Abstract

In the mid-1980s, many European countries introduced fixed-term contracts. Since then their labor markets have become more dynamic. This paper studies the implications of such reforms for the duration distribution of unemployment, with particular emphasis on the changes in the duration dependence. I estimate a parametric duration model using cross-sectional data drawn from the Spanish Labor Force Survey from 1980 to 1994 to analyze the chances of leaving unemployment before and after the introduction of fixed-term contracts. I find that duration dependence has increased since such reform. Semi-parametric estimation of the model also shows that for long spells, the probability of leaving unemployment has decreased since such reform.

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Fixed-Term Contracts and the Duration Distribution of Unemployment

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Fixed-term Contracts and the Duration Distribution of Unemployment

Maia Güell

1. Introduction

In the mid-1980s, many European countries introduced fixed-term contracts in order to fight the high and persistent levels of unemployment that they had been suffering since the mid-1970s. European labor markets have typically been characterized by a wide use of permanent contracts with high regulated firing costs. The idea behind this policy was to increase flexibility in the labor market by allowing employers the option of hiring workers under shorter contracts with negligible firing costs.[1]

Since their introduction, fixed-term contracts have been widely used and an increasing number of new jobs are fixed-term (see OECD, 1993). European labor markets have become *more dynamic* in terms of inflows and outflows from unemployment to employment, but the unemployment rate has remained very much unchanged. Much of the existing research on fixed-term contracts (or temporary contracts, TCs)[2] has focused on their effect on unemployment. There is a wide consensus among economists that the introduction of such contracts does not necessarily increase employment despite the emergence of a dual labor market among employed workers.[3] In this paper, I study the potential effects of the increased labor flows due to TCs on the duration distribution of unemployment and the possibility that the unemployed pool becomes segmented.

Along with the high rates of unemployment, another worrisome feature of European labor markets is the high proportion of unemployed workers who have been unemployed for a long period of time (see Machin and Manning, 1999). In Europe, on average, between 1983 and 1994, 48% of the unemployed had been in unemployment for more than 12 months (the long-term unemployed, LTU), while in the US this proportion was only 9% (see table 1)[4]. Therefore, it is important to investigate whether the introduction of TCs has improved the functioning of the labor market in these terms.

This paper provides some theoretical considerations of the effects of introducing TCs on the duration distribution of unemployment and then it presents an application to Spain, which is a striking case in this context. In particular, I analyze the effects of TCs on the incidence of LTU, on the duration dependence of unemployment and on the outflow rate of the LTU workers.

In the mid-1980s, Spanish unemployment was around 20% of the labor force, the highest among OECD countries. In 1984, Spain introduced TCs in an extreme way compared to other European countries. In particular, while in some countries TCs are restricted to some

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

[2] The terms fixed-term contract and temporary contract (TC) will be used interchangeably throughout this paper.

[3] See, among others, Aguirregabiria and Alonso-Borrego (1999), Alonso-Borrego *et al.* (1999), Bentolila and Dolado (1994), Güell (2000) and Saint-Paul (1996).

[4] Calculations based on countries for which data were available.

type of workers or sectors, there are not such restrictions in Spain. The fact that Spain was also considered among the most regulated labor markets in the OECD motivated this type of reform (see OECD, 1994). Figure 1 shows the evolution of the unemployment rate as well as the increase of the share of TCs for the 1980s and mid-1990s.

A decade after the introduction of fixed-term contracts, the unemployment rate in Spain was back to pre-reform levels. Moreover, the share of fixed-term employees became the highest among Europe, around 33%, while the European average was 11% (see table 1). As a consequence, in 1994, a second reform that restricted the use of TCs was implemented. So, the Spanish experience between 1980 and 1994 appears to be particularly useful to draw some lessons about the effects of these types of policies.

Although the effects of TCs on unemployment have been unsatisfactory, there have been other changes in other dimensions of the labor market that can reasonably be attributed to these flexibility measures. First, inflows and outflows from unemployment have increased substantially over this period (see figures 2 and 3). This increase in turnover during this period has been mainly driven by TCs. After the 1984 reform, on average, as much as 94% of all newly registered contracts have been TCs (see figure 4); also Bover *et al.* (1997) and García-Pérez (1997) find that TCs increase the employment chances of the unemployed in Spain. As for the inflows back to unemployment, as much as 75% of the unemployed who have been unemployed for less than 3 months were separated from their jobs because their fixed-term contract came to an end. Another supporting fact is that, on average, the renewal rate of TCs into permanent ones has been very low, around 15%, which implies a large flow from non-renewed workers into unemployment.[5] Finally, García-Serrano (1998) studies the role of TCs in worker turnover in Spain and concludes that these contracts account for the largest portion of the hiring and separations rates.

The increase of outflows from unemployment has implied a second important change in the Spanish labor market that has to do with the proportion of unemployed workers with long unemployment spells, which is also among the highest in Europe (see table 1). The incidence of LTU typically displays anti-clockwise loops over the business cycle. Figure 5 shows that, for a given unemployment rate, the incidence of LTU in Spain in the early 1990s is lower than in the mid-1980s. Comparing periods which are at the same point in the cycle, say from 1983 to 1985 and from 1992 to 1994, it can be seen that there has been a shift in the unemployment rate-LTU relationship. In fact, this seems to be a common feature in some European countries (see table 1). As mentioned in Machin and Manning (1999), when the outflow rate increases at any duration of unemployment, the incidence of LTU tends to lower. Therefore, the lower incidence of LTU can also be attributed to the increased outflows since the introduction of TCs.

Previous studies that estimate the probability of leaving unemployment in Spain find that there is a very strong duration dependence.[6] In other words, *ceteris paribus*, unemployed workers with shorter unemployment spells have higher probabilities of leaving unemployment than those with longer spells. An important question remains open. Whether the introduction of TCs has changed the duration distribution of unemployment through changes in

[5] See Güell and Petrongolo (2000).

[6] See, for instance, Alba (1999), Bover *et al.* (1997), García-Pérez (1997), Jenkins and García-Serrano (2000) and Machin and Manning (1999).

duration dependence. This paper aims to analyze the changes in duration dependence of the unemployed *before and after* the introduction of TCs in Spain.

Panel data from the Labor Force Survey in Spain is available only after 1987. Therefore, to analyze the changes in duration dependence before and after the introduction of TCs, I will use the cross-sectional data drawn from the same survey for the years 1980 to 1994. I exploit these data following the parametric duration model suggested by Nickell (1979a). In order to further study if employment chances of TCs have benefited only some of the unemployed at the expense of others, I estimate a semi-parametric version of Nickell's model. I discuss under which conditions such a model can be estimated.

The rest of the paper is organized as follows. Section 2 provides some theoretical considerations of the introduction of TCs on the duration distribution of unemployment. Section 3 describes the data used. Section 4 provides a duration model for the transition from unemployment to employment. Section 5 presents the empirical results, and section 6 concludes.

2. Theoretical Considerations

In this section, I consider different hiring rules of firms under TCs and their implications for the duration distribution of unemployment. Firms can either choose randomly among the pool of unemployed workers or, alternatively, firms can rank applicants by their spells of unemployment, hiring first those workers with shortest duration of unemployment (see Blanchard and Diamond, 1994).

The introduction of TCs increases outflows from unemployment since these contracts are less costly than permanent contracts. As Machin and Manning (1999) show, when the outflow rate increases at *any* duration of unemployment, the incidence of LTU tends to lower. This implies that independently of the hiring rule adopted by firms, the incidence of LTU will be reduced with the introduction of TCs (as figure 5 shows). The intuition of this result is that even if TCs did not increase the outflow rate of the LTU, as long as other unemployed workers with shorter spells get employed, then there is less build-up into longer spells.

However, the different hiring rules adopted by firms can have different effects on the duration dependence of unemployment. It is important to note that, to the extent that firms do not hire randomly, there are strong arguments for the fact that duration dependence might have increased with the introduction of TCs. As Blanchard and Diamond (1994) show, if firms rank unemployed workers and hire those with the shortest spells of unemployment, then the exit rate from unemployment is a decreasing function of duration. In the extreme model where all unemployed workers were homogeneous and only duration of unemployment influenced workers' chances to leave unemployment, then the short-term unemployed would exit first after the introduction of TCs. And when their TC would finish, they would go back to unemployment and they would again be the unemployed with the shortest spell and with higher chances of being re-employed than the others.

The econometric model proposed to analyze the changes in duration dependence allows for several workers' characteristics to influence the chances of getting a TC and leaving unemployment. It is worth noting than in the other extreme model where only certain key characteristics would make unemployed workers more likely to be re-employed with a TC (for instance, gender, age, education), then, after the reform, workers with such characteristics would enjoy higher exit rates than workers without such characteristics. And, as long as they would maintain these same characteristics, they would continue to have higher exit rates when they returned to unemployment after the end of their contract. So, workers without such characteristics would tend to experience longer spells of unemployment than the others.[7]

Therefore, to the extent that firms do not hire randomly, TCs will tend to be enjoyed always by the same group of unemployed workers which implies that the duration dependence of unemployment will increase. Despite this effect of TCs, this type policy could still be Pareto efficient if the probability of exiting unemployment of those who do not get a TC remains unaffected. This will be investigated empirically in the next sections by estimating a semi-parametric version of the econometric model. If that is not the case, then those workers who do not benefit from the higher employment chances given by TCs will remain stuck in unemployment, experiencing lower chances of leaving unemployment than before the reform. Thus, TCs can generate a segmented unemployment pool. That is, some unemployment workers will be constantly churning from unemployment to employment under TCs, while the other unemployed workers will not exit unemployment, experiencing longer and longer durations of unemployment.

3. The Data

I use the Spanish Labor Force Survey (*Encuesta de la Población Activa, EPA*), which is carried out quarterly on a sample of some 60,000 households. It is designed to be representative of the total Spanish population and contains very detailed information about the labor force status of individuals.

My sample contains data from all the second quarters of each year from 1980 to 1994. The time span of the sample is an important feature of the data because it will allow me to analyze the characteristics of the unemployed *before and after* the introduction of TCs. All the unemployed people in the sample are asked how long they have been looking for a job. This search time will be used as the individual's uncompleted duration of unemployment. Results will be based on this variable. In steady state, the average uncompleted duration of unemployment is proportional to the average completed duration of unemployment (see Layard *et al.*, 1991).

There have been several methodological reforms in the EPA which have implied changes in the way some questions have been asked as well as the inclusion of more variables over time. In particular, the way the surveyed unemployed workers are asked about their duration in unemployment and the possible answers given as options by the EPA questionnaires have changed three times (see the Appendix A, for details). Also, in the earlier years, the possible answers were designed in the form of a band (for example, 1 to 3 months). The econometric model specified will deal with this grouping of the data.[8]

[7] I am implicitly assuming that there are no changes in the unobserved characteristics of the population that influence the chances of leaving unemployment.

[8] In the previous version of this paper, I aggregated the data after 1987 into the same duration groups as the earlier data. The results obtained there qualitatively are the same as those presented here (see Güell, 1999).

As far as the number of variables available is concerned, until 1987 there was no information on unemployment benefits or on the reason for previous job loss. These variables can be particularly important for the analysis of the probability of leaving unemployment. In order to fully exploit all the relevant information contained in the data, my analysis will be carried out in two parts. First, I use all the years of the sample, from 1980 to 1994. The analysis is undertaken with those variables common to all sample years. This first part of the analysis thus exploits information for a very long time period at the expense of some relevant variables only available in the most recent years. These additional variables will be exploited in the second part of the analysis for those years for which they are available, from 1987 to 1994.

Before the 1984 reform, TCs were only allowed for seasonal jobs. One key feature of the reform was that it allowed the use of TCs for jobs that are not necessarily seasonal. The higher share of TCs in total employment after the reform can be mainly attributed to their widespread use in non-seasonal jobs.[9] The variable for the reason of previous job loss distinguishes between the end of a TC and other reasons. Although this variable was not available before 1987, the workers who were unemployed because their TC ended could only be those who had held a seasonal TC. Therefore, it can be thought that the reform generated an exogenous increase in the number of workers who lost their job due to the expiration of a (non-seasonal) TC.[10] Therefore, the reason-for-job-loss is a potential source of identification of the change in duration dependence after the introduction of TCs.

As will be discussed, one main assumption of the econometric model is that the composition flow into unemployment is fixed over time. For this reason, I have excluded women from my sample. Thus, my sample includes all men who are unemployed and who report how long they have been searching for a job. I will exclude men aged 65 or older because transitions to non-employment are more likely for this group. Since I want to focus on the effects of TCs on the *existing* distribution of unemployment, I will also exclude first-job seekers.[11] This leaves me with a sample of 80,790 unemployed male workers.

Explanatory variables available for the whole sample period include personal characteristics of the individual such as age, education and marital status, as well as some household characteristics such as the number of kids and the number of working adults in the household. Finally, the local unemployment rate is also included to capture business cycle effects.[12] This quarterly regional unemployment rate will be the only time-varying regressor.[13] For the second part of the analysis, two more variables are available: a dummy variable that indicates if the worker receives unemployment benefits (UI); and a variable that indicates the reason for job loss. From this variable, I construct a dummy variable indicating whether the reason for separation from the previous job was the ending of a TC (end TC). This variable is very important for my purpose since it can potentially capture all the unemployed workers that

[9] See Güell and Petrongolo (2000).

[10] This variable distinguishes between the end of a seasonal TC and a general TC (the TCs introduced in 1984). Between 1987 and 1994, the number of people who have finished a seasonal TC have remained constant; while the number of people who have finished a general TC has increased substantially: from 13 percent of the unemployed in 1987 to 26 percent in 1994.

[11] Alba (1998) estimates the determinants of the employment probabilities of first-job seekers.

[12] See Bover *et al.* (1997) for a more detailed study of business cycle effects on unemployment duration.

[13] Unemployment rate at the regional level includes 50 provinces within Spain.

enjoy the greater employment chances provided by this type of contract. Separate estimation of the model will be done for these workers.

Table 2 reports average sample values for the whole sample (column 1) as well as for each subsample for which a different model will be allowed (columns 2 to 7). Column 2 corresponds to the sample for the years *before* the reform. Columns 3 to 7 correspond to different samples for the years *after* the reform.[14] First, the whole period after the reform (column 3); column 4 only considers the years 1992 to 1994, which correspond to the years which are at comparable points of the business cycle as the years before the reform (see figure 5). Column 5 reports the values for the years from 1987 to 1994, for which more variables are available, and finally, columns 6 and 7 correspond to the unemployed who ended a TC and the ones who were separated for other reasons for the years 1987 to 1994. The (uncompleted) duration of unemployment for the different sub-samples is also reported. As mentioned, this should be compared carefully for sample years corresponding to different points of the cycle. For the last two columns, however, it can be seen that people who are unemployed because of the ending of a TC have about half the (uncompleted) duration of the workers who lost their job for other reasons.

As will be discussed in the next section, in order to estimate the method proposed by Nickell (1979a), it is necessary to complement these cross-sectional data with time series of the inflows into unemployment. I use the monthly registered unemployed and monthly registered new contracts from the Spanish Employment Office (INEM) to construct monthly inflows into unemployment. Since only those unemployed who have worked before can claim unemployment insurance, first-job seekers generally do not register at the Employment Office. This reinforces the exclusion of this group in my analysis.

4. Econometric Specification

My sample has only cross-sectional data on uncompleted spells of unemployment. I will estimate the hazard rate of leaving unemployment following the method proposed by Nickell (1979a). The main requirement in order to implement this method is to have historical data on the inflows into unemployment. The intuition behind this duration model is that the cross-sectional data represent the unemployed that have "survived" with different durations at time t, while the inflow data represent the population "at risk" at different points in time. Generally, these data are easily available at the aggregate level. As Nickell shows, assuming that the composition of the flow into unemployment is fixed over time, the model can be estimated. As it will be discussed later, the frequency of these inflow data is an important issue to be considered in order to estimate such a model, especially semi-parametrically.

Suppose that the probability of leaving unemployment from time t to time $t+1$ for an unemployed individual i, conditional on having entered unemployment at time $t-s$ and on being unemployed at t is given by

$$h_i(t, s) = h(x_i(t_i, s), t, s) \tag{4.1}$$

[14] In the previous version of the paper, the period after the reform was separated into two periods according to the importance of TC in the economy. The results obtained there qualitatively are the same as those presented here (see Güell, 1999).

where t_i is the date in which the interview took place (in my case, the second quarter of every different year considered) and x_i are the relevant characteristics of the individual i, which include the individual's regional unemployment rate during all the spell of unemployment. I have specified h to depend on t. More precisely, I allow the hazard function to be different for different time periods. For example, the hazard for the years before the reform can be different from the hazard for the years after the reform. However, within a sub-period, h does not depend on t. That is, for example, the same function is assumed for the different years prior to the reform (as in Nickell, 1979a).

To write the likelihood, it is necessary to derive the probability for an individual of being unemployed at time t. First, let $S_i(t, v)$ be the probability individual i of and being (remained) unemployed at time t conditional on having entered unemployed at time $t - v$. Therefore

$$S_i(t, v) = \prod_{\tau=1}^{v}(1 - h(x_i(t, \tau), t, \tau)), \text{ for } v \geq 1 \tag{4.2}$$

Suppose that the probability of an individual i of having entered unemployment at time τ is given by $u_i(\tau)$. Then the probability of being unemployed at time t, $U_i(t)$, is given by

$$U_i(t) = \sum_{\tau=0}^{\infty} u_i(t - \tau)S_i(t, \tau) \tag{4.3}$$

It is then possible to write the likelihood for an unemployed individual in my sample, that is, the probability of having entered unemployment at time $t - v$ conditional on being unemployed at time t as

$$L_i = \frac{u_i(t_i - v)S_i(t_i, v)}{\sum_{v=0}^{\infty} u_i(t_i - v)S_i(t_i, v)} \tag{4.4}$$

As mentioned above, the duration of unemployment is presented in my sample in the form of bands. That is, given the date of the interview, t_i, the individual could have entered unemployment at any time between $t_i - a_i$ and $t_i - b_i$. Therefore given my data, the likelihood becomes

$$L_i = \frac{\sum_{v=a_i}^{b_i} u_i(t_i - v)S_i(t_i, v)}{\sum_{v=0}^{\infty} u_i(t_i - v)S_i(t_i, v)} \tag{4.5}$$

For instance, for someone with unemployment duration between 3 to 6 months, the numerator of L_i has 3 terms. Obtaining prior estimates of u_i, say $\widehat{u}_i$, I can then write down the likelihood for my unemployed sample, of individuals, $i = 1, ..., I$ as

$$L = \prod_{i=1}^{I}\left(\frac{\sum_{v=a_i}^{b_i} \widehat{u}_i(t_i - v)S_i(t_i, v)}{\sum_{v=0}^{\infty} \widehat{u}_i(t_i - v)S_i(t_i, v)}\right) \tag{4.6}$$

The probability of individual i, with current characteristics x_{ci}, of having entered unemployment at time τ is defined by

$$u_i(\tau) = k(x(x_{ci}, \tau), \tau) \frac{\text{aggregate flow into unemployment in month } \tau}{\text{aggregate employment in month } (\tau - 1)} \tag{4.7}$$

where $k(x(.), \tau)$ is the proportion of the inflow into unemployment at time τ with characteristics x. Assuming that k is independent of time, this probability can be estimated by

$$\widehat{u}_i(\tau) = constant \times (aggregate\ flow\ into\ unemployment\ in\ month\ \tau) \tag{4.8}$$

There are two mechanisms by which $k(x(., \tau), \tau)$ is affected over time. First, to assume that $k(x(., \tau), .)$ is constant means to assume that any changes in relevant characteristics over time are small. This corresponds to the standard assumption of time unvarying regressors. Second, assuming that $k(x(.), \tau)$ is constant also means that there are small changes in the proportions of individuals with particular characteristics in the inflow into unemployment. This point is more difficult to test mainly because, at least to my knowledge, in Spain there is no data on inflows for the different relevant characteristics.[15] Nickell points out that the stability of the aggregate inflow over time can (partly) justify this assumption. This is also the case within the different sub-periods for which the model is estimated.[16] In any case, the lack of detailed breakdowns of the inflow data does not allow to test directly this assumption.

The solution adopted regarding this point has been to exclude women for whom the stationarity assumption seems particularly strong, specially over the period of study. The only disaggregation of the inflow data in Spain is by gender. This has allowed me to concentrate on male inflow data that match my sample from the cross-sectional data.

There is one last thing to be specified in order to compute the likelihood function given by (4.6). This has to do with the infinite sum in the denominator. I will assume that for long enough durations, the conditional probability specified in (4.1) does not depend on duration and that the estimated probability of having entered unemployment is a constant. In particular, I make these assumptions for durations greater than 36 months. The corresponding $\widehat{u}$ is the average inflow rate of the calendar year corresponding to 36 months of duration of unemployment for every individual (u_{36}). Finally, the likelihood to be maximized is as follows

$$L = \prod_{i=1}^{I} \left(\frac{\sum_{v=a_i}^{b_i} \widehat{u}_i(t_i - v) S_i(t_i, v)}{\sum_{v=0}^{36} \widehat{u}_i(t_i - v) S_i(t_i, v) + \frac{u_{36}}{h_i(36)} S_i(t_i, 36)} \right) \tag{4.9}$$

[15] Nickell (1979a and 1979b) points out the same problem for the UK.

[16] For the years before the reform, the average male inflow was 129.3 thousand with standard deviation of 19.9. For the years after the reform, the average male inflow was 238.8 thousand (s.d. of 49.2); from 1985 to 1989, the average male inflow was 195.1 thousand (s.d. of 22.5) and for the years 1990-1994 it was 282.9 (s.d. of 14.8). The latter breakdown for the years after the reform corresponds to the one analyzed in the previous version of the paper, where the results obtained there qualitatively are the same as those presented here.

I will first specify $h(t, s)$ following a proportional hazard model where the underlying base-line is a Weibull distribution. That is,

$$h_i(t, s) = 1 - \exp\left(\int_s^{s+1} \lambda(u)du\right), \text{ where } \lambda(s) = -\exp(x_i'\beta)\lambda_0(s) \tag{4.10}$$

and

$$\lambda_0(s) = \alpha s^{\alpha-1} \tag{4.11}$$

Note that I have assumed that the frequency of the inflow data allows to make a one–to-one correspondence between each duration of unemployment (v) at time t and its flow into such state $(t - v)$. This is because the available inflow data has monthly periodicity.

As mentioned, intuitively, the introduction of TCs would tend to increase the duration dependence in unemployment. A further question remains to be answered empirically. That is, whether the introduction of TCs has also implied an increase of employment chances for some individuals at the expense of some others (or, alternatively, if the chances of the latter group have remained unchanged in respect to the situation before the reform). The parametric estimation does not allow us to answer a such question since the base-line hazard (see (4.11)) monotonically decreases with duration, so changes in the duration dependence parameter (α) will imply shifts of the whole base-line function. Therefore, this question needs to be investigated semi-parametrically. Therefore, I will re-estimate the above model allowing a flexible base-line hazard, that is, I do not specify any functional form for $\lambda_0(s)$.

Given the grouping of the duration data, a step-wise base-line hazard will be estimated where every step corresponds to each duration group (see the Appendix A, first column).[17] It is important to note how the frequency of the inflow data plays a role in the estimation. In order to estimate the model, it is crucial that each duration group (the population that has survived and that we observe at time t in the cross-section with duration s) can be matched univocally with its population at risk, namely the inflow at time $t - s$. If this condition did not hold (for instance, if the inflow at $t - s$ could be attached to more than one survival group) an identification problem would arise since, a given inflow point could correspond to more than one duration group. In this case, it would not be possible to estimate a separated step for such group. Let s_i be the frequency of the inflow data. That is, we observe the inflow data at period t, $t - s_i$, $t - 2s_i$, etc. In the cross-section each duration group has duration s (which depends on how the aggregation of durations is done). It is then crucial that $s_i \leq s$ so that a different step can be estimated for each duration group.

When the inflow is less frequent than the duration groups, then the step-wise assumptions (or even the parametric) will not suffice to estimate such model. Further assumptions could be made to recover, for instance, monthly inflows from quarterly inflows. However, this would seem to be less appropriate in the semi-parametric case.

In this paper, the inflow data is monthly. The duration groups of the cross-sectional data vary over time. Before 1987, the grouping of the data is quarterly (except for the first one),

[17] For comparison reasons, the same steps will be estimated for the years after 1987. Since I want to concentrate on the last step, I have imposed the same step for the first two duration groups.

then 6 months and then yearly. After 1987, the grouping is monthly (if duration is less than 2 years) and then yearly. Therefore, there is no identification problem.

5. Empirical Results

I now estimate the hazard of leaving unemployment as modeled in the previous section. First, I concentrate on changes in the duration dependence and I estimate the Weibull base-line hazard specified in (4.11). Table 3 reports the estimates for the whole sample, 1980—1994. Every variable is interacted with a post-reform dummy ($d8594$, which takes value 1 for years from 1985 to 1994). The duration dependence parameter is statistically different before and after the reform. Table 8 reports the duration dependence estimates for these subsamples. And figure 6 plots the hazard of leaving unemployment for the reference category estimated by this regression for the years before and after the reform. As can be seen, in the years after the reform the duration dependence of unemployment is much higher than before. For durations of less than 5 months, the probability of leaving unemployment are much higher than before. But the reverse is true for durations of 6 months or more.

The effects of the individual characteristics on the probability of exiting unemployment are fairly standard and consistent with previous studies (see Alba, 1999, and Bover *et al.*, 1997). The re-employment probability decreases with age. Being married substantially increases the probability of finding a job. This has to do with lower reservation wages of these individuals given their household responsibilities and for the same reason their attachment to the labor market is strong. Similarly, the effect of the number of kids is positive, but small. Also, the effect of the number of working adults in the household is negative, but again, not very large. The estimated coefficients on education are negative for the pre-reform years, but positive afterwards. Bover *et al. (*1997) find that secondary education has no significant effect while a university degree has a positive effect on the re-employment probabilities. My result may be partially explained by the fact that very few people with a university degree are among the unemployed before the reform. Alba (1999) finds that the variable education increases the likelihood of re-employment only for workers with vocational education.

In the period after the reform, there are some years of expansion (from 1985 to 1991) and some years of recession (from 1992 to 1994). As mentioned before, the LTU typically displays anti-clockwise loops over the cycle, and this can imply that the duration dependence is higher in expansion years because the proportion of LTU is higher. Indeed, when estimating the probability of leaving unemployment for the post reform period with each variable interacted with a recession dummy ($d9294$, which takes value 1 for the recession years), I find that this is the case (see table 4). However, comparing the estimated parameter of the duration dependence for the recession years, it is still lower than in the pre-reform period (see table 8, column 4).[18]

A further check of the increase of duration dependence after the introduction of TCs, despite the fact there are some expansion years in the post reform period, is to compare the years 1983 and 1992, which are the most comparable in terms of unemployment rates. Table

[18] Testing that the duration dependence parameter in these recession years is the same as in the pre-reform period (that is, $\alpha = 0.849$) gives the test statistic $z = 13.51$. Therefore, the null hypothesis is rejected at standard levels of significance.

5 reports these estimates, where $d92$ takes value 1 for the year 1992. As can be seen, the main result still holds.

The number of variables available in the Spanish Labor Force Survey has increased over time. Therefore, I estimate a second set of regressions in which more variables are included for the period 1987—1994. The inclusion of more variables can affect the estimated duration dependence parameter. Therefore, it is important to check whether the above result is affected by the exclusion of these variables. Table 6 displays the results of the estimations without the UI dummy and the end-of-temporary-contract dummy; the estimations including only one of the two variables; and the estimation including both variables. As expected, the parameter of duration dependence increases with the different specifications. However, it is always lower than in the pre-reform period.[19]

The effects of the variables also included in the previous regressions remain very similar. As can be seen, the effect of UI is positive. At first glance, this result may be surprising if one has in mind the standard *disincentive effect* from job search theory (see Mortensen, 1970 and 1977).[20] There are several possible reasons for this result. First, the UI variable is only an indicator of whether the unemployed person is receiving benefits when being interviewed. There is wide consensus that the effects of unemployment benefit levels are far from robust, being in general not very significant and of small size, and that other dimensions of unemployment compensation may be more important, such as duration of benefits.[21] Moreover, given the nature of my data, it is possible that some unemployed people are interviewed once their UI has already expired and they report they are not recipients of UI. This biases downwards the negative effect of UI.[22]

Secondly, Alba (1999) and Bover *et al.* (1997) find that the effect of receiving unemployment is significant and quite sizeable, but that this effect is reduced over the spell of unemployment. This can be seen by simply calculating the correlation between UI receipt and duration of unemployment for different durations (see the Appendix B). Indeed, this correlation is higher at shorter durations, suggesting that the disincentive effect is present for short durations. But, for long durations not only it is not negligible, but it is negative. Wadsworth (1990) and Schmitt and Wadsworth (1993) exploit the idea that UI facilitate search by providing income with which to finance job search efforts (*the job offers effect*). These studies compare the search behavior of benefit claimants and non-claimants. They find that non-claimants search harder during the initial stages of unemployment when benefits may provide a temporary leisure subsidy to benefit claimants. As unemployment duration lengthens, search activities fall for both groups, but benefit recipients are able to maintain a higher level of search effort and therefore have a relatively higher probability of receiving a job offer. This *job offers* effect seems to be very strong in my data. A possible reason

[19] Although these regressions are not strictly comparable, since the inclusion of additional variables modifies the estimated α upwards, this comparison is more restrictive than it should be.

[20] However, Toharia (1997) reviews different studies on the disincentive effects of UI in Spain and concludes that, on the whole, the studies available are not conclusive.

[21] See Atkinson and Micklewright (1991) for a review.

[22] On top of that, as pointed out by Alba (1999), this dummy variable takes value zero both for people who are not registered in the public employment office and those who are registered but do not receive benefits. The problem of this data is that some unemployed people may have applied for unemployment benefits but may not have received a response at the moment of the interview.

why the cited works on Spain may differ on the effect of UI is the treatment given to the unemployment duration variable.[23] Finally, there was an important reform of the unemployment benefit system in 1992 which reduced both the level and duration of UI.[24] Alba (1999) finds that the negative incentive effect of benefits on re-employment probabilities is reduced starting in 1993 and becomes more important from 1994 onwards.

The other new variable included in this second part of the analysis is a dummy that equals 1 if the reason for separation from the previous job was the ending of a TC (*endtc*). As can be seen in table 6, the estimated coefficient on this variable is positive and significant. This result accords with the idea that TCs have isolated some of the unemployed, making them more employable than the other unemployed. I investigate this issue further by estimating a model in which every variable is interacted with this dummy (see table 7). As can be seen, those jobless workers who are unemployed because their TC came to an end have less duration dependence than the other unemployed (this includes voluntary quits, redundancy, retirement, illness, etc.). Canziani and Petrongolo (2000) estimate a semi-parametric duration model using the panel version of the Spanish Labor Force Survey data for the years 1987—1996 and also find that those jobless workers whose TC ended have higher re-employment probabilities. They also find that job-quitters have higher chances to leave unemployment than those unemployed who were dismissed in their previous job.[25] Jenkins and García-Serrano (2000) using data from the national unemployment benefit administration database find that those who entered UI from a TC have much higher re-employment probabilities than those whose contract was a permanent one.

Table 8, columns 5 and 6, report the duration dependence estimates for those unemployed for which the reason of separation in their last job was the ending of a TC and for those for which there was another reason. Figure 7 plots the hazard of leaving unemployment for these two groups of unemployed workers from these regressions. Unemployed workers who came from a TC have greater chances of leaving unemployment at any duration than the others. Secondly, the hazard for those who became jobless because of the ending of a TC is flatter than for the other groups of individuals. That is, although there is negative duration dependence, it is much smaller than for those individuals that lost their jobs for other reasons. These results are also suggestive of the idea that TCs have increased the employment chances for a group of the unemployed that churns from employment to unemployment frequently. The remaining unemployed have lower chances of re-employment and these chances get worse at longer durations.

It is interesting to note that education has an insignificant effect those who ended a TC while it has a positive significant effect for those who became jobless for other reasons than the ending of a TC. One possible explanation is that since people who became unemployed because of the ending of a TC have greater chances of leaving unemployment they are more attached to the labor market and therefore having a university degree or not does not

[23] Alba (1999) excludes from his sample unemployed people of more than 36 months. Bover *et al.* (1997) treat durations of more than 14 months as censored at 14 months.

[24] The motivation for this reform was the increased inflows and outflows from unemployment though temporary contracts, which generated an important deficit in the Spanish unemployment benefit system. Also, the minimum duration of a temporary contract was raised from six months to one year and consequently increased the minimum job tenure required for benefits eligibility.

[25] Kugler and Saint-Paul (2000) find similar results using US data.

substantially affect the probability of finding a job. Instead, people that became jobless for other reasons are less attached to the labor market and therefore having a university degree can improve their chances of becoming employed.

Back to the effects of UI, the coefficient for both groups appears to be positive although larger for those who became unemployed for reasons other than ending a TC. This is consistent with the mentioned effect of unemployment insurance along the duration of unemployment because the unemployed workers who became unemployed for other reasons experience longer durations of unemployment.

All the results found above are suggestive of the fact that the distribution of the duration of unemployment has become more unequal in the mid-1990s compared to the late 1980s. A possible explanation for this fact is the introduction of TCs. As mentioned above, these contracts have implied an increase in the average outflow rate. To the extent that these contracts have implied that higher employment chances have not been shared equally among all the unemployed, then the outflow rate of unemployment of those who have not benefited from TCs will not have increased. The above results suggest that the re-employment chances of the latter have been reduced. To further investigate this, I will re-estimate the above model semi-parametrically in order to compare each step of the base-line hazard before and after the reform without imposing any functional form.

Tables 9a and 9b report the estimates for the whole sample where every variable is interacted with a post-reform dummy ($d8594$, which takes value 1 for the years 1980—1994). This regression is like the one reported in Table 3, except that four different steps of the base-line hazard are allowed. Figure 8 plots the hazard of leaving unemployment for the reference category estimated by this regression for the years before and after the reform. As can be seen, in the years after the reform the two last steps of the base-line are lower than in the years before the reform. This result goes in line with the results that were suggested from the previous parametric estimations. That is, the fact that conditional on being long-term unemployed, the chances of leaving unemployment after the reform are lower than before.

6. Conclusion

In this paper I have analyzed the effects of the introduction of TCs on the duration distribution of unemployment in Spain, with particular emphasis on the changes in duration dependence. The motivation was, on the one hand, to study whether this type of policy had an impact through different dimensions of the labor market, given the rather unsuccessful effect of this type of policy in reducing unemployment. And, on the other hand, since the introduction of TCs has made the labor market *more dynamic*, the motivation was also to study the impact of the increase in inflows and outflows from unemployment to employment as captured in the duration of the unemployed.

The paper has exploited cross-sectional data available for a very long period of time (from 1980 to 1994) that allows an analysis of the chances of leaving unemployment *before and after* the introduction of TCs in Spain. In particular, the idea that the paper has explored is that, even if the incidence of LTU may be lower due to the increased (average) outflow rate, if the greater chances given by TCs are not equally distributed among all the unemployed workers, then the duration of those who remain stuck in unemployment will be higher and higher. I have found evidence of this effect.

I also find that the chances of finding a job at any duration are significantly higher for those unemployed workers who became unemployed due to the ending of a TC in the previous job than for those unemployed workers who became unemployed for other reasons. And there is a stronger duration dependence for this latter group. These results again suggest that TCs have generated an important increase in the (average) outflow from unemployment, but that only some of the unemployed have enjoyed these greater chances at the expense of the others. It seems plausible that these changes are driven by the introduction of TCs, since this was the major institutional change in the time period studied.

It is often argued that a high proportion of LTU is a possible cause of high unemployment itself. Although this causality has to be analyzed with caution (see Machin and Manning, 1999), in the case of Spain it is possible that the limited success of flexibility measures in reducing unemployment could be linked to the fact that TCs have not helped to reduce the duration dependence in unemployment.

Table 1: Unemployment rate, incidence of LTU and share of TCs for several countries

		1983	1984	1985	1986	1987	1988	1989	1990	1991	1992	1993	1994
Denmark	u rate	9.0	8.5	7.1	5.4	5.4	6.1	7.3	7.7	8.4	9.2	10.1	8.2
	LTU	33.0	-	39.3	32.9	30.6	28.7	25.9	33.7	31.2	27.0	-	32.1
	TCs	-	12.5	12.3	-	11.2	11.5	10.0	10.8	11.9	11.0	10.7	12.0
France	u rate	8.1	9.7	10.1	10.2	10.4	9.8	9.3	8.9	9.5	10.4	11.7	12.3
	LTU	42.2	-	46.8	47.8	45.5	44.8	43.9	38.3	37.3	36.1	34.2	38.3
	TCs	3.3	-	4.7	-	7.1	7.8	8.5	10.5	10.2	10.5	10.9	11.0
Germany*	u rate	6.9	7.1	7.2	6.5	6.3	6.2	5.6	4.8	5.6	6.6	7.9	8.4
	LTU	39.3	-	47.9	48.9	48.2	46.7	49.0	46.3	45.5	33.5	-	44.3
	TCs	10.0	-	10.0	-	11.6	11.4	11.0	10.5	10.1	10.5	10.3	10.3
Ireland	urate	14.0	15.5	16.8	16.8	16.6	16.1	14.7	13.4	14.8	15.4	15.6	14.3
	LTU	36.9	-	64.7	65.2	66.4	66.0	67.3	67.2	60.3	-	-	64.3
	TCs	6.2	-	7.3	-	8.6	9.1	8.6	8.5	8.3	8.7	9.4	9.5
Italy	urate	7.7	8.1	8.5	9.2	9.9	10.0	10.0	9.1	8.8	9.0	10.3	11.4
	LTU	57.7	-	65.8	66.1	66.4	69.0	70.4	71.1	67.1	58.2	-	61.5
	TCs	6.6	-	4.8	-	5.4	5.8	6.3	5.2	5.4	7.5	6.0	7.3
Portugal	u rate	7.8	8.5	8.7	8.4	6.9	5.5	4.9	4.6	4.0	4.2	5.7	7.0
	LTU	-	-	56.0	56.0	56.6	51.2	48.3	48.1	38.3	30.9	-	43.4
	TCs	-	-	-	14.4	16.9	18.5	18.7	18.3	16.4	11.0	9.8	9.4
Spain	u rate	17.5	20.3	21.7	21.0	20.1	19.1	16.9	16.2	16.4	18.5	22.8	24.1
	LTU	52.4	-	56.7	57.6	62.0	61.5	58.5	54.0	51.1	47.4	50.1	56.1
	TCs	-	-	-	-	15.6	22.4	26.6	29.8	32.2	33.5	32.2	33.7
UK	u rate	11.1	11.1	11.5	11.5	10.6	8.7	7.3	7.1	8.8	10.1	10.5	9.6
	LTU	47.0	-	47.0	45.9	45.9	44.7	40.8	36.0	28.1	35.4	-	45.4
	TCs	5.5	-	7.0	-	6.3	6.0	5.4	5.2	5.3	5.5	5.9	6.5
US	u rate	9.6	7.5	7.2	7.0	6.2	5.5	5.3	5.5	6.7	7.4	6.8	6.1
	LTU	13.3	-	9.5	8.7	8.1	7.4	5.7	5.6	6.3	11.2	11.7	12.2

Notes: (1) urate is the unemployment rate; LTU is the share of unemployed with spells$\geq$12 months and TCs is the share of workers under a TC among employed; (2)*Since 1991, data on Germany and EU include the new German Länder; (3) Source: OECD (1993, 1996 and 1999).

Table 2: Sample characteristics

	(1) 1980-94	(2) 1980-84	(3) 1985-94	(4) 1992-94	(5) 1987-94	(6) End TC	(7) Other reasons
age	35.738	35.010	36.041	36.667	35.926	34.632	36.317
	(13.440)	(13.267)	(13.502)	(14.141)	(13.603)	(13.490)	(13.378)
married	0.527	0.551	0.517	0.492	0.497	0.452	0.566
	(0.499)	(0.497)	(0.499)	(0.499)	(0.500)	(0.498)	(0.495)
second. or	0.326	0.197	0.379	0.476	0.417	0.435	0.268
univ. ed.	(0.469)	(0.398)	(0.485)	(0.499)	(0.493)	(0.496)	(0.443)
n. of kids	1.003	1.219	0.914	0.808	0.865	0.881	1.068
	(1.253)	(1.416)	(1.168)	(1.047)	(1.124)	(1.133)	(1.307)
n. of working	0.718	0.699	0.728	0.751	0.762	0.766	0.694
adults	(0.886)	(0.872)	(0.891)	(0.884)	(0.906)	(0.911)	(0.871)
UI				0.463	0.430	0.441	0.216
				(0.499)	(0.495)	(0.496)	(0.411)
end TC				0.713	0.629	1.000	0.000
				(0.452)	(0.483)	(0.000)	(0.000)
log local	2.692	2.423	2.803	2.691	2.781	2.800	2.747
unemployment	(0.484)	(0.512)	(0.425)	(0.443)	(0.441)	(0.447)	(0.428)
duration of	14.492	12.724	15.227	13.657	15.026	11.897	20.347
unemployment[1]	(12.887)	(11.676)	(13.227)	(12.333)	(13.286)	(11.603)	(14.241)
Total No. of spells	80,790	23,720	57,070	18,991	44,053	27,740	16,313

[1] Uncompleted duration of unemployment. Reported duration for years after 1987 has been grouped as in previous years.

Note: (1) Standard deviations in brackets; (2) Source: EPA.

Table 3: Maximum likelihood estimates of the probability of leaving unemployment, Weibull specification: full sample, 1980-1994

	Coefficient
α	0.841
	(0.023)
$\alpha \times d8594$	-0.346
	(0.026)
constant	-0.865
	(0.104)
constant $\times$ $d8594$	0.952
	(0.124)
age	-0.011
	(0.001)
age $\times$ $d8594$	0.002
	(0.001)
married	0.197
	(0.026)
married $\times$ $d8594$	-0.215
	(0.029)
second. or univ. ed.	-0.094
	(0.027)
second. or univ. ed. $\times d8594$	0.151
	(0.029)
n. of kids	0.021
	(0.007)
n. of kids $\times$ $d8594$	0.026
	(0.009)
n. of working adults	-0.042
	(0.012)
n. of working adults $\times$ $d8594$	0.018
	(0.014)
log local unemployment	-0.304
	(0.021)
log local unemployment $\times$ $d8594$	0.081
	(0.025)
mean log-likelihood	-2.402
No. of obs.	80,790

Note: (1) Standard errors in brackets; (2)The variable $d8594$ is equal to 1 for the years 1985 to 1994 and zero otherwise; (3) Source: EPA.

Table 4: Maximum likelihood estimates of the probability of leaving unemployment, Weibull specification: post reform sample, 1985-1994

	Coefficient
α	0.402
	(0.010)
$\alpha \times d9294$	0.257
	(0.017)
constant	0.428
	(0.061)
constant $\times$ $d9294$	-1.128
	(0.101)
age	-0.017
	(0.001)
age $\times$ $d9294$	0.013
	(0.001)
married	0.156
	(0.014)
married $\times$ $d9294$	-0.254
	(0.021)
second. or univ. ed.	-0.070
	(0.012)
second. or univ. ed. $\times$ $d9294$	0.143
	(0.019)
n. of kids	0.011
	(0.005)
n. of kids $\times$ $d9294$	0.081
	(0.008)
n. of working adults	-0.014
	(0.006)
n. of working adults $\times$ $d9294$	0.001
	(0.010)
log local unemployment	-0.115
	(0.013)
log local unemployment $\times$ $d9294$	-0.084
	(0.020)
mean log-likelihood	-2.668
No. of obs.	57,070

Note: (1) Standard errors in brackets; (2) The variable $d9294$ is equal to 1 for the years 1992 to 1994 and zero otherwise; (3) Source: EPA.

Table 5: Maximum likelihood estimates of the probability of leaving unemployment, Weibull specification: 1983 and 1992

	Coef.
α	0.738
	(0.034)
$\alpha \times d92$	-0.269
	(0.039)
constant	-0.384
	(0.168)
constant $\times$ $d92$	0.390
	(0.211)
age	-0.014
	(0.001)
age $\times$ $d92$	0.009
	(0.002)
married	0.183
	(0.037)
married $\times$ $d92$	-0.344
	(0.046)
second. or univ. ed.	0.023
	(0.036)
second. or univ. ed. $\times$ $d92$	0.105
	(0.045)
n. of kids	0.017
	(0.011)
n. of kids $\times$ $d92$	0.051
	(0.015)
n. of working adults	-0.047
	(0.017)
n. of working adults $\times$ $d92$	0.028
	(0.022)
log local unemployment	-0.282
	(0.037)
log local unemployment $\times$ $d92$	0.175
	(0.048)
mean log-likelihood	-2.353
No. of obs.	9,974

Note: (1) Standard errors in brackets; (2) The variable $d92$ is equal to 1 for the year 1992 and zero otherwise; (3) Source: EPA.

Table 6: Maximum likelihood estimates of the probability of leaving unemployment, Weibull specification: post reform sample, 1987-1994

	(1) Coef.	(2) Coef.	(3) Coef.	(4) Coef.
α	0.495	0.536	0.615	0.651
	(0.009)	(0.009)	(0.009)	(0.010)
constant	0.133	-0.133	-0.975	-1.181
	(0.050)	(0.051)	(0.055)	(0.056)
age	-0.008	-0.009	-0.005	-0.006
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
married	-0.030	-0.124	0.050	-0.032
	(0.010)	(0.011)	(0.001)	(0.011)
second. or univ. ed.	0.054	0.067	0.034	0.046
	(0.010)	(0.010)	(0.010)	(0.011)
n. of kids	0.057	0.064	0.048	0.053
	(0.004)	(0.004)	(0.004)	(0.004)
n. of working adults	-0.015	-0.008	-0.008	-0.002
	(0.005)	(0.005)	(0.005)	(0.006)
log local unemployment	-0.232	-0.234	-0.027	-0.273
	(0.010)	(0.010)	(0.011)	(0.011)
unemployment insurance		0.308		0.271
		(0.010)		(0.010)
end of temporary contract			0.819	0.813
			(0.012)	(0.012)
mean log-likelihood	-2.973	-2.963	-2.911	-2.904
No. of obs.	44,053	44,053	44,053	44,053

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

Table 7: Maximum likelihood estimates of the probability of leaving unemployment, Weibull specification: reason separation in the last job, 1987-1994

	Coefficient
α	0.503
	(0.020)
$\alpha\times$ endtc	0.180
	(0.023)
constant	0.416
	(0.104)
constant $\times$ endtc	-1.281
	(0.121)
age	-0.017
	(0.001)
age $\times$ endtc	0.015
	(0.001)
married	-0.039
	(0.021)
married $\times$ endtc	0.019
	(0.025)
second. or univ. ed.	0.219
	(0.019)
second. or univ. ed. $\times$ endtc	-0.248
	(0.023)
n. of kids	0.062
	(0.008)
n. of kids $\times$ endtc	-0.017
	(0.009)
n. of working adults	-0.002
	(0.010)
n. of working adults $\times$ endtc	0.006
	(0.012)
unemployment insurance	0.492
	(0.019)
unemployment insurance $\times$ endtc	-0.325
	(0.023)
log local unemployment	-0.516
	(0.020)
log local unemployment $\times$ endtc	0.355
	(0.024)
mean log-likelihood	-2.891
No. of obs.	44,053

Note: (1) Standard errors in brackets; (2) The variable *endtc* is equal to 1 if the reason of last job loss was the ending of a TC and zero if other reasons; (3) Source: EPA.

Table 8: Summary of duration dependence estimates

		(1) 1980-1984	(2) 1985-1994	(3) 1985-1991	(4) 1992-1994
(I)	α	0.841	0.495	0.402	0.659
		(0.023)	(0.008)	(0.010)	(0.014)
	Source	Table 3	Table 3	Table 4	Table 4
		(5) End TC	(6) Other reasons		
(II)$^{(*)}$	α	0.683	0.503		
		(0.011)	(0.020)		
	Source	Table 7	Table 7		

$^{(*)}$Includes same regressors as in (I) as well as UI dummy.
Note: (1) Standard errors in brackets; (2) Source: EPA.

Table 9a: Baseline hazard estimates

	spell months	Coefficient
step1	*1 to 6*	-0.569
		(0.046)
step1 × d8594		-0.047
		(0.054)
step2	*6 to 12*	-2.321
		(0.153)
step2 × d8594		0.422
		(0.160)
step 3	*12 to 24*	-1.222
		(0.049)
step3 × d8594		-0.542
		(0.059)
step 4	*24 to 36*	-2.602
		(0.070)
step4 × d8594		-0.376
		(0.766)

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

Table 9b: Maximum likelihood semi-parametric estimates of the probability of leaving unemployment: full sample, 1980-1994

	Coefficient
age	-0.080
	(0.005)
age $\times$ $d8594$	0.014
	(0.006)
married	0.147
	(0.014)
married $\times$ $d8594$	-0.155
	(0.016)
second. or univ. ed.	-0.050
	(0.014)
second. or univ. ed. $\times d8594$	0.081
	(0.016)
n. of kids	0.012
	(0.004)
n. of kids $\times$ $d8594$	0.023
	(0.005)
n. of working adults	-0.031
	(0.006)
n. of working adults $\times$ $d8594$	0.016
	(0.008)
log local unemployment	-0.204
	(0.012)
log local unemployment $\times$ $d8594$	0.036
	(0.014)
mean log-likelihood	-2.280
No. of obs.	80,790

Note: (1) Standard errors in brackets; (2)The variable $d8594$ is equal to 1 for the years 1985 to 1994 and zero otherwise; (3) Source: EPA.

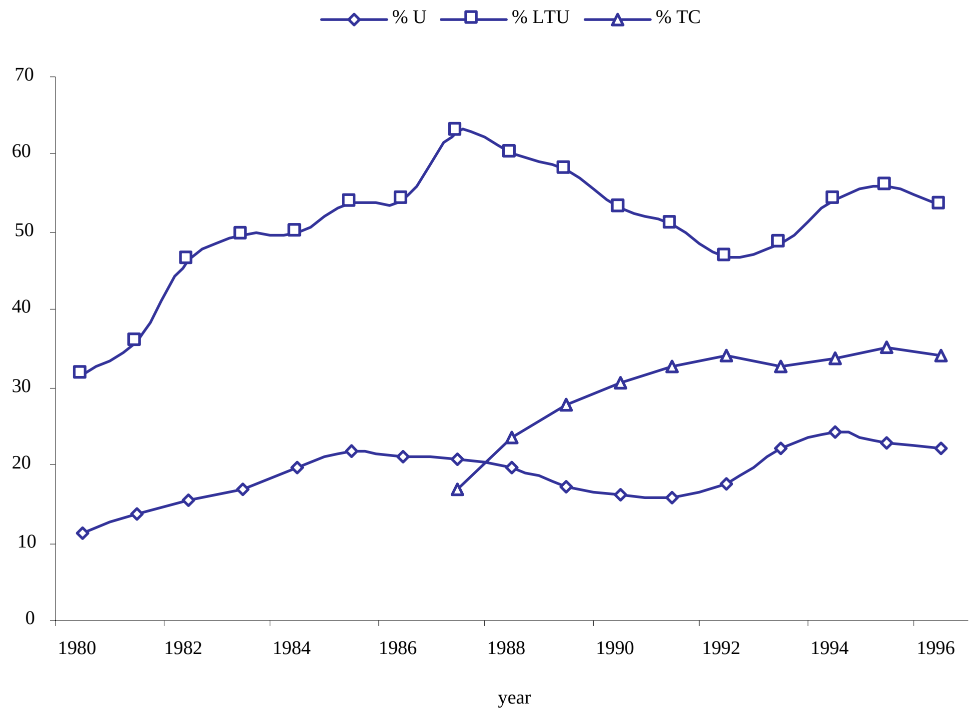

Figure 1: Unemployment rate, proportion of LTU and share of temporary contracts, 1980-1996. Source: EPA.

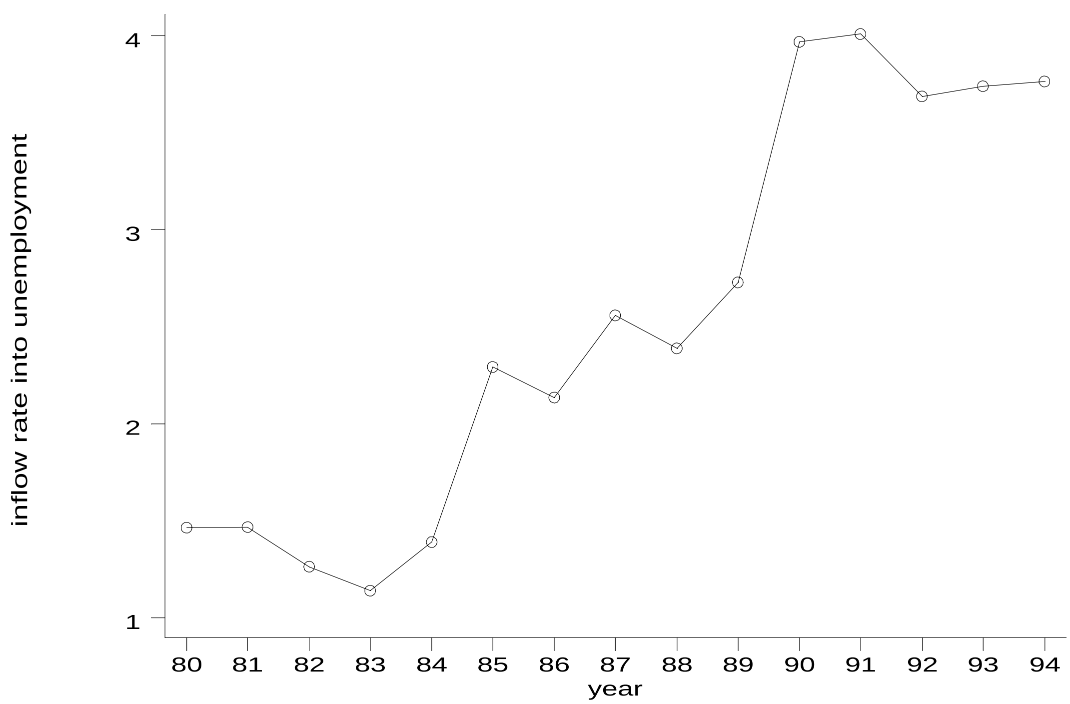

Figure 2: Evolution of inflow rates from employment into unemployment, 1980-1994. Source: MLR.

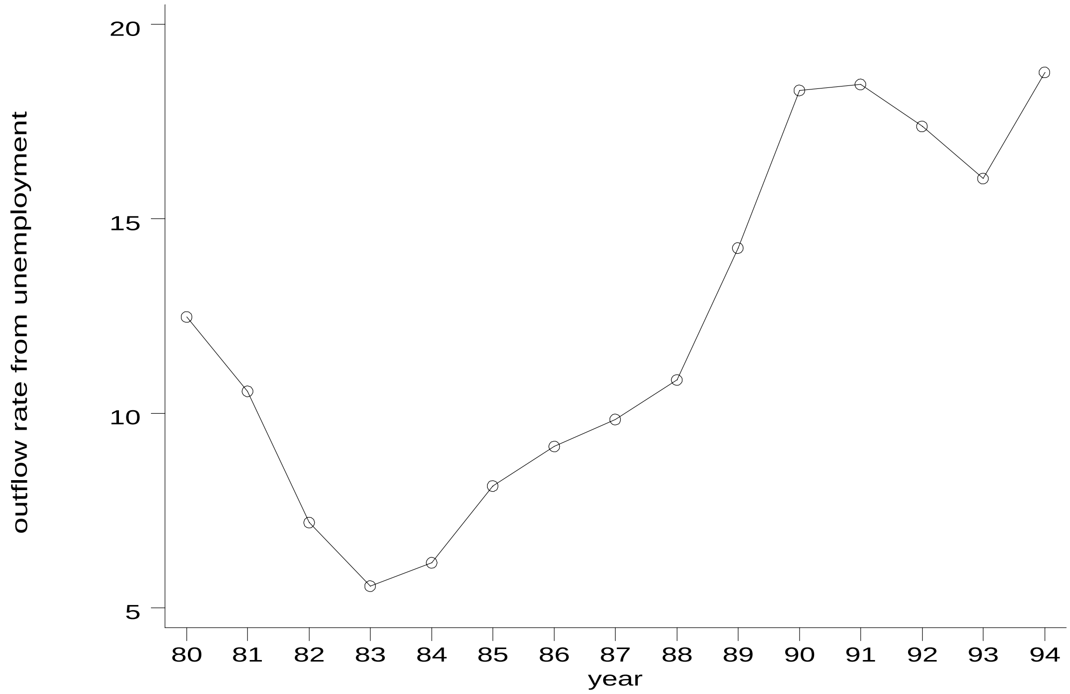

Figure 3: Evolution of outflow rates from unemployment into employment, 1980-1994. Source: MLR.

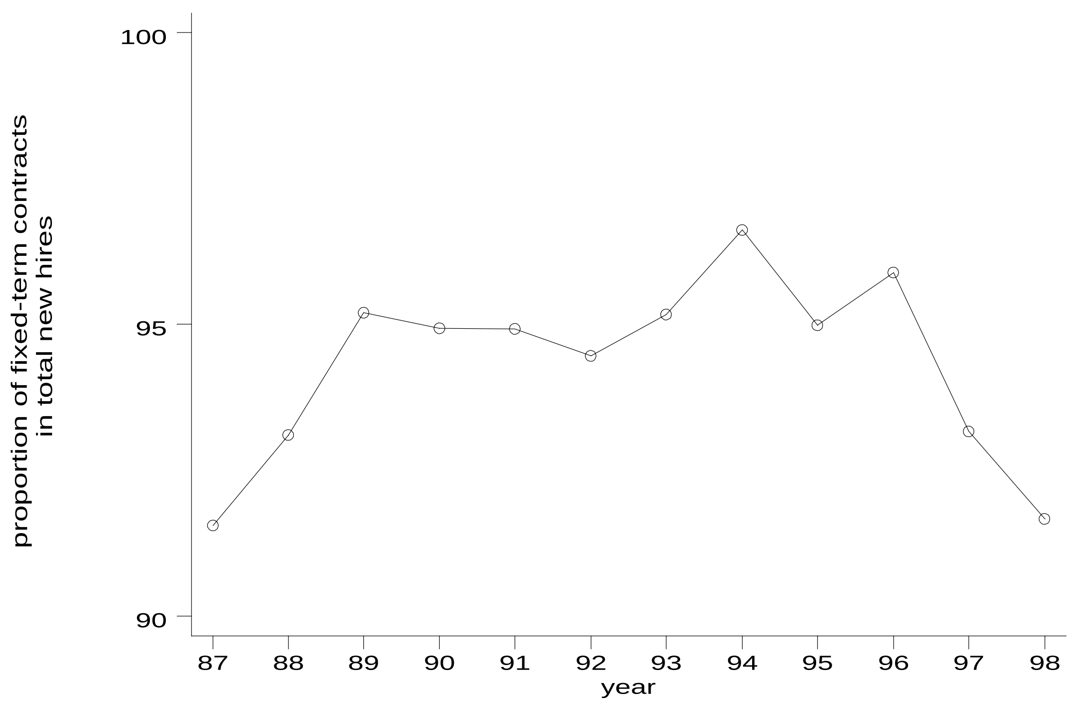

Figure 4: Evolution of the share of fixed-term contracts in new hires, 1987-1998. Source: MLR.

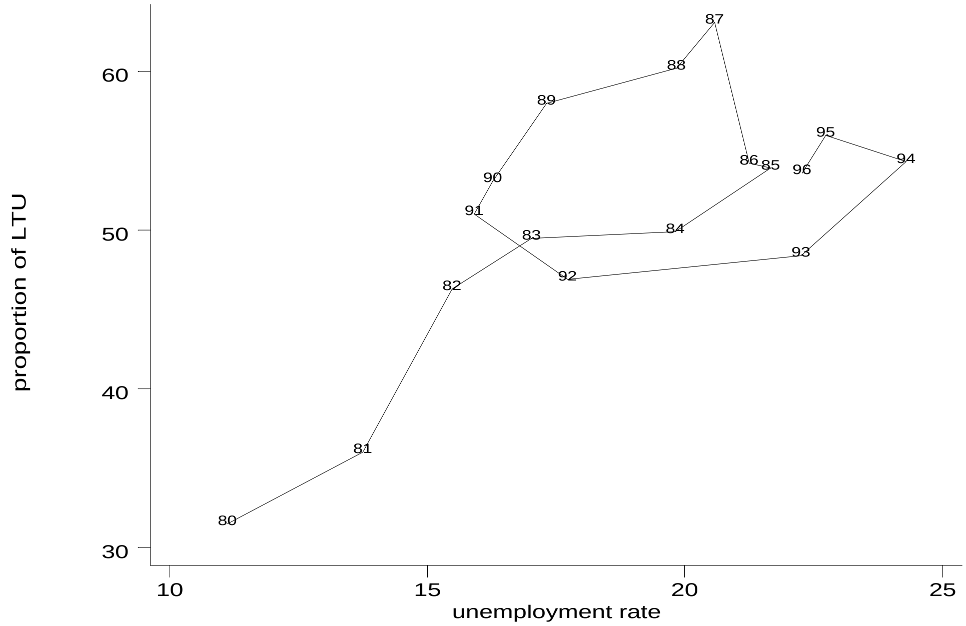

Figure 5: The incidence of LTU and the unemployment rate, 1980-1996. Source: EPA.

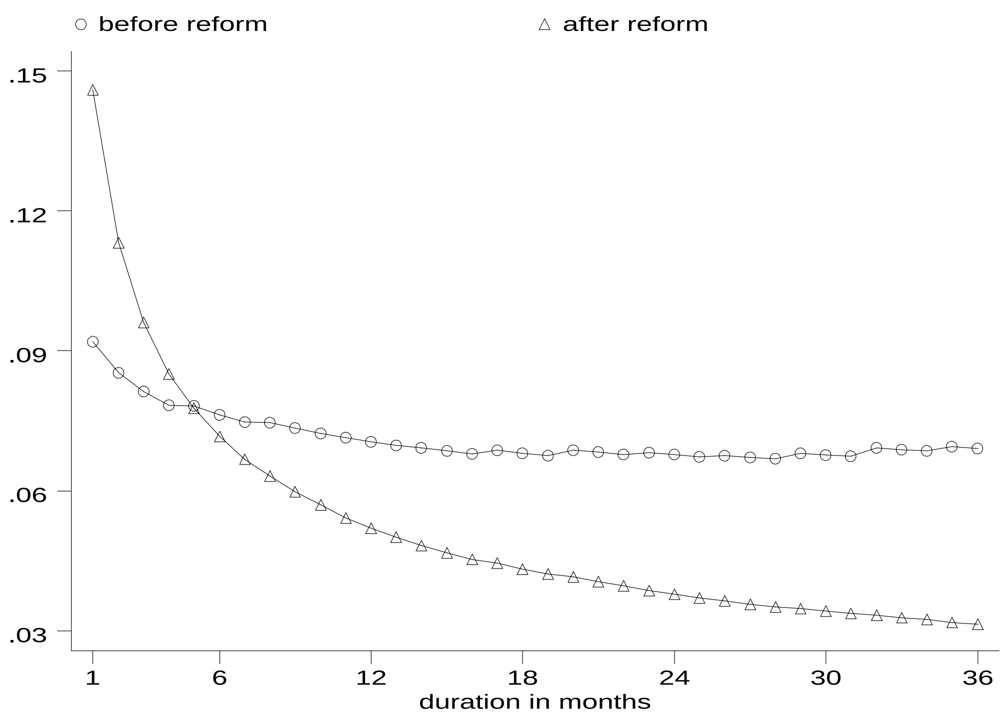

Figure 6: Hazard of leaving unemployment for the reference worker before and after the introduction of TC. Ref. category: age 35, not married, primary ed. or below, no kids, no working adults in household (see Table 3).

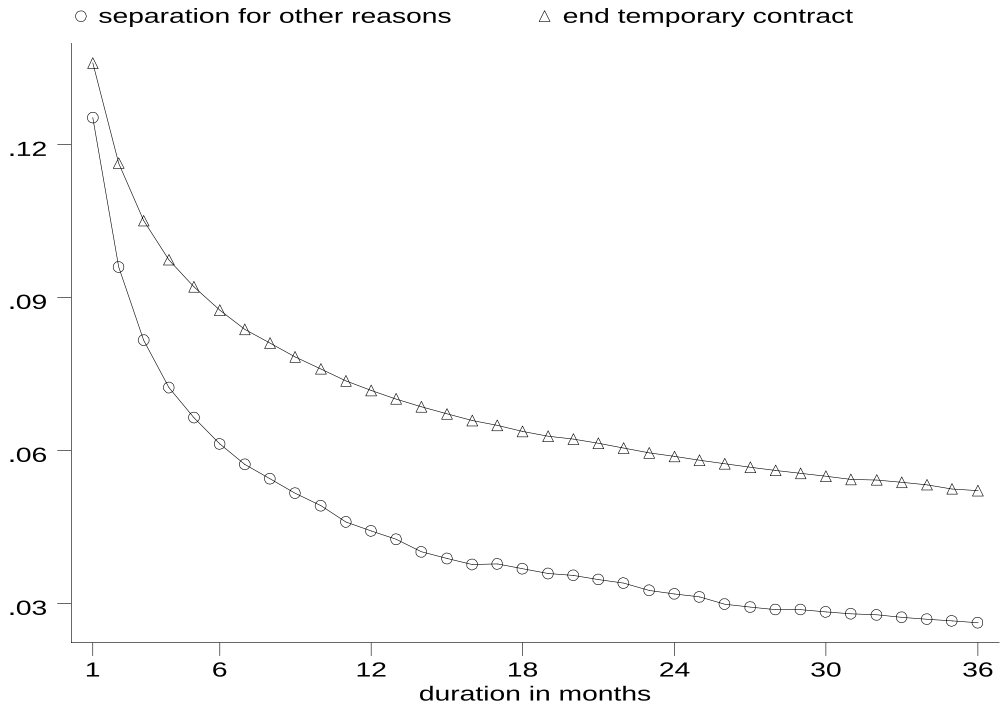

Figure 7: Hazard of leaving unemployment for workers who became jobless because ending of TC and for other reasons. Ref. category: age 35, not married, primary ed. or below, no kids, no working adults in household, no UI (see Table 7).

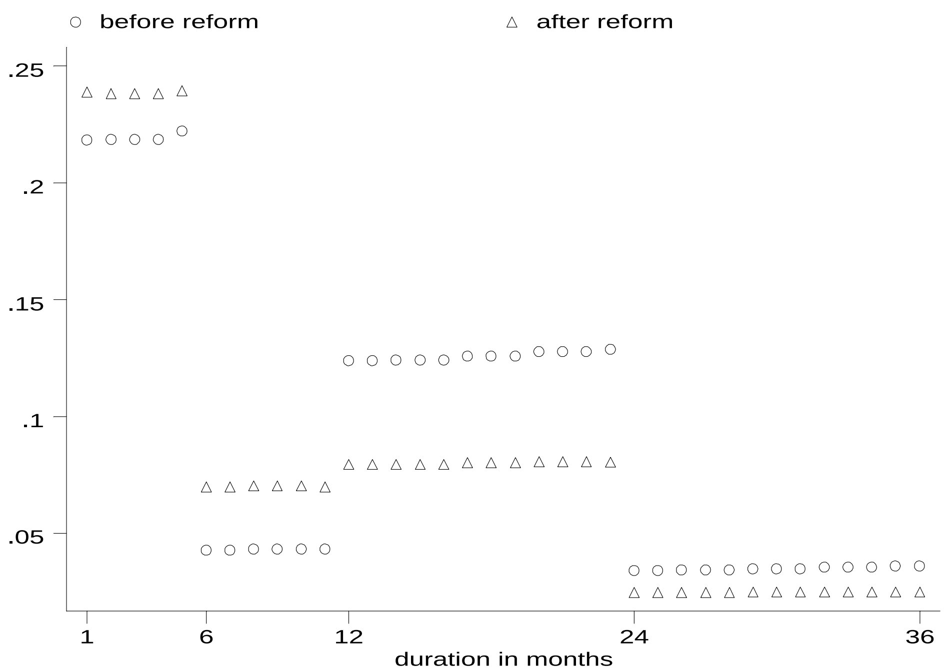

Figure 8: Hazard of leaving unemployment for the reference worker before and after the introduction of TC. Ref. category: age 35, not married, primary ed. or below, no kids, no working adults in household (see Table 9a, 9b).

References

[1] Aguirregabiria, Victor and César Alonso-Borrego (1999), "Labor contracts and flexibility: evidence from a labor market reform in Spain", Universidad Carlos III Working Paper No 99-27.

[2] Alba, Alfonso (1998): "Re-employment probabilities of young workers in Spain." *Investigaciones Económicas,* 22(2).

[3] Alba, Alfonso (1999): "Explaining the transitions out of unemployment in Spain: the effect of unemployment insurance", *Applied Economics,* 31, 183-193.

[4] Alonso-Borrego, César, Jesús Fernandez-Villaverde and José E. Galdón-Sánchez (1999), "A general dynamic equilibrium model of temporary contracts", Federal Reserve Bank of Minneapolis mimeo.

[5] Atkinson, Anthony B. and John Micklewright (1991): "Unemployment compensation and labour market transitions: a critical review", *Journal of Economic Literature*, 29, 1679-1727.

[6] Bentolila, Samuel and Juan J. Dolado (1994), "Labour flexibility and wages: lessons from Spain", *Economic Policy*, 18, 53-99.

[7] Blanchard, Olivier J. and Peter Diamond (1994), " Ranking, Unemployment Duration, and Wages", *Review of Economic Studies*, 61, 417-434.

[8] Bover, Olympia, Manuel Arellano and Samuel Bentolila (1997),"Unemployment duration, benefit duration, and the business cycle", *CEMFI working paper* No. 9717.

[9] Canziani, Patrizia and Barbara Petrongolo (2000): "Firing costs and Stigma: A theoretical analysis and evidence from micro data", *forthcoming European Economic Review.*

[10] García-Pérez, José Ignacio (1997): "Las tasas de salida del empleo y el desempleo en España (1978-1993)", *Investigaciones Económicas*, vol XXI (1), 29-53.

[11] García-Serrano, Carlos (1998): "Worker turnover and job reallocation:the role of fixed-term contracts", *Oxford Economic Papers* 50, 709-725.

[12] Grubb, David and William Wells (1993), "Employment Regulation and patterns of work in EC countries", OECD Economic Studies, 21, 7-58.

[13] Güell, Maia (1999): "The effects of fixed-term contracts on the duration distribution of unemployment: the Spanish case", mimeo London School of Economics.

[14] Güell, Maia (2000): "Fixed-term Contracts and Unemployment: an Efficiency Wage Analysis", Industrial Relations Section, Princeton University, Working Paper No. 433.

[15] Güell, Maia and Barbara Petrongolo (2000): "Workers' transitions from temporary to permanent employment: the Spanish case", Centre for Economic Performance, London School of Economics, Discussion Paper 438.

[16] Jenkins, Stephen P. and Carlos García-Serrano (2000): "Re-employment probabilities for Spanish men: what role does the unemployment benefit system play?" Institute for Labour Research Discussion Paper 00/55, University of Essex.

[17] Kugler, Adriana D. and Gilles Saint-Paul (2000): "How do firing costs affect worker flows in a world with adverse selection", mimeo.

[18] Layard, Richard, Stephen J. Nickell and Richard Jackman (1991) *Unemployment: Macroeconomic Performance and the Labour Market,* Oxford University Press.

[19] Machin, Stephen and Alan Manning (1999): "The causes and consequences of long-term unemployment in Europe" in Orley Ashenfelter and David Card (eds.), *Handbook of Labor Economics*, Volume 3. Amsterdam: North-Holland.

[20] Mortensen, Dale T. (1970): "Job search, the duration of unemployment and the Phillips curve", *The American Economic Review*, 60, 847-862.

[21] Mortensen, Dale T. (1977): " Unemployment insurance and job search decisions", *Industrial and Labor Relations Review*, 30 (4), 505-517.

[22] Nickell, Stephen J. (1979a): "Estimating the probability of leaving unemployment", *Econometrica,* 47 (5), 1249-1266.

[23] Nickell, Stephen J. (1979b): "The effect of unemployment and related benefits on the duration of unemployment", *The Economic Journal,* 89, 34-49.

[24] OECD (1993), *Employment Outlook*, Paris.

[25] OECD (1994), *The OECD Jobs Study*, Paris.

[26] OECD (1996), *Employment Outlook*, Paris.

[27] OECD (1999), *Employment Outlook*, Paris.

[28] Saint-Paul, Gilles (1996), "Dual labor markets, a macroeconomic perspective", Cambridge, MA: MIT Press.

[29] Schmitt, John and Jonathan Wadsworth (1993): "Unemployment benefit levels and search activity", *Oxford Bulletin of Economics and Statistics*, 55 (1), 1-24.

[30] Toharia, Luis (1997), *Spain,* Labour Market Studies series No. 1, Employment and Social Affairs, European Commission, Brussels.

[31] Wadsworth, Jonathan (1990): " Unemployment benefits and search effort in the UK labour market", *Economica*, 58, 17-34.

Appendix

A. The duration of unemployment in the EPA

The way in which the unemployed workers have been asked about their duration in unemployment and the possible answers given by the EPA questionnaires has changed over time. The table below summarizes these changes:

Table 10: The duration of the unemployment in the EPA

until 1987 (I)	1987 (II) - 1991(I)	from 1992 (I)
How long have you been looking for a job?	*How long have you been looking for a job?*	*Which day did you start looking for a job?*
Less than 1 month *1 to 3 months* *3 to 6 months*	If less than 2 years, *number of months*	*Month*
6 months to 1 year *1 to 2 years* *2 years or more*	If 2 years or more, *number of years*	*Year*

B. Unemployment benefits and duration

Table 11: Correlation of UI receipt and duration of unemployment

	1987-1994
all durations	-0.109
less than 3 months	0.093
less than 6 months	0.152
more than 6 months	-0.184
more than 12 months	-0.214

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