Available online at www.sciencedirect.com
ScienceDirect
Labour Economics 14 (2007) 153–183
LABOUR ECONOMICS
www.elsevier.com/locate/econbase

How binding are legal limits? Transitions from temporary to permanent work in Spain

Maia Güell [a], Barbara Petrongolo [b,*]

[a] *Department of Economics and Business, Universitat Pompeu Fabra, CREA, CEP (LSE), CEPR and IZA, Barcelona, Spain*
[b] *Department of Economics, London School of Economics, CEP (LSE), CEPR and IZA, London, UK*

Received 28 July 2005; received in revised form 21 August 2005; accepted 20 September 2005
Available online 9 November 2005

Abstract

This paper studies the duration pattern of fixed-term contracts and the determinants of their conversion into permanent ones in Spain, where the share of fixed-term employment is the highest in Europe. We estimate a duration model for temporary employment, with competing risks of terminating into permanent employment versus alternative states, and flexible duration dependence. We find that conversion rates are generally below 10%. Our estimated conversion rates roughly increase with tenure, with a pronounced spike at the legal limit, when there is no legal way to retain the worker on a temporary contract. We argue that estimated differences in conversion rates across categories of workers can stem from differences in worker outside options and thus the power to credibly threat to quit temporary jobs.

JEL classification: C41; J41; J60
Keywords: Temporary contracts; Duration models

1. Introduction

Several European labor markets have been characterized by a wide use of permanent contracts with stringent and costly firing regulations. In the mid-1980s, in order to fight the high and persistent levels of unemployment, some European countries enhanced the flexibility of their labor markets by allowing employers to hire workers on a fixed-term basis, with negligible termination costs upon expiry of contract. Typically, there exists a legal duration limit in the use

* Corresponding author.
E-mail addresses: maia.guell@upf.edu (M. Güell), b.petrongolo@lse.ac.uk (B. Petrongolo).

0927-5371/$ - see front matter
doi:10.1016/j.labeco.2005.09.001

of these contracts, after which an employer can either offer the worker a contract of undetermined duration or dismiss her. Since their introduction, fixed-term contracts have been widely used and they accounted for most new hirings in all sectors and occupations, especially in countries characterized by high levels of employment protection (OECD, 1993). European labor markets have become more dynamic in terms of higher inflows and outflows between unemployment and employment, but the overall level of unemployment did not seem largely affected by the introduction of fixed-term contracts.

The consequences of the introduction of temporary (or fixed-term) contracts have raised interest and concern among both academics and policy-makers (see Booth et al., 2002a; OECD, 2002). Some consensus has formed among economists that the introduction of temporary contracts does not necessarily increase employment, while creating dualism in the labor market (see, among others, Bentolila and Dolado, 1994; Blanchard and Landier, 2002; Güell, 2000; Cahuc and Postel-Vinay, 2002). An important aspect of the use of temporary contracts is their pattern of promotion into regular contracts of indefinite duration. Mixed employment effects of the introduction of temporary contracts and rising dualismprovide some clear signal that temporary contracts largely failed to provide workers with effective "stepping stones" to permanent employment.

In this paper we study the determinants of the conversion of temporary contracts (TCs) into permanent contracts (PCs) as well as the duration pattern of TCs. In doing this, we focus on one country, Spain, mostly because it represents an extreme experience in several labor market dimensions. Compared to other OECD countries, Spain has had until recently the highest rate of unemployment, and ranks second in terms of strictest employment protection legislation (OECD, 1999). This situation triggered an experiment of "flexibility at the margin", started in 1984 with the introduction of TCs. This reform was somewhat more radical than in other European countries. In particular, while in some countries TCs are restricted to some type of workers or sectors,[1] the Spanish 1984 reform did not limit in any way the applicability of TCs. At the same time, the 1984 reform set an "up or out" clause after three years of continuous employment in a TC. Upon expiry of this legal limit a temporary employee has to be promoted to a permanent contract or dismissed.

Soon after their introduction, coinciding with the economic expansion of the late 1980s, more than 90% of newly created contracts have been fixed-term,[2] and this translated into a rapidly growing stock of temporary employment, from 11% in 1983 to approximately 35% by the early 1990s, which is more than three times the European average (see OECD, 1987, 1993). However, during the same time span, unemployment remained as high as before the reform. Within a decade, the Spanish labor market had experienced record rates of gross job creation, but little permanent employment had been created as only a small fraction of TCs had been converted into PCs. The labor market had gradually evolved towards a dual structure, with two thirds of employees retaining a permanent status and the rest working in a highly mobile market. Interestingly enough, once these effects became evident, Spanish policy makers restricted the applicability of TCs and offered fiscal incentives for their conversion into PCs (1994 reform). Later reforms (in 1997 and 2001) continued to limit the applicability of TCs as well as offering incentives to convert TCs into PCs (see Appendix A for more institutional details).

[1] See Grubb and Wells (1993) and OECD (1993, 1994, 1999) for a detailed description of fixed-term employment regulations in Europe.

[2] Bover and Gomez (2004) find that exit rates from unemployment into temporary employment are ten times larger than exit rates into permanent employment.

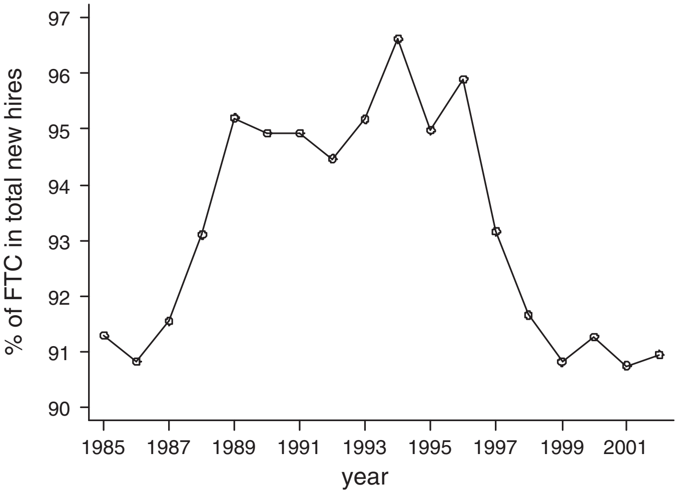

Fig. 1. Evolution of the share of fixed-term contracts in new hires, 1985–2002. Source: MLR (Spanish Ministry of Labor).

There exists a growing literature which studies several aspects of the impact of TCs on labor markets in OECD countries, with special reference to the Spanish case (see Dolado et al., 2002a for a comprehensive survey). However, there is an important aspect which is to date largely underexplored in this literature, namely the study of the conversion of TCs into PCs and its timing. This paper tries to shed light on the economic mechanisms behind conversions of TCs and their implications for the dualism of the labor market.[3] We argue that some of the patterns found in the timing of conversion rates of TCs may be suggestive of variation in temporary workers' outside option and their ability to threat employers to quit current temporary jobs in search for better matches.

In order to understand dualism in the labor market, it is useful to distinguish between entry into and exit from temporary jobs. It can certainly be argued that in several institutional settings the entry into temporary employment is a first stepping-stone into permanent employment, and indeed the probability of accession to permanent employment is higher for those on TCs than for the unemployed (see Farber, 1999 for evidence for the US). This statement holds trivially for Spain, in which over 90% of accessions to permanent employment happen as conversions of TCs, as shown in Fig. 1. As the entry margin displays little variation, the main source of dualism in the Spanish labor market lies in the exit margin, i.e. in the conversion of TCs into PCs. This will be the focus of this paper.

We estimate a duration model of temporary employment using the panel version of the Spanish Labor Force Survey (EPA), started in 1987. We believe that duration models best describe the dynamics of the transition process between temporary and permanent employment by exploiting the strength of a panel data, which is the possibility of being able to track individuals over time and observe exactly how long they take to make an employment change. Moreover, the use of individual information on worker characteristics that can be obtained from the EPA shows how the prospect of permanent employment is shared among temporary workers, and to what extent

[3] Closely related to our work is the recent growing literature on the role of Temporary Help Agencies, and thus a specific typology of TCs, as potential springboards towards permanent employment, see Ichino et al. (2005); Autor and Houseman (2005), and references therein.

there are some categories that are more likely than others to remain trapped in temporary jobs. The additional advantage related to the use of EPA data is the length of the period covered by the survey. We use data for the period 1987–2002, which allows us to assess the conversion pattern of TCs introduced in 1984, as well as analyze the effects of the later reforms.

The existing literature contains only a few contributions on conversion rates for Spain. Amuedo-Dorante (2001) examines the determinants of Spanish employers' conversions of temporary contracts into permanent ones using information on the composition of firm level employment. She finds that dismissal costs hardly affect contract conversions, which mostly respond to employment expectations and union pressure for increased employment stability. In our study we focus on individual rather than firm-level conversion rates, in order to estimate the time pattern of conversions. Most existing studies on the determinants of individual conversion rates use logit specifications (Toharia, 1996; Alba, 1998), which may prove rather inflexible when applied to the analysis of the dynamic path of transition rates. To our knowledge, the only duration study on Spanish conversion rates is Amuedo-Dorante (2000), who estimates transitions out of temporary employment using EPA individual records from 1995:2 through 1996:2, and finds that conversion rates are very low, regardless of job tenure. Our paper uses a longer sample period to study the time pattern of permanent conversions. In doing this, we allow for variation of conversion rates both over tenure levels (within job matches) and across different categories of workers.

The paper is organized as follows. Section 2 proposes a simple framework for the use and conversion of TCs, which should guide our empirical analysis. Section 3 describes the data set used. Section 4 illustrates the duration model to be estimated. Section 5 presents the results and Section 6 finally concludes.

2. A simple framework

This section proposes a simple theoretical framework that illustrates firm use of TCs and their conversion into PCs. Our goal here is to derive testable implications for the relationship between conversion rates and contract duration for different categories of workers. The model proposed in this section is the simplest possible that would deliver our testable predictions. A number of extensions to this bare-bone model are described in Appendix B.

TCs can firstly be used by employers for covering seasonal or casual jobs- and, with limited exceptions, this was indeed the only use of TCs that was permitted in Spain until 1984. As shown in Fig. 2, the proportion of TCs represented by seasonal jobs is fairly low, and has been virtually unaffected by the 1984 reform. What the reform has greatly affected is the incidence of TCs in non-seasonal jobs.

When covering general, non-seasonal jobs, TCs may be used as a screening device in cases in which the productivity of a job-worker pair is not directly observable upon hiring. In this perspective, job matches are interpreted as "experience goods", in the tradition of Jovanovic (1979, 1984). In a high-firing-cost scenario, the introduction of TCs would therefore provide employers with the adequate instrument for experiencing the quality of a match within the legal limit.[4]

But there are also reasons why, even in the presence of perfectly observable match quality, employers may rely on TCs simply as a cheaper and more flexible factor of production within the whole legal duration limit, as TCs involve lower termination costs and generally pay lower wages

[4] PCs also allow for a legal probation period free of firing costs, which ranges between two weeks and 6 months for different categories of workers. TCs allow de facto a probation period of 3 years.

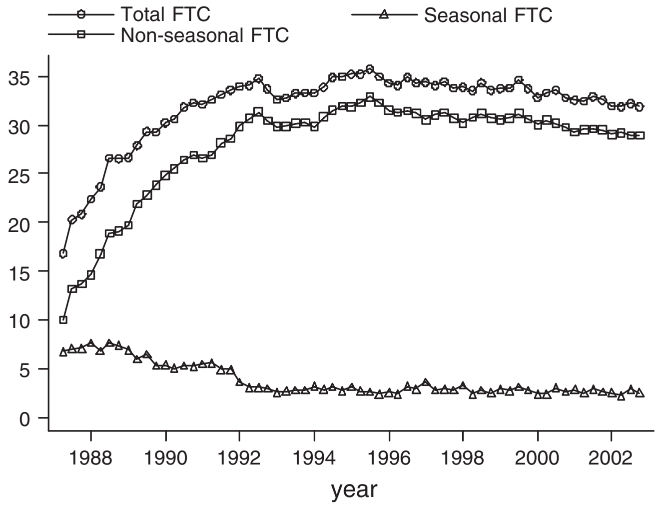

Fig. 2. The share of fixed-term contracts (%) in total employment, 1987–2002. Source: EPA.

than PCs.[5] However, as a worker on a TC is more likely to quit in order to accept a better match, firms using TCs are trading off lower labor costs with a higher quit rate and the risk of losing productive matches. Clearly, such trade-off depends on whether a temporary employee can exert a credible threat to quit, and thus her outside options, and on how easily she can be replaced.[6]

Below we characterize the optimal conversion rate of TCs, taking into account that a higher conversion rate increases future termination costs but at the same time prevents quits. In doing this, we assume for simplicity that worker productivity is observable upon start of contract, so as to abstract from screening motives in the use of TCs. This allows us to focus on the impact of worker outside options and binding legal limits on conversion rates. At the end of this section we will briefly discuss how the testable implications of our simple model can be affected in the presence of unobservable productivity and screening.

We assume that the productivity of a job-worker pair is match specific, and independent of the type of contract used. On the one hand, PCs are more expensive for the firm as they can only be destroyed with costly firing procedures. On the other hand, TCs cannot be used indefinitely, and the law establishes a legal duration limit upon which firms have the choice of converting a TC into a PC or destroying the job. But firms may choose to convert a TC into a PC well before the legal limit, depending on a worker's threat to quit.

We model firm decisions in discrete time. Firms post vacancies and meet workers according to some contact technology. Upon contact, a firm and a worker observe the productivity of a potential match, p_i, which can be thought of as a realization from a probability distribution. High enough realizations of pi become productive matches.[7] Any match is created as temporary,[8] and

[5] Jimeno and Toharia (1993) and De La Rica (2004) find that temporary workers in Spain earn approximately 10% less than permanent ones, after controlling for observable personal and job characteristics.

[6] See also Booth et al. (2002b) for a discussion of alternative roles of TCs and an application to UK data.

[7] In general equilibrium the acceptance rule would be endogenous (see Pissarides, 1990, chapter 6). Our partial equilibrium predictions would be robust to this endogenization.

[8] This assumption is empirically grounded (see Fig. 1). However, a richer model in which matches with high enough productivity are created as permanent from the start would simply provide a further reason for quits on TCs without affecting our comparative statics.

can last as such for at most two periods. The up-or-out clause applies at the end of the second period.

In the first period, a TC produces pi and pays a wage w_{1i}^{T}. At the end of the first period, the match may be hit by an idiosyncratic shock and become unprofitable, with an exogenous probability s, in which case the worker is dismissed with zero firing costs. If the match is still profitable, the firm can convert the TC into a PC (thus before the legal limit), and this happens with probability R_1. If a worker is not offered a permanent conversion, she may decide to quit the firm with probability q. Quits on TCs are modeled as a function of a worker's outside options a,

$$q = q(a), \tag{1}$$

with $\partial q/\partial a > 0$.[9] If the worker does not quit, the job remains temporary in the second period, during which it still produces pi and pays a wage w_{2i}^{T}. At the end of the second period, the match can again be hit by a negative shock with probability s. If not, the TC is either converted into a PC, with probability R_2, or it is destroyed, as expiry of the legal limit prevents the firm from further renewing the job on a temporary basis. If a contract is not renewed at this stage, from the firm's point of view it does not matter to further distinguish whether for the worker this is a voluntary separation (quit) or an involuntary one (layoff). In other words, the quit threat no longer matters to the firm at the end of the second period, when match continuation on a temporary basis is no longer an option. When a TC is converted into a PC, its productivity remains unchanged and the wage paid to the worker is w_i^P. PCs can still be destroyed with probability s, having paid a firing cost F. We finally assume that the quit rate on permanent jobs is zero[10] and that wages are exogenous.

Having said this, the values to the firm of temporary jobs of tenure 1 and 2 periods are defined by the following Bellman equations, respectively:

$$\begin{aligned} J_{1i}^{T} = p_i - w_{1i}^{T} + \frac{1}{1+r}\big[&sV + (1-s)R_1 J_i^P + (1-s)(1-R_1)qV \\ &+ (1-s)(1-R_1)(1-q)J_{2i}^{T}\big] \end{aligned} \tag{2}$$

$$J_{2i}^{T} = p_i - w_{2i}^{T} + \frac{1}{1+r}\left[sV + (1-s)R_2 J_i^P + (1-s)(1-R_2)V\right] \tag{3}$$

where V represents the value of a vacancy, J_{Pi} represents the value of the job if covered by a PC and r is the discount rate.

Similarly, the value to the firm of a permanent job is

$$J_i^P = p_i - w_i^P + \frac{1}{1+r}\left[s(V-F) + (1-s)J_i^P\right]. \tag{4}$$

[9] We are assuming that quits are unaffected by current wages, and that conversion rates are the only instrument employers can use in order to prevent quits (see Güell, 2000 for a similar framework).

[10] In principle, workers on permanent jobs may be searching for better job worker matches, but as new matches would start as temporary, the quit rate on permanent jobs would be lower than that on temporary ones. As the qualitative conclusions of the model would not be affected, we assume for simplicity that such quit rate is zero.

We finally assume free entry of vacancies in this economy, so that in equilibrium $V=0$; which in turn yields the following expressions for J_{1i}^T, J_{2i}^T and J_i^P:

$$J_{1i}^T = p_i - w_{1i}^T + \frac{1}{1+r}\left[(1-s)R_1 J_i^P + (1-s)(1-R_1)(1-q)J_{2i}^T\right] \tag{5}$$

$$J_{2i}^T = p_i - w_{2i}^T + \frac{1}{1+r}(1-s)R_2 J_i^P \tag{6}$$

$$J_i^P = \frac{1+r}{r+s}\left(p_i - w_i^P\right) - \frac{1}{r+s}sF. \tag{7}$$

Upon creation of a match, the firm maximizes its lifetime value with respect to the conversion rates R_1 and R_2. At the end of the first period, a firm computes the optimal "early" renewal rate R_1^*. This is defined by the following first order condition, obtained by differentiating (5) with respect to R_1:

$$J_i^P - (1-q)J_{2i}^T = 0. \tag{8}$$

As neither J_i^P nor J_{2i}^T depend on R_1, this delivers a corner solution for R_1, i.e. $R_1^*=0$ if $J_i^P<(1-q)J_{2i}^T$ and $R_1^*=1$ if $J_i^P\geq(1-q)J_{2i}^T$. Intuitively, if the value of a PC is higher than the value of a period-2 TC, weighted by its survival probability $(1-q)$; then it is optimal to offer the worker a permanent conversion. In other words, $\Pr(R_1^*=1)=\Pr\,[J_i^P\geq(1-q)J_{2i}^T]$. It can be shown that $\Pr\,[J_i^P\geq(1-q)J_{2i}^T]$ increases with p_i, as match productivity has a stronger impact on the value of permanent rather than temporary jobs- the former being more likely to be destroyed than the latter. Therefore the probability of an early conversion also increases with p_i, i.e. firms are willing to prevent quits on more productive jobs by offering workers an early conversion. Firing costs F affect R_1^* by reducing $\Pr\,[J_i^P\geq(1-q)J_{2i}^T]$, and thus the probability of an early conversion. Finally, workers' outside options, a, affect R_1^* through the behavior of quits. In particular, an increase in a raises q and thus $\Pr\,[J_i^P\geq(1-q)J_{2i}^T]$, which also raises the probability of an early conversion, R_1^*.

One period later, a similar decision is taken with respect to R_2, which is defined by the following first order condition, obtained by differentiating (7) with respect to R_2:

$$J_i^P = 0. \tag{9}$$

As J_i^P does not depend on R_2, this would also give a corner solution $R_2^*=0$ if $J_i^P<0$ and $R_2^*=1$ if $J_i^P\geq 0$. Clearly, if the lifetime value of a PC is positive, it is more profitable to covert a TC into a PC rather than opening a new vacancy, whose equilibrium value is zero. In other words $\Pr(R_2^*=1)=\Pr(J_i^P\geq 0)$. As $\Pr(J_i^P\geq 0)$ increases with match productivity pi and decreases with firing costs F, the probability of a later renewal increases with match productivity and decreases with firing costs.

To summarize, if $J_i^P<0$; conditions (8) and (9) imply $R_1^*=R_2^*=0$. If $J_i^P\geq 0$, then $R_2^*=1$. In particular, if $J_i^P\geq(1-q)J_{2i}^{\mathrm{T}}$, $R_1^*=1$ and viceversa. Therefore, three scenarios can arise. First, matches with $J_i^P<0$ are never converted. Second, matches with $0\leq J_i^P<(1-q)J_{2i}^T$ are only converted at the legal limit. Finally, matches with $J_i^P\geq(1-q)J_{2i}^T$ are converted before the legal limit.

Having said this, we may expect the following predictions:

1. There should be significant heterogeneity in the timing of conversion that we observe in the data. Depending on the values of J_i^P and J_{2i}^T, temporary contracts that are never converted could coexist with both early and late conversions.

2. Higher productivity increases the likelihood of a conversion at all, but in particular it raises the likelihood of an early conversion (scenario three above) with respect to the one of a late conversion (scenario two). This effect is due to the higher impact of p_i on J_i^P than on J_{2i}^T.
3. Following a similar argument, higher firing costs reduce the likelihood of a conversion at all, and, conditional on conversion, they make late renewals more likely than early renewals.
4. Finally, better worker outside options and bargaining power increase the likelihood of early conversions, and leave unchanged that of late conversions.

These predictions will be tested by estimating a duration model of temporary employment with flexible duration dependence, and comparing early and late conversion rates for the same type of workers and across workers with different characteristics, namely productivity, firing costs and outside options. In our data, productivity can be proxied by skills, firing costs are determined by the institutional environment, and outside options are proxied by skills and sectoral unemployment rates.

The simple framework presented here can be extended in a number of ways (see Appendix B for details). The extension that is probably most insightful for our purposes would allow for imperfect information about worker productivity and thus the use of TCs as screening devices (see also Engellandt and Riphahn (2005) for an application on Swiss data). In a high-firing-cost scenario, TCs would provide employers with the adequate instrument for experiencing the quality of a match during the maximum legal limit. Under this hypothesis, TCs that display high productivity are later converted on a permanent basis. Permanent conversions due to successful screening may happen at any time during the first three years of an employer-worker relationship, although we expect "early" conversions (well before expiry of the three years legal limit) to be more likely, since presumably the screening period should not take as long as three years. In other words, as soon as a job match is perceived to be productive enough, a firm may have a sufficient incentive to promote a temporary worker, instead of keeping him/her in a TC for the entire legal duration. Using the notation of our model, early renewals for successful screening can be modeled by an increase in R_1, everything else held constant. Firing for screening reasons could be modeled by raising the probability of exogenous layoffs on TCs (say $\tilde{s}$) above the corresponding layoff probability on PCs (s). Conditional on a worker successfully completing the screening period, the role of p_i, F and a on conversion rates remains the same as in the simple model illustrated above.

3. The data

The data used in this paper is drawn from the Spanish Labor Force Survey (Encuesta de la Población Activa), which is carried out every quarter on a sample of some 60,000 households. Since the second quarter of 1987, the EPA is a rotating panel, in which each household can be surveyed for a maximum of six consecutive quarters. Each quarter a new cohort of households is selected, and one sixth of existing households leave the sample. The EPA is designed to be representative of the total Spanish population, and contains very detailed information on labor force status of individuals within each household. Labor force transitions can be studied by linking consecutive information on the same individuals, available for all cohorts selected since 1987:2.[11]

[11] For a more detailed description of the EPA see: http://www.ine.es/dacoin/dacoinci/epalsti02.htm.

Table 1
Quarterly transitions across labor market states

		quarter $t+1$			
		NE	PC	new TC	same TC
quarter t	NE	96.62	0.48	2.91	
	PC	2.20	96.32	1.48	
	TC	16.26	5.70	13.93	64.11

Transition rates are computed according to the distribution of individuals across labor market states at quarter $t+1$, conditional on their status at quarter t. Source: EPA.

Our sample includes individuals belonging to cohorts that entered the survey between 1987:2 and 2002:4, covering more than a full cycle of the Spanish economy. We select all respondents who completed six quarterly interviews, and declared to hold a TC in any of the interviews.

In order to give a flavor of labor market transitions in our sample, Tables 1 and 2 report quarterly and yearly transition probabilities across three labor market states: non-employment, permanent employment, and temporary employment. Both tables display extremely strong persistence in the non-employment and the permanent employment states. As expected, the temporary employment category displays significant turnover, although most of such mobility represents reshu-ing across TCs, as shown in the bottom row of Table 2. In our duration model, we concentrate on individual transitions out of the first TC that is observed during the survey period. This leaves us with 162,092 temporary employment spells. The duration of each contract is constructed using self-reported information from the various quarterly interviews. Given that no contract identifier is supplied, in order to follow each single TC across interviews we rely on information concerning (i) the type of contract held; and (ii) the uncompleted duration of the present contract. The type of contract held can be permanent or temporary. The uncompleted duration of the present contract is expected to rise across interviews with calendar time, and to drop to zero whenever there is a contract switch. We therefore consider a spell of temporary employment as completed when either we observe a change in the type of contract or a drop in the uncompleted duration of the present contract.[12]

Roughly two thirds of temporary employment spells that we observe started during the survey period. The remaining third started before the worker was selected for the survey, so that we need to condition on the length of temporary employment at the first interview date, using once more the information on the elapsed duration of the current contract that is reported at the first interview.[13] Until the end of 1998, the self-reported elapsed duration up to the interview date is measured in months if it is lower than one year, and in years otherwise. Starting in 1999, such information is directly reported in months.

Either method has clear drawbacks. For the period 1987–1998, reported uncompleted durations are simply equal to the integer of $m/12$, where m represents the true duration in months, so that whenever the reported elapsed duration is 1 year, this means anything between

[12] We also computed the duration of fixed term contracts according to a more restrictive definition of a single spell. In particular, we considered a spell as completed when either (i) there is a change in the type of contract, or (ii) there is a drop in the uncompleted duration of the present contract, or (iii) there is a change in the sector where the worker is employed. No appreciable change was detected with respect to the definition given in the main text, which is the one we adopt in the empirical analysis reported here.

[13] See also Güell and Petrongolo (2005, Section 3) for a more detailed discussion of limitations in using the self-reported elapsed duration for constructing spells of temporary employment.

Table 2
Yearly transitions across labor market states

		year $t+1$			
		NE	PC	new TC	same TC
year t	NE	93.50	1.18	5.31	
	PC	6.01	91.15	2.85	
	TC	22.98	12.30	44.01	20.71

Transition rates are computed according to the distribution of individuals across labor market states at quarter $t+4$, conditional on their status at quarter t. Source: EPA.

12 and 23 months; whenever it is 2 years, this means anything between 24 and 35 months, and so on. Such data bunching problem could be eliminated by focusing only on entrants into temporary employment, who do not have any rounded measure of elapsed duration attached. However, this would only allow us to observe the time pattern of the conversion probability for at most six quarters of duration, and would leave us without any information on the behavior of the hazard towards the legal duration limit of TCs.

We therefore choose to exploit information on all spells, and correct for bunching in the following way. We convert all durations in quarters, which implies that any individual whose elapsed duration is 4 quarters or longer reports contract duration $\tilde{j}$, which is a multiple of 4, and to which corresponds a non-rounded duration $j \in \{\tilde{j}, \tilde{j}+1, \tilde{j}+2, \tilde{j}+3\}$. We thus assume that j is a random draw from a uniform distribution with discrete support $\{\tilde{j}, \tilde{j}+1, \tilde{j}+2, \tilde{j}+3\}$.[14] All observations with $\tilde{j} \geq 4$ are therefore assigned an elapsed duration $\tilde{j}$, $\tilde{j}+1$, $\tilde{j}+2$, or $\tilde{j}+3$ with equal 1/4 probabilities.

While for the period 1987–1998, elapsed durations are heavily bunched but we are given a clear rounding method, for the later period elapsed durations are in principle not bunched, as they are directly reported in months, but probably subject to some form of subjective rounding, whose magnitude is unknown ex ante. Indeed, we observe some small heaps in the distribution of uncompleted durations in correspondence of multiples of twelve months, and in particular at 12, 24 and 36 months. On the one hand, aggregating monthly durations up to quarters alleviates this problem. On the other hand, heaps in correspondence of 12, 24 and 36 months would not systematically bias our estimates of the baseline hazard towards multiples of 12 months, as what may be rounded is only the elapsed duration at the first interview date, to which one needs to add the non rounded ongoing duration during the survey period in order to obtain the total contract duration. We therefore simply measure elapsed contract duration at the first interview date converting the reported duration in quarters. Given that different rounding methods apply to our data before and after 1998, and that we deal with them in different ways, we estimate our duration models separately for the periods 1987–1998 and 1999–2002.

Each spell of temporary employment can terminate with a new TC, a PC, joblessness, or it can be censored if the worker is last observed holding the TC at the sixth interview. The proportion of TCs that terminates with a permanent conversion started around 18% at the beginning of our sample period and has declined monotonically until 1997 (6%), experiencing some recovery thereafter, as depicted in Fig. 3.

[14] Note that the assumption of uniform distribution is not restrictive, as j measures the elapsed uncompleted contract duration, and not the duration for which the contract is initially signed.

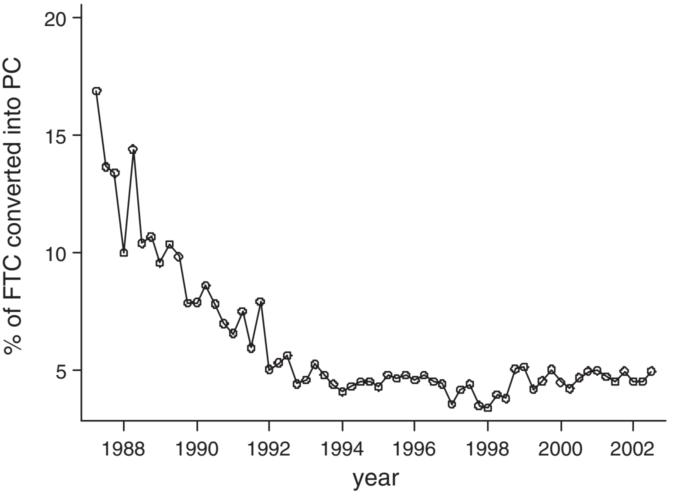

Fig. 3. The proportion of fixed-term contracts converted into permanent ones, 1987–2002. Source: EPA.

Given that we cannot use an employer identifier, we are not sure that new PCs observed in the survey are conversions of previous TCs with the same employer, rather than newly-created jobs elsewhere in the economy. However, as Fig. 1 shows, the fact that over 90% of all new contracts registered at employment offices have been fixed-term would suggest that the vast majority of PCs that we observe in the survey are created through conversions of TCs.

Table 3 reports the distribution of observed spells, according to their destination state. The figures reported suggest that, at relatively short durations, TCs are more likely to end up into non-employment. As duration proceeds, the probability of non-employment decreases, while the chances of permanent employment increase. The table also shows evidence of some TCs continuing beyond the legal limit of 3 years.[15] While there may be some imperfect compliance by employers shortly after the 3 years legal limit, we believe that durations much longer than 3 years should mostly reflect measurement error. We therefore treat all durations longer than 14 quarters as censored at 14 quarters.

It should be noted at this stage that not all TCs in Spain are subject to a legal limit of three years. In particular, contracts that fall into this category are "General" TCs (also known as "Employment promotion TCs") and "Training and Apprenticeship" TCs. Unfortunately the EPA does not allow to single out these two categories among other types of TCs ("Seasonal", "Probation", "Substitution" or "Task or Service" contracts) for the whole sample period. In particular, for the period 1987–1991, TCs are disaggregated into three categories ("Training", "Seasonal", and "Other"). Starting in 1992, more detail is given for the "Other" category, which is further disaggregated into "Probation", "Substitution", "Task or Service", and "Other". The "General" TCs are classified in the category "Other" (see also Dolado et al., 2002a, 2002b).

[15] In 1993, TCs could be extended for a fourth year (see Appendix A, Table A, note 7).

Table 3
The duration distribution of fixed-term contracts, by state of exit

Duration (quarters)	NE	PC	New TC	Same TC	Total No. of spells
1	54.33	10.13	13.69	21.85	47,622
2	34.67	8.73	38.80	17.80	38,684
3	28.81	10.67	37.92	22.59	20,751
4	19.53	11.92	46.28	22.27	16,295
5–8	15.53	12.89	27.03	44.55	23,101
9–12	15.90	20.78	22.68	40.64	7,775
>12	13.16	13.63	21.95	51.26	7,864
Total No. of spells	54,306	18,023	46,673	43,090	162,092

Each row sums to 100, with each entry giving the probability to exit into any of the four states, conditional on the contract duration. All our rounded elapsed durations $\tilde{j}$ are replaced with random draws from a uniform distribution with discrete support $\{\tilde{j}, \tilde{j}+1, \tilde{j}+2, \tilde{j}+3\}$. Source: EPA.

The crucial question for the interpretation of our estimates is then how the inclusion of contracts other than “General” and “Training” TCs in our sample may affect our estimates. Clearly, the inclusion of other types of contracts that do not have a three-year legal limit would lower our estimate of the three-year spike. This can be seen more clearly by looking at the conversion pattern of each type of contract for the period 1992–2002 (for which a relatively more disaggregate information on type of contract is available in the EPA). Table 4 shows raw conversion rates by duration for each type of TC. The categories that have clear spikes at 9–12 quarters (approaching the 3-year duration limit) are the seasonal and the probation ones (that do not account for a large share of temporary employment anyway) and the “Other” category. We are thus being conservative in estimating the three-year spike. Were we able to single out “General” TCs for the whole sample period, we would have found an even higher spike.

Explanatory variables included in our regressions are individual characteristics such as gender, age, education, and marital status. Year dummies (referring to the year in which the individual obtained a conversion or, in case of censoring, to the year in which she was last interviewed) are also included in order to capture any time pattern in conversion probabilities across the Spanish business cycle. Finally, sector dummies and the sectoral unemployment rate

Table 4
Permanent conversion rates by duration and type of contract, 1992–2002

duration (quarters)	Training and Apprent	Seasonal	Other	Probation	Substitution	Task or Service
1	5.55	3.04	9.16	16.77	5.28	6.08
2	4.86	5.57	8.54	12.47	7.31	6.14
3	5.81	6.83	10.70	10.44	9.42	7.22
4	10.54	9.81	11.98	12.96	9.77	8.14
5–8	10.89	7.55	13.82	8.70	9.61	7.51
9–12	13.67	24.75	21.48	20.00	14.64	10.17
>12	9.83	6.80	6.82	0.00	2.10	4.71
All durations	7.12	4.79	10.63	13.48	7.12	6.91

All our rounded elapsed durations $\tilde{j}$ are replaced with random draws from a uniform distribution with discrete support $\{\tilde{j}, \tilde{j}+1, \tilde{j}+2, \tilde{j}+3\}$. Source: EPA.

Table 5
Sample characteristics of temporary workers

	NE	PC	new TC	same TC	Total sample
Female	45.38	39.99	35.23	41.32	40.95
Age 16–24 yrs	41.24	35.96	41.74	41.51	40.87
Age 25–34 yrs	26.87	33.12	30.64	28.01	29.08
Age 35–44 yrs	15.94	16.49	15.76	16.53	16.21
Age 45+yrs	15.95	12.86	11.86	13.22	13.85
No qualification	14.97	8.66	8.05	10.52	11.17
Primary education	28.84	28.27	26.87	26.76	27.80
Secondary education	46.39	47.47	54.73	45.97	48.92
University education	9.52	13.98	10.28	15.82	11.95
Married	40.09	40.57	37.93	36.90	38.95
Agriculture	17.66	4.96	7.29	5.29	10.03
Manufacturing	15.66	22.23	22.15	18.48	19.06
Construction	15.93	12.92	18.86	19.68	17.48
Services	50.75	58.31	51.69	55.81	53.44
Average unemp. rate	12.54	10.89	13.09	11.13	12.19
Total No. of spells	54,306	18,023	46,673	43,090	162,092

All entries (except the average unemployment rate) indicate the percentage of workers with a given characteristic in the sample. Standard deviations in parenthesis. Source: EPA.

(measured at the start of the survey period or at the start of the TC if this happened later) should capture the effect of overall labor market performance, if any, on the conversion of contracts. Average sample values of these variables are reported in Table 5, for both the whole sample and each type of destination.

4. Econometric specification

The panel structure of the data set described requires a discrete time hazard function approach, as outlined in Narendranathan and Stewart (1993) and Jenkins (1995). Suppose that the transition out of temporary employment is a continuous process with hazard

$$\theta_i(t|x_i) = \lambda(t)\ \exp\ (x_i{}'\beta), \tag{10}$$

where $\lambda(t)$ denotes the baseline hazard, x is a vector of time-invariant explanatory variables, and β is a vector of unknown coefficients. The discrete time hazard denotes the probability of a spell of temporary employment being completed by time $t+1$, given that it was still continuing at time t. The discrete time hazard is therefore given by

$$h_i(t|x_i) = 1-\exp\left\{-\int_t^{t+1}\theta_i(u|x_i)du\right\} = 1-\exp\{-\exp(x_i{}'\beta)\gamma(t)\} \tag{11}$$

where

$$\gamma(t) = \int_t^{t+1}\lambda(u)du \tag{12}$$

denotes the integrated baseline hazard. We do not specify any functional form for $\gamma(t)$, and estimate the model semiparametrically.

The (log) likelihood contribution of a spell of length d_i is

$$L_i = c_i \ln h_i(d_i|x_i) + \sum_{t=1}^{d_i-1} \ln\,[1 - h_i(t|x_i)] = c_i \ln\{1 - \exp[-\exp(x_i'\beta)\gamma(d_i)]\} - \sum_{t=1}^{d_i-1} \exp\,(x_i'\beta)\gamma(t), \quad (13)$$

where c_i is a censoring indicator that takes the value 1 if d_i is uncensored and zero otherwise.

We need to adapt the likelihood contribution (13) to our stock sample. As we observe spells of temporary employment that started before the survey period, we use self-reported information to find out the quarter in which these spells begun, and we condition transition rates on the length of temporary employment at the first interview date. Suppose that an individual i enters the survey after j_i quarters of temporary employment and holds the TC for another k_i quarters, for a total duration $d_i = j_i + k_i$, that can be either censored or uncensored. The individual likelihood contribution becomes

$$L_i = c_i \ln h_i(j_i + k_i x_i) + \sum_{t=j_i+1}^{j_i+k_i-1} \ln\{1 - h_i(t|x_i)\}$$

$$= c_i \ln(1 - \exp[-\exp\{x_i'\beta\}\gamma(j_i + k_i\,|)]) - \sum_{t=j_t+1}^{j_i+k_i-1} \exp\{x_i'\beta\}\gamma(t). \quad (14)$$

The baseline hazard can be estimated non-parametrically by maximizing the log-likelihood $L = \sum_{i=1}^{n} L_i$ with respect to the $\gamma(\cdot)$ terms and the β vector. The vector of controls x_i includes a number of individual and job-related characteristics, which are treated as time invariant. Appendix C explains in detail how this empirical specification can be brought to the data when the available measure of duration is bunched.

We next make standard extensions to the econometric model outlined. First, as TCs can terminate with the conversion into a PC or alternative states, we need to consider a competing risk model, that distinguishes exits into permanent employment from exits into alternative states. It can be shown that, if distinct destinations depend upon disjoint subsets of parameters, the parameters of a given cause-specific hazard can be estimated by treating durations finishing for other reasons as censored at time of exit (see Narendranathan and Stewart, 1993). We therefore treat all temporary employment spells that end in a new TC or in non-employment as censored at the time the first contract is terminated. Having said this, the semi-parametric hazard specification (14) used for the single-risk model can be applied for the permanent job hazard.

Finally, we control for the effect of possibly omitted regressors in the exit from fixed-term employment by conditioning the hazard rate on an individual's unobserved characteristics, summarized into a random disturbance v. The conditional (discrete time) hazard rate is then written as

$$h_i(t|x_i, v_i) = 1 - \exp[-\exp(x_i'\beta + v_i)\gamma(t)] \quad (15)$$

with v_i independent of x_i and t. Note however that, in a competing risk framework, allowing for a random disturbance term in each of the cause-specific hazards requires an additional assumption, namely the independence of these disturbance terms across the cause-specific hazards.[16]

The conditional likelihood contribution for the ith individual is the given by $L_i|v_i = c_i \ln h_i(j_i + k_i|x_i, v_i) + \sum_{t=j_i+1}^{j_i+k_i-1} \ln\{1 - h_i(t|x_i, v_i)\}$. The unconditional likelihood contribution (that depends on observable regressors only) is obtained by integrating the conditional one over v_i:

$$L_i = \int \left\{ c_i \ln h_i(j_i + k_i|x_i, v_i) + \sum_{t=j_i+1}^{j_i+k_i-1} \ln[1 - h_i(t|x_i, v_i)] \right\} f(v_i) dv_i. \quad (16)$$

Among potential functional forms for *f(vi)*, a very convenient candidate is the gamma distribution, which delivers a closed form solution for (16) and therefore avoids numerical integration (see Lancaster, 1979; see also Han and Hausman, 1990; Dolton and O'Neill, 1996, for an application of gamma-distributed unobserved heterogeneity to discrete time hazard models).

Under these assumptions the individual likelihood contribution is given by

$$L_i = \ln\left\{ \left[1 + \sigma^2 \sum_{t=j_i+1}^{j_i+k_i-1} \exp\ (x_i'\beta)\gamma(t) \right]^{-1/\sigma^2} - c_i \left[1 + \sigma^2 \sum_{t=j_i+1}^{j_i+k_i} \exp\ (x_i'\beta)\gamma(t) \right]^{-1/\sigma^2} \right\}, \quad (17)$$

where σ^2 is an extra parameter to be identified.

5. Empirical results

We move on to estimating the econometric model outlined in the previous Section, for the determinants of worker transitions from temporary to permanent employment. The results of our estimates are reported in Table 6. These estimates refer to the sample period 1987–1998, for which we have a consistent measure of contract duration. Separate estimates for the later period are reported further down in Table 10. Two specifications of our regression equation are provided. In the first one we do not allow for unobserved heterogeneity among individuals. In the second one we control for the effect of possibly omitted regressors by allowing for a Gamma-distributed disturbance term.

The effect of several individual characteristics on conversion probabilities are fairly standard, and consistent with previous results obtained from logit estimates (see Alba, 1998). Column I of Table 6 shows that the probability of a permanent conversion increases with age up to prime age and stays constant afterwards. Being married positively affects the probability of obtaining a permanent contract. Gender and education have the expected effects on conversion rates, although they are not significantly different from zero. Industry dummies show that conversion rates are highest in services and lowest in construction. Time fixed-effects imply in turn a roughly monotonically decreasing trend in the proportion of TCs being converted on a permanent basis. Such trend is stronger in the first half of the sample period and then fades away

[16] The alternative approach would be to assume perfect correlation (as opposed to zero correlation) between the cause-specific disturbance terms (see Narendranathan and Stewart, 1993, for a discussion of advantages and disadvantages of the two methods).

Table 6
Maximum likelihood estimates of the transition from temporary to permanent employment: 1987:2–1998:4

	I	II
Characteristics		
Female	−0.019 (0.018)	−0.015 (0.021)
Age 25–34 yrs	0.194 (0.023)	0.225 (0.025)
Age 35–44 yrs	0.152 (0.030)	0.191 (0.036)
Age 45+yrs	0.135 (0.033)	0.170 (0.041)
Secondary education	−0.014 (0.021)	−0.022 (0.025)
University education	0.015 (0.032)	0.015 (0.037)
Married	0.101 (0.022)	0.120 (0.026)
Manufacturing	0.108 (0.037)	0.085 (0.056)
Construction	−0.216 (0.023)	−0.280 (0.052)
Services	0.231 (0.037)	0.252 (0.055)
Year 1988	−0.085 (0.047)	−0.138 (0.058)
Year 1989	−0.333 (0.045)	−0.456 (0.058)
Year 1990	−0.520 (0.047)	−0.693 (0.057)
Year 1991	−0.490 (0.048)	−0.707 (0.058)
Year 1992	−0.678 (0.040)	−0.896 (0.056)
Year 1993	−0.675 (0.042)	−0.885 (0.072)
Year 1994	−0.765 (0.044)	−1.005 (0.075)
Year 1995	−0.729 (0.044)	−0.958 (0.069)
Year 1996	−0.863 (0.040)	−1.109 (0.062)
Year 1997	−1.091 (0.047)	−1.372 (0.064)
Year 1998	−1.122 (0.047)	−1.414 (0.059)
Year 1999	−1.099 (0.071)	−1.350 (0.085)
Log log unemployment rate	−0.271 (0.057)	−0.337 (0.103)
Base line hazard steps		
Step 1	0.075 (0.007)	0.082 (0.018)
Step 2	0.074 (0.007)	0.090 (0.020)
Step 3	0.068 (0.007)	0.091 (0.020)
Step 4	0.094 (0.009)	0.138 (0.029)
Step 5	0.078 (0.008)	0.124 (0.028)
Step 6	0.061 (0.007)	0.097 (0.023)
Step 7	0.072 (0.008)	0.110 (0.024)
Step 8	0.105 (0.013)	0.111 (0.026)
Step 9–11	0.055 (0.006)	0.095 (0.023)
Step 12	0.147 (0.017)	0.214 (0.050)
Step 13–14	0.068 (0.007)	
σ^2		1.421 (0.110)
Mean log-likelihood	−0.358	−0.353
No. of obs.	125,077	125,077

(1) Standard errors in parenthesis; (2) Source: EPA.

in the late 1990s, consistently with what we observed in the raw data of Fig. 3. Finally, sectoral unemployment has a negative and significant impact on conversion rates. As lower unemployment implies better outside opportunities for temporary workers in search for better jobs, it enables them to more credibly threat their employer in case of low conversion prospects. This is in line with prediction (4) of Section 2. Very similar results (nor reported here) were obtained when using time-varying unemployment rates instead of time invariant. This is not surprising, in the light of the relatively strong persistence of Spanish unemployment at the quarterly frequencies.

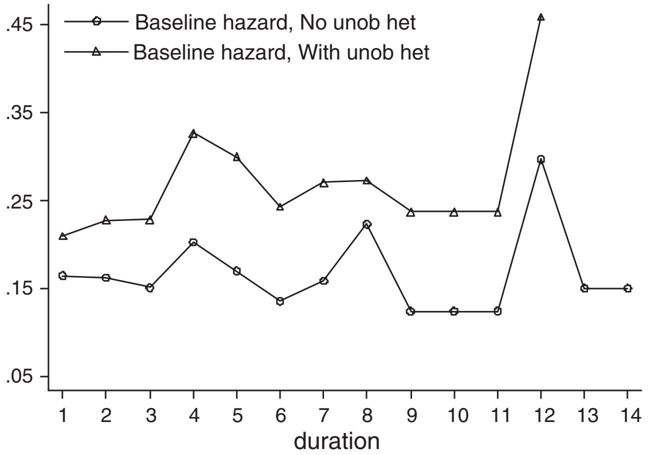

Fig. 4. Predicted hazard for the transition from TC to PC, 1987–1998 (see Table 6). Reference category: male, not married, aged 16–24, secondary education, employed in services, started TC in 1987.

The quarterly steps of the baseline hazard are reported at the bottom of Table 6. In the estimates provided we impose a constant hazard across steps 9–11 and steps 13–14, respectively.[17] Above 8 quarters of contract duration, step 12 was the only one that was individually identified. As step 12 coincides with the 3-year legal limit of TCs, the relatively higher density of completed spells at this duration allowed us to identify this step separately from adjacent ones.

The parallel estimation that controls for the effect of unobserved heterogeneity is represented in column II of Table 6. The positive and significant variance of the Gamma distributed disturbance shows that there is some residual heterogeneity among individuals, which is not properly accounted for by included regressors. However, the partial effect of most regressors remains practically unchanged if compared with the case where no unobserved heterogeneity is accounted for, as does the global fit of the regression. As there is no major difference between the estimates of column I and II,[18] and the additional restrictions embodied in specification II seem largely unnecessary, in the regressions that follow we do not allow for unobserved heterogeneity in our estimates.

The predicted hazards corresponding to regressions I and II of Table 6 are plotted in Fig. 4 for a typical temporary worker (single male, aged 16–24, with completed secondary education, employed in the service sector). Controlling for the presence of unobserved heterogeneity in regression II simply scales upward the whole hazard, as it is reasonable to expect, but hardly changes its overall time pattern. It can be noted that, with both specifications, the hazard has some spikes at durations around one, two and three years. This denotes substantial heterogeneity in the time pattern of conversion rates and is consistent with prediction (1) of

[17] We first attempted to estimate the fully unrestricted model with 14 baseline steps and found that steps 9–11 were not separately identifiable, and similarly for steps 13 and 14. See Appendix C for a formal discussion of identification problems.

[18] The only change from column I is that step 13 and 14 are not even jointly identified (and when we attempted to identify them, the corresponding coefficient was virtually zero and the others very close to those reported in column II of Table 6).

Table 7
Maximum likelihood estimates of the transition from temporary to permanent employment: 1987:2–1998:4; High and Low education

	I	II
	High education	Low education
Characteristics		
Female	−0.043 (0.021)	0.027 (0.032)
Age 25–34 yrs.	0.190 (0.025)	0.166 (0.047)
Age 35–44 yrs.	0.125 (0.043)	0.140 (0.050)
Age 45+yrs.	0.212 (0.062)	0.114 (0.050)
University education	0.038 (0.027)	–
Married	0.124 (0.028)	0.070 (0.036)
Manufacturing	−0.106 (0.078)	0.199 (0.070)
Construction	−0.419 (0.074)	−0.101 (0.060)
Services	−0.004 (0.078)	0.364 (0.070)
Year 1988	−0.174 (0.060)	0.027 (0.071)
Year 1989	−0.380 (0.060)	−0.272 (0.071)
Year 1990	−0.575 (0.057)	−0.449 (0.071)
Year 1991	−0.518 (0.058)	−0.458 (0.075)
Year 1992	−0.738 (0.064)	−0.588 (0.071)
Year 1993	−0.755 (0.088)	−0.545 (0.096)
Year 1994	−0.771 (0.093)	−0.749 (0.096)
Year 1995	−0.729 (0.080)	−0.741 (0.085)
Year 1996	−0.864 (0.073)	−0.869 (0.090)
Year 1997	−1.091 (0.062)	−1.122 (0.077)
Year 1998	−1.109 (0.060)	−1.252 (0.080)
Year 1999	−1.135 (0.093)	−1.036 (0.145)
Log unemployment rate	−0.338 (0.142)	−0.210 (0.146)
Base line hazard steps		
Step 1	0.078 (0.024)	0.082 (0.025)
Step 2	0.082 (0.025)	0.073 (0.020)
Step 3	0.078 (0.024)	0.063 (0.019)
Step 4	0.115 (0.035)	0.076 (0.023)
Step 5	0.085 (0.026)	0.078 (0.024)
Step 6	0.070 (0.021)	0.055 (0.017)
Step 7	0.075 (0.023)	0.0793 (0.026)
Step 8	0.114 (0.036)	0.108 (0.035)
Step 9–11	0.063 (0.020)	0.049 (0.018)
Step 12	0.152 (0.050)	0.168 (0.054)
Step 13	0.074 (0.023)	0.065 (0.021)
Step 14	0.074 (0.023)	0.065 (0.021)
Mean log-likelihood	−0.360	−0.353
N. of obs.	79,598	45,478

(1) Standard errors in parenthesis; (2) Source: EPA.

Section 2. As one would expect, TCs are more likely to be converted at integer yearly durations than otherwise.[19]

We checked the significance of those spikes using a Wald test for the equality of adjacent baseline hazard steps. Using the estimates from column 1 of Table 6, we found that, at durations

[19] Note that minimum durations of TCs are always multiple of quarters, and multiple of years for “General” TCs from 1992 onwards (see Appendix A). Moreover, starting in 1992, the EPA contains information on the length of contracts being signed, which displays clear spikes at 1, 2, and 4 quarters.

around one year, the baseline hazard at 4 quarters is significantly higher than both the one at 3 quarters ($\chi^2=70.97$, against the critical value $\chi^2(1, 0.05)=3.84$), and the one at 5 quarters ($\chi^2=27.69$). At durations around two years, the baseline hazard at 8 quarters is significantly higher than both the one at 7 and the one at 9–11 quarters ($\chi^2=13.68$ and $\chi^2=37.30$, respectively). Finally, at duration around three years, the baseline hazard at 12 quarters is significantly higher than both the previous and the later one ($\chi^2=37.30$ and $\chi^2=33.57$; respectively). Also, while the spikes at one and two years are not significantly different from each other ($\chi^2=2.25$), the one at three years is significantly higher than both of them ($\chi^2=13.09$ and $\chi^2=25.23$; respectively). Using the estimates from column 2 of Table 6, which control for unobserved heterogeneity, the spike at two years disappears, as the step at 8 quarters is not significantly different from adjacent ones, and we are left with an early and a late spike in permanent conversions, around durations of one and three years respectively. As with the previous estimates, the baseline hazard at three years is significantly higher than at both one and two years. Substantial time variation in conversion rates, and especially the coexistence of early and late spikes, are consistent with prediction (1) of Section (2).

Different population groups have different employment prospects and unemployment rates, which affect their outside options and thus their bargaining power on temporary jobs. In particular, skilled workers have lower unemployment rates than the less-skilled (see Dolado et al., 2002b), and Spanish women have higher unemployment rates than males (see Azmat et al., in press). We thus estimate separate duration models of temporary employment for men and women, the skilled and the unskilled.

We first split our sample along the educational dimension, and define as skilled all workers who have completed secondary education. Table 7 shows that while skilled women have lower conversion rates than skilled men, no significant gender differences can be detected among the less-skilled. The steps of the baseline hazard are shown in the lower part of the Table, and the predicted hazard is plotted in Fig. 5. As expected, the predicted hazard at most durations is higher for educated workers than for the less-skilled. However, the later spikes, especially the one at three years, are relatively more important for the less-skilled than for the skilled. In

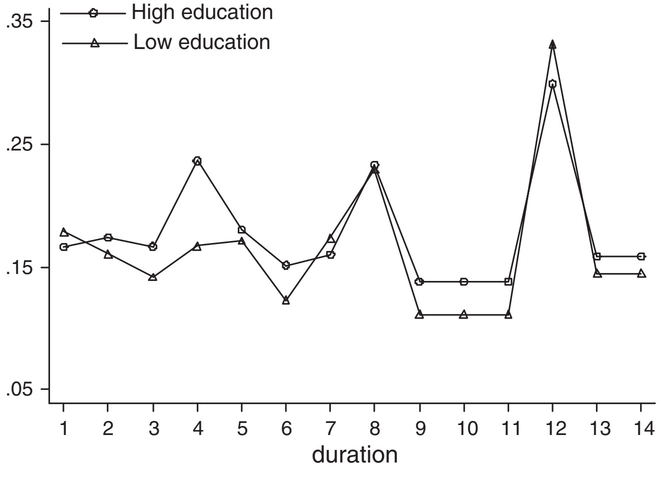

Fig. 5. Predicted hazard for the transition from TC to PC, high and low education samples, 1987–1998 (see Table 7). Reference category: secondary education (for high education), male, not married, aged 16–24, employed in services, started TC in 1987.

particular, there is really no early spike for the less-skilled, as the predicted hazard at 4 quarters is not significantly different from the one at 5 quarters and the one at 8 quarters is not significantly different from the one at 7 quarters. The fact that the time pattern of renewals is everywhere lower and more strongly increasing for the less-skilled than for the skilled is in line with prediction (2) of Section (2): skilled workers tend to occupy more productive job matches, which are thus more likely to be converted before the legal limit. Also, one would expect that the less skilled are generally in a weaker bargaining position than the skilled, as they may be more easily replaced. Moreover, in a high unemployment scenario, the skilled may take up unskilled jobs,

Table 8
Maximum likelihood estimates of the transition from temporary to permanent employment: 1987:2–1998:4; Males and Females

	I	II
	Males	Females
Characteristics		
Age 25–34 yrs	0.201 (0.029)	0.165 0.034
Age 35–44 yrs	0.171 (0.039)	0.085 (0.047)
Age 45+yrs	0.108 (0.044)	0.141 (0.054)
Secondary education	0.039 (0.027)	−0.117 (0.033)
University education	0.164 (0.046)	−0.153 (0.038)
Married	0.149 (0.028)	0.047 (0.032)
Manufacturing	0.120 (0.057)	0.052 (0.109)
Construction	−0.235 (0.049)	0.282 (0.145)
Services	0.194 (0.057)	0.260 (0.107)
Year 1998	−0.001 (0.059)	−0.226 (0.077)
Year 1989	−0.285 (0.060)	−0.413 (0.075)
Year 1990	−0.490 (0.060)	−0.560 (0.074)
Year 1991	−0.408 (0.061)	−0.605 (0.071)
Year 1992	−0.700 (0.057)	−0.633 (0.076)
Year 1993	−0.669 (0.083)	−0.649 (0.107)
Year 1994	−0.780 (0.080)	−0.701 (0.120)
Year 1995	−0.778 (0.074)	−0.625 (0.108)
Year 1996	−0.914 (0.064)	−0.758 (0.086)
Year 1997	−1.116 (0.063)	−1.021 (0.087)
Year 1998	−1.143 (0.057)	−1.066 (0.078)
Year 1999	−1.092 (0.105)	−1.103 (0.113)
Log unemployment rate	−0.261 (0.122)	−0.351 0.185
Base line hazard steps		
Step 1	0.071 (0.018)	0.069 (0.028)
Step 2	0.072 (0.018)	0.066 (0.026)
Step 3	0.066 (0.017)	0.062 (0.025)
Step 4	0.087 (0.022)	0.092 (0.037)
Step 5	0.073 (0.018)	0.074 (0.030)
Step 6	0.054 (0.014)	0.063 (0.026)
Step 7	0.062 (0.016)	0.076 (0.031)
Step 8	0.117 (0.032)	0.075 (0.032)
Step 9–11	0.047 (0.012)	0.059 (0.025)
Step 12	0.173 (0.049)	0.095 (0.042)
Step 13–14	0.071 (0.018)	0.053 (0.021)
Mean log-likelihood	−0.362	−0.351
No. of obs.	75,527	49,550

(1) Standard errors in parenthesis; (2) Source: EPA.

crowding out the less-skilled of their usual occupations (see Dolado et al., 2002b). In this sense, these results empirically support prediction (4) of Section 2. Screening and early conversions for successful workers are also more likely to apply to the skilled rather than the less-skilled, and this is again confirmed in our estimates.

Some gender differences in conversion rates are detected in Table 8. While age effects are similar for men and women, education has a positive effect on male conversion rates, but a negative effect on female ones, and this could explain the non-significant effect found in Table 6. As education presumably enhances productivity and a worker's outside options, we find that it has the expected impact on male conversion rates but not on female ones, as if other unmeasured factors such as, say, labor market attachment, were more relevant than observable human capital for women's promotions. It seems moreover that, in the interim period between the two reforms, conversion rates keep falling for males, while stabilizing for females. The unemployment rate has similar qualitative impact on conversion rates across genders, if anything stronger for females.

The baseline hazard steps for the regressions by gender are reported in the second half of Table 8, and the corresponding predicted hazards are plotted in Fig. 6. In general, the baseline hazard is higher for males than for females. This would be consistent with higher outside options for males, and thus more credible threats to quit the current TC for better opportunities in the labor market. However, the fact that the three-year spike in conversion rates is significantly higher than both the one-and the two-year spike for men, while for women all three spikes are not significantly different from one another, is not entirely consistent with this story. This would be probably easier to rationalize as the consequence of lower human capital accumulation on the job by females than by men.

Finally, we assess how the 1994 and 1997 reforms affected the time pattern of conversions. The 1994 reform was aimed at reducing the applicability of "General" TCs and at enhancing the conversion rates for labor market groups with supposedly poorer labor market prospects. This mostly took the form of payroll tax reductions for newly-hired workers under permanent contracts (see Table A). Such incentives to the conversion of TCs can be modeled in our framework as an increase in the productivity of PCs, or a reduction in firing costs. The 1997 reform reinforced the 1994 trends by further restricting the use of "General" TCs, and

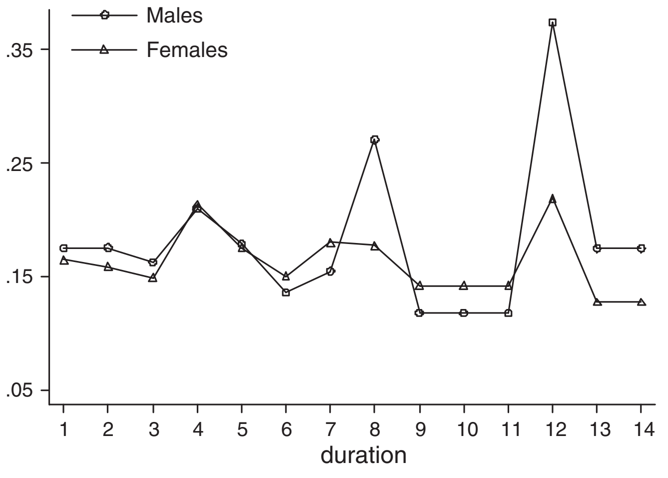

Fig. 6. Predicted hazard for the transition from TC to PC, male and female samples, 1987–1998 (see Table 8). Reference category: male/female, not married, aged 16–24, secondary education, employed in services, started TC in 1987.

Table 9
Maximum likelihood estimates of the transition from temporary to permanent employment: Pre 1994 and Post 1994

	I	II
	1987:2–1994:1	1994:3–1998:4
Characteristics		
Female	−0.053 (0.021)	0.056 (0.029)
Age 25–34 yrs	0.199 (0.026)	0.133 (0.031)
Age 35–44 yrs	0.207 (0.032)	0.007 (0.030)
Age 45+yrs	0.181 (0.041)	0.020 (0.041)
Secondary education	−0.026 (0.027)	0.036 (0.035)
University education	0.203 (0.040)	−0.234 (0.051)
Married	0.100 (0.027)	0.099 (0.030)
Manufacturing	0.094 (0.060)	0.009 (0.100)
Construction	0.067 (0.065)	−0.705 (0.043)
Services	0.225 (0.064)	0.098 (0.085)
Year 1988	−0.108 (0.046)	
Year 1989	−0.364 (0.046)	
Year 1990	−0.544 (0.046)	
Year 1991	−0.521 (0.044)	
Year 1992	−0.682 (0.056)	
Year 1993	−0.646 (0.083)	
Year 1994	−0.728 (0.113)	
Year 1995		0.025 (0.045)
Year 1996		−0.110 (0.049)
Year 1997		−0.357 (0.061)
Year 1998		−0.438 (0.082)
Year 1999		−0.309 (0.114)
Unemployment rate	−0.432 (0.149)	−0.378 (0.185)
Base line hazard steps		
Step 1	0.059 (0.019)	0.026 (0.008)
Step 2	0.046 (0.015)	0.039 (0.011)
Step 3	0.044 (0.014)	0.035 (0.010)
Step 4	0.051 (0.016)	0.063 (0.018)
Step 5	0.043 (0.014)	0.050 (0.015)
Step 6	0.032 (0.011)	0.044 (0.013)
Step 7	0.047 (0.016)	0.036 (0.012)
Step 8	0.085 (0.028)	0.042 (0.016)
Step 9–11	0.027 (0.009)	0.039 (0.013)
Step 12	0.135 (0.045)	0.027 (0.019)
Step 13–14	0.049 (0.016)	0.026 (0.009)
Mean log-likelihood	−0.430	−0.280
N. of obs.	63,113	59,257

(1) Standard errors in parenthesis; (2) Source: EPA.

introducing a new typology of PC with lower firing costs.[20] We noted above that, despite the reforms, the share of temporary employment did not fall after 1994 (see Fig. 2), but at least stabilized after one decade of sustained increase. Also, the proportion of TCs being converted into PCs stabilized in 1994 and slightly increased since 1998 (see Fig. 3). We next document this

[20] For the effects of the 1997 reform on permanent employment, see Kugler et al. (2003). See also García-Pérez and Muñoz-Bullón (2003) for an analysis of employment transitions in the 1990s for the youth labor market.

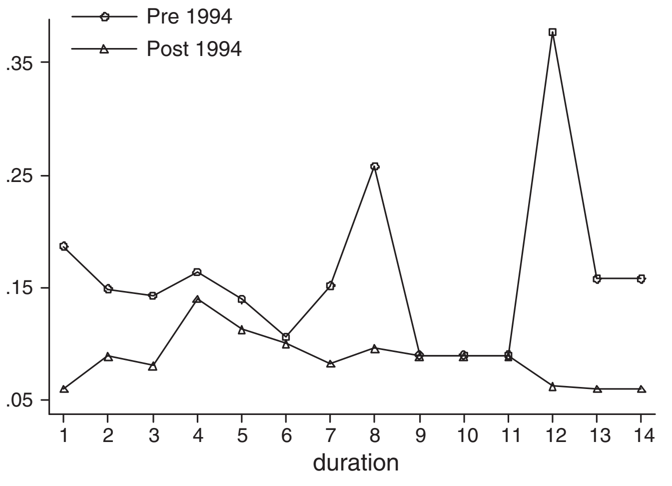

Fig. 7. Predicted hazard for the transition from TC to PC, contracts started before and after 1994 (see Table 9). Reference category: male, not married, aged 16–24, secondary education, employed in services, started TC in 1987 (for pre 1994 sample) and started TC in 1994 (for post 1994 sample).

trend in conversion rates, and check whether such overall tendency conceals diverging patterns for different labor market segments.

We split our sample into two subperiods, corresponding to different institutional environments. These are 1987:2-1994:1 and 1994:3-1998:4.[21] Temporary spells are allocated to these subperiods according to their starting quarter, or the first survey quarter if the contract had already started at the first survey date. Although there was a reform in 1997, we provide pooled estimates for the post 1994 period for two reasons. First, the 1997 reform did not imply any major discontinuity with respect to the 1994 reform as far as the conversion of TCs was concerned, and basically strengthened the incentives to permanent conversions of TCs. Second, the post 1997 period would be rather short, from 1998:1 to 1998:4, and would not allow us to identify the baseline hazard steps for durations longer than one year.

In Table 9 we report results for the pre and the post 1994 periods. Our estimates clearly show that permanent conversion prospects of women, the less educated and younger workers have improved after 1994. The female dummy switches from negative and significant in the first sub-period, to positive and significant in the second one, and the reverse is true for the university education dummy. Conversion rates are reduced for those aged 25–34 and even more older workers. Interestingly, before 1994 conversion rates are highest for the middle age category 35–44, but they drop to the same level as for the 16–24 category with the reform. Targeting subsidies to the conversion of contracts for women (in occupations in which they are under-represented) and young workers seems to have been effective in enhancing their prospects of accessing permanent employment. Also, conversion rates after 1994 have strongly deteriorated in construction.

Clearly, the time pattern of conversions is greatly affected after the 1994 reform, as shown in the lower part of Table 9 and in Fig. 7. Before 1994, clear spikes can be detected

[21] As the reform was passed in May 1994, it is not clear whether the old or the new legislation applies to contracts signed exactly in 1994:2, and we therefore drop them from our sample.

in conversion rates around 1, 2 and 3 years, each of them being higher than the previous one at conventional significance levels. In particular, the permanent conversion probability for the reference worker after 3 years of temporary employment is twice as high as the one at one year. Interestingly, after the 1994 reform, there is a small spike in conversion rates at one year, and after that conversion rates decline steadily, without any later spike. One the one hand, it can be concluded that the 1994 reform has successfully affected the use of TCs in the sense of inducing employers to earlier rather than later conversions- consistently with predictions (2) and (3) of Section 2. On the other hand, it can be clearly noted that, except at durations of 9–11 quarters, the conversion rates after 1994 are always lower than the ones for the earlier period. While affecting the time pattern of conversions, the 1994 reform failed quite badly at pushing higher their average level. One possible reason for this can be found in the changing composition of temporary employment after 1994. The 1994 reform limited substantially the applicability of “General” TCs, which were those initially conceived by the legislator for being converted into PCs. With a declining share of “General” TCs in total temporary employment, and an increasing share of “Task or Service” contracts, it is probably not surprising that average renewal rates also declined.

Table 10
Maximum likelihood estimates of the transition from temporary to permanent employment: 1999:1–2002:4

Characteristics	
Female	−0.090 (0.034)
Age 25–34 yrs	0.035 (0.038)
Age 35–44 yrs	−0.228 (0.056)
Age 45+yrs	−0.255 (0.063)
Secondary education	0.112 (0.039)
University education	0.035 (0.041)
Marrried	0.079 (0.042)
Manufacturing	0.751 (0.265)
Construction	−0.418 (0.168)
Services	0.636 (0.213)
Year 2000	0.073 (0.048)
Year 2001	0.085 (0.088)
Log unemployment rate	−0.121 (0.290)
Base line hazard steps	
Step 1	0.016 (0.008)
Step 2	0.026 (0.013)
Step 3	0.023 (0.012)
Step 4	0.038 (0.020)
Step 5	0.031 (0.016)
Step 6	0.026 (0.014)
Step 7	0.026 (0.013)
Step 8	0.042 (0.022)
Step 9	0.032 (0.017)
Step 10	0.016 (0.008)
Step 11	0.015 (0.008)
Step 12	0.022 (0.012)
Step 13	0.016 (0.008)
Step 14	0.011 (0.006)
Mean log-likelihood	−0.402
N. of obs.	37,015

(1) Standard errors in parenthesis; (2) Source: EPA.

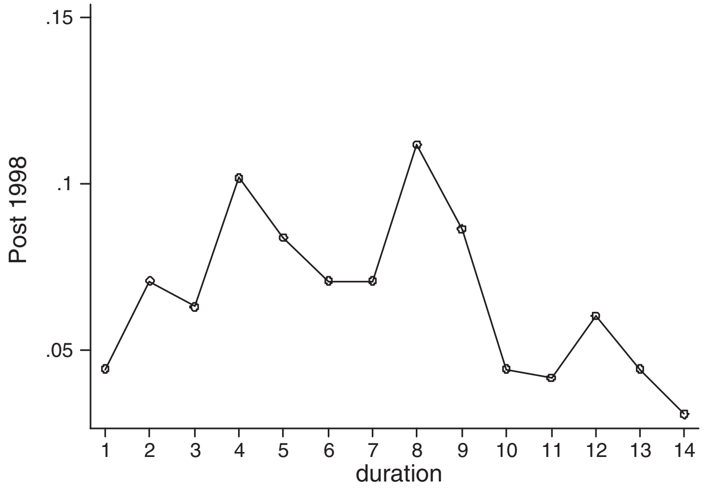

Fig. 8. Predicted hazard for the transition from TC to PC, 1999–2002 (see Table 10). Refeerence category: male, not married, aged 16–24, secondary education, employed in services, started TC in 1999.

For the last 3 years of our sample, corresponding to 1999–2002, the duration of temporary employment spells is measured differently from the previous period, as explained in detail in Section 3, and duration data are therefore not directly comparable. In particular, as durations are measured more precisely, we manage to separately identify all quarterly steps in the baseline hazard. We therefore provide separate estimates for this later period in Table 10 and Fig. 8. The most noticeable difference from the 1994–1998 period is age effects turning strongly negative from age 35, possibly due to the impact of the 1990s reforms, targeted at permanent employment prospects of youths. Also, the gender dummy is now negative and significant, while the impact of the unemployment rate becomes non-significantly different from zero. Finally, the level of conversion rates is lower than in the earlier subsample in correspondence of all durations. At the same time, the third spike becomes lower than the previous two. The same tendency towards lower but flatter conversion rates that we found for the 1994–1998 period is also detected for this final subsample.

6. Conclusions

Given that most accessions to permanent employment in Spain happen through TCs, the conversion of TCs into PCs is a key aspect of labor market segregation among Spanish workers and of the overall performance of the Spanish labor market. This paper has studied the determinants and the timing of the conversion of TCs into PCs in Spain using panel data for the period 1987–2002, to shed light on the potential of temporary employment as a stepping stone for stable, regular jobs. Specifically, we estimated a duration model for temporary employment, with flexible duration dependence for the permanent employment hazard.

We argued that the level and the timing of permanent promotions of TCs can be suggestive of different levels of workers' outside options. Using a simple framework for firm's choice of conversion rates, we have argued that conversion rates should be higher for workers who can credibly threat their employer to quit a temporary job. Finally, the pattern of renewals should be

Table A
Summary of Legislation on TCs in Spain

Month/Year	Contracts introduced[a]	Limits on duration: Minimum	Limits on duration: Maximum	Eligible workers[b]	Indemnities at termination	Subsidies to firms
03/1980	PC	indefinite			20 (45) days'wage per year worked if fair (unfair)[c]	
	"Causal" TC	limited[d]				
	TrainingTC	–	1 year	Secondary edu. degree or above obtained in thelast 2yrs.		
	Apprenticeship TC	–	2 years	Aged 16-18 (except for disabled) with < secondary edu.		
08/1984	General TC	6 months	3 years	Any unemployed worker[e]	12 days' wage per year worked[f]	
	Training TC	3 months	3 years	Secondary edu. degree or above obtained in the last 4yrs.		
	Apprenticeship TC	3 months	3 years	Aged 16-20 (except for disabled) with < secondary edu.		
07/1992	General TC	1 year	3 years[g]	Same as in 1984		
05/1994	General TC	1 year	3 years	Unemployed aged >45, disabled, LTU[h] Any firm starting a new economic activity	Same as in 1984	Fiscal incentives to hire eligible-UI recipients and to transform TC into PC for aged <25 or >45, females,[i] disabled
	Training TC	6 months	2 years	Same as in 1984		If they transform TC into PC
	Apprenticeship TC	6 months	3 years	Same as in 1984		
12/1997	New PC	indefinite		Unemployed aged 18-29 or >45, LTU, disabled, TC since 05/1997[j]	20 (33) days' per year worked if fair (unfair)[c]	Fiscal incentives to hire eligible workers, and to transform TC into New PC
	General TC	1 year	3 years	Disabled workers		If they transform TC into New PC
	Training TC	6 months	2 years	Same as in 1984		If they transform TC into New PC
	Apprenticeship TC	6 months	2 years	Same as in 1984		If they transform TC into New PC
07/2001	New PC	indefinite		Aged 16-29 o r>45, females,[h] unemp. w/ dur. >6months, disabled	Same as in 1997	Fiscal incentives to hire eligible workers; transform TC into New PC
	General TC	1 year	3 years	Disabled workers	8 days per year worked[f]	If they transform TC into New PC
	Training TC	6 months	2 years	Same as in 1984		If they transform TC into New PC
	Apprenticeship TC	6 months	2 years	Same as in 1984		If they transform TC into New PC

steeper for workers with lower outside options. We find that these predictions are broadly confirmed by the estimates of our duration model for temporary employment, which delivers a clear spike at the legal limit, and higher conversion rates in cases where workers supposedly have higher outside options.

The likely role of worker outside options is more clearly visible when we estimate separate conversion rates by education, as conversion rates are lower and more steeply increasing with tenure on temporary jobs for the less-skilled. The results are less clear-cut when we estimate separate conversion rates by gender. Men have higher conversion rates, and presumably better employment prospects than women in the Spanish labor market. However, the fact that the time pattern of conversions is less steep for women would point in the opposite direction if one were to assess conversion rates solely on the basis of workers' threat to quit. When comparing conversion rates for men and women it is likely that other factors play an important role over and above outside labor market options, namely noneconomic reasons for quits (see Manning, 2003, chapter 7), and thus different job attachment for the two genders.

Acknowledgements

We wish to thank Manuel Arellano, Samuel Bentolila, Olimpia Bover, Andrea Ichino, Alan Manning, Steve Nickell, Gilles Saint-Paul, the editors Andrea Ichino and Juan Dolado and two anonymous referees for very useful comments. M. Güell acknowledges financial support from the Spanish Ministry of Science and Technology & FEDER (Project BEC2003-00412) and B. Petrongolo acknowledges financial support from the Spanish Ministry of Education (Grant No. DGES PB97-0091).

Notes to Table A:

[a]If not stated otherwise, contracts introduced previously remain available.

[b]As a general principle for eli gibility for all types of TCs introduced in 1984, workers have a maximum 3 year limit of continuous temporary employment with one or several employers. If a worker has been continuously employed on TCs for 3 years, she needs to wait 12 months before being eligible for a new one. Similarly, for firms this limit binds for a given job, in the sense that they cannot cover the same position for more than 3 years with TCs. Also, firms cannot hire a temporary worker if they have reduced their workforce for economic reasons or their dismissals have been declared unfair in the previous 12 months (in practice, it is difficult to assess whether these rules have been effectively enforced).

[c]The fair (unfair) indemnity can be paid for a max. of 12 months (24 months). In case of unfair, forgone wages are al so paid. Workers can be fired for disciplinary reasons (without indemnity) or economic reasons (indemnites indicated in table). Workers always have the right to sue the employer if she disagrees with the dismissal case. Once the case is taken to court, it can be declared fair or unfair. Around 72% of cases that go to court are declared unfair (see Galdón-Sánchez and Güell, 2000).

[d]Different limits apply to different types of causal TCs. For a specific project: indeterminated but limited. For replacement: durati on of leave. For transitory production ci rcumstances: max 6 months in a year.

[e]Workers can be hired to undertake the normal activity of the firm or a any new economic activity.

[f]Indemnities to be paid upon expiry of contract. There is no right to sue the employer for unfair dismissal in this case.

[g]TCs which lasted 3 years and expire between march 3rd and december 31st, 1993, can be extended to a 4th year. After 4 years, firms get a subsidy if they convert the TC into a PC.

[h]LTU refers to long-term unemployed (12 or more months in unemployment).

[i]Females in professions or jobs in which they are underepresented.

[j]Royal Decree Laws 8/97 and 9/97 were approved in May before the December law. The respecti ve laws from each reform are: law 8/80 (Estatuto de los Trabajadores), law 32/84, law 22/92, laws 10/94 and 11/94; laws 63/1997, 64/97 and 66/97; law 12/2001.

Appendix A. The institutional background

Current legislation on labor contracts is contained in the Worker's Statute (Estatuto de los Trabajadores, ET) of 1980, which has since been modified on four occasions with the 1984, the 1994, 1997 and 2001 reforms. The ET of 1980 established priority to contracts of indefinite duration and allowed TCs only for jobs that were temporary in their nature (like for particular projects, e.g. constuction; or seasonal jobs, e.g. tourism). Some forms of training contracts for young, first-time job seekers were also allowed (apprenticeship contracts and training contracts). Other situations in which TCs were allowed was for eventual increases in demand or replacement of a permanent worker in case of absence or temporary suspension of contract. The ET also established the possibility for the Government to use TCs as an incentive to promote employment. Except in this last situation, a specific cause was generally required in order to sign a TC ("causal" TC).

The 1984 reform exploits this last possibility in an extreme way, and introduces flexibility by extending the applicability of TCs. It introduces a new "General" TC, as well as making training contracts more flexible. After the reform, any worker can be hired on a temporary basis without the requirement of a specific cause. This implies that for any job, employers can freely choose between a PC or a TC. The 1994, 1997 and 2001 reforms have restricted the applicability of TC and introduced subsidies for their conversion into PCs. Table A summarizes the relevant aspects of the Spanish legislation on temporary employment during the past 20 years. TCs can be characterized according to: i) limits on their duration (upon expiry, it is not possible to retain the worker under a TC: either the worker is promoted to a PC or dismissed); ii) eligibility conditions for workers; iii) indemnities at their termination; and iv) subsidies to firms.

Appendix B. Possible extensions to the theoretical frame-work

The simple framework of Section 2 can be extended in a number of ways. First, if wagesare endogenized, they presumably respond to productivity, firing costs, and workers' outside options. In this case changes in productivity induce changes in wages. In order to ensure $\partial R_2^* = \partial p_i > 0$ one simply needs wages adjusting less than one for one to productivity changes. In order to ensure $\partial R_1^* = \partial p_i > 0$, one also needs the impact of productivity on wages on PCs to be not too large with respect to the corresponding impact on wages on TCs. A natural way to endogenize wages in this framework would be to introduce a Nash sharing rule, in which workers appropriate a given share of the total surplus generated by each job match. In this case, the first condition is always satisfied, as wages have a component that is totally inelastic to productivity, and simply depends on unemployment income (see Pissarides, 2000, chapter 1). The second condition mostly rests on the magnitude of workers' share of total surplus: if this is the same on PCs and TCs (or lower on PCs than TCs), then the condition of interest is satisfied. Otherwise, one needs such difference to be not too large.

Higher firing costs lower wages to be paid on newly-created permanent jobs (see Pissarides, 2000, chapter 9), and in order to ensure $\partial R_1^*/\partial F$, $\partial R_2^*/\partial F < 0$ one needs again wages adjusting less than one for one to changes in F . Finally, higher outside options also increase wages, and $\partial R_1^*/\partial a > 0$ if w_{2i}^T is more responsive to outside options than w_i^P, such condition being sufficient but not necessary. Under the above mentioned Nash sharing rule, better outside options increase wages on each type of contract by increasing the value of unem-ployment. As temporary workers experience a higher probability of future unemployment (through expiry of legal limits and exogenous separations) than permanent workers, the value of unemployment (and thus the value of a) would have a stronger impact on w_{2i}^T than on w_i^P. Having said this, our comparative statics results would be robust to the endogenization of wages.

Second, one could allow labor productivity on temporary jobs to depend on R_1 and R_2, to take into account the effect of permanent employment prospects on worker effort and productivity. While this does not affect our comparative statics results with respect to F and a, a sufficient condition for this not to alter the comparative statics results with respect to p is that p is concave in both R_1 and R_2.

Finally, learning on-the-job and specific human capital accumulation could be introduced. In this case the value of job matches and conversion rates would rise over time and be highest at the legal limit.

Appendix C. Data bunching problems in estimated like-lihood functions

The generic $\gamma(j_i+k_i)$ term in Eq. (14) is represented by the product between a vector of duration dummies and a vector of associated coefficients, each of dimension 14, given that we identify at most 14 quarterly baseline hazard steps. Typically, individuals with contract duration equal to $d_i=j_i+k_i$ would have the d_i^{th} element in the vector of duration dummies equal to one, and the remaining 13 elements equal to zero. Non-censored spells of length d (with $c_i=1$ and $d_i=d$), allow therefore to identify the d^{th} element in the vector of coefficients, which represents the d^{th} baseline hazard step (see the second row of Eq. (14)). And this holds for $d^{th}=1, \ldots 14$, in principle allowing us to identify the whole baseline hazard.

In order to compute d_i^{th} for each individual i, we need to know j_i (the elapsed contract duration at the first interview date) and k_i (the contract duration during the survey period). While k_i is precisely observed during our whole sample period, j_i is precisely measured only in the 1999–2002 subsample.[22]

In the 1987–1998 subsample, we know j_i precisely only for those individuals who report $\tilde{j}\leq 3$, and for them the true j_i is simply equal to the reported value $\tilde{j}_i$. For those who report $\tilde{j}\geq 4$, the true j_i can be any integer between $\tilde{j}_i$ and $\tilde{j}_i+3$, and specifically we assume that j_i is a random draw from a uniform distribution with discrete support $\{\tilde{j}, \tilde{j}+1, \tilde{j}+2, \tilde{j}+3\}$, as described in Section 3: Thus for them total contract duration $d_i=j_i+k_i$ is also a random draw from a uniform distribution with discrete support $\{\tilde{d}_{i,}\ \tilde{d}_i+1, \tilde{d}_i+2, \tilde{d}_i+3\}$, with $\tilde{d}_i=\tilde{j}+k_i$. The corresponding vector of duration dummies will have four non-zero values, equal to bc each, in correspondence of $\tilde{d}$, $\tilde{d}+1$, $\tilde{d}+2$ and $\tilde{d}_i+3$. If some of these values are higher than 14, we censor them at 14 quarters, which implies adjusting the censoring indicator accordingly. Consider for example an uncensored spell with $k_i=4$ and $\tilde{j}_i=8$. The implied spell duration is therefore 12, 13, 14 or 15 quarters, with equal bc probabilities. In particular, this spell would be longer than 14 quarters with probability bc, and the associated censoring indicator is reduced from 1 to be.

This treatment of spells whose duration is bunched has consequences for identification of baseline steps associated to durations of 9–14 quarters, for which we need to rely on relatively long (and therefore bunched) elapsed durations at the first interviewdate. If completed, these spells all have non-integer duration dummies, and possibly non-integer censoring indicators. In other words, the γ (9)– γ(14) terms become more collinear than they would otherwise be, and their associated censoring indicator may become smaller, which makes it harder to identify them separately. We will come back to this issue in Section 5.

[22] Abstracting here from subjective rounding, which we discussed in Section 3.

References

Alba, A., 1998. How Temporary is Temporary Employment in Spain? Journal of Labor Research 19, 695–710.

Amuedo-Dorante, C., 2001. From "Temp-to-Perm: Promoting Permanent Employment in Spain. International Journal of Manpower 22, 625–647.

Amuedo-Dorante, C., 2000. Work Transitions into and out of Involuntary Temporary Employment in a Segmented Market: Evidence from Spain. Industrial and Labor Relations Review 53, 309–325.

Autor, D.H., Houseman, S.N., 2005. Do Temporary Help Jobs Improve Labor Market Outcomes for Low Skilled Workers? Evidence from Random Assignments, mimeo MIT, available at http://econ-www.mit.edu/faculty/download_pdf.php?id=1185.

Azmat, G., Güell, M., Manning, A., in press. Gender Gaps in Unemployment Rates in OECD Countries. Journal of Labor Economics.

Bentolila, S., Dolado, J., 1994. Labour Flexibility and Wages: Lessons from Spain. Economic, 54–99. (April).

Blanchard, O., Landier, A., 2002. The Perverse Effects of Partial Labor Market Reforms: Fixed Duration Contracts in France. Economic Journal, Features: Symposium on Temporary Work, F214–F244.

Booth, A., Dolado, J.J., Frank, J., 2002a. Symposium on temporary work. Introduction. Economic Journal Features: Symposium on Temporary Work, F181–F188.

Booth, A.L., Francesconi, M., Frank, J., 2002b. Temporary Jobs: Stepping Stones or Dead Ends? Economic Journal Features: Symposium on Temporary Work, F189–F213.

Bover, O., Gomez, R., 2004. Another Look at Unemployment Duration: Exit to a Permanent vs. a Temporary Job. Investigaciones Economicas XXVIII, 285–314.

Cahuc, P., Postel-Vinay, F., 2002. Temporary Jobs, Employment Protection and Labor Market Performance. Labour Economics 9, 63–91.

De La Rica, S., 2004. Wage Gaps between Workers with Indefinite and Fixed-term Contracts: The Impact of Firm and Occupational Segregation, mimeo, University of Bilbao. Available at: http://www.ehu.es/SaradelaRica/docs/fe2-final.pdf.

Dolado, J.J., García-Serrano, C., Jimeno, J.F., 2002a. Drawing Lessons from the Boom of Temporary Jobs in Spain. Economic Journal, Features: Symposium on Temporary Work, F270–F295.

Dolado, J.J., Jansen, M., Jimeno, J.F., 2002b. A Matching Model of Crowding-Out and On-the-Job Search (with an Application to Spain). CEPR Discussion Paper 3466.

Dolton, P., 1996. The Restart Effect and the Return to Full-Time Stable Employment. Journal of the Royal Statistical Society Series A 159 (2), 275–288.

Engellandt, A., Riphahn, R.T., 2005. Temporary Contracts and Employee Effort. Labour Economics 12 (3), 281–299.

Farber, H.S., 1999. Alternative and Part-Time Employment Arrangements as a Re-sponse to Job Loss. Journal of Labor Economics 17, S142–S169.

Galdón-Sánchez, J.E., Güell, M., 2000. Let's go to court! Firing costs and dismissal conflicts. Princeton University-IR Section Working Paper No. 444.

García-Pérez, J.I., Muñoz-Bullon, F., 2003. The nineties in Spain: too much Flexibility in the Youth Labour Market?" Universidad Carlos III, Madrid, Business Economics Series, Working Paper No. 3.

Grubb, D., Wells, W., 1993. Employment Regulation and Patterns of Work in EC Countries. OECD Economic Studies 0 (21), 7–58.

Güell, M., 2000. Fixed-term Contracts and Unemployment: An Efficiency Wage Analysis. Princeton University-IR Section Working Paper 433.

Güell, M., Petrongolo, B., 2005. "How binding are legal limits? Transitions from temporary to permanent work in Spain", UPF Working Paper 682. Available at http://www.econ.upf.es/~mguell/july05WP.pdf.

Han, A., Hausman, J., 1990. Flexible Parametric Estimation of Duration and Competing Risk Models. Journal of Applied Econometrics 5, 1–28.

Ichino, A., Mealli, F., Nannicini, T., 2005. Temporary Work Agencies in Italy: A Sringboard Toward Permanent Employment? Mimeo European University Institute. Available at http://www.iue.it/Personal/Ichino/springboard13.pdf.

Jenkins, S., 1995. Easy Estimation Methods for Discrete Time Duration Models. Oxford Bulletin of Economics and Statistics 57 (1), 129–137.

Jimeno, J., Toharia, L., 1993. The Effects of Fixed-term Employment on Wages: Theory and Evidence from Spain. Investigaciones Económicas 27, 475–494.

Jovanovic, B., 1979. Job Matching and the Theory of Turnover. Journal of Political Economy 87 (5), 972–990.

Jovanovic, B., 1984. Matching, Turnover, and Unemployment. Journal of Political Economy 92 (1), 108–122.

Kugler, A., Jimeno, J.F., Hernanz, V., 2003. Employment Consequences of Restrictive Permanent Contracts: Evidence from Spanish Labor Market Reforms. CEPR Discussion Paper 3724.
Lancaster, T., 1979. Econometric Methods for the Duration of Unemployment. Econometrica 47, 939–956.
Manning, A., 2003. Monopsony in Motion, Imperfect Competition in Labour Markets. Princeton University Press, Princeton, NJ.
Narendranathan, W., Stewart, M., 1993. Modelling the Probability of Leaving Unemployment: Competing Risks Models with Flexible Base-line Hazards. Applied Statistics 42, 63–83.
OECD, 1987. Employment Outlook.
OECD, 1993. Employment Outlook.
OECD, 1999. Employment Outlook.
OECD, 2002. Employment Outlook.
OECD, 1994. The OECD Jobs Study.
Pissarides, C., 1990. Equilibrium Unemployment Theory. Basil Blackwell, Oxford.
Pissarides, C., 2000. Equilibrium Unemployment Theory. Second edition. MIT Press, Cambridge, MA.
Toharia, L., 1996. Empleo y Paro en España: Evolución, Situación y Perspectivas. Economiaz 35, 36–61.