Centre de Referència en Economia Analítica

Barcelona Economics Working Paper Series

Working Paper nº 75

How binding are Legal Limits? Transitions from Temporary to Permanent Work in Spain

Maia Güell and Barbara Petrongolo

December 2003

HOW BINDING ARE LEGAL LIMITS? TRANSITIONS FROM TEMPORARY TO PERMANENT WORK IN SPAIN[¤]

MAIA GUELL
Universitat Pompeu Fabra
CEP (LSE), CEPR, CREA and IZA

BARBARA PETRONGOLO
London School of Economics
CEP (LSE), CEPR and IZA

Barcelona Economics WP n^0 75

December 2003

Abstract

In the mid-1980s, most accessions to permanent employment in several European countries have been through …xed-term contracts. This paper studies the duration pattern of …xed-term contracts and the determinants of their conversion into permanent ones in Spain, where the share of …xed-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 ‡exible duration dependence. Our estimated conversion rate has a clear, pronounced spike at 3 years of duration, coinciding with the legal maximum duration of these contracts, suggesting that some …xed-term contracts are only converted when there is no other way to retain the worker. Also, there is a spike around 1 year of duration, indirectly suggesting that some …xed-term contracts may be used as screening devices: workers who successfully pass the screening may obtain a conversion much before the legal duration limit.

Keywords: Fixed-term contracts, Duration models.
JEL Classi…cation: C41, J41, J60.

[¤]We wish to thank Manuel Arellano, Samuel Bentolila, Olimpia Bover, Alan Manning, Steve Nickell and Gilles Saint-Paul for very useful comments. M. Güell acknowledges …nancial support from the Spanish Ministry of Science and Technology (Project No. BEC2000-1026) and B. Petrongolo acknowledges …nancial support from the Spanish Ministry of Education (Grant No. DGES PB97-0091). Email addresses: maia.guell@upf.edu and b.petrongolo@lse.ac.uk.

Several European labor markets have been characterized by a wide use of permanent contracts with stringent and costly …ring regulations. In the mid-1980s, in order to …ght the high and persistent levels of unemployment, some European countries enhanced the ‡exibility of their labor markets by allowing employers to hire workers on a …xed-term basis, with negligible termination costs upon contract expiry. Typically, there exists a legal duration limit in the use of these contracts, after which an employer can either o¤er the worker a contract of undetermined duration or dismiss her. Since their introduction, …xed-term contracts have been widely used and they account 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 in‡ows and out‡ows between unemployment and employment, but there has not been a sizeable bene…cial impact on aggregate unemployment.

The consequences of the introduction of …xed-term contracts have raised interest and concern among both academics and policy-makers (see Booth et al. 2002 and OECD 2002). Some consensus has formed among economists that the introduction of …xed-term contracts does not necessarily increase employment, while creating dualism in the labor market (see, among others, Bentolila and Dolado 1994, Blanchard and Landier 2000 and Güell 2000). An important aspect of the use of …xed-term contracts is their pattern of promotion into regular contracts of inde…nite duration. Mixed employment e¤ects of the introduction of …xed-term contracts and rising dualism provide some clear signal that …xed-term contracts largely failed to provide workers with e¤ective "stepping stones" to permanent employment.

In this paper we study the determinants of the conversion of …xed-term contracts (FTCs) into permanent contracts (PCs) as well as the duration pattern of FTCs. 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 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 FTCs. This reform was somewhat more radical than in other European countries. In particular, while in some countries FTCs are restricted to some type of workers or sectors[1], the Spanish 1984 reform did not limit in any way the applicability of FTCs. At the same time, the 1984 reform set an "up or out" clause after three years of continuous employment in a FTC. 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 expansion of the late 1980s, more than 90% of newly created contracts have been FTCs, and this translated in a rapidly growing stock of temporary jobs, from 11% in 1983 to approximately 35% by the early 1990s, which is more than three times the European average (see OECD 1987, 1993 and Toharia 1997). But, at the same time, unemployment has 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 because only a small fraction of FTCs has 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 FTCs and offered fiscal incentives for their conversion into PCs (1994 reform). Later reforms (in 1997 and 2001) continued to limit the applicability of FTCs as well as offering incentives to convert FTCs into PCs (see table A in the Appendix for more institutional details).

There exists a growing literature which studies several aspects of the impact of FTCs on labor markets in OECD countries, with special reference to the Spanish case (see Dolado, García-Serrano and Jimeno 2002 for a comprehensive survey). However, there is an important

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

aspect which is to date largely underexplored in this literature, namely the study of the conversion of FTCs into PCs and its timing. This paper concentrates on these issues trying to shed some light on the kind of use that employers make of these contracts and the implications for the dualism of the labor market.

In order to understand dualism in the labor market, it is useful to distinguish between entry into and exit from "bad" jobs (i.e. temporary contracts). Given that most employer-worker relationships in Spain start on a temporary basis, the main source of dualism lies in the exit margin, i.e. the promotion of FTCs into PCs, which is 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 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 FTCs introduced in 1984, as well as analyze the effects of the later reforms.

The existing literature contains only a few contributions on renewal rates for Spain. Amuedo-Dorantes (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 shed light on the time pattern of conversions. Most existing studies on the determinants of individual conversion rates use logit speci…cations (Toharia 1996 and Alba 1998), which may prove rather in‡exible 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-Dorantes (2000), who estimates transitions out of temporary employment using EPA individual records from 1995:2 through 1996:2, and …nds 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 for di¤erent categories of workers, and assess the impact of both the 1994 and the 1997 temporary employment reforms.

1. Hypothesis

We consider below alternative uses of FTCs, and the implied time pattern of their renewal into PCs.[2] In doing this we implicitly assume that both the use of FTCs and their timing of promotion (if any) is driven by employer choices rather than worker preferences for temporary jobs, as a permanent contract is at least as desirable as a temporary one from a worker's point of view, for both job stability considerations and wage gains.[3] This is also clearly con…rmed by information on the reason for holding a FTC, contained in the EPA: between 1987 and 2002, as much as 85% of temporary workers reported that they were holding a FTC because they could not …nd a PC, and only 1% reported that it was because they did not want a PC.

FTCs can …rstly be used by employers for covering seasonal or casual jobs - and, with limited exceptions, this was indeed the only use of FTCs that was permitted in Spain until 1984. As shown in Figure ??, the proportion of FTCs represented by seasonal jobs is fairly low, and has been virtually una¤ected by the 1984 reform (if anything, it has fallen slightly since 1987). What the reform has greatly a¤ected is the incidence of FTCs in non-seasonal jobs.

[2] See Booth et al. (2002) for a discussion of the implications for wage di¤erentials between temporary and permanent workers. This paper cannot provide evidence on wage di¤erentials, as the EPA does contain information on wages.

[3] Jimeno and Toharia (1993) and De La Rica and Felgueroso (1999) …nd that temporary workers earn approximately 10% less than permanent ones, after controlling for observable personal and job characteristics.

When covering general, non-seasonal jobs, FTCs 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 FTCs would therefore provide employers with the adequate instrument for experiencing the quality of a match during the maximum legal limit of three years.[4] Under this hypothesis, temporary job-worker pairs which display high productivity are later renewed on a permanent basis. Permanent renewals due to successful screening may happen at any time during the first three years of an employer-worker relationship, although we expect "early" renewals (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.[5] 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 FTC for the entire legal duration. While cheaper in some respects (because of lower termination costs and wages), a continuous use of FTCs may discourage or delay any investment in specific human capital and reduce motivation and worker retention when the worker may effectively threat the employer.

But there are also reasons why employers may rely on FTCs simply as a cheaper and more flexible factor of production for the whole legal period of three years. This happens when the worker can exert no credible threat on the employer (due for example to the availability of close substitutes for her skills), and the job does not require significant investment in specific human capital. Under these circumstances, FTCs are likely to be renewed, if anything, upon expiry of the three year limit, after which there is no other legal way to retain the worker.

To summarize, we would expect therefore a significant spike in renewal rates around three years of contract duration.[6] This spike should be particularly important for "low quality"

[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. FTCs allow de facto a probation period of 3 years.

[5] See Varejao and Portugal (2002) for an alternative way of assesing the screening role of FTCs.

[6] Note that from 1995 onwards we may also find these types of spikes around 2 years, since the 1994 reform

temporary jobs (in which no specific human capital is required and/or the worker can easily be replaced). Also, earlier spikes are expected especially for "higher quality" jobs, as soon as the job match is perceived to be productive enough, as in "genuine" probation contracts. As alternative uses of FTC may apply differently to different categories of workers, to shed some light into these issues we separately estimate renewal rates for a number of labor market segments. Also, we analyze to what extent the limits to the use of FTCs, and the subsidies to their conversions into PCs that were introduced with the 1994, the 1997 and 2001 reforms have affected the time pattern of renewal rates.

2. 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 1987 (second quarter), 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.[7]

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 FTC in any of the interviews.

In order to give a flavor of labor market transitions in our sample, Tables 5.1 and 5.2 report quarterly and yearly transition probabilities across three labor market states: non-employment, permanent employment, and temporary employment. Both tables display extremely strong

limited at 2 years the maximum duration of some type of FTC (see table A in the Appendix for more details).

[7] For a more detailed description of the EPA see: http://www.ine.es/dacoin/dacoinme/inotepa.htm

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 reshuffling across FTCs, as shown in the bottom row of Table 5.2.

In our duration model, we concentrate on individual transitions out of the first FTC 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 FTC 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 fixed-term. 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.[8]

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. 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 and

[8] 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.

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 FTCs.

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 $j \in \left\{\tilde{j}; \tilde{j}+1; \tilde{j}+2; \tilde{j}+3\right\}$: Given this, we assume that j is a random draw from a uniform distribution with discrete support $\left\{\tilde{j}; \tilde{j}+1; \tilde{j}+2; \tilde{j}+3\right\}$[9]. 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 bc 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

[9] 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.

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 FTC, a PC, joblessness, or it can be censored if the worker is last observed holding the FTC at the sixth interview. The proportion of FTCs that terminated with a permanent renewal started around 18% at the beginning of our sample period and has declined monotonically until 1997 (6%), experiencing some recovery thereafter, as depicted in Figure ??. These proportions look slightly lower than those computed in Toharia (1996, Table 4), although they follow exactly the same trend up to the early 1990s. It is worth noticing however that the renewal rates computed here refer to the proportion of workers that hold a FTC at some point in time and hold a permanent one at the next interview, i.e. direct transitions from temporary to permanent employment. Toharia (1996) computes instead the proportion of permanent workers that held a FTC one year back. We prefer to look at direct switches between two subsequent interviews because yearly renewals may conceal additional labor market transitions.

Given that we cannot use an employer identifier, we are not sure that new PCs observed in the survey are renewals of previous FTCs with the same employer, rather than newly-created jobs elsewhere in the economy. However, as Figure ?? 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 renewals of FTCs.

Table 5.3 reports the distribution of observed spells, according to their destination state. The figures reported suggest that, at relatively short durations, FTCs 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 FTCs continuing beyond the legal limit of 3 years.[10] While there may be some imperfect

[10] In 1993, FTCs could be extended for a forth year (see Appendix, Table A, note 7).

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.

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 entered the survey, or the starting year of the FTC if this happened later) are also included in order to capture any time pattern in renewal probabilities across the Spanish business cycle. Finally, sector dummies and the sectoral unemployment rate (also measured at the start of the survey period or at the start of the FTC if this happened later) should capture the effect of overall labor market performance, if any, on the renewal of contracts. Average sample values of these variables are reported in Table 5.4, for both the whole sample and each type of destination.

3. 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

$$\theta_i(t|x_i) = \lambda(t)\exp(x_i'\beta)\,; \tag{3.1}$$

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

$$h_i(t|x_i) = 1 - \exp\left\{-\int_t^{t+1}\theta_i(u|x_i)du\right\} = 1 - \exp\{-\exp(x_i'\beta)\gamma(t)\} \tag{3.2}$$

where

$$\gamma(t) = \int_t^{t+1}\lambda(u)du \tag{3.3}$$

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

$$\begin{aligned} L_i &= c_i \ln h_i(d_i|x_i) + \sum_{t=1}^{d_i-1} \ln[1 - h_i(t|x_i)] \\ &= c_i \ln\{1 - \exp[-\exp(x_i'\beta)\gamma(d_i)]\} - \sum_{t=1}^{d_i-1} \exp(x_i'\beta)\gamma(t); \end{aligned} \tag{3.4}$$

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 (3.4) to our stock sample. As we observe spells of temporary employment that started before the survey period, and we can use self-reported information to find out the quarter in which these spells begun, we to 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 FTC 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

$$\begin{aligned} L_i &= c_i \ln h_i(j_i + k_i|x_i) + \sum_{t=j_i+1}^{j_i+k_i-1} \ln\{1 - h_i(t|x_i)\} \\ &= c_i \ln(1 - \exp[-\exp\{x_i'\beta\}\gamma(j_i + k_i)]) - \sum_{t=j_t+1}^{j_i+k_i-1} \exp\{x_i'\beta\}\gamma(t). \end{aligned} \tag{3.5}$$

The baseline hazard can be estimated non-parametrically by maximizing the log-likelihood $L = \sum_{i=1}^{n} L_i$ with respect to the $\gamma(.)$ terms and the β vector. The vector of controls x_i includes a number of individual and job-related characteristics, that are treated as time invariant, and are measured at the start of the fixed-term contract (or at the time of the first interview if the contract had previously started).

Note that when bringing this empirical specification to our data, the generic $\gamma(j_i + k_i)$ term 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 equation (3.5)). 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.[11]

In the 1987-1998 subsample, we know j_i precisely only for those individuals who report $\tilde{j}_i \leq 3$, and for them the true j_i is simply equal to the reported value $\tilde{j}_i$. For those who report $\tilde{j}_i \geq 4$, 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 $\left\{\tilde{j}_i, \tilde{j}_i + 1, \tilde{j}_i + 2, \tilde{j}_i + 3\right\}$, as described in Section 2. Thus for them total contract duration $d_i = j_i + k_i$ is also a random draw from a uniform distribution with discrete support $\left\{\tilde{d}_i, \tilde{d}_i + 1, \tilde{d}_i + 2, \tilde{d}_i + 3\right\}$, with $\tilde{d}_i = \tilde{j}_i + k_i$. The corresponding vector of duration dummies will have four non-zero values, equal to $\frac{1}{4}$ each, in correspondence of $\tilde{d}_i, \tilde{d}_i + 1, \tilde{d}_i + 2$ 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 $\frac{1}{4}$ probabilities. In particular, this spell would be longer than 14 quarters with probability $\frac{1}{4}$, and the associated censoring indicator is reduced from 1 to $\frac{3}{4}$.

This treatment of spells whose duration is bunched has consequences for identification of

[11] Abstracting here from subjective rounding, which we discussed in Section 2.

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 interview date. If completed, these spells all have non-integer duration dummies, and possibly non-integer censoring indicators. In other words, the $\gamma(9) - \gamma(14)$ 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 when commenting our estimation results in the next Section.

With these caveats in mind, we make standard extensions to the econometric model outlined. First, as FTCs can terminate with the conversion into a PC or alternative states, we need to consider a competing risk model, that distinguishes exit into permanent employment from exit 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 FTC or in non-employment as censored at the time the first contract is terminated. Having said this, the semi-parametric hazard specification (3.5) 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

$$h_i(t|x_i; v_i) = 1 - \exp[-\exp(x_i'\beta + v_i)\gamma(t)] \tag{3.6}$$

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.[12]

[12] The alternative approach would be to assume perfect correlation (as opposed to zero correlation) between

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:

$$L_i = \int \left\{ c_i \ln h_i(j_i + k_i|x_i; v_i) + \sum_{t=j_i+1}^{j_i+k_i-1} \ln[1 - h_i(t|x_i; v_i)] \right\} f(v_i)dv_i. \tag{3.7}$$

Among potential functional forms for $f(v_i)$, a very convenient candidate is the gamma distribution, which delivers a closed form solution for (3.7) and therefore avoids numerical integration (see Lancaster 1979; see also Han and Hausman, 1990, and 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

$$L_i = \ln \left\{ \left[1 + \sigma^2 \sum_{t=j_i+1}^{j_i+k_i-1} \exp(x_i'\beta)\,\gamma(t)\right]^{-1/\sigma^2} - c_i \left[1 + \sigma^2 \sum_{t=j_i+1}^{j_i+k_i} \exp(x_i'\beta)\,\gamma(t)\right]^{-1/\sigma^2} \right\}, \tag{3.8}$$

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

4. Empirical results

We move on to estimating the econometric model outlined in Section 3, for the determinants of worker transitions from temporary to permanent employment. The results of our estimates are reported in Table 5.5. 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 5.9. 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

the cause-specific disturbance terms (see Narendranathan and Stewart, 1993, for a discussion of advantages and disadvantages of the two methods).

we control for the e¤ect of possibly omitted regressors by allowing for a Gamma-distributed disturbance term.

The e¤ect of several individual characteristics on renewal probabilities are fairly standard, and consistent with previous results obtained from logit estimates (see Alba, 1998). Column I of Table 5.5 shows that the probability of a permanent renewal increases with age up to prime age and stays constant afterwards. Being married positively a¤ects the probability of obtaining a permanent contract, while gender and education do not. Industry dummies show that renewal rates are highest in services and lowest in construction. Time …xed-e¤ects imply in turn a roughly monotonically decreasing trend in the proportion of FTCs being renewed on a permanent basis. Such trend is stronger in the …rst half of the sample period and then fades away in the late 1990s, consistently with what we observed in the raw data of Figure ??. Finally, sectoral unemployment has a negative and signi…cant impact on renewal rates. As low unemployment implies better outside opportunities for temporary workers, it enables them to more credibly threat their employer in case of low renewals. This evidence is in line with a use of FTCs mainly driven by …rms' choices rather than workers' preferences for temporary employment.

The quarterly steps of the baseline hazard are reported at the bottom of Table 5.5. In the estimates provided we impose that the hazard is constant across steps 9-11 and across steps 13-14, respectively.[13] Above 8 quarters of contract duration, step 12 was the only one that was individually identi…ed. As step 12 coincides with the 3-year legal limit of FTCs, 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 e¤ect of unobserved heterogeneity is represented in column II of Table 5.5. The positive and signi…cant variance of the Gamma-distributed

[13] We …rst attempted to estimate the fully unrestricted model with 14 baseline steps and found that steps 9-11 were not separately identi…able, and similarly for steps 13 and 14. See Section 3 for a formal discussion of identi…cation problems.

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,[14] and the additional restrictions embodied in specification II seem largely unnecessary[15], 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 5.5 are plotted in Figure ?? 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.

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 5.5, we found that, at durations around one year, the spike 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 spike 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 spike 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).

[14] 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 as those reported in column II of Table 5.5).

[15] The restrictions are that (i) omitted heterogeneity can be adequately captured by a gamma-distributed disturbance, uncorrelated with observed regressors, and (ii) such disturbance is uncorrelated across risks.

Using the estimates from column 2 of Table 5.5, 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 renewals, 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.

As one would expect, FTCs are more likely to be renewed at integer yearly durations than otherwise.[16] Also, evidence on the baseline hazard may suggest that some FTCs are plausibly used as a screening device, and "successful" workers obtain a permanent renewal much before the legal limit. A spell of roughly one year seems in fact reasonable for adequately assessing the performance of a worker, and in order to retain those who pass the screening employers choose not to wait until the maximum legal limit of the contract. But there also seem to exist contracts that are only renewed upon expiry of the legal limit of three years: such contracts are probably used as a cheaper/more flexible option to adjust employment, and are only renewed when there is no other legal way to retain the worker. If anything, our estimates suggest that late renewals are more frequent than early renewals.

As alternative uses of FTCs may affect difference categories of workers in different ways, we run our estimates separately for men and women, the skilled and the unskilled. Some gender differences in renewal rates are detected in Table 5.6. While age effects are similar for men and women, education has a positive effect on male renewal rates, but a negative effect on female ones. In other words, the human capital accumulated through education does not enhance permanent promotions for females at is does for men, as if other unmeasured factors such as labor market attachment, were more relevant than observable human capital for women' promotions. It seems moreover that, in the interim period between the two reforms, renewal rates keep falling for males, while stabilizing for females. The unemployment rate has similar

[16] Note that minimum durations of FTCs are always multiple of quarters, and multiple of years for general FTCs from 1992 onwards (see Table A in the Appendix). 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.

qualitative impact on renewal rates across genders, if anything stronger for females.

The baseline hazard steps for these two regressions are reported in the second half of Table 5.6, and the corresponding predicted hazards are plotted in Figure ??. While the three-year spike in renewal rates is signi…cantly higher than both the one- and the two-year spike for men, for women all three spikes are not signi…cantly di¤erent from one another. If anything, this suggests that the screening use of FTCs applies more to female than male employment. Given low participation rates and high turn-over of Spanish women, a temporary employment spell may be used by employers in order to assess the degree of labor market attachment of their female employees.

We next split our sample along the educational dimension, and de…ne as skilled all workers who have completed secondary education. Table 5.7 shows that while skilled women have lower renewal rates than skilled men, no signi…cant gender di¤erences 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 Figure ??. 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 particular, there is really no early spike for the less-skilled, as the predicted hazard at 4 quarters is not signi…cantly di¤erent from the one at 5 quarters ($\hat{A}^2 = 0{:}21$) and the one at 8 quarters is not signi…cantly di¤erent from the one at 7 quarters ($\hat{A}^2 = 1{:}89$). One would expect that the less skilled are generally in a weaker bargaining position than the skilled, as they may be more easily replaced. Also, in a high unemployment scenario, the skilled may take up unskilled jobs, crowding out the less-skilled of their usual occupations (see Dolado, Jansen and Jimeno 2002). Screening and early renewal for successful workers therefore plausibly applies to the skilled rather than the less-skilled, and this is con…rmed in our estimates.

Finally, we assess how the 1994 and 1997 a¤ected renewals for targeted groups, and whether they have altered their time pattern. Recall that the 1994 reform was aimed at reducing the applicability of general FTCs and enhancing the renewal rates for labor market groups with supposedly poorer labor market prospects. The 1997 reform reinforced the 1994 trends, by introducing new subsidies for permanent renewals further restricting the use of general FTCs.[17] We noted above that, despite the reforms, the share of temporary employment did not fall after 1994 (see Figure ??), but at least stabilized after one decade of sustained increase. Also, the proportion of FTCs being converted into permanent ones stabilized in 1994 and slightly increased since 1998 (see Figure ??). We next document this trend in renewal rates, and check whether such overall tendency conceals diverging patterns for di¤erent labor market segments.

We split our sample into two subperiods, corresponding to di¤erent institutional environments. These are 1987:2-1994:1 and 1994:3-1997:3. Temporary spells are allocated to these subperiods according to their starting quarter, or the …rst survey quarter if the contract had already started at the …rst 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, and basically strengthened the incentives to permanent renewals of FTCs. 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 5.8 we report results for the pre and the post 1994 periods. Our estimates clearly show that permanent renewal prospects of women, the less educated and younger workers have improved after 1994. The female dummy switches from negative and signi…cant in the …rst sub-period, to positive and signi…cant in the second one, and the reverse is true for the university dummy. Renewal rates are reduced for those aged 25-34 and even more older workers. Interestingly, before 1994 renewal rates are highest for the middle age category 35-44, but they drop at the same level as for the 16-24 category with the reform. Targeting subsidies to the

[17] For the e¤ects of the 1997 reform on permanent employment, see Kugler et al. (2002). See also García-Pérez and Muñoz-Bullón (2003) for an analysis of employment transitions in the 1990s for the youth labor market.

renewal of contracts for women 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 renewals is greatly affected after the 1994 reform, as shown in the lower part of Table 5.8 and in Figure ??. Before 1994, clear spikes can be detected in renewal rates around 1, 2 and 3 years, each of them being higher than the previous one at conventional significance levels. In particular, the permanent renewal probability for the reference worker after 3 years of temporary employment is twice as high the one at one year. Interestingly, after the 1994 reform, there is a small spike in renewal rates at one year, and after that renewal rates decline steadily: any later spike has completely disappeared. One the one hand, it can be concluded that the 1994 reform has successfully affected the use of FTCs in the sense of inducing employers to earlier rather than later renewals. On the other hand, it can be clearly noted that, except at durations of 9-11 quarters, the renewal rates after 1994 are always lower than the ones for the earlier period. While affecting the time pattern of renewals, the 1994 reform failed quite badly at pushing higher their average level.

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 5.9. The most noticeable difference from the 1994-1998 period is age effects turning strongly negative from age 35, hinting once more at the impact of the 1990s reforms, targeted at permanent employment prospects of the youth. Also, the impact of the sectoral unemployment rate on renewal rates becomes non significantly different from zero, being already somewhat reduced after 1994 (see Table 5.8). The responsiveness of renewal rates to local labor market conditions is thus getting weaker throughout our sample period. Finally, comparing the predicted hazard rates in Figures ?? and ??, it can be noted that the level of renewal probabilities is further reduced in correspondence of all durations in the later period, consistently with the trend already observed during 1987-1998. Having acknowledged this, the negative duration dependence in renewal rates already observed for 1994-1998 is maintained during 1999-2002. In particular, the renewal spikes around one and two years of temporary employment are signi…cantly higher than the one at three years.

5. Conclusions

Given the record incidence of temporary employment in Spain and the low conversion rates of …xed-term contracts into permanent ones, temporary employment is the major source of labor market segregation among Spanish workers. This paper has studied the determinants and the timing of the conversion of FTCs 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. Speci…cally, we estimated a duration model for temporary employment, with ‡exible duration dependence for the permanent employment hazard.

We argued that the timing of permanent promotions of FTCs can be suggestive of alternative reasons why …rms opt for temporary hirings, other than for covering jobs whose underlying nature is temporary, as has been typically the case for Spain before 1984. On the one hand, FTCs can be used as a screening device when the productivity of a job-worker pair is not initially observable, and may be renewed into PCs as soon as the uncertainty is resolved. This happens when …rms perceive a real trade-o¤ between using FTCs and PCs. In other words, while FTCs are cheaper in several respects, they may discourage worker motivation, retention, and speci…c human capital investment if the worker can credibly threat the employer. On the other hand, for workers who cannot credibly threat their employers, and for jobs which do not require speci…c human capital, FTCs may simply be used as a cheaper alternative to PCs up to their legal duration limit of three years. Low conversion rates, mostly concentrated around the legal limit, would be in line with this second explanation, while earlier spikes in renewal would be more consistent with a screening story in the use of FTCs.

In our estimates, we find both early and late spikes in the renewal rates of FTCs, around durations of one and three years, respectively. The later spike is relatively more important for men and for the less skilled. If anything, the screening use of FTCs seems to apply more to women rather than men, most likely to assess the degree of job attachment of women, and to the skilled rather than the less-skilled, who can be more easily replaced by new temporary workers at the legal duration limit of their contracts.

Also, we detect some effects of the 1994 reform, which restricted the applicability of general FTCs and introduced incentives to firms for their renewals. After 1994, renewal prospects improve for women, the youth and for the less-skilled. Targeted subsidies seem to have been effective in enhancing transition to permanent employment. Finally, the 1994 reform successfully induced firms to earlier renewals: after 1994 predicted renewal hazards display a spike around one year of duration and monotonically fall afterwards, with no evidence of any later spike. However, the reform failed quite badly at raising the average renewal rate across durations, which remained around 5% for the whole post 1994 period.

Table 5.1: Quarterly transitions across labour 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

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

Table 5.2: Yearly transitions across labour 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

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

Table 5.3: 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

Notes. 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 $\tilde{d}$ are replaced with random draws from a uniform distribution with discrete support $\{\tilde{d}; \tilde{d}+1; \tilde{d}+2; \tilde{d}+3\}$: Source: EPA.

Table 5.4: 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

Notes. 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.

Table 5.5: Maximum likelihood estimates of the transition from temporary to permanent employment: 1994-1998.

	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)
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)		
¾^2			1.421	(0.110)
mean log-likelihood	-0.358		-0.353	
No. of obs.	125,077		125,077	

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

Table 5.6: Maximum likelihood estimates of the transition from temporary to permanent employment: 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)
constuction	-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)
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	

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

Table 5.7: Maximum likelihood estimates of the transition from temporary to permanent employment: 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 educaction	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)
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	

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

Table 5.8: Maximum likelihood estimates of the transition from temporary to permanent employment: pre 1994 and post 1994.

	I Pre 1994		II Post 1994	
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	

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

Table 5.9: Maximum likelihood estimates of the transition from temporary to permanent employment: Full sample after 1998.

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)
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	

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

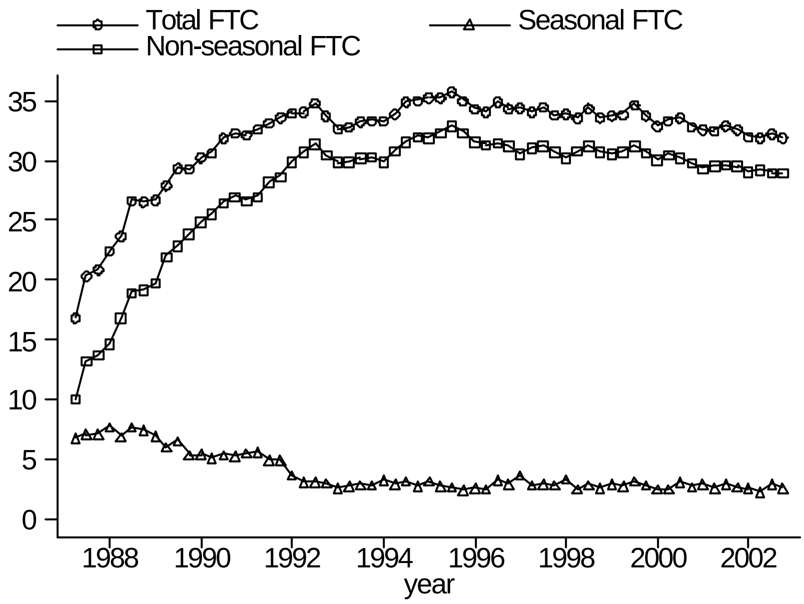

Figure 5.1: The share of …xed-term contracts (%) in total employment, 1987-2002. Source: EPA.

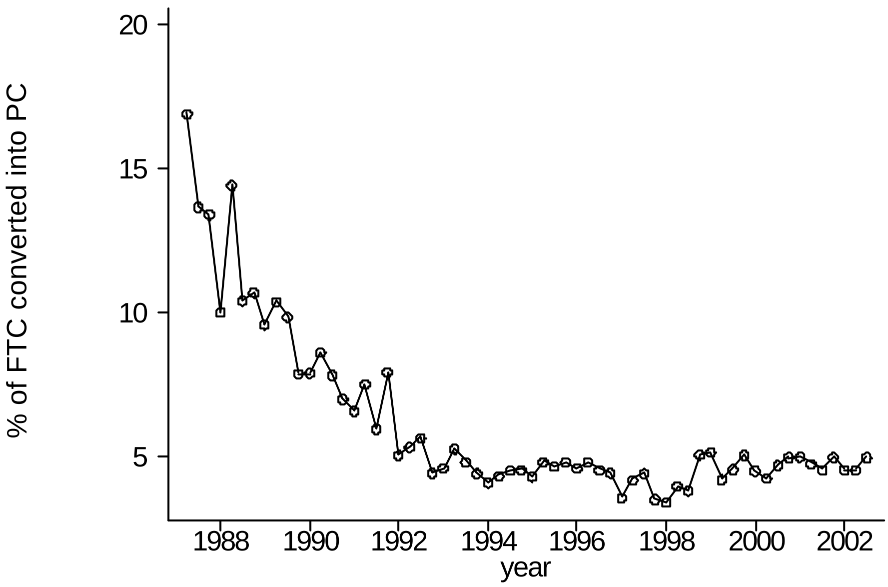

Figure 5.2: The proportion of …xed-term contracts converted into permanent ones, 1987-2002. Source: EPA.

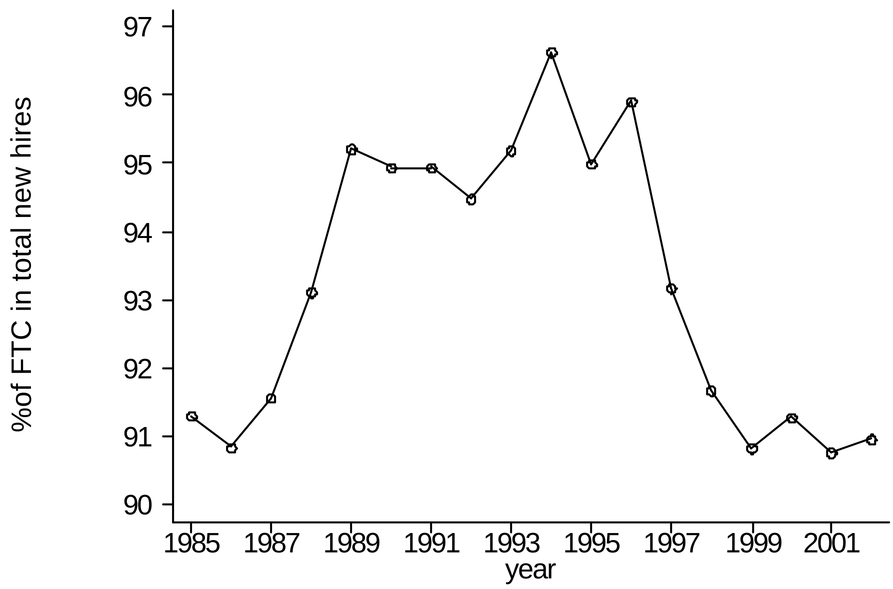

Figure 5.3: Evolution of the share of …xed-term contracts in new hires, 1985-2002. Source: MLR (Spanish Ministry of Labor).

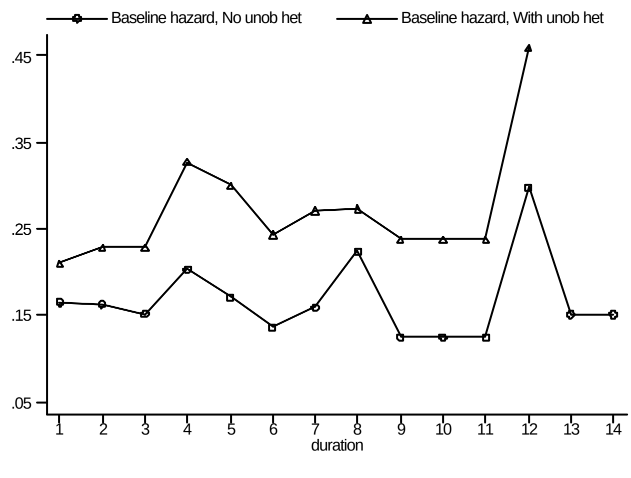

Figure 5.4: Predicted hazard of transition from FTC to PC, full sample until 1998 (see table 5.5). Reference category: male, not married, age 16-24 yrs., secondary education, employed in services, started TC in 1987.

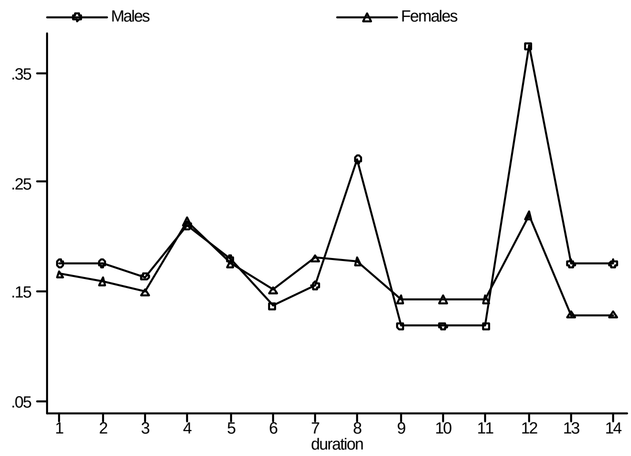

Figure 5.5: Predicted hazard of transition from FTC to PC, male and female samples until 1998 (see table 5.6). Reference category: male/female, not married, age 16-24 yrs., secondary education, employed in services, started TC in 1987.

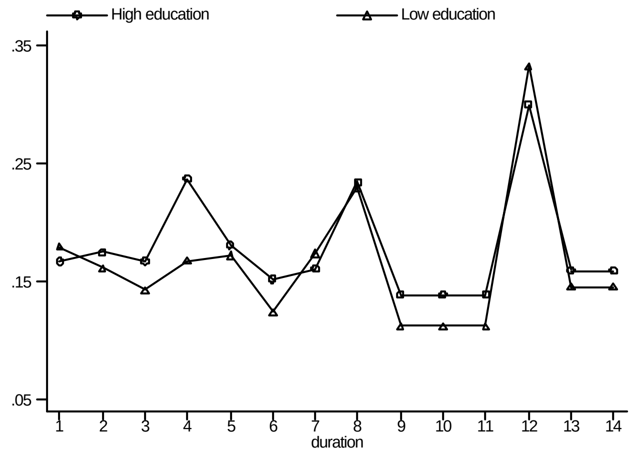

Figure 5.6: Predicted hazard of transition from FTC to PC, high and low education samples until 1998 (see table 5.7). Reference category: high/low education, not married, age 16-24 yrs., employed in services, started TC in 1987.

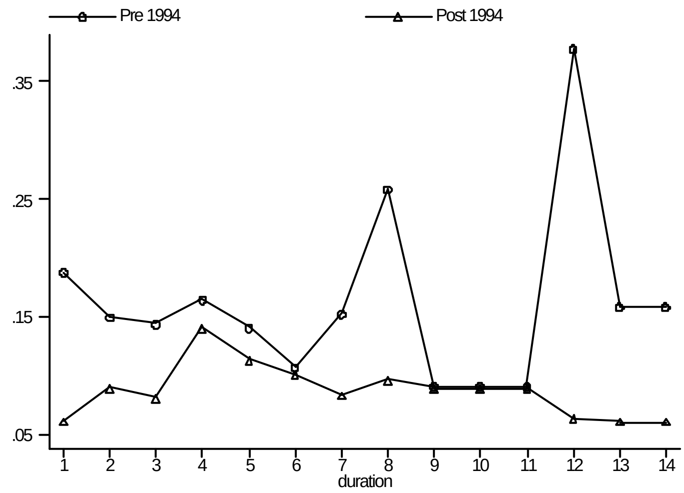

Figure 5.7: Predicted hazard of transition from FTC to PC, contracts started before and after 1994 (see table 5.8). Reference category: male, not married, age 16-24 yrs., secondary education, employed in services.

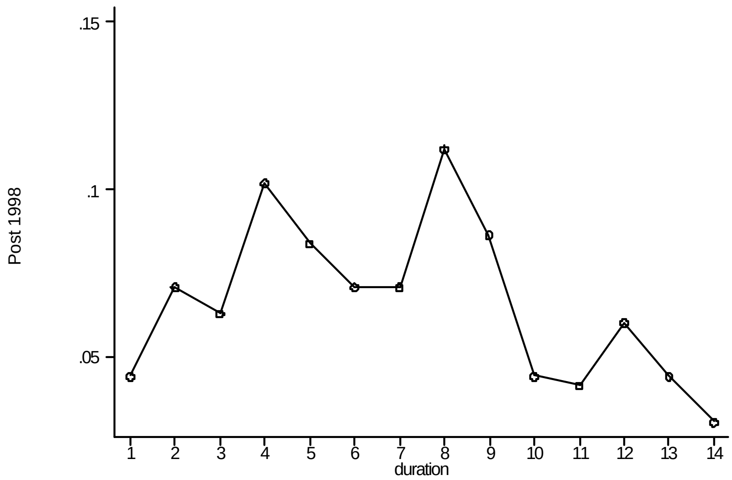

Figure 5.8: Predicted hazard of transition from FTC to PC, contracts started after 1998 (see table 5.9). Reference category: male, not married, age 16-24 yrs., secondary education, employed in services, started TC after 1999.

References

[1] Alba, A. (1998), "How Temporary is Temporary Employment in Spain?", Journal of Labor Research 19, pp. 695-710.

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

[3] 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.

[4] Bentolila, S. and J. Dolado (1994), "Labour Flexibility and Wages: Lessons from Spain", Economic Policy, April 1994, pp. 54-99.

[5] Booth, A., J.J. Dolado and J. Frank (2002), "Symposium on temporary work. Introduction", Economic Journal, Features: Symposium on Temporary Work, F181-F188.

[6] Booth, A. L., Francesconi, M. and Frank, J. (2002), "Temporary Jobs: Stepping Stones or Dead Ends?", Economic Journal, Features: Symposium on Temporary Work, F189-F213.

[7] De La Rica, S. and F. Felgueroso (1999), "Wage Di¤erentials Between Permanent and Temporary Workers: Further Evidence", mimeo, University of Bilbao.

[8] Dolado, J.J., C. García-Serrano and J.F. Jimeno (2002), "Drawing Lessons from the Boom of Temporary Jobs in Spain", Economic Journal, Features: Symposium on Temporary Work, F270-F295.

[9] Dolado, J.J., M. Jansen and J.F. Jimeno (2002), "A Matching Model of Crowding-Out and On-the-Job Search (with an Application to Spain)", CEPR Discussion Paper No. 3466.

[10] Dolton, P. and D. O'Neill (1996), "The Restart E¤ect and the Return to Full-Time Stable Employment", Journal of the Royal Statistical Society, Series A, 159(2), pp. 275-88.

[11] Galdón-Sánchez, J.E. and M. Güell (2000), "Let's go to court! Firing costs and dismissal con‡icts", Princeton University-IR Section working paper 444.

[12] García-Pérez, J. I. and F. Muñoz-Bullón (2003), "The nineties in Spain: too much Flexibility in the Youth Labour Market?", mimeo.

[13] Grubb, D. and W. Wells (1993), "Employment Regulation and Patterns of Work in EC Countries", OECD Economic Studies, 0(21), pp. 7-58.

[14] Güell, M. (2000), "Fixed-term Contracts and Unemployment: An E¢ciency Wage Analysis", Industrial Relations Section Working Paper No 433.

[15] Han, A. and J. Hausman (1990), "Flexible Parametric Estimation of Duration and Competing Risk Models", Journal of Applied Econometrics 5, pp. 1-28.

[16] Jenkins, S. (1995), "Easy Estimation Methods for Discrete Time Duration Models", Oxford Bulletin of Economics and Statistics 57(1), pp. 129-137.

[17] Jimeno, J and L. Toharia (1993), "The Effects of Fixed-term Employment on Wages: Theory and Evidence from Spain", Investigaciones Económicas 27, pp. 475-494.

[18] Jovanovic, B. (1979), "Job Matching and the Theory of Turnover", Journal of Political Economy 87(5), pp. 972-990.

[19] Jovanovic, B. (1984), "Matching, Turnover, and Unemployment", Journal of Political Economy 92(1), pp. 108-122.

[20] Kugler, A., J.F. Jimeno and V. Hernanz (2002), "Employment Consequences of Restrictive Permanent Contracts: Evidence from Spanish Labor Market Reforms", mimeo, Universitat Pompeu Fabra.

[21] Lancaster, T. (1979), 'Econometric Methods for the Duration of Unemployment', Econometrica 47, pp. 939-956.

[22] Narendranathan, W. and M. Stewart (1993), "Modelling the Probability of Leaving Unemployment: Competing Risks Models with Flexible Base-line Hazards", Applied Statistics 42, pp. 63-83.

[23] Nickell, S. (1979), "Estimating the Probability of Leaving Unemployment", Econometrica 47, 1249-1266.

[24] OECD (1987), Employment Outlook.

[25] OECD (1993), Employment Outlook.

[26] OECD (1999), Employment Outlook.

[27] OECD (2002), Employment Outlook.

[28] OECD (1994), The OECD Jobs Study.

[29] Toharia, L. (1996), "Empleo y Paro en España: Evolución, Situación y Perspectivas", Economiaz 35, pp. 36-61.

[30] Toharia, L. (1997), "The Labour Market in Spain", mimeo, University of Alcalá.

[31] Varejao, J. M and P. Portugal (2002), "Why do firms use Fixed-term contracts?", mimeo, University of Oporto.

Appendix: 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 modi…ed on four occasions with the 1984, the 1994, 1997 and 2001 reforms. The ET of 1980 established priority to contracts of inde…-nite duration and allowed FTCs only for jobs which 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 …rst job seekers were also allowed (apprenticeship contracts and training contracts). Other situations in which FTCs 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 FTCs as an incentive to promote employment. Except in this last situation, a speci…c cause was generally required in order to sign a FTC ("causal" FTC).

The 1984 reform exploits this last possibility in an extreme way, and introduces ‡exibility by extending the applicability of FTCs. It introduces a new general FTC, as well as making training contracts more ‡exible. After the reform, any worker can be hired on a temporary basis without the requirement of a speci…c cause. This implies that for any job, employers can freely choose between a PC or a FTC. The 1994, 1997 and 2001 reforms have restricted the applicability of FTC 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. FTCs can be characterized according to: i) limits on their duration (upon expiry, it is not possible to retain the worker under a FTC: either the worker is promoted to a PC or dismissed); ii) eligibility conditions for workers; iii) indemnities at their termination; and iv) subsidies to …rms.