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ABSTRACT

This paper uses the "Encuesta de Condiciones de Vida Y Trabajo" (ECVT)--a survey of the labor force activity of over 61,000 persons in Spain in 1985 when unemployment exceeded 20%--to examine the effect of unemployment insurance (UI) and family status on longrun joblessness. It finds that (1) duration of joblessness is some 30% longer for those eligible for UI benefits than for those ineligible for UI; (2) the longterm unemployed are disproportionately secondary workers for whom the family serves as a form of welfare; (3) hazard rates linking the chances of jobfinding to duration of unemployment in the 1981-85 period of massive joblessness did not decline with duration; (4) the length of unemployment spells reduces wages moderately but has huge effect on the probability that re-employed workers take secondary sector jobs; (5) the UI eligible earn more and are more likely to gain regular full-time jobs than those ineligible for UI, congruent with the additional months of job search associated with UI.

The estimated effects of duration on the hazard and on earnings are consistent with the implications of labor supply and search analysis but not with the view that long unemployment spells create a class of unemployables. Our results imply a sizeable reduction in longterm unemployment with economic recovery.

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As the OECD country with the most severe unemployment problem in the 1980s[1] Spain offers an especially fruitful case for assessing the determinants and consequences of longrun unemployment. If unemployment insurance (UI) benefits are a major cause of longrun unemployment,[2] one would expect the Spanish UI system to be one of the most generous in Western Europe. If work skills deteriorate significantly as unemployment lengthens and/or if the most able obtain jobs first, one would expect the probability of escaping unemployment to fall with duration of joblessness. Finally, if the longterm unemployed are noncompetitive with the employed the wages and jobs they eventually obtain ought to differ significantly from those of the "insiders" who maintain their jobs. In this paper we use the "Encuesta de Condiciones de Vida y Trabajo" (ECVT) -- a survey of the labor force activity of over 61,000 persons in Spain in 1985 when unemployment exceeded 20% -- to examine these expectations. We find that:

1. Unemployment insurance increases the length of unemployment substantially: hazard rates for escaping unemployment fall with duration of benefits and rise sharply after benefits are exhausted, so that duration of joblessness is some 30% longer for those eligible for UI benefits than for those ineligible for UI.

2. The longterm unemployed are disproportionately secondary workers for whom the family serves as a form of unemployment insurance: women, older workers, and nonhousehold heads.

3. Hazard rates linking the chances of jobfinding to duration of unemployment in the 1981-85 period of massive joblessness were

constant or rising (depending on specification), ruling out a story of longterm unemployment in terms of its adverse impact on job-finding.

4. Wages fall moderately while the probability of entering the secondary sector increases markedly with length of unemployment for the displaced workers who obtain jobs. Re-employed persons eligible for UI earn more and are more likely to end up with regular full-time jobs than those ineligible for UI. The improved earnings associated with UI imply a reasonable payoff to the additional months of job search.

Data: The ECVT Survey

We derive these conclusions from analysis of the ECVT Survey which the Spanish government conducted in the fourth quarter of 1985 to obtain information on work history, underground economic activity, and earnings unavailable in the standard Spanish labor force survey, the "Encuesta de Población Activa" (EPA). Covering more than 61,000 persons aged 14 and over, the ECVT is the largest and potentially best cross section data set with which to analyze longrun unemployment in Europe in the 1980s. It includes over 4,500 workers with a spell of unemployment, most lasting over a year, and contains questions on work histories that provide data on complete as well as incomplete spells of unemployment and on workers who changed jobs without unemployment. The ECVT's major weakness is the absence of information on wages on the job prior to unemployment. To determine duration of unemployment on the ECVT, we use a series of questions regarding mobility and work status.

The key question asks respondents, "How many times have you changed firms?" An answer of once or more triggers questions about whether the most recent change involved a spell of unemployment, the reasons for the change, and characteristics of the previous job. By excluding persons unemployed and recalled by the same firm, the ECVT bypasses problems created by the distinct search behavior of persons on temporary layoff (Katz 1986). The survey differentiates between involuntarily displaced workers (those whose firm closed or who were fired or permanently laid off, and those whose contract ended);[3] and those who left their job for voluntary reasons. A follow-up question, "After leaving this job, how long did you spend (or are you) out of work and actively looking for a job?," allows us to identify those who sought work after losing their job.

We concentrate on the last spell of unemployment of displaced workers, giving us a sample of 4,517 workers who lost their job.[4] From the question on labor force status in the previous week, we determined that 1,896 of those who suffered joblessness were working, 1,365 were looking for work in that week,[5] and 1,256 were neither working nor looking for work in the reference week. As comparison groups, we also look at employed workers who were never unemployed during the period: 858 respondents who changed jobs without a spell of unemployment, and 7,237 who never changed their job and were employed in the reference week.[6] In several tabulations we distinguish between workers who lost their jobs during the 1981-85 recession and those who lost their jobs in earlier years, which reduces our samples moderately.

The unemployment data in the ECVT is consistent with published data on unemployment in Spain from the EPA survey. The rate of unemployment in the ECVT is approximately the same as in the EPA.[7] In our sample of displaced workers, 54% of all unemployment spells and 63% of incomplete spells lasted one year or more. This compares to EPA-based estimates of incomplete spells that 57% of the unemployed were out of work for one year or longer (OECD).[8] More detailed analyses of the ECVT (Muro et al. 1989) show that it is also broadly consistent with other information on the Spanish labor market, making it a valid survey from which to assess longterm unemployment in the country.

Table 1 provides a descriptive overview of the samples from the ECVT that we use. It records the duration of joblessness, demographic characteristics, unemployment insurance eligibility, and reason for joblessness for: all displaced workers who suffered unemployment; those who were displaced but did not become unemployed; persons who remained at the same job during the period;[9] and the shortterm and longterm unemployed. In this table and throughout our analysis, we define the longterm unemployed as persons with 12+ months of unemployment if they were employed in the reference week or 6+ months if they were unemployed at that time.[10]

Comparing workers who lost their jobs (column 1) to those who were employed throughout the period and never changed firms (column

2), we see that the displaced jobless were more likely to be men, young, less educated and blue collar workers. They are less likely to have other family members working than workers with jobs. Comparing workers who became displaced and experienced unemployment (column 1) with those who were displaced without unemployment (column 3), we see that younger workers, females, nonhousehold heads, and those displaced in 1981 and after are more likely to have suffered joblessness. Finally, comparing the longterm and shortterm displaced jobless (columns 4 and 5), we see that the longterm jobless are older (many young unemployed persons never obtained a first job and are thus excluded from our sample); are more likely to be female; and are more likely to have lost their job as a result of their firm shutting down or being laid off than other displaced workers.

The Spanish UI System

Economic analyses of longrun unemployment often stress the importance of UI in providing income for job losers that allows them to search for better jobs and remain jobless for extended periods of time (Burdett 1979). Spain has a generous unemployment insurance system compared to most OECD countries (see OECD, 1988), which makes UI a potential cause of the long duration of Spanish unemployment. According to Spanish labor law, an individual involuntarily unemployed who contributed social security payments on the previous job for 6 or more months is eligible for UI benefits. The magnitude and duration of UI benefits have varied over time. Prior to 1980 a worker received 80% of the average contributory wage in the previous six months of employment during the first 6 months of unemployment, 70% of that wage for the next six months, and 60% of that wage for an additional 6 months. After 1980 the length of eligibility for UI depended on tenure in the previous job. If the previous job lasted between 6 and 12 months, the period of eligibility is 3 months, while for every 6 additional months of tenure in the previous job, eligibility increases by 3 additional months to a maximum of 18 months. After 1984 the maximum duration of entitlement to receive UI under the contributory system was extended to 24 months. Workers over 55 were eligible for early retirement benefits, leading us to distinguish them from others in the analyses.

The legally mandated differences in length of eligibility among workers with different work histories over time provides us with the variation in potential benefits needed to estimate how UI duration affects length of joblessness. Column 1 of Table 1 shows that in our sample 70% of the displaced jobless were eligible for UI; 17% were eligible for 3-6 months; 7% were eligible for 9-12 months, and 46% were eligible for 15-18 months. The first indication that UI may affect longterm unemployment can be seen in columns 4 and 5 of the table, which show that 77% of the longterm unemployed compared to only 58% of the shortterm unemployed were eligible for UI--a huge 19 point differential.

Determinants of Longterm Unemployment

To assess the effect of UI eligibility and other factors on longterm unemployment in Spain,[11] we decompose the probability of being longterm unemployed using conditional probabilities into: the probability of being longterm unemployed given that the worker is displaced and jobless; the probability of being displaced and jobless conditional on being displaced; and the probability of being displaced. The probit equations in Table 2 estimate the determinants of these probabilities and also give the probability of being currently employed conditional of having been displaced and jobless. The column 1 calculation shows that, consistent with the means in Table 1, eligibility for UI substantially raises the probability of longterm unemployment given that one becomes unemployed.[12] According to the estimated coefficients, a person with mean characteristics who was displaced but not eligible for UI had a probability of being longterm unemployed of 54%; the .27 coefficient in the table raises the probability by 11 percentage points to 65%. This is smaller than the 19 point difference in the probability of being longterm unemployed found in the table 1 means but still sizeable. While UI is not the cause of longterm unemployment (some 23% of the longterm unemployed were not eligible to receive UI benefits; an additional 22% of the displaced workers were unemployed for more than two years, exhausting their eligibility)[13] the calculation suggests that it is an important contributing factor.

The probit in column 1 also shows that nonhousehold heads are more likely to be longterm unemployed than others. This is consistent with the notion that the longterm unemployed rely on their family for financial support -- i.e. that the strong family system substitutes for unemployment insurance. In fact, if we distinguish between the longterm unemployed by eligibility for UI, we find a remarkable difference in household status. While a majority (62%) of the longterm unemployed who are eligible for UI are household heads, just 31% of the longterm unemployed ineligible for UI are household heads. We interpret this to mean that dependents can afford to be longterm unemployed without UI but that household heads cannot.

The probit in table 2 gives the probability of being currently employed conditional on having been displaced and jobless. It shows that those eligible for UI are less likely to be currently employed while those who are heads of households are more likely to be currently employed, consistent with the column 1 findings on longrun unemployment. The education and previous job tenure coefficients obtain oppositely-signed coefficients consistent with the column 1 findings.

In contrast to the estimated significant effect of UI eligibility on being longterm unemployed in column 1, the column 2 estimate of the conditional probability of suffering unemployment given displacement shows no UI impact. The implication is that UI lengthens durations of unemployment but does not affect the occurrence of unemployment given displacement. Two factors appear to explain this: one is that most displaced workers without unemployment (61%) were displaced before 1981, when unemployment just began to grow in Spain; the other is that a large proportion of those displaced without unemployment become self-employed (27% compared to 15% of displaced jobless workers). Finally, the column 3 estimate of the probability of being displaced shows that older, more educated and white collar workers are less likely to have lost their job than other workers, but that household heads had neither a greater nor lesser chance of losing their job than others. As nearly all of those never displaced are eligible for UI, UI cannot be cited as a cause of displacement, though in some seasonal employment sectors it may be part of the industry's normal adjustment pattern.

UI and Duration of Joblessness

To get a better fix on the effect of UI on duration of joblessness, we estimated Kaplan-Meier empirical survival functions[14] for persons differentiated by months of UI eligibility. If UI affects the hazard rate, the survival function for those without UI should fall below the survival function for persons eligible for UI, and be lowest for those with the least eligibility. Figure 1, which records our estimated survival functions, supports both expectations: the lowest survival function is for the non-UI eligible; the next lowest is for those with 3-6 months eligibility; the next is for those with 9-12 months eligibility; and the highest is for those with 15-18 months eligibility. While the absence of covariates leaves open the possibility that the differences reflect measured heterogeneity among the four groups, the figure makes a strong prima facie case that the duration of UI benefits in Spain has played a role in extending spells of unemployment.

Figure 2 takes the analysis a step further by recording the hazard rate for escaping unemployment by the length of time until a person with UI exhausts his eligibility. It groups individuals by three month intervals to deal with the problem that people remember and report retrospective losses of jobs and job-finding in focal intervals of 1 month, half a year, a year, etc. A positive number on the horizontal axis means that the person has that number of months of eligibility remaining; a negative number implies that they used up their eligibility that many months ago. If UI affects the hazard rate, we would expect lower hazards during the period of eligibility followed by a jump in the hazard around the time the individual loses eligibility and possibly greater hazards after loss of eligibility than during eligibility. The data support these expectations: the average hazard in the period when the persons have more than two months eligibility is .051; it jumps to .104 when eligibility falls from two months to minus six months, then averages .045 per month.

To obtain a better fix on the shape of the hazard function for escaping joblessness, we calculated empirical hazards for all displaced persons and for those displaced in the 1981-85 period of high unemployment. We used 3-month grouped data and 1 year grouped data to reduce the problem of reporting spikes. These hazards are displayed in Figure 3. The 3-month grouped data show spikes around 12, 18, 24 months, which makes it difficult to determine the shape of the hazard. For persons displaced in 1981-85 the 1-year grouped data gives a declining hazard for the first year and a relatively flat hazard thereafter; for all displaced persons, it gives a declining hazard for three years followed by a relatively flat hazard. Whether the initial drop in hazards is due to observed heterogeneity or other factors requires analysis of the effect of covariates on duration of joblessness, to which we turn next.

Weibull and Cox Proportional Hazard Models

To control for covariates correlated with UI eligibility and with duration of joblessness, we first estimate a hazard model using the Wiebull distribution: $h(t:x) = h0(t) \exp (X'b)$ where $h0(t)$ is an arbitrary baseline hazard for the duration of unemployment from the Wiebull family at^{a-1}, and where the vector X reflects characteristics and b measures the effect of X on the hazard. In this formulation, the coefficient a reflects potential duration dependence, with a>1 implying an increasing escape rate. We estimated the model by maximum likelihood, taking into account both uncompleted (censored) spells and completed spells.[15]

The results of the analysis are shown in Table 3 for the entire sample (column 1); for workers whose contract ended and who were thus likely to have anticipated job loss (column 2); for other job losers (column 3); and for persons who lost their job in the period 1981-1985 (column 4). The dummy variables for UI eligibility -- ELIG36, ELIG912 and ELIG1518 -- obtain significant coefficients in all of the calculations save for workers who lost their job for a reason other than their contract ending. One possible explanation is that workers on temporary jobs use UI benefits to adjust to intermittent spells of joblessness, moving from one short run job to another, while those whose displacement comes as a complete surprise do not make such plans. The estimated scale coefficients are around unity in columns 1-3 and exceed unity for persons displaced in 1981-85 in column 4. All of these estimates reject the notion that the hazard rate falls with duration.

How important is the estimated UI effect on duration of joblessness? The magnitude of the coefficients on the UI dummy variables in columns 1, 2 and 4 of Table 3 indicate that the effect is substantial in social as well as statistical terms. Though the coefficients increase only moderately between the eligibility categories, they all differ greatly from the 0 for the deleted UI-ineligible group: for the entire sample duration of joblessness ranges from .20 to .26 ln points higher for those eligible for UI than for those ineligible for UI (column 1); for those displaced in 1981-85, duration of joblessness is .21 to .34 ln points higher for those eligible for UI (column 4). Translated into months of unemployment, these estimates indicate that the UI eligible are unemployed on the order of 4 to 6 months more than the ineligible! This result is, we stress, found in the raw means of our data and is thus not an artefact of the Weibull model. Dividing the sample between those eligible for UI and those ineligible for UI and calculating mean durations of joblessness for the two groups we find that the mean duration of completed spells was 17 months for persons eligible for UI compared to 11.4 months for those ineligible for UI; and that the mean duration of uncompleted spells was 26 months for the UI eligible compared to 18.2 months for the UI ineligible -- large differences by any reckoning.[16]

The coefficients on the other variables in the Table 3 calculations provide additional insight into the determinants of longrun joblessness in Spain. Columns 1 and 4 of the table show that workers displaced as a result of a cutback in work and those individually fired had long duration of joblessness compared to workers whose contract ended (the deleted group). Workers whose contract ended essentially have advance notice of displacement, and thus are likely to begin their search process earlier. Workers who are fired are likely to be negatively selected and have a harder time finding work thereafter (Gibbons and Katz 1989). The variables that reflect human capital -- education and previous job tenure -- have very different effects on duration of joblessness. Education reduces duration while previous tenure raises it.[17] If we interpret tenure as reflecting specific human capital, and education as reflecting general human capital, these differing effects suggest that specifically trained workers have greater problems coping with displacement, possibly because they suffer greater losses of earnings capacity and adjust their reservation wages relatively slowly.

Additional Estimates

We made two further estimates of the effect of UI on longterm unemployment and of the shape of the underlying hazard rate, using more complicated econometrics. First, relaxing the Weibull assumption of a constant hazard rate, we estimated a non-parametric Cox proportional hazard model. In this specification, the dependent variable is the escape rate rather than the log duration of joblessness, so that the signs of the estimated parameters have the opposite meaning to those in the Wiebull formulation: a covariate with positive sign indicates an increase in the escape rate and thus a reduction in duration of joblessness. Column 1 of Table 4 gives the estimated parameters in the Cox model. They confirm our principal findings: UI eligibility substantially increases the length of time jobless, while being a household head reduces it. They also confirm that education and tenure have opposite effects on duration of joblessness.

Second, to obtain the strongest possible test of the UI effect, we estimated a proportional hazard model with time varying covariates for eligibility for UI. To do this we organized our data into a three-month-persons file, in which individuals enter the file several times to reflect their time unemployed. Here, UI eligibility is measured by a person's status at each interval so that a person unemployed for six months but eligible for three months of UI is coded eligible the first time he is in the file but not the second time. This approach provides an especially useful test of the effect of UI as it distinguishes the behavior of people during and after eligibility. In this analysis we use two different measures of the UI incentive. In column 2 of Table 4 our measure (TRUI) is the months of eligibility that a person had in each interval: for instance, TRUI takes the value 5 months if the person is in month 1 of unemployment and is eligible for 6 months of UI; it is 2 months if they are in month 4 of unemployment and eligible for 6 months of UI; and it is 0 if they have not been or became ineligible. The results of these estimates confirm that UI significantly reduces the probability of leaving unemployment. In column 3 of the table we use a slightly different specification of the UI variable to take account of the declining replacement rate as duration of eligibility increases. The variable here, BENEF, takes the value 0.8 for the first six months of eligibility (when people receive 80% of their previous wage); 0.7 for the next six months; 0.6 for the remaining period of eligibility; and zero for the rest of time of unemployment. It has an even stronger effect on the estimated hazard. The stronger estimated relation between UI and duration of joblessness as we improve our measure of the economic incentive of UI supports a behavioral interpretation of our findings.

What is the shape of the hazard rate in the Cox proportional hazards model? To calculate this we estimated survival functions using the model in column 1 of Table 4.[18] Once the survival function was recovered, we regressed log(-log (S)) on a constant and logarithm of months of unemployment spell. The coefficient on the log of unemployment is our estimate of the duration dependence parameter, corrected for the relevant covariates.[19] The results of this exercise are given in Table 5, which records estimated duration dependence parameters for the sample as a whole, for those who became unemployed in the period 1981-85, and for those who became unemployed earlier. With no covariates the estimated relation between the hazard and duration is .65 for the entire sample and below one for the subsamples as well, consistent with the declining hazard shown in Figure 3. Inclusion of the covariates in the Cox model raises the estimated duration parameter to .79, indicating measured heterogeneity in our sample.

When we decompose the sample between persons who lost their job in 1981-85, and those who lost their job earlier, however, we see that the overall sample result is due to very different patterns for the two groups. The estimated duration dependence parameter is .97 after controlling for observed heterogeneity for those who became jobless in 1981-85, implying little or no duration dependence, but is significantly downward sloping for those who lost their job prior to 1981. This pattern mirrors the differences in the Weibull model duration dependence estimates, also given in the table: a rising hazard for people who lost their job in 1981-85 and a falling hazard for people who lost their job earlier. The difference in the shapes may reflect differences in unobserved heterogeneity between the periods, with greater heterogeneity in the earlier period when few people were displaced compared to the latter period when relatively many were displaced.[20] Since unobserved heterogeneity produces declining hazards,[21] the overall implication is that **the hazard rate did not decline for persons displaced in 1981-85**. This casts doubt on any explanation of longrun unemployment as resulting from an extraordinary adverse effect of longrun joblessness on labor skills and the probability of finding a job.

Acceptance Wages and Jobs

The other side of the search analysis of unemployment relates to the wages and jobs the unemployed eventually obtain. To what extent does longterm unemployment induce workers to accept less desireable jobs and reduced wages? Does UI enable workers to find better employment as a result of extended job search?

Table 6 seeks to answer these questions.[22] Columns 1-3 record regression coefficients of the effect of UI eligibility and duration of joblessness on the log wages of unemployed workers who obtained jobs at the time of the survey, controlling for standard demographic and human capital variables.[23] As the re-employed are a select group of the jobless we include an inverse Mills ratio (based on the probit in column 2 of Table 2 for whether someone who became unemployed obtained a job at the time of the survey) in the column 3 calculation: it had little effect on the results. Column 4 turns from wages to the type of job a person obtains: it records probit coefficients for the probability that a re-employed worker obtains a regular full-time job.

The estimates show substantial effects for both UI and duration of joblessness on wages and the type of job the unemployed obtain. The column 1 regression indicates that eligibility for UI has a moderate 3.6% positive effect on earnings, which rises to 5.7% when we control for duration of unemployment and selectivity bias. While lack of data on the individuals' earnings on the previous job makes inferences about the impact of longterm unemployment on wages shaky, these estimates are consistent with rational search behavior as the jobless with UI wait until they find a higher paying job even exclusive of any value placed on leisure during the unemployment spell. Consider, for example, a worker whose spell of joblessness is increased by 6 months by UI. During this time he receives approximately 80% of his previous pay, so that his annual earnings are about 10% ((100-80) x 1/2 year) lower. If the workers' pay is 3.6% higher as a result of the extended search, the payback period is less than 3 years.

Column 2 shows that a year of joblessness reduces earnings of workers by about 3%. We interpret this in the framework of the search theory model as reflecting depreciation of work skills with joblessness. In the standard search model the reservation wage (W*) is related to the value of search (V) and the relevant interest rate (r) (Mortensen 1986): $W^* = rV$. An n% depreciation in earnings capacity or human capital due to joblessness decreases the value of search proportionately, and thus has a proportionate impact on W*. If the sole reason for the decline in earnings with duration is depreciation of human capital, the 3% coefficient would be equal to the coefficient on experience in an earnings equation for never displaced workers. In fact, in an earnings equation for never displaced workers,[24] we obtain a coefficient on experience of .018, which suggests that the reservation wage of workers falls for other reasons as well (i.e. liquidity constraints (Mortensen 1986); learning (McCall 1989; Burdett and Vishwanath 1988)).

The coefficients on other variables in the wage regressions of Table 6 are consistent with a human capital interpretation of the effect of longterm joblessness on earnings. The variable SAMEOCC, which takes the value 1 if a person is re-employed in the same occupation, is positive, as one would expect if there are occupation-specific skills. Returns to experience and to education are lower for workers who were displaced than for workers who were not displaced (the experience coefficient for the not displaced was .018 while the coefficient on schooling was .05 -- see note 22); which suggests greater losses in human capital for those with greater past investments. Again, however, the absence of wages for individuals prior to displacement means that we cannot rule out the possibility that the results reflect unobserved skill differences.

The final and most striking result in Table 6 is found in the column 4 probit equation. It shows that UI eligibility has a huge effect on obtaining a regular full-time job, while duration of joblessness has, by contrast, a substantial negative effect. The coefficient on UI implies that a person in the sample with the mean characteristics and who is eligible for UI had a 12 percent higher chance of obtaining a full-time regular job than that person would have had absent UI. The coefficient on duration implies that a year of unemployment reduces the chances of obtaining a regular-full-time job by some 7 percentage points. That UI operates largely through the type of job it enables people to find is indicated further in the reduced coefficient of UI in the ln earnings equation when we add the regular full-time variable to the earnings regression in column 3.

Conclusion

Our findings -- that UI substantially prolongs the duration of unemployment in Spain; that longterm unemployment is concentrated among secondary workers for whom the family serves as a form of UI; that the hazard rate was relatively flat among those displaced in 1981-85 when unemployment skyrocketed; and that duration of unemployment reduces wages and the chances of obtaining a regular full-time job -- are broadly consistent with the implications of labor supply and search analysis for understanding the behavior of the longterm unemployed. The estimated magnitudes of the effects of duration on the hazard and on earnings are, however, inconsistent with the view that long durations of unemployment creates a class of unemployables. To the contrary, they suggest that longterm unemployment is more a problem of a slow rate of receipt of job offers relative to reservation wages (which are raised by UI), than one of extraordinary loss of labor market skills as time unemployed proceeds. If our results are correct, economic expansion in Spain will reduce longterm unemployment at a reasonable pace. In addition, the evidence of improved earnings with UI eligibility suggests that the UI-induced increased job search has an economic payoff for those workers.

Endnotes

1. In 1985 the rate of unemployment in Spain was above 20% and incomplete spells of unemployment averaged more than two years. Fina 1988 analyses the causes of this high unemployment.

2. Studies have addressed the effect of UI on duration of unemployment. See, for example, Ehrenberg and Oaxaca (1977), Moffitt (1985) and Katz and Meyer (1988) for estimates for the United States; and Atkinson et al. (1984) and Narendranathan et al. (1985) for estimates for the United Kingdom. But they have deal with much shorter spells of unemployment than in Spain.

3. We classify persons who lost their job when their contract ended as involuntary job losers. We differentiate them from other job losers in some calculations.

4. We deleted 62 persons with over 120 months of unemployment as likely to be out of the work force. Including them had no effect on our results.

5. Some workers who said they were actively looking for a job since becoming displaced were not looking for work in the reference week. We did some calculations excluding them from our sample, obtained the same results as in the text, and included them in the calculations.

6. To determine the number of workers actually employed, we applied further minor restrictions to allow for persons not working because of illness, vacation, etc., and for problems of similar nature. Moreover, as the survey contains questions for those who did no work in the previous week aimed at revealing occasional work done in the last three months, we impose having worked a minimum of 13 hours per week in the last three months to consider a worker as currently employed.

7. The detailed ECVT questions on underground economic activity reduce the rate of unemployment by just 3 or so percentage points.

8. We attribute the modest difference between the estimates to inclusion of youths and those who leave their jobs voluntarily in the EPA but not in our data.

9. This sample is composed of workers who reported never changing firms; some may have been out of work at one time but were recalled by the same firm. The sample excludes workers who have moved without experiencing displacement (voluntary turnover) or who were displaced without unemployment.

10. This definition assumes that persons with incomplete spells of 6+ months will average complete spells of one year or more. We also examined those with complete or incomplete spells of 12+ months and obtained similar results to those reported.

11. See Dell'Aringa and Lodovici (1988) for an analysis of the determinants of unemployment duration for displaced workers in Italy.

12. We investigated whether non-eligibility because the worker did not make payments to the UI system or because the worker held a job for too short a period had different effects on duration of unemployment. The differences were insignificant.

13. In addition, many unemployed youths never held a job and thus were never eligible for UI benefits in the first place.

14. In an uncensored sample, the survival function is the cumulative survival rate -- proportion of persons who remain unemployed at the end of each interval of time. As our sample involves right censoring, the product limit or Kaplan-Meier estimate of the survival function provides the correct statistics. Let n_i be the number of workers unemployed (risk-set) in the beginning of the interval i, and d_i the number of workers who find jobs in the course of that interval. Then $p_i=(n_i-d_i)/n_i$ is the proportion of workers who remain unemployed at the end of the interval i. The product limit estimate is $P_i=p_1.p_2...p_i$. Notice that $1-p_i$ is the empirical hazard rate. See Kalbfleisch and Prentice 1980, pp. 10-16 for more details.

15. The model in reduced form is $Y = \mathbf{x}\beta + \sigma\epsilon$, where $Y=\log(T)$, $\beta=-\sigma b$ is a vector of unknown regression parameters, $\sigma = 1/a$ is an unknown Weibull scale parameter, and ϵ is a vector of errors with an extreme-value distribution (Kalbfleisch and Prentice 1980, p. 24). The model is estimated by maximum likelihood using a Newton-Raphson algorithm through SAS program. See Kalbfleisch and Prentice, 1980, pp. 84-86, and Lancaster 1979, for details about the form of the likelihood function.

16. In addition, we estimated an exponential model with a 0-1 dummy variable for UI eligibility and calculated the expected mean duration for those eligible and for those ineligible for UI. The coefficient on the dummy for UI eligibility was .22. The expected mean duration for those ineligible was 27 months, so that eligibility raised duration of joblessness by nearly 6 months, controlling for other differences between workers. Thus, this model gives comparable results to those in the text.

17. Given the definition of UI eligibility, it is highly correlated with previous job tenure. This might be an explanation why eligibility is not significant among workers displaced for

reasons other than ending of contract. Note that tenure has a strong effect on duration of unemployment for the same group of workers.

18. After estimating the time fixed Cox model, we recovered the survival function given the covariates using the STATA program. See Kalbfleisch and Prentice 1980, pp. 84-87.

19. In the Weibull distribution, the survival function is $S(t)=\exp[-(ht)^{a}]$. Hence, $\log[-\log S(t)]=a(\log h+\log t)$ and the coefficient of log t, given S(t) as an estimate of the survival function, measures duration dependence.

20. Some workers displaced before 1981 were still unemployed at the end of 1985. Those who found a job are clustered either around short spells or around very long ones. This implies that most people found a new job quickly and kept them thereafter. Many of those who found a job in the period, and moved voluntarily afterward, are not in the sample of displaced workers. If they were displaced again their latest spell has precedence in our sample since the survey only reports the last spell of unemployment.

21. Those with higher propensities of finding jobs exit the sample earlier than those with innately lower hazard rates.

22. See Addison and Portugal (1989) and Podgursky and Swaim (1987) for evidence for the U.S.

23. As wage variable in the ECVT is coded, we have converted it into a continuous variable by taking the median of the intervals.

24. The estimates for workers not displaced are the following:

	Coeff.	t
CONST.	9.745	(233.2)
SEX	0.139	(12.84)
EXPER	0.018	(15.37)
EXPER2	-0.0004	(-13.01)
EDUC	0.053	(39.10)
ADJ.R-SQ.	.50	
N	7140	

Other variables included are: WCOLL, NCHIL, OTHERW, REGULF, and dummies for regions and sectors.

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TABLE 1.- DEFINITIONS AND MEANS OF VARIABLES IN ECVT, BY EMPLOYMENT/UNEMPLOYMENT STATUS

		Displ. jobless	Never displ.	Displ. non-jobless	Shortrun unempl.	Longrun unempl.
DURATION	Months unemployed	19.419			3.536	29.923
LONG DUR	=1 if longterm unempl.	0.602			0.0	1.0
AGE	Age at survey date	36.301	38.046	40.613	33.990	37.831
AGEDISP	Age at displacement	33.390		36.889	32.075	34.249
SEX	=1 if male	0.690	0.588	0.808	.751	.650
EDUC	Years of schooling	6.993	8.647	6.959	7.182	6.869
YDISP0	=1 if disp. before 1981	0.263		0.609	.219	.293
YDISP13	=1 if disp. 1981-1983	0.272		0.139	.134	.364
YDISP4	=1 if displaced in 1984	0.173		0.072	.107	.217
YDISP5	=1 if displaced in 1985	0.262		0.171	.504	.102
FIRED	=1 if individually fired	0.172		0.163	.136	.196
DOWN	=1 if firm shut down	0.261		0.306	.181	.315
CUTBACK	=1 if cut back on work	0.075		0.066	.053	.090
PROV.P.	=1 if fired after provisionary period	0.017		0.017	.021	.015
END	=1 if end of contract	0.471		0.446	.607	.382
WCOLL	=1 if White collar	0.179	0.494	0.194	.165	.188
HEAD	=1 if household head	0.552	0.520	0.719	.563	.545
NCHIL	Dependent children	1.083	0.990	1.259	1.092	1.076
OTHERW	Others working in the family	0.499	0.612	0.453	.512	.491
ELIGUI	=1 if eligible for UI	0.697		0.753	.582	.773
ELIG36	=1 if eligible for 3-6 months of UI	0.173		0.107	.200	.155
ELIG912	=1 if eligible for 9-12 months of UI	0.066		0.047	.057	.072
ELIG1518	=1 if eligible for 15-18 months of UI	0.458		0.598	.325	.547
AD55	=1 if AGEDISP>=55	0.104		0.103	.090	.114
TENURE	Previous job tenure	4.599		5.303	3.181	5.541
EXP	Age-Education-6	23.306	23.370	27.634	20.799	24.963
REGULF	=1 if re-empl. in a regular full-time job	0.591	0.778	0.745	.577	.609
LWAGE	Log monthly net earnings in current job	10.586	10.686	10.668	10.599	10.570
SAMESECT	=1 if re-employed in the same sector	0.486		0.489	.555	.403
SAMEOCCU	=1 if re-employed in the same occupation	0.539		0.465	.621	.442
SAMPLE SIZE		4517	7237	858	1798	2719

Source: Tabulated from ECVT survey.

TABLE 2.- ESTIMATES OF PROBIT COEFFICIENTS FOR CONDITIONAL PROBABILITIES OF JOBLESS STATUS. (T STATISTICS)

	(1)	(2)	(3)	(4)
Depend. var.:	Longrun unempl. =1	Currently employed =1	Displaced jobless =1	Displaced =1
Sample:	All displaced jobless	All displaced jobless	All displaced workers	All workers
ELIGUI	0.274 (4.92)	-0.137 (-2.53)	-0.046 (-0.79)	
SEX	-0.401 (-6.75)	0.536 (9.63)	-0.172 (-2.70)	.068 (2.23)
AGEDISP	0.008 (2.98)	-0.017 (-6.43)	-0.010 (-3.74)	
AGE				-.008 (-7.60)
EDUC	-0.007 (-0.98)	0.027 (3.69)	-0.026 (-3.53)	-.034 (-9.70)
TENURE	0.008 (2.03)	-0.011 (-2.83)	0.001 (0.43)	
YDISP0	1.156 (18.74)	1.242 (19.60)	-0.684(-11.18)	
YDISP13	1.578 (24.98)	0.420 (7.20)	0.115 (1.63)	
YDISP4	1.413 (20.78)	0.188 (2.93)	0.251 (3.02)	
FIRED	0.282 (4.31)	-0.123 (-1.95)	0.082 (1.21)	
DOWN	0.189 (2.99)	0.031 (0.50)	0.044 (0.69)	
CUTBACK	0.232 (2.48)	-0.125 (-1.40)	0.135 (1.40)	
PROV.P.	-0.028 (-0.16)	0.008 (0.04)	-0.056 (-0.32)	
WCOLLAR	0.044 (0.67)	0.092 (1.43)	0.030 (0.44)	-.352(-12.63)
HEAD	-0.343 (-5.39)	0.519 (8.51)	-0.186 (-2.73)	.034 (0.99)
NCHIL	-0.003 (-0.19)	0.008 (0.48)	0.005 (0.31)	.018 (1.89)
OTHERW	-0.025 (-0.73)	0.006 (0.14)	-0.059 (-1.64)	-.108 (-6.09)
AD55	0.325 (2.94)	-0.599 (-5.49)	0.101 (0.99)	
CONST.	-1.160 (-5.62)	-0.230 (-1.07)	2.062 (9.23)	.235 (2.11)
LOG LIKEL.	-2190	-2380	-1983	-7312
N	4201	4201	5010	12440

Note: All the regressions include 17 region and 4 sector dummies

TABLE 3.- ESTIMATED EFFECT OF VARIABLES ON UNEMPLOYMENT DURATION USING WEIBULL MODEL. (ASYMPTOTIC STANDARD ERRORS)

	(1) Entire sample		(2) End of contract		(3) Other reasons		(4) Displ. 1981-85	
ELIG36	0.185	(0.074)	0.315	(0.091)	0.003	(0.137)	0.211	(0.063)
ELIG912	0.193	(0.103)	0.354	(0.147)	-0.019	(0.150)	0.234	(0.087)
ELIG1518	0.255	(0.065)	0.336	(0.102)	0.164	(0.093)	0.343	(0.084)
AGEDISP	0.011	(0.002)	0.015	(0.004)	0.008	(0.003)	0.023	(0.003)
SEX	-0.714	(0.064)	-0.635	(0.095)	-0.783	(0.088)	-0.422	(0.064)
TENURE	0.013	(0.004)	0.001	(0.008)	0.016	(0.005)	0.008	(0.005)
EDUC	-0.034	(0.007)	-0.024	(0.010)	-0.041	(0.009)	-0.017	(0.007)
YDISP0	1.100	(0.073)	1.208	(0.107)	0.943	(0.115)	1.396	(0.057)
YDISP13	1.357	(0.069)	1.397	(0.095)	1.286	(0.112)		
YDISP4	0.864	(0.079)	0.806	(0.102)	0.879	(0.131)	0.831	(0.064)
FIRED	0.227	(0.068)			0.198	(0.071)	0.119	(0.072)
DOWN	0.007	(0.064)					-0.028	(0.070)
CUTBACK	0.233	(0.097)			0.222	(0.093)	0.128	(0.109)
PROV. P.	0.020	(0.167)			0.005	(0.173)	0.158	(0.166)
WCOLL	-0.03	(0.065)	-0.073	(0.110)	-0.018	(0.080)	-0.110	(0.068)
HEAD	-0.58	(0.063)	-0.579	(0.089)	-0.629	(0.089)	-0.475	(0.064)
NCHIL	-0.00	(0.017)	-0.001	(0.027)	0.000	(0.022)	-0.016	(0.020)
OTHERW	-0.01	(0.036)	0.046	(0.052)	-0.081	(0.050)	0.040	(0.036)
AD55	1.015	(0.152)	1.097	(0.263)	0.954	(0.184)	0.656	(0.166)
CONST.	2.423	(0.208)	1.924	(0.311)	2.932	(0.289)	1.771	(0.231)
SCALE	1.029	(0.018)	1.045	(0.027)	1.000	(0.024)	0.824	(0.019)
LOG LIKEL.	-4961		-2358		-2574		-2998	
N	4201		1982		2219		3062	

Note: All the regressions include 17 region and 4 sector dummies

TABLE 4.- ESTIMATES OF EFFECT OF VARIABLES ON UNEMPLOYMENT DURATION USING A COX PROPORTIONAL HAZARD MODEL. (T STATISTICS)

	(1) Time fixed covariates		(2) Time varying covariates		(3) Time varying covariates	
ELIG36	-.176	(-2.43)				
ELIG912	-.179	(-1.78)				
ELIG1518	-.243	(-3.82)				
TRUI			-.030	(-6.13)		
BENEF					-.670	(-8.55)
AGEDISP	-.011	(-4.34)	-.011	(-4.34)	-.010	(-4.24)
SEX	.643	(10.40)	.630	(10.22)	.635	(10.29)
TENURE	-.011	(-2.56)	-.007	(-1.78)	-.007	(-1.82)
EDUC	.028	(4.08)	.028	(4.12)	.028	(4.03)
YDISP0	-.747	(-9.57)	-.613	(-7.81)	-.658	(-8.82)
YDISP13	-1.13	(-14.99)	-1.052	(-13.74)	-1.090	(-14.44)
YDISP4	-.757	(-9.38)	-.718	(-8.87)	-.741	(-9.16)
FIRED	-.217	(-3.27)	-.181	(-2.73)	-.176	(-2.67)
DOWN	-.018	(-0.29)	.007	(0.12)	.017	(0.28)
CUTBACK	-.204	(-2.17)	-.175	(-1.87)	-.167	(-1.79)
PROV.P.	-.034	(-0.21)	-.019	(-0.12)	-.039	(-0.24)
WCOLL	.035	(0.56)	.031	(0.50)	.036	(0.57)
HEAD	.527	(8.62)	.526	(8.63)	.542	(8.87)
NCHIL	.005	(0.33)	.008	(0.48)	.007	(0.43)
OTHERW	.010	(0.29)	.009	(0.26)	.008	(0.24)
AD55	-.955	(-6.50)	-.943	(-6.43)	-.964	(-6.57)
LOG LIKEL.	-16187		-16176		-16158	
N	4201		17371		17371	

Note: All the regressions include 17 region and 4 sector dummies.

TABLE 5.- ESTIMATED DEPENDENCE PARAMETERS (S.D.) BY PERIOD OF DISPLACEMENT AND MODEL

	Whole sample	1981-1985	<1979-1980
Empirical hazard (Kaplan-Meier)	.65 (.005)	.70 (.010)	.60 (.007)
Non-parametric (Cox model)	.79 (.007)	.97 (.012)	.71 (.014)
Parametric (Weibull model)	.972 (.019)	1.212 (.020)	.81 (.034)
N	4517	3326	1191

Note: See text for the way these values have been calculated.

TABLE 6.- ESTIMATES OF THE EFFECT OF UI ELIGIBILITY AND DURATION OF JOBLESSNESS ON WAGES AND TYPE OF JOB OBTAINED BY RE-EMPLOYED WORKERS

	Ordinary Least Squares regressions Dependent variable: log monthly net earnings			Probit model If regular full-time job =1
	(1)	(2)	(3)	(4)
CONST	9.989 (140.1)	9.933 (140.6)	9.957 (143.1)	-0.929 (-3.03)
SEX	0.130 (6.34)	0.073 (3.42)	0.123 (6.12)	0.053 (0.61)
EXPER	0.007 (2.55)	0.007 (2.84)	0.007 (2.49)	0.009 (0.79)
EXPER2	-0.0001 (-2.18)	-0.0001 (-1.81)	-0.0001 (-2.10)	-0.0001 (-0.85)
TENURE	-0.0002 (-0.13)	0.001 (0.91)	-0.0001 (-0.08)	0.0001 (0.02)
EDUC	0.029 (10.64)	0.028 (10.39)	0.028 (10.61)	0.027 (2.27)
YDISP0	0.066 (3.03)		0.036 (1.69)	0.811 (8.27)
YDISP13	-0.007 (-0.32)		-0.028 (-1.27)	0.550 (5.71)
YDISP4	0.026 (1.05)		0.006 (0.26)	0.408 (3.87)
FIRED	-0.039 (-1.81)	-0.021 (-1.00)	-0.047 (-2.23)	0.210 (2.27)
DOWN	-0.004 (-0.23)	-0.007 (-0.39)	-0.018 (-0.97)	0.328 (3.75)
CUTBACK	-0.020 (-0.64)	0.003 (0.09)	-0.029 (-0.97)	0.284 (2.06)
PROV.P.	-0.028 (-0.54)	-0.030 (-0.57)	-0.045 (-0.89)	0.381 (1.70)
WCOLL	0.131 (7.13)	0.118 (6.49)	0.117 (6.53)	0.136 (1.45)
HEAD	0.079 (3.79)	0.023 (1.05)	0.054 (2.63)	0.442 (4.97)
NCHIL	0.009 (1.47)	0.008 (1.35)	0.010 (1.72)	-0.025 (-0.92)
OTHERW	-0.003 (-0.33)	-0.005 (-0.49)	-0.010 (-0.94)	0.131 (2.69)
SAMESECT	0.008 (0.47)	0.005 (0.30)	0.009 (0.57)	-0.035 (-0.49)
SAMEOCCU	0.093 (5.46)	0.079 (4.67)	0.083 (4.96)	0.148 (2.05)
ELIGUI	0.036 (1.97)	0.057 (3.13)	0.022 (1.22)	0.311 (4.00)
REGULF			0.158 (10.18)	
DURATION		-0.034 (-6.65)		-0.177 (-7.75)
LAMBDA		0.127 (6.25)		
ADJ. R-SQ.	.25	.28	.29	
LOG LIKEL.				-1138
N	2001	2001	2001	2001

Notes: All the regressions include 17 region and 4 sector dummies.
In regression (4) sector and occupation are those of the previous job.
Unemployment duration is considered in years.

Figure 1 .–Survival function by UI elig.
(Kaplan–Meier estimates)

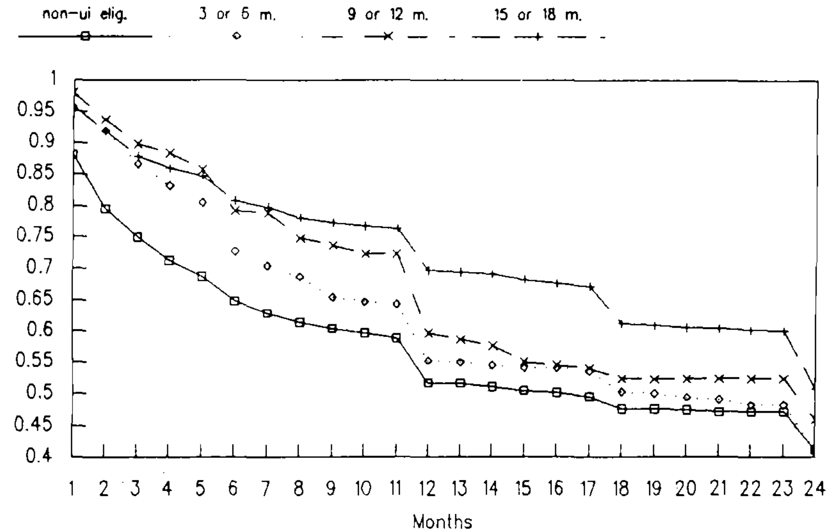

Figure 2.– Time until exhaustion hazard
Kaplan–Meier estimates
(Three month intervals)

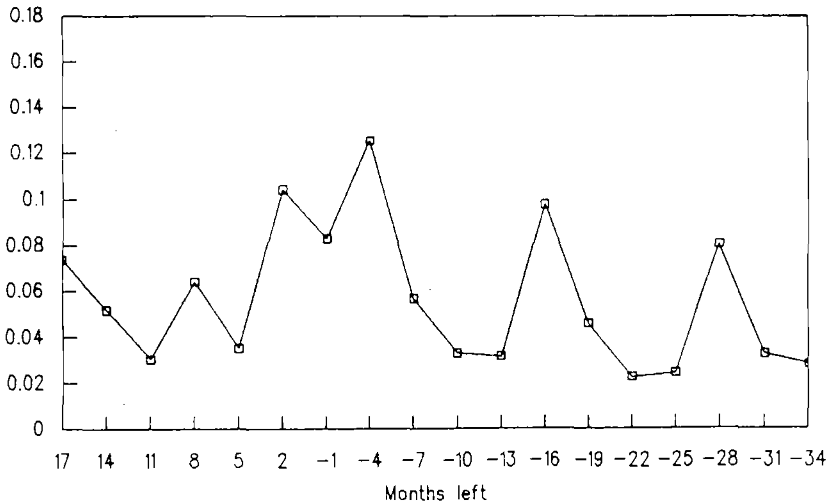

Figure 3.- HAZARD RATE
KAPLAN-MEIER ESTIMATES