was done before, we estimate the model considering different specifications and splitting the sample by gender or age. Nevertheless, we only show a summary of the results in Table 19 where we display the results relative to the policy parameter β_0 for the different specifications.

We obtain that the effect of the wage subsidy is positive but since we also get a negative coefficient for the square term we have to conclude that the impact of the policy is positive but at a decreasing rate. We have computed the maximum in each case and for the full sample this is around 8.000€. In this model the effect is larger for males than for females, that is, the probability of working under a permanent contract for men is more sensible to a change in the incentives than the one for females. As before, the effect is also larger for young workers. [23]

Finally, we have computed in Table 20 the elasticity of the probability of being working under a permanent contract with respect to the quantity of the subsidy (measured in thousand euros). We can see that this elasticity is quite small, being only substantial for young workers. For these workers the policy seems to be quite effective given the reaction to an increase in the subsidy is quite large.

5.2 The Effects of Wage Subsidies over the Transition Rate to Permanent Employment

To be completed.

6 Conclusions

Spain is one of the countries with the highest rate of temporary contracts and given this might bring negative costs in terms of efficiency and equity conditions, national and regional governments have designed policies to foster the creation of permanent employment. Since the labor reform of 1997 the national government offered discounts in payroll taxes for new permanent contracts. Simultaneously, since 1997, different regional governments have also 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. One interesting point of this exercise is that we take advantage from all the variability derived from these regional policies, that is, regional, time and individual eligibility criteria. Since we have longitudinal data we apply a Difference-in-Differences 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 we also estimate the model splitting the whole sample by age or gender. Two main empirical approaches are applied. Firstly we estimate the incidence of regional wage subsidies on the

[23] In this model we do not consider the discounts in the payroll taxes offered by the national government.

permanent employment rate. Secondly, we examine their impact on the transition rate to a permanent employment.

From our preliminary results, we can conclude that the policy of subsidizing permanent contracts shows a small positive effect over the rate of permanent contracts. Moreover, the estimated average treatment effect shows a highly decreasing pattern over the duration of the policy implementation. The effect of being subject to this policy is larger for females than for males (the estimated marginal effect for females is 3.9 percentage points of increase over the rate of permanent contracts) but the predicted elasticity to the amount of the subsidy is not larger for females than for males. Maybe, having larger subsidies for females than for males is, given this result, not very useful.

Finally, we can evaluate the predicted elasticity with the figures of 2002. We obtain that an increase of 1.000€ for the mean value of the subsidy in Spain that year, would have made the number of permanent contracts to increase in around 40.000 more contracts. Not much if we could evaluate the cost of such a measure.

One final conjecture for the decreasing pattern found in the treatment effect is that firms quickly adapt their strategies to these policies that are not well controlled once the money has been transferred to the firm. Hence, some of them could be using these subsidies to substitute previous permanent contracts for new ones and hence, the total number of permanent contracts will not go up substantially. Furthermore, the larger effect found for young workers, which is exactly what finds other papers evaluating the same labor market reform, could be working as reducing the number of temporary contracts the young workers need in order to be tested by the firm.

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Main Statistics

Table 1: Percentage of Permanent Contracts by regions 1995-2002 (EPA)

	1995	1996	1997	1998	1999	2000	2001	2002	1995-2002*
Spain	**65.6%**	**66.9%**	**67.4%**	**67.8%**	**67.9%**	**68.6%**	**69.2%**	**69.7%**	**4.1**
Andalusia	60.2%	60.3%	61.4%	60.7%	59.1%	58.5%	57.9%	56.8%	-3.4
Aragon	70.9%	71.5%	71.1%	71.5%	70.9%	73.4%	74.6%	76.6%	5.7
Asturias	76.4%	74.9%	76.2%	75.9%	75.4%	72.8%	73.1%	71.1%	-5.3
Balearic Islands	67.0%	69.2%	70.1%	70.2%	70.3%	71.3%	70.7%	72.8%	5.8
Canary Islands	55.8%	56.1%	58.8%	58.5%	58.7%	63.9%	60.6%	59.6%	3.8
Cantabria	69.3%	70.4%	70.4%	70.1%	72.6%	71.5%	73.8%	74.8%	5.5
Castilla-León	68.1%	70.9%	69.5%	70.7%	70.8%	73.1%	72.0%	71.5%	3.4
Castilla-La Mancha	56.1%	62.5%	60.1%	61.0%	62.8%	64.1%	63.7%	64.0%	7.9
Catalonia	65.2%	65.7%	66.9%	69.1%	69.3%	71.2%	74.9%	76.1%	10.9
Valencia	56.7%	59.8%	58.6%	62.1%	63.0%	65.3%	65.5%	64.1%	7.4
Extremadura	64.6%	70.5%	67.4%	67.8%	66.3%	65.9%	63.5%	64.8%	0.2
Galicia	67.1%	67.6%	67.4%	67.7%	66.8%	66.1%	65.7%	66.1%	-1.0
Madrid	78.2%	77.7%	78.2%	77.8%	78.9%	78.1%	79.6%	79.7%	1.5
Murcia	61.2%	62.7%	62.1%	62.8%	61.8%	62.5%	62.3%	62.1%	0.9
Navarra	70.1%	73.5%	73.6%	74.3%	75.7%	74.1%	74.6%	73.8%	3.7
Basque Country	66.5%	68.8%	68.1%	68.9%	68.0%	68.9%	68.9%	70.3%	3.8
Rioja	73.6%	70.2%	69.7%	72.6%	73.0%	75.1%	77.3%	77.4%	3.8

*Differences in percentage points

Figure 1: Rate of Permanent Employment, Spanish Regions (EPA, 1995-2002)

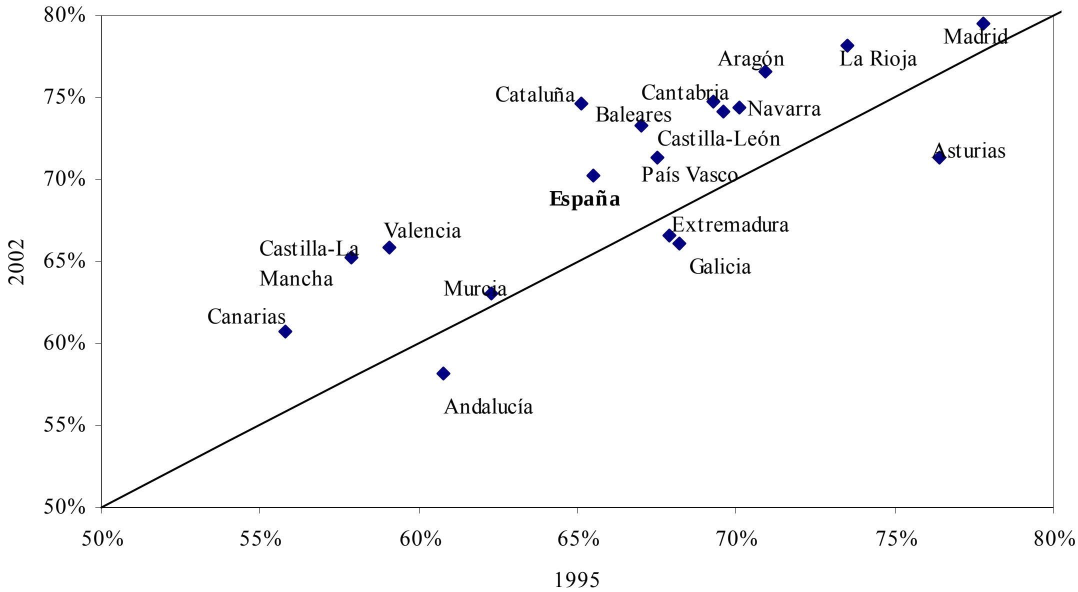

Table 2: Percentage of permanent contracts by personal characteristics, Spanish Regions (average 1995-2002, EPA)

	Age				Gender		Studies				
	18-24	25-30	40-55	55-65	Men	Woman	No Studies	Primary	Secondary	Superior (short-term)	Superior (long term)
Spain	**29.6%**	**66.1%**	**83.3%**	**88.0%**	**68.8%**	**64.6%**	**62.8%**	**63.3%**	**66.9%**	**77.4%**	**75.9%**
Andalusia	32.6%	57.1%	78.3%	85.0%	60.8%	57.6%	55.8%	53.1%	59.9%	76.4%	72.9%
Aragon	34.4%	72.3%	86.6%	90.7%	74.8%	68.3%	69.3%	70.7%	71.3%	77.7%	76.9%
Asturias	31.5%	70.3%	89.8%	90.5%	75.4%	72.7%	82.0%	72.0%	73.3%	83.9%	77.9%
Balearic Islands	34.3%	73.8%	84.1%	89.7%	58.9%	61.2%	73.0%	69.2%	66.2%	80.7%	76.2%
Canary Islands	25.0%	58.7%	77.8%	87.5%	61.1%	58.5%	63.9%	53.9%	60.2%	75.4%	71.6%
Cantabria	27.2%	67.0%	89.5%	93.3%	72.8%	70.3%	74.3%	79.1%	69.9%	69.8%	82.9%
Castilla-León	30.3%	69.5%	87.7%	90.4%	73.8%	69.4%	63.9%	69.9%	70.7%	79.6%	77.0%
Castilla-La Mancha	28.3%	63.9%	79.6%	84.2%	63.5%	60.2%	53.1%	56.6%	65.5%	77.9%	74.9%
Catalonia	31.3%	71.9%	83.6%	88.2%	70.7%	66.2%	71.1%	65.6%	67.5%	77.9%	76.9%
Valencia	27.1%	63.8%	77.6%	82.5%	63.9%	58.5%	60.9%	57.3%	62.9%	72.7%	71.2%
Extremadura	37.9%	66.6%	81.4%	87.1%	67.2%	67.9%	61.2%	61.6%	68.4%	79.7%	80.1%
Galicia	26.2%	65.1%	82.3%	90.4%	67.8%	66.4%	71.6%	63.0%	65.1%	77.6%	76.7%
Madrid	44.8%	77.3%	90.2%	93.6%	78.9%	75.6%	78.0%	76.3%	76.3%	82.1%	83.3%
Murcia	26.2%	63.8%	80.8%	90.2%	63.8%	59.2%	63.0%	57.9%	60.2%	76.4%	71.4%
Navarra	34.8%	72.5%	90.1%	94.3%	76.7%	70.3%	80.9%	75.5%	73.4%	74.3%	68.9%
Basque Country	24.1%	65.8%	87.6%	89.5%	72.8%	63.5%	79.9%	70.6%	64.9%	73.5%	67.1%
Rioja	36.2%	73.0%	86.1%	92.7%	74.0%	71.7%	71.1%	70.8%	72.5%	81.1%	73.5%

Table 3: Percentage of permanent contracts by labor characteristics, Spanish Regions (average 1995-2002, EPA)

	Firm Size			Occupation Skill			Hours of work		Firm's Ownership	
	Small	Medium	Big	High	Medium	Low	Full-Time	Part-Time	Public	Private
Spain	**59.6%**	**68.3%**	**79.6%**	**80.1%**	**68.8%**	**59.2%**	**69.2%**	**43.8%**	**86.6%**	**70.9%**
Andalusia	67.6%	72.1%	80.3%	78.8%	62.3%	48.9%	62.4%	32.1%	79.9%	51.7%
Aragon	49.6%	63.0%	76.2%	80.8%	73.1%	66.9%	74.7%	53.5%	83.9%	68.5%
Asturias	66.6%	73.1%	83.3%	85.9%	71.4%	70.9%	75.9%	52.0%	87.9%	69.2%
Balearic Islands	67.3%	76.8%	79.1%	81.0%	72.2%	63.2%	71.7%	56.5%	85.2%	66.8%
Canary Islands	53.0%	63.7%	70.8%	76.9%	63.9%	59.2%	61.4%	39.3%	77.6%	54.1%
Cantabria	64.0%	72.7%	82.8%	82.3%	70.1%	67.9%	73.1%	51.2%	82.9%	68.4%
Castilla-León	64.3%	72.5%	82.7%	81.9%	73.9%	65.7%	74.6%	48.1%	84.0%	67.0%
Castilla-La Mancha	56.3%	63.9%	77.6%	53.1%	56.6%	65.5%	64.9%	42.4%	78.9%	56.%
Catalonia	61.3%	71.9%	78.6%	80.1%	71.6%	51.5%	71.9%	43.7%	82.9%	66.3%
Valencia	58.1%	64.8%	72.6%	80.9%	68.3%	63.9%	63.7%	43.6%	78.2%	58.5%
Extremadura	63.9%	73.6%	80.4%	76.2%	64.6%	54.4%	69.7%	44.1%	76.1%	62.2%
Galicia	61.2%	68.1%	77.3%	82.6%	73.0%	55.1%	68.6%	50.1%	83.7%	61.7%
Madrid	71.8%	78.3%	85.2%	86.0%	77.3%	71.3%	79.1%	50.5%	90.8%	74.2%
Murcia	55.2%	66.8%	75.8%	77.0%	65.9%	53.2%	64.4%	41.8%	82.4%	56.1%
Navarra	70.8%	77.5%	80.1%	78.9%	72.5%	71.4%	75.3%	60.1%	77.9%	72.9%
Basque Country	57.1%	72.8%	79.6%	75.9%	65.6%	65.9%	71.5%	38.4%	65.1%	79.8%
Rioja	72.2%	69.0%	78.1%	83.1%	74.8%	71.5%	74.1%	51.9%	85.5%	69.9%

Table 4: New Contracts, 1995-2002 (Social Security Records)

	Permanent Contract	Temporary Contract
Average 1995-2004	20%	80%
1996	18%	82%
2000	20%	80%
2002	19%	81%

Table 5: Transitions From a Temporary Contract

	Permanent		Temporary		Unemployment	
	Total	Same Firm	Total	Same Firm	Total	Same Firm
Media 1995-2004	7%	25%	41%	36%	52%	-
1996	7%	15%	36%	24%	57%	-
2000	8%	27%	41%	37%	50%	-
2003	5%	30%	43%	39%	52%	-

Table 6: Type of transitions and its relation with the duration of the temporary contract

	Total	Permanent		Temporary		Unemployed	
		Total	Same Firm	Total	Same Firm	Total	Same Firm
< 3 Months	53%	4%	40%	42%	55%	53%	-
3-6 Months	22%	8%	63%	40%	50%	51%	-
6-12 Months	14%	10%	59%	38%	46%	52%	-
12-24 Months	7%	15%	52%	43%	38%	41%	-
> 24 Months	3%	19%	43%	42%	36%	38%	-

Figure 2: Exist Rate from a Temporary Contract by Type of Transition

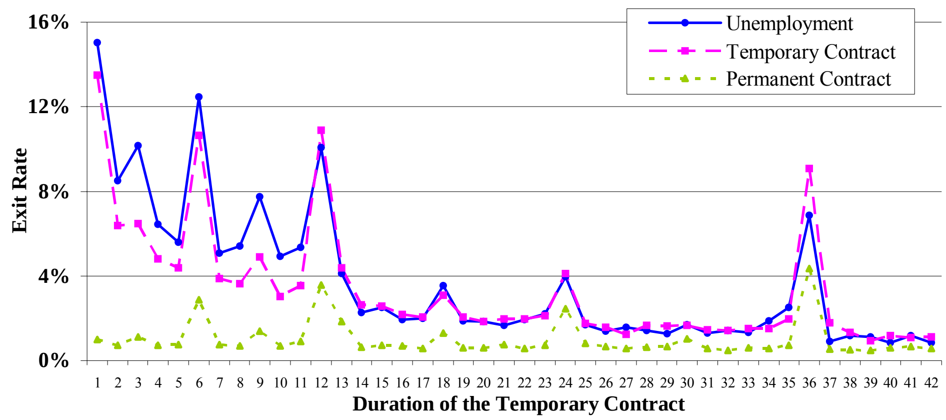

Figure 3: Exit from Unemployment, Men (1995-2004)

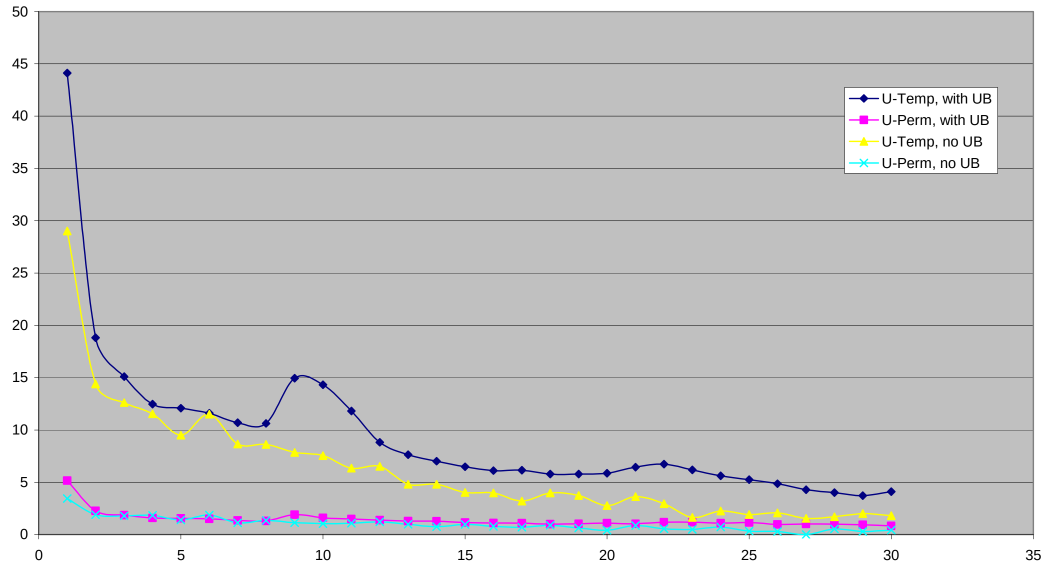

Figure 4: Exit from Unemployment, Women (1995-2004)

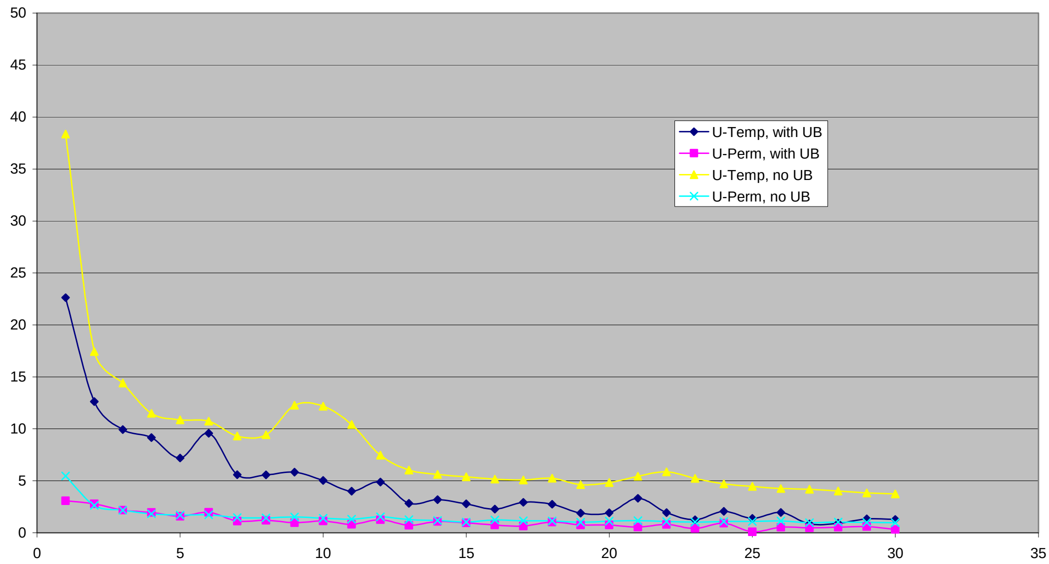

Table 0: Regional Incentives to the creation of Indefinite Contracts

	1995	1996	1997	1998	1999	2000	2001	2002
Andalusia								
Aragon								
Asturias								
Balearic Islands								
Canary Islands								
Cantabria								
C. Leon								
C. Mancha								
Catalunya								
Valencia								
Extremadura								
Galicia								
Madrid								
Murcia								
Navarra								
Basque Country								
Rioja								

Table 12: Regional Incentives to the creation of Indefinite Contracts: eligibility conditions

	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	16-25	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	16-30 & 45 or more	1998	all ages	1.998
	all ages	2003-2004	all ages	2003-2004
Madrid	all ages	1998-2004	all ages	1998-2004
Murcia	all ages	1998-2003	all ages	1998-2003
Navarra	all ages	1998-2004	all ages	1998-2004
Basque Country	all ages	1998-2004	all ages	1998-2004
Rioja	all ages	1998-2004	all ages	1998-2004

Table 13: Regional Incentives to the creation of Indefinite Contracts

	Minimum	Mean	Maximum
Andalusia	2.400	3.844	6.012
Aragon	1.200	3.684	5.160
Asturias	3.600	4.100	4.500
Balearic Islands	1.653	1.726	1.800
Canary Islands	3.000	3.300	3.600
Cantabria	1.803	3.604	4.808
C. Leon	1.800	3.605	5.115
C. Mancha	3.000	3.300	3.600
Catalunya	0	0	0
Valencia	1.800	4.513	7.466
Extremadura	4.166	10.076	14.028
Galicia	3.000	3.600	4.200
Madrid	3.600	7.971	12.000
Murcia	3.000	4.838	7.200
Navarra	3.000	3.900	4.800
Basque Country	3.273	4.459	7.512
Rioja	4.491	5.001	6.011

Table 14: Regional Incentives to the creation of Indefinite Contracts by personal characteristics

	Minimum	Mean	Maximum
Males	1.200	5.916	14.028
Females	1.200	6.411	14.028
Aged 18-30	1.200	5.520	12.000
Aged 31-44	1.200	5.652	14.028
Aged 45-64	1.653	6.256	14.028

Table 15: Probability of Having a Permanent Contract (N =1276447)

	No regressors		**+ region & year dummies**		**Logit**		**Logit (with Mundlack Corr.)**		**Random Effects (Normal D.)**	
	Coef.	Std. Er.	Coef.	Std. Er.	Coef.	Std. Er.	Coef.	Std. Er.	Coef.	Std. Er.
FULL SAMPLE										
Dijt	0,007	0,012	0,146 **	0,013	0,158 **	0,015	0,152 **	0,015	0,241 **	0,062
Dur x Dijt					-0,041 **	0,003	-0,039 **	0,003	-0,116 **	0,016
Change of sign					4		4		2	
MALES										
Dijt	-0,040 **	0,014	0,144 **	0,016	0,210 **	0,018	0,185 **	0,019	0,243 **	0,079
Dur x Dijt					-0,041 **	0,004	-0,038 **	0,004	-0,120 **	0,021
Change of sign					5		5		2	
FEMALES										
Dijt	0,084 **	0,025	0,190 **	0,031	0,242 **	0,032	0,248 **	0,033	0,364 **	0,124
Dur x Dijt					-0,044 **	0,005	-0,041 **	0,005	-0,165 **	0,026
Change of sign					5		7		2	
18-30 YEARS										
Dijt	-0,180 **	0,021	0,144 **	0,024	0,197 **	0,025	0,204 **	0,026	0,213 **	0,098
Dur x Dijt					-0,047 **	0,005	-0,049 **	0,005	-0,089 **	0,024
Change of sign					4		5		2	
31-44 YEARS										
Dijt	0,102 **	0,016	0,126 **	0,020	0,113 **	0,022	0,100 **	0,023	0,100 *	0,098
Dur x Dijt					-0,035 **	0,005	-0,029 **	0,005	-0,029 *	0,025
Change of sign					3		4		3	
45-65 YEARS										
Dijt	0,324 **	0,035	0,122 **	0,037	0,029	0,042	0,010	0,044	-0,225 *	0,176
Dur x Dijt					-0,044 **	0,007	-0,052 **	0,008	-0,076 *	0,041
Change of sign					1		1		3	

Table 16:The Probability of Having a Permanent Contract (Mundlack´s Correction)

	Whole Sample		Subsamples by Gender				Subsamples by Age					
			Men		Woman		Young		Middle Age		Old	
	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.
Gender	0.314	0.005					0.174	0.008	0.395	0.009	0.349	0.014
Labor Experience	0.136	0.001	0.133	0.001	0.139	0.002	0.304	0.005	0.163	0.007	-0.078	0.013
Labor Experience^2	-0.002	0.000	-0.002	0.000	-0.002	0.000	-0.012	0.000	-0.004	0.000	0.001	0.000
No Studies	-0.674	0.017	-0.840	0.022	-0.438	0.027	-0.069	0.051	-0.757	0.031	-1.129	0.041
Primary Studies	-0.308	0.010	-0.343	0.015	-0.295	0.014	-0.175	0.016	-0.451	0.017	-0.762	0.038
Secondary Studies	0.152	0.010	0.163	0.014	0.090	0.013	0.155	0.015	0.120	0.016	-0.267	0.038
Superior Studies (short duration)	0.107	0.010	0.083	0.016	0.135	0.013	-0.027	0.016	0.194	0.016	0.148	0.039
Civil Status	-0.337	0.006	-0.469	0.008	-0.182	0.008	-0.389	0.009	-0.384	0.009	0.037	0.019
Relative Sector Size (employee)	-3.521	0.087	-4.063	0.118	-2.964	0.136	-2.204	0.134	-2.053	0.154	-2.904	0.235
Big Firm	0.243	0.014	0.105	0.019	0.383	0.022	0.325	0.021	0.157	0.024	0.241	0.035
Full Time Job	0.952	0.008	1.430	0.018	0.887	0.009	0.768	0.012	0.918	0.012	0.985	0.018
Private Firm	-0.484	0.016	-0.846	0.023	-0.169	0.024	-0.439	0.027	-0.149	0.028	-0.173	0.041
Medium Skill	-0.191	0.008	-0.110	0.011	-0.263	0.011	-0.006	0.011	-0.206	0.012	-0.504	0.024
Low Skill	-0.655	0.008	-0.670	0.011	-0.661	0.013	-0.306	0.012	-0.737	0.013	-0.984	0.023
Seasonal Job	0.489	0.025	0.959	0.040	0.164	0.033	0.419	0.043	0.511	0.040	0.568	0.052
Share of Small Firms	-1.832	0.062	-0.910	0.083	-2.721	0.097	-1.171	0.093	-1.214	0.107	-1.490	0.160
D_{ijt}	0.152	0.015	0.185	0.018	0.248	0.032	0.204	0.026	0.100	0.022	0.010	0.043
D_{ij}	0.021	0.013	0.000	0.016	0.265	0.028	0.158	0.039	-0.015	0.020	0.369	0.065
D_{jt}	-0.061	0.014	-0.074	0.017	-0.165	0.030	-0.109	0.024	-0.034	0.020	0.115	0.043
D_j	0.245	0.022	0.281	0.030	-	-	0.199	0.047	0.253	0.036	-0.175	0.077
D_{ijt}*Dur	-0.039	0.003	-0.038	0.004	-0.041	0.005	-0.049	0.005	-0.029	0.005	-0.052	0.008
Constant	-1.397	0.045	-1.800	0.065	2.907	0.105	-2.657	0.072	-1.689	0.093	2.370	0.241

We also control for time, regional and sectorial dummies, relative regional specialization and time average regional and individual variables .

Table 17 : The Probability of Having a Permanent Contract (Random Effects Estimation)

	Whole Sample		Subsamples by Gender				Subsamples by Age					
			Men		Woman		Young		Middle Age		Old	
	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.	Coef	Err.Est.
Gender	1.664	0.032					0.702	0.018	0.395	0.009	0.349	0.014
Labor Experience	0.616	0.007	0.604	0.009	0.615	0.009	0.953	0.011	0.163	0.007	-0.078	0.013
Labor Experience^2	-0.009	0.000	-0.009	0.000	-0.009	0.000	-0.032	0.001	-0.004	0.000	0.001	0.000
No Studies	-3.221	0.096	-4.027	0.126	-2.488	0.139	-1.442	0.107	-0.757	0.031	-1.129	0.041
Primary Studies	-2.114	0.064	-2.518	0.086	-1.794	0.094	-1.370	0.035	-0.451	0.017	-0.762	0.038
Secondary Studies	-0.110	0.060	-0.317	0.080	-0.056	0.086	-0.101	0.033	0.120	0.016	-0.267	0.038
Superior Studies (short duration)	0.070	0.063	-0.092	0.087	0.465	0.083	-0.137	0.036	0.194	0.016	0.148	0.039
Civil Status	-1.032	0.033	-1.650	0.047	-0.387	0.048	-0.993	0.020	-0.384	0.009	0.037	0.019
Relative Sector Size (employee)	-10.060	0.410	-12.450	0.530	-8.702	0.688	-4.208	0.277	-2.053	0.154	-2.904	0.235
Big Firm	0.559	0.039	0.259	0.051	0.895	0.061	0.539	0.037	0.157	0.024	0.241	0.035
Full Time Job	2.198	0.036	3.150	0.076	1.980	0.041	1.584	0.025	0.918	0.012	0.985	0.018
Private Firm	-0.384	0.054	-1.352	0.078	0.173	0.078	-0.030	0.054	-0.149	0.028	-0.173	0.041
Medium Skill	-0.840	0.046	-0.298	0.067	-1.218	0.065	-0.067	0.025	-0.206	0.012	-0.504	0.024
Low Skill	-2.373	0.048	-2.202	0.064	-2.356	0.075	-0.770	0.028	-0.737	0.013	-0.984	0.023
Seasonal Job	1.064	0.089	1.879	0.135	0.488	0.111	0.722	0.084	0.511	0.040	0.568	0.052
Share of Small Firms	-3.297	0.264	-0.586	0.334	-6.607	0.431	-1.380	0.191	-1.214	0.107	-1.490	0.160
D_{ijt}	0.241	0.059	0.243	0.074	0.364	0.116	0.213	0.054	0.100	0.022	0.010	0.043
D_{ij}	0.099	0.074	-0.086	0.099			1.096	0.085	-0.015	0.020	0.369	0.065
D_{jt}	-0.024	0.053	-0.118	0.068	-0.053	0.106	-0.140	0.052	-0.034	0.020	0.115	0.043
D_{j}	1.877	0.122	2.188	0.159	1.964	0.143	0.668	0.103	0.253	0.036	-0.175	0.077
D_{ijt}*Dur	-0.116	0.015	-0.120	0.020	-0.165	0.025	-0.089	0.011	-0.029	0.005	-0.052	0.008
Constant	-6.672	0.210	-5.879	0.282	-8.521	0.436	-9.123	0.151	-1.689	0.093	2.370	0.241
sigma_u	6.395	0.019	6.472	0.026	6.377	0.029	3.657	0.008				
rho	0.926	0.000	0.927	0.001	0.925	0.001	0.803	0.001				

Table 18: Marginal Effects on the Probability of Having a Permanent Contract (Percentage Points)

Duration	**Full Simple**	**Males**	**Females**	**18-30 Years**	**31-45 Years**	**46-65 Years**
1	0,019	0,023	0,039	0,038	0,011	-0,003
2	0,013	0,017	0,032	0,026	0,006	-0,007
3	0,006	0,012	0,024	0,014	0,002	-0,012
4	-0,001	0,006	0,016	0,002	-0,003	-0,016
5	-0,008	0,000	0,008	-0,010	-0,008	-0,021
6	-0,015	-0,007	0,000	-0,022	-0,012	-0,026

Table 19: Probability of Having a Permanent Contract (N =1276447)

	Logit		Logit (with Mundlack Corr.)		Random Effects (Normal D.)	
	Coef.	Std. Er.	Coef.	Std. Er.	Coef.	Std. Er.
FULL SAMPLE						
subsidy	0,045 **	0,004	0,046 **	0,004	0,077 **	0,018
subsidy(squared)	-0,003 **	0,000	-0,003 **	0,000	-0,005 **	0,001
Max. Effect	7,581		7,682		8,051	
MALES						
subsidy	0,077 **	0,005	0,076 **	0,006	0,135 **	0,024
subsidy(squared)	-0,005 **	0,000	-0,005 **	0,000	-0,010 **	0,002
Max. Effect	7,168		7,160		6,784	
FEMALES						
subsidy	0,028 **	0,007	0,025 **	0,007	0,047 *	0,029
subsidy(squared)	-0,001 *	0,001	-0,001 *	0,001	-0,003 *	0,002
Max. Effect	13,508		17,255		7,833	
18-30 YEARS						
subsidy	0,043 **	0,007	0,046 **	0,007	0,016	0,017
subsidy(squared)	-0,002 **	0,001	-0,002 **	0,001	-0,0004	0,001
Max. Effect	12,746		12,260		21,475	
31-44 YEARS						
subsidy	0,018 *	0,007	0,015 **	0,007	0,017	0,029
subsidy(squared)	-0,001	0,001	-0,001 **	0,001	-0,003 *	0,002
Max. Effect	6,721		6,936		2,525	
45-65 YEARS						
subsidy	0,015 *	0,010	0,009	0,010	-0,013	0,045
subsidy(squared)	-0,002 **	0,001	-0,002 **	0,001	-0,002	0,003
Max. Effect	3,921		2,886		4,361	

Table 20: Elasticity of the Probability of Having a Permanent Contract Relative to the Subsidy

Full Sample	Males	Females	18-30 Years	31-45 Years	46-65 Years
0,018	0,022	0,021	0,075	0,004	-0,002