MEASURING THE EFFECT OF SPELL RECURRENCE ON POVERTY DYNAMICS(*)

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ABSTRACT

The analysis of poverty dynamics yields important insights about the expected effectiveness of alternative social policies on poverty reduction. This paper analyses the effect of spell recurrence on poverty dynamics taking into account multiple poverty and non-poverty spells by spell order. Using data for Spain during a seven year period, we obtain that the poverty exit and re-entry rates vary not only with personal or household characteristics but also with spell accumulation and with the duration of past spells. Results indicate that the effect of duration dependence is significant and turns out to be different by spell order.

Keywords: poverty dynamics, hazard models, multiple states, multiple spells, unobserved heterogeneity.

JEL Classification: D1, D31, I32.

1. INTRODUCTION

The literature centred on the analysis of the lowest part of the income distribution has produced a large amount of work on the dynamics of poverty in recent years. A first interesting result of this research is the proposal of a new dimension in the measurement of poverty which refers to the time individuals spend below the poverty line. Certainly, in the analysis of poverty it is of great interest to be able to characterise the complete low income pattern of individuals along time.

The advantage of distinguishing the characteristics of individuals that suffer from *persistent* poverty in contrast with those that experience poverty for a relatively short time, is that different policies may be designed in fighting against each of these situations. Fighting against long-term or *persistent* poverty will imply designing educational and health policies for poor children and offering stable complementary monetary transfers for poor adults. In contrast, *transitory* poverty could benefit from short-term labour market policies promoting employment stability and some short-period money transfers working as income substitutes. For this second group, we believe that it certainly becomes important to investigate the relevance of poverty spell recurrence and, most importantly, to measure to what extent the probability of leaving a poverty spell depends on having had a previous experience in poverty with a varied length. The relevance of distinguishing *recurrence* within the transitory poor becomes particularly evident after one of the key findings of the OECD (2001) report: "…the typical year spent in poverty is lived by persons who experience multiple years of poverty and whose long-term incomes are below the poverty threshold on average, even though their yearly income may periodically exceed the poverty threshold" (Chapter 2, 1st page).

The literature on poverty dynamics has largely focussed on the analysis of spells and the estimation of entry and re-entry hazards after the seminal works of Allison (1982) and Bane and Ellwood (1986) which have recently been fostered by Stevens (1999), Devicienti (2001) or Biewen (2006). These papers study the extent and composition of chronic poverty in a variety of countries using a hazard rate approach that accounts for multiple spells of poverty and incorporates spell duration, individual and household characteristics and unobserved heterogeneity. In general, previous hazard rate approaches assume that the consideration of individual unobserved heterogeneity captures the correlation across individual spells and thus identifies various types of individuals in the sample through a joint distribution of individual specific effects with respect to spells of poverty and non-poverty. This assumption imposes the estimation of a single exit and re-entry hazard rate for each individual independent of the number of poverty spells previously experienced. In contrast, our approach stems from believing that the complete individual poverty history may play a relevant role, in itself, in determining the likelihood of experiencing a new poverty or non-poverty spell. Therefore, predicted exit and re-entry hazards should incorporate the information on both the duration and the accumulation of spells. In fact, it could be the case that the relative importance of unobserved characteristics in determining poverty exit and re-entry hazard rates may, in part, hide a genuine spell accumulation effect that can be distinguished if we allow poverty exit and re-entry hazards to vary as spells accumulate.

Even if this particular issue is virtually unexplored in the literature on the dynamics of poverty and social exclusion, the type of analysis proposed has been commonly utilized since long ago in a variety of other subjects. For example, demography researchers use these methods for the analysis of the life cycle fertility in order to know if the age of the marriage, the occurrence of births inside or outside the marriage, the age of first birth and the durations of previous birth intervals significantly affect the timing of subsequent births over the life cycle, see Heckman *et al.* (1985) and Heckman and Walker (1990). In marketing, it is often used to analyse the purchase timing and brand switching decisions of households for a frequently purchased product in order to check if the price, feature advertisements, special displays and household specific characteristics affect the probability of buying products in the future - see Jain and Vilcassim (1991) and Vilcassim and Jain (1991). Furthermore, within the literature on labour economics there is now already an important number of papers devoted to the analysis of recurrent unemployment and its effects on the individual's probability of leaving unemployment in a forthcoming spell i.e. the relevance of unemployment history on a current unemployment spell. These models were introduced by Lancaster (1979, 1990) and are popular because they easily incorporate censored spells and variables that change over time while they also allow one to examine how the probability of leaving poverty changes with spell duration and when spells accumulate - see Heckman and Borjas (1980), Trivedi and Alexander (1989), Bonnal *et al.* (1997), Omori (1997), Roed *et al.* (1999) or Arranz and Muro (2001, 2004).

Following these approaches we tackle the complete analysis of the influence of poverty history on exit and re-entry hazards in Spain by estimating hazard rates that consider all the information available on the individual poverty situation using the European Community Household Panel dataset (hereinafter, ECHP) for the period 1994-2000. Our methodology is in line with event history analysis given that we estimate a mixed proportional hazard model with multiple states and multiple spells that allows for lagged duration dependence explanatory variables[1].

[1] Multiple spells may occur because there are multiple observations of the same kind while multiple states occur because individuals may experience poverty or non-poverty at any interview moment.

The case of Spain is of particular interest for the measurement of poverty recurrence given that this country registers a relatively large percentage of individuals who suffer from transitory poverty (see OECD, 2001) in comparison with other developed countries. In particular, if one focuses on the working-age population, Spain stands out as the European country with one of the highest levels of short-term poverty (poor at least once in three years) with figures that are even slightly over those observed in the US and Canada.

Thus, our paper contributes to the literature on poverty dynamics in several ways: First, we aim to contribute to the debate on the determinants of the probability of leaving poverty by trying to disaggregate the distinct effects of unobserved heterogeneity and duration dependence. Secondly, our approach allows for different poverty exit, entry and re-entry hazards when spells accumulate, challenging previous studies based on poverty persistence that estimate one exit and one re-entry hazard rate independent of the number and duration of individual poverty experiences. Thirdly, since omitting the left-truncated spells would lead to serious selection bias (see Iceland 1997) our work takes the option of running a model with and without left-censored data and comparing their results. This strategy allows us discuss the substantive differences in results when analysing exit hazard rates for individuals with left-censored first-spells (for whom the poverty spell is in progress) compared to individuals who are new entrants to poverty. Finally, the use of time-varying covariates within each spell in the estimation of the exit and re-entry hazards, allows us to aim for the best possible fit of the individual's complete poverty experience.

The paper is organized as follows. Section 2 presents the most important previous results in the literature on the analysis of poverty dynamics in general and poverty outflow in particular. In Section 3, we describe the longitudinal data set used, detailing the definition of the variables and undertaking a thorough descriptive analysis of the observed poverty and non-poverty spells in the Spanish dataset. Section 4 presents the econometric model while Section 5 discusses the main estimation results. Finally, we present our main findings in the conclusions.

2. SOME PREVIOUS APPROACHES IN THE LITERATURE

The analysis of the dynamics of poverty was initiated in the United States during the eighties, mainly as a result of the availability of a mature and reliable longitudinal data survey: the Panel Survey of Income Dynamics (PSID), ongoing since 1968[2]. In the European context it is only in the beginning of the nineties that

[2] One of the most relevant papers in this literature published in that period are Allison (1982) and Bane and Ellwood (1986). Other interesting papers are Hill (1981), Plotnick (1983), Duncan (1984) or Sawhill (1988).

Duncan *et al.* (1993) try to compare, for the first time, the duration of poverty in a group of countries using a variety of data sources[3]. Fortunately for the development of this literature, in 1994 the European Statistical Office decided to obtain accurate and comparable longitudinal data information for most countries in the European Union initiating the ECHP Survey which, after some years, has become a basic tool for the analysis of social cohesion dynamics in the European Union. The exploitation of this dataset, together with some nationally based panels available for some particular countries, has allowed a large list of researchers to present plausible answers to important issues related to the duration and persistence of poverty in Europe[4].

The development of new statistical techniques in the estimation of transition probabilities that take state dependence into account, as Aassve *et al.* (2005) note in their literature revision, has produced an important number of ways to estimate transition risks. In the first place some papers have used *components of variance* models to capture the dynamics of income using a complex error structure[5]. These models are able to predict the fraction of the population likely to be in poverty for different lengths of time. This methodology has the advantage of including all individual income information in time while avoiding the ex-ante definition of poverty using a binary indicator. Its main disadvantage, however, is that it assumes that the dynamics of the income process is identical for all individuals in the sample, whatever their income level. Clearly, this does not seem to match reality and, in fact, Stevens (1999) and Devicienti (2001) conclude that, in comparison with duration models, components of variance models perform worse in fitting observed patterns of poverty in the US and the UK respectively[6].

Some other recent approximations to the estimation of outflow hazard rates propose the estimation of *Markovian transition models* (first order Markov) taking simultaneously into account that individuals are not randomly distributed either within the poor at first interview (initial conditions problem) or within the effectively observed at second interview (attrition problem) - see Cappellari and Jenkins (2004). In some sense, this type of approach focuses on sample heterogeneity and avoids assigning relevance to spell duration or persistence in

[3] This is first the paper on poverty dynamics in Europe that we know of, authors compare the duration of poverty in Germany, Sweden, The Netherlands and Luxembourg and in the Lorena region (France).

[4] Examples of these are Jarvis and Jenkins (1996), Cantó (1998), Antolín *et al.* (1999), Jenkins and Rigg (2001), Devicienti (2001), Aassve *et al.* (2005), Biewen (2006) or Cantó *et al.* (2007). These papers use information on the ECHP or the British Household Panel Survey (BHPS) for the U.K., the German Socioeconomic Panel (GSOEP) for Germany or the *Encuesta Continua de Presupuestos Familiares* (ECPF) for Spain.

[5] Examples of this approach are Lillard and Willis (1978), Stevens (1999) and Devicienti (2001).

[6] However, this result does not necessarily imply that these models perform worse in the estimation of the effects of explanatory variables on transition risk.

the determination of the outflow rate, which is in line with the arguments raised by Blumen *et al*. (1955) in order to explain why empirical transition matrices underestimate the main diagonal of the matrix thus biasing downwards any estimation of persistence.

Some more recent approaches to the analysis of poverty transitions choose to use binary dependent variable dynamic random effects models where the poverty situation at moment *t* depends on the fact that you were poor at *t-1*, a list of covariates and some unobserved individual effect. In this case state dependence is summarized by the coefficient estimated for lagged poverty. In this way Wooldridge (2005) models the distribution of the unobserved effect conditional on the initial value of poverty (initial conditions) and a group of exogenous variables[7].

A different view on the matter was offered by Shorrocks (1976) who attributed the phenomenon to a violation of the first-order Markov assumption which implies that the extension of the Markov process, in as much as the longitudinal information allows us to, is the way to proceed in the accurate estimation of the outflow rate. In this second line of argument, a long-standing approach to model poverty transitions has been the use of *duration models* in a hazard rate framework. Since the main contributions to this literature due to Kalbfleisch and Prentice (1980), Allison (1982), Duncan (1984) or Bane and Ellwood (1986), a large list of papers have developed single-spell duration models based on Markov chains that allow for the estimation of the transition probability taking into account all the relevant longitudinal information offered by panel datasets. A very relevant contribution to the easy estimation of hazard rates as an n-Markov chain by using a simple logit or probit model was proposed by Jenkins (1995)[8].

[7] Examples of this approach to the analysis of the dynamics of poverty appear in Biewen (2004), Hansen *et al*. (2006) and Poggi (2007). In the same line a very recent paper by Devicienti and Poggi (2007) analyses the dynamic relationship between two supposedly intimately related concepts of deprivation such as income poverty and social exclusion. In their paper these authors propose the use of a dynamic random-effects bivariate probit following the recent contribution by Stewart (2007) who generalizes Wooldridge (2005) proposal to the bivariate case. Interestingly, their consideration of higher order dynamics (second-order Markov) makes the effect of initial conditions lose strength in determining poverty in successive periods, which seems to underline the usefulness of the information on individual poverty previous to the t-1 moment.

[8] Empirical applications of this methodology selecting a poverty inflow sample of spells for Spanish data referred to the 1985-1995 period appear in Cantó (2002, 2003). Finnie and Sweetman (2003) use this methodology on administrative data to estimate exit and re-entry transitions for Canadian individuals not including unobserved heterogeneity but stratifying the sample by family status (single, couple with/without children and lone parents) and including events as a relevant reason for transitions occurrence. Cantó *et al*. (2007) use a similar strategy to estimate exit transitions for Spanish households distinguishing between households with and without children and including a large list of events as explanatory

However, a list of recent papers have highlighted the limitations of the use of single spells approaches in the fitting the observed pattern of poverty persistence and have proposed a new methodology that allows for the consideration of multiple poverty and non-poverty spells simultaneously. This methodology was first proposed by Stevens (1999) and then used by Devicienti (2001), Hansen and Wahlberg (2004) and Biewen (2006)[9]. These papers do not only consider the estimation of the probability of leaving a poverty spell but are able to estimate the hazard rate for multiple spells while controlling for unobserved heterogeneity, an important source of bias for the estimated coefficients for duration[10]. However, this approach has an important disadvantage

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variables of the transition equation. A single-spell approach is also used by Fouarge and Layte (2005) who estimate the exit probability including unobserved heterogeneity in an individual specific error term with Gaussian distribution. Much more simple is the methodology in Valletta (2006) who uses a pool of transitions to estimate the probability of leaving and entering poverty for a sample of individuals living in working-age households including labour market and demographic events as explanatory variables but without considering the effect of duration, unobserved heterogeneity or past poverty occurrences.

[9] Devicienti (2001) adds the consideration of the initial conditions problem by adjusting the contribution to the exit hazard rate of those individuals who are already below the poverty line at first household interview. In general, in most available datasets the lack of information on the previous household socio-economic situation implies that most papers must either avoid to use left-censored observations or avoid the inclusion of duration as a explanatory variable. However, Devicienti (2001) reports that fitting his type of model is computationally too demanding for a relatively short panel dataset of eight years like the ECHP. Hansen and Wahlberg (2004) use this same approach for data from Swedish administrative records which imply no attrition and a most accurate income measurement. Biewen (2006) shows how correct standard errors (not affected by the correlation between individuals from the same household) can be computed for the hazard rate model by taking into account clustering of observations at the household level. These standard errors are relevant in order to construct confidence intervals to compare the hazard rate with the components-of-variance approach. Other recent papers as Fertig and Tamm (2007) have followed a similar methodology to that of Devicienti (2001) including Biewen's correct standard errors. These authors analyse the duration of child poverty in Germany trying to contribute to the literature by reducing the effect of left-censored spells on results. In order to do so they select a sample of newly born children. However, the problem these authors face is that left-censoring in poverty is a household matter and not an individual specific problem so their strategy, in our point of view, is hardly able to improve our research knowledge of how results would change if we could use a genuine non-left-censored spells sample.

[10] These approaches recognise explicitly the existence of two processes that can generate persistence: unobserved heterogeneity and true state dependence. In the first process, individuals could be heterogeneous with respect to characteristics that reduce their probability of leaving poverty (think for example that some unobserved persistent individual characteristic reduces the probability that the individual experiences a positive event between t-1 and t, i.e. being lazy, being an alcoholic or drug consumer, etc.). In the second process, experiencing poverty in a specific time period, in itself, increases the probability of undergoing poverty in subsequent periods.

in order to study recurrence given that it only allows for the estimation of a single exit and re-entry hazard rate independent of the number of poverty experiences that the individual may have accumulated in time. This means that, virtually, the recurrence of poverty spells is assumed not to affect the estimated probability of transition.

The analysis presented here tries to improve our knowledge on to what extent the accumulation of poverty spells and the individual poverty history (lagged poverty and non-poverty durations) has a relevant role in determining future poverty risks within a duration model framework. Therefore, we aim to relax the assumption on the independence of the recurrent poverty experiences while controlling for unobserved heterogeneity and allowing for the inclusion of time-varying covariates[11].

We examine the persistence of poverty during a seven year period using data for Spain from 1994 to 2000 estimating by various hazard exit and re-entry rates jointly by spell order while including lagged spell durations as explanatory variables. In order to check the robustness of our main results, we replicate all relevant calculations for a sample of *new-entrants* to poverty thus restricting the analysis to individuals who are observed to become poor within the observation window[12].

3. THE ECHP DATA SET

3.1. A short description of the ECHP data set

The dataset we use is constructed using the information for Spain from the European Community Household Panel (hereinafter, ECHP) for the period

[11] An alternative methodology that tries to account for the way in which past poverty can have an effect on future poverty and thus aims to relax the assumption on the independence of the recurrent poverty experiences has been developed by Biewen (2004). This paper is related to the attempts to model poverty transitions in a more structural way initially proposed by Burgess and Propper (1998) and recently simplified by Aassve *et al.* (2005). These approaches propose a comprehensive model of poverty dynamics by modelling demographic and employment processes that underlie the poverty outcome. The main problem these models face is the need to simplify the large number of simultaneous risks for each household member given the limited number of equations and parameters one can identify. Further, a series of assumptions are to be made in terms of the fertility and employment process and this becomes most difficult in socioeconomic contexts where these processes and their income effects are largely unexplored.

[12] Checking the robustness of results in this way is referred to as particularly convenient in Iceland (1997). The author refers to the fact that, even if this does not solve the problem of left-truncation it sheds light on the issues of interest. Moffitt and Rendall (1993) also run a model with and without left-censored data in their study on lifetime distribution of female headship in the U.S. and discuss the substantive differences in the results.

1994-2001. The dataset used was designed by *Eurostat* in order to obtain country-comparable statistics on many demographic and socio-economic aspects of the European population related to labour market issues, income, living standards, education, employment and not employment-related satisfaction, health and migration, among others. The ECHP collects information about individual age, sex, education, income and labour market status together with an important amount of family composition variables.

The ECHP is annually based and has a longitudinal structure that allows following individuals during eight years, unless they voluntarily leave the survey earlier. Most of the variables included in the survey describe the individual's and household's situation at the moment of the interview (1994 along to 2001) or refer to information on the current month. However, the information on annual individual income from a variety of sources refers to that obtained by the individual during the year previous to the interview. Thus, in the construction of the relevant income variable to determine the individual situation of poverty in a given year we believe that it is important to make demographic and income information contemporaneous, especially if we wish to include time-varying covariates or events as explanatory variables of poverty flows. This means that we have to drop the information on incomes for 1993 (declared in 1994) and on characteristics for 2001 which implies that we finally use the information from seven complete waves instead of eight. The advantage of this procedure is that the definition of poor is based on contemporaneous information on incomes and needs which becomes crucial when we aim to correctly measure the effect of demographic or socioeconomic events on the individual's probability of experiencing a transition.

3.2. Sample selection and descriptive analysis

The period of observation in our study is from 1994 to 2000 and our sample includes individuals with a complete interview in the survey and whose household reports previous year income information[13]. As noted earlier, our sample reduces slightly when we match demographic and socioeconomic characteristics with yearly income so that we have contemporaneous information on both. Thus, our sample includes 19,044 individuals of which 15,042 (79 percent) are adults and 4,002 (21 percent) are children below 16 years of age (see table A2)[14].

[13] We eliminate between a 1 and 2 percent of individuals due to the lack of complete interview – see table A1.

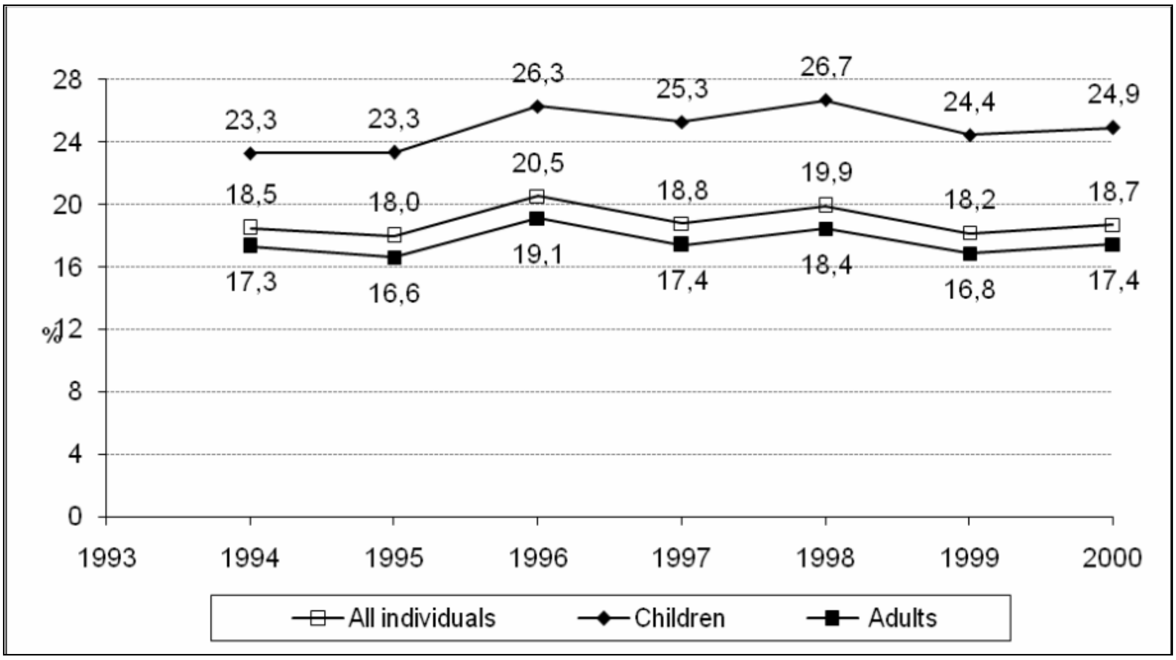
[14] It is important to note here that given that individuals change households by creating a new one between two consecutive interviews (emancipation, divorce or separation), we must make some adjustments to household income so that individuals that change household

For the purposes of our research, we use the standard definition of poverty, thus an individual is poor if total household income of the household he or she lives in is less than 60 per cent of the contemporary household median income. The results on static poverty for this sample are reported in figure 1 showing that individual adult poverty rates in Spain were quite stable during the period under study.

Figure 1

RELATIVE INDIVIDUAL POVERTY INCIDENCE SPAIN

ECHP 1994-2001, using contemporaneous income



Note: These results are obtained using the ECHP using contemporary income and characteristics and a modified OECD scale. Calculations are made for individuals weighted by their population weight each particular year.

Regarding the particular characteristics of poverty dynamics in Spain, and limiting our sample to those individuals that are observed at all interviews, Table 1 provides a variety of measures of chronic, transitory and recurrent poverty in order to contrast our results with those in Valletta (2006) or OECD (2001). In particular, the OECD (2001) chapter on poverty dynamics calculates that the always poor in Spain were 8.3 per cent of the sample in the first three years of data available from the ECHP, one of the highest percentages in the European Union context. Only Greece, Italy and Portugal register a larger percentage while Denmark registers the lowest percentage with a 2.6 percent. Thus, the characteristics of poverty duration in Spain are reported there to be significantly

effectively contribute to the income of the household where they were when household characteristics were observed. Clearly, with attrition, this strategy implies that we lose information on some individuals and our sample reduces. Indeed, our final sample reduces between a 9 to 14 percent with respect to a non-contemporaneous sample depending on the year considered – see table A2.

different for individuals living within working-age households compared to the rest. Our results, using the complete ECHP data period, confirm this result.

Table 1 reports that the mean annual poverty rate in Spain is significantly higher for households whose head is below 64 years of age than otherwise. In terms of dynamics, results show also that working-age households in Spain suffer a similar level of chronic poverty compared to the rest but a significantly higher level of transitory poverty[15]. In particular, while having one poverty spell takes place with a similar probability in both population groups, having two, three or more spells happens much more often to those individuals in working-age households than to the rest. Thus, recurrent poverty in Spain seems to be linked to working-age households whose main members are active in the labour market.

Table 1

POVERTY DYNAMICS IN SPAIN: SOME DESCRIPTIVE RESULTS.

ECHP 1994-2000. Balanced Panel. (weighted for attrition)

	Working-age households			Non-working age households		
		95% confidence interval		value	95% confidence interval	
	value	min	max	value	min	max
Static poverty:						
Annual poverty rate (mean of the period)	16.9	—	—	13.8	—	—
Chronic versus transitory poverty:						
Never poor	51.8	50.3	53.3	60.5	57.6	63.6
Always poor	3.0	2.6	3.4	3.3	2.6	4.1
Poor at least once	45.2	43.7	46.6	36.1	33.4	39.2
Poor 5 out of 7 years	11.2	10.4	12.1	8.5	7.2	9.8
Permanent-income poverty*	15.4	—	—	11.6	—	—

(Sigue)

[15] Note that this result is different depending on the indicator used. Those indicators which are not affected by transitory poverty, such as the "percentage of always poor", show a similar level of chronic poverty in both groups. If one uses the "percentage of poor 5 waves out of 7" or the "poor at least once" indicators, which are contaminated by transitory poverty, one finds a higher level of poverty in working-age households due to the contribution of short-term or transitory poverty. A similar effect is captured by the "permanent-income poor" indicator which sums all households incomes along the period and calculates the percentage of households whose total income in the seven-year period is below the 60% median income of the total sample during the whole period.

(Continuación)

	Working-age households			Non-working age households		
		95% confidence interval		value	95% confidence interval	
	value	min	max	value	min	max
Recurrent versus transitory poverty:						
One poverty spell	60.3	58.4	62.3	72.9	69.2	76.3
Two poverty spells	33.0	31.0	34.9	23.3	20.1	26.8
Three or more poverty spells	6.6	5.8	7.6	3.8	2.3	5.5
One spell, left-censored	14.5	13.1	16.0	16.6	13.3	20.3
Two spells, first left-censored	10.9	9.6	12.3	4.4	3.0	5.9
Three or more spells, first left-censored	5.9	5.1	6.9	2.3	1.3	3.5
First spell not left-censored	68.6	66.7	70.4	76.6	72.8	79.9
Total number of individuals	**7,992**			**1,830**		

Note: These results are obtained for individuals present in all seven waves of the panel. The definition of working-age household is that in which the head of household is 15 to 64 years of age as in OECD (2001), measured at first interview in the panel. In all dynamic results we use last interview longitudinal weights while in static poverty calculations these are multiplied by cross-sectional weights. Confidence intervals are obtained using bootstraps of 1000 repetitions and are bias corrected. Permanent-income poverty is calculated as the percentage of the sample for who average (equivalent) income over the seven years falls below the poverty line over this period.
(*) This indicator is calculated summing up all household incomes along the period and calculating the percentage of households whose total income in the seven-year period is below the 60% median income of the total sample during the whole period.

In order to construct our final sample for dynamic analysis, we select individuals present in the 1994 sample and in all consecutive interviews until either the survey ends or they suffer from attrition (they leave prior to the end of the survey)[16]. Thus individuals joining the survey after 1994 are not included in our sample. After this selection we have an unbalanced panel of individuals that we will use for a preliminary descriptive analysis of poverty incidence, poverty persistence and poverty transition rates using conditional probabilities calculations – see table 2[17]. Results from table 2 indicate that, as in the cross-sectional sample,

[16] Note that if the individual leaves prior to the end of the survey the ongoing spell will be treated as right-censored.

[17] This first unbalanced sample includes 19,044 individuals (a total of 99,507 person-year observations) and, as it would be expected due to attrition, the sample size falls along the

the pattern of static poverty in our sample is quite flat: the incidence of poverty is pretty stable between 19-21 per cent of the households in the sample.

Some preliminary results on transitions are also reported in table 2, these are obtained by calculating the conditional probability that the individual is in a certain situation at moment *t* given his/her situation in the previous year, *t-1*. The first row of these conditional probabilities indicates the individual probability of remaining in poverty in two consecutive interviews i.e. two-year poverty persistence. Some 52 per cent of individuals who were poor in 1994 continue to be in poverty in 1995. In subsequent waves, this conditional probability fluctuates only slightly from 49 per cent in 1996 to just 55 per cent in 2000. Thus, for the entire period, these results show that there is a considerable persistence in poverty, a mean of almost 53 per cent of individuals who where poor at time *t-1* are also poor one year later. As expected, transition probabilities from poor to non-poor are higher than from non-poor to poor but entry and exit from poverty do not seem to have a clear pattern along the period[18]. Interestingly, the probability of attrition does not appear to be determined by the individual poverty situation. Indeed, even if in 1995 and 1997 the probability of attrition was slightly higher for the group of the poor (approx. 12.5 versus 10.5 per cent), this difference is not observable at any other moment.

Table 2

POVERTY INCIDENCE AND SHORT-TERM PERSISTENCE.

ECHP 1994-2000. Unbalanced Panel. (weighted for attrition**)**

	1994	1995	1996	1997	1998	1999	2000	Mean 1994-2000
Incidence								
Headcount index (% poor)	19.88	19.17	17.98	20.99	18.90	19.60	18.76	19.32
Conditional probabilities								
Poverty short-term persistence Prob ($y_t=1/y_{t-1}=1$)		52.31	49.39	54.48	51.89	54.3	54.95	52.89
Poverty entry occurs Prob ($y_t=1/y_{t-1}=0$)		7.87	7.76	10.66	7.23	8.86	8.54	8.49

(Sigue)

period. Table A3 in the appendix contains a similar analysis for the total sample from the panel without any selection. As one can easily check it appears that results are remarkably similar.

[18] Our results in this table match those obtained for the period 1994-1996 by the OECD (2001) report where using an OECD equivalent scale and a 60% of the median income poverty line the headcount index is 19.2, the entry rate is 8.3 and the exit rate is 39.7.

(Continuación)

	1994	1995	1996	1997	1998	1999	2000	Mean 1994-2000
Poverty exit occurs Prob (y_t=0/y_{t-1}=1)		35.17	39.22	32.91	38.23	36.27	38.62	36.74
Persistence out of poverty Prob (y_t=0/y_{t-1}=0)		81.58	80.45	79.08	82.98	80.48	84.5	81.51
Atrittion								
Prob (y_t=mis/y_{t-1}=0)		10.54	11.79	10.27	9.79	10.66	6.96	10.00
Prob (y_t=mis/y_{t-1}=1)		12.51	11.39	12.61	9.88	9.42	6.44	10.38
Sample Size (weighted)	19,044	16,959	15,016	13,853	12,871	11,979	11,588	19,044

Note: These results are obtained using the ECHP contemporary income and characteristics information and using a modified OECD equivalence scale. Calculations of headcound index are made for individuals weighted by their population weight each particular year. The sample here is that of all individuals present in 1994 and in consecutive interviews in the ECHP panel until the survey ends or they suffer from attrition. Note that y_t=1 if the individual is poor in time t and 0 if the individual is non-poor, "*mis*"means that attrition occurred.

Our econometric estimations of transitions rates will require individual consecutive observations (to allow for current and lagged poverty and re-entry spells) and a common date of entry to the panel - see Heckman (1981). Therefore, our final sample of individuals is an unbalanced panel of 3,664 individuals who were poor in 1994 and have consecutive observations in the panel[19]. Given that the incidence, short-term persistence and recurrence remained quite constant across the period we believe that this sample selection is particularly adequate in this context (see Table 3). This choice allows us to use the longest observation window possible and provides us with a stock of 3,664 individuals in poverty whose first poverty spell is, by definition, in progress at the start of the sample period[20]. This sample's first spells are all left-censored poverty spells for which duration is unknown because the start date is missing. By definition, second, third or fourth spells are non-left-censored. However, some of the poverty and non-poverty spells will be right-censored because they

[19] This means that we require an individual not to have missing observations in between interviews to be included in this dataset. This sample amounts to a total of 19,219 person-year observations.

[20] Note here that we cannot distinguish if the spell began precisely in 1994 or was in progress before the start of the sampling period. This sample doest not include individuals who started the ECHP and may temporarily exit the ECHP presenting missing values across several years (because we do not know their status of poverty and non-poverty). There are 333 individuals of this type who experience 1,496 poverty spells.

will be still in progress at the end of the ECHP time window. For the latter we only know that the elapsed time of the spell was longer than the interval between the spell start date and the end of the ECHP observation window (1994-2000): censored durations. Clearly, ending spell dates are only known for those spells that are observed to finish within the observation window[21].

This sample includes left-censored spells given that erasing left-censored spells in progress at the start of the sample (even if considering unobserved heterogeneity) provokes a form of sample selection bias as Stevens (1997) and Iceland (1997) indicate[22]. Indeed, if we select only those individuals who are observed to enter poverty in 1995 and follow their future movements into and out of poverty across the period 1994-2000 –see table A4 in the Appendix–, the conditional transitions from non-poor to poor obtained from this sample, are significantly higher than those reported for the complete sample in table 2 or even the first sample in table A3 (they range from 14 per cent to 32 per cent while those for the first sample are almost always below a 10 percent)[23]. However, we will return to this issue later in order to analyse, in detail, the consequences of keeping left-censored spells in the sample on our results[24].

Results on poverty incidence and persistence using our selected sample are reported in table 3[25]. We observe that, in comparison with our first sample, the individual's probability of stepping into poverty is now significantly higher (more than a double the risk of transition than our first sample in table 2: 7.8 to 8.5 compared to 24.2 to 19.2)[26]. In contrast, individuals in this second sample show

[21] Note here that the duration of the spells (of poverty or non-poverty) of those individuals who leave prior to the end of the survey (attrition) are also considered as right censored observations.

[22] Note that including left-censored spells is also the choice in the descriptive analysis undertaken by Stevens (1999) or Jenkins and Rigg (2001). It is easy to see that if we were to eliminate them from the sample and only consider individuals who begin a new spell after 1994, thus entering poverty in 1995 for the first time, we would overstate transition probabilities given that the selected sample would have experienced at least one transition (non-poor to poor) since 1994.

[23] However, note that the estimations of models ignoring unobserved heterogeneity or omitted variables which only include spells that begin after the start of the sample, give consistent estimates of poverty transition rates for the population, Heckman and Singer (1984).

[24] It is not possible for us to model the hazard rate of an individual's first entry into poverty or first exit from poverty (initial conditions) because we do not have information on the pre-1994 income histories of those who were poor (or not) before 1994. Our only control for left-censoring is to estimate a separate baseline hazard for left-truncated spells as Callens *et al.* (2005) suggest.

[25] In the Appendix we include a similar table for three types of individuals: those below 16 years of age, those between 16 and 65 and those above 65 (See table A5.-A7.).

[26] This appears a reasonable effect of selecting individuals already touched by poverty in 1994.

a higher persistence in poverty (53 to 68 per cent depending on the year compared to 49.4 to 55 percent) and thus also a lower probability leaving poverty at any two subsequent interviews. Also they show a lower persistence out of poverty which implies that they are more likely to suffer poverty recall. In sum, this sample includes more chronic poor but also more short-term recurrently poor individuals than our first sample.

Table 3

POVERTY INCIDENCE AND SHORT-TERM PERSISTENCE: MAXIMUM OBSERVATION WINDOW. ECHP 1994-2000. (weighted for attrition)

	1994	1995	1996	1997	1998	1999	2000	Mean 1994-2000
Incidence								
Headcount index (% poor)	100	59.98	51.35	52.94	51.36	45.58	40.39	57.37
Conditional probabilities								
Poverty short-term persistence Prob (y_t=1/y_{t-1}=1)		53.87	60.42	65.29	68.06	58.03	61.08	61.13
Poverty entry occurs Prob (y_t=1/y_{t-1}=0)			24.2	29.83	24.72	22.96	19.15	20.14
Poverty exit occurs Prob (y_t=0/y_{t-1}=1)		35.94	30.64	22.93	25.36	32.99	33.03	30.15
Persistence out of poverty Prob (y_t=0/y_{t-1}=0)			63.36	62.95	67.88	65.27	73.69	55.53
Atrittion								
Prob (y_t=mis/y_{t-1}=0)			12.44	7.22	7.4	11.78	7.15	7.67
Prob (y_t=mis/y_{t-1}=1)		10.19	8.94	11.78	6.58	8.98	5.89	8.73
Sample Size (weighted)	3,664	3,291	2,964	2,685	2,523	2,270	2,154	3,664

Note: These results are obtained using the ECHP contemporary income and characteristics information and using a modified OECD equivalence scale. Calculations of headcound index are made for individuals weighted by their population weight each particular year. The sample here is that of all individuals present in 1994 and in consecutive interviews in the ECHP panel until the survey ends or they suffer from attrition. Note that y_t=1 if the individual is poor in time t and 0 if the individual is non-poor, "*mis*"means that attrition occurred.

Results on spell duration pooling the data for all the poverty and non-poverty periods without considering the order of each occurrence are reported in table 4. The last row of this Table reflects the long-term persistence of poverty in Spain in the period under analysis: 9.4 per cent of individuals who are poor in 1994 continue to be below the poverty line seven years later, this percentage rises to 12 per cent for those who are able to step out of poverty for a year but come back to it suffering a second poverty spell of, at least, 6 years of length. We find that nearly 26 per cent of the individuals remain four years or more in poverty and 35.5 per cent of the individuals who exit poverty (thus enter a non-poverty period) remain one year in non-poverty, 22.5 per cent two years and 12 per cent at least six.

Table 4

FREQUENCY DISTRIBUTIONS OF ELAPSED DURATIONS, ALL SPELLS.

Sample restricted to individuals who are poor in 1994 and consecutive observation in panel ECHP 1994-2000

Elapsed duration	all poverty spells		All non-poverty spells	
	Freq	%	Freq	%
1	1,395	38.1	865	35.5
2	796	21.7	549	22.5
3	514	14.0	361	14.8
4	277	7.6	232	9.5
5	214	5.8	145	5.9
6	123	3.4	288	11.8
7	345	9.4	—	—
Total individuals	3,664	100	2,440	100
Mean (Std. Dev.)	2.69 (1.94)		2.63 (1.70)	

Tables 5 and 6 focus on the frequency distribution of elapsed poverty and non-poverty spells by order of occurrence of the particular spell. Our findings in these tables highlight the importance of considering multiple spells in the analysis of poverty dynamics: out of the 3,664 individuals who are in poverty since 1994, 27 per cent have two occurrences along the time of observation and nearly 7.5 per cent have three or more occurrences. This implies that a 35 per cent of the individuals re-enter poverty during the seven year period and, out of these, a 21 per cent actually re-enter twice or three times. In terms of duration, first-spells have a mean duration of two and a half years while the duration of second and third poverty spells is slightly shorter (1.7 and 1.3 years respectively)[27].

[27] This result could be affected by the seven year interview structure of the dataset.

Indeed, Table 6 shows that 47 per cent of first poverty spells have an elapsed duration of one year and this percentage increases up to 54 per cent if we are in a second occurrence and to 73 per cent in a third one, meaning that if one has a second or third poverty spell, these spells are likely to be particularly short, in fact they are most likely to be one year periods. A similar result is obtained for non-poverty spells. In sum, there seem to be certain groups of individuals that are particularly prone to exit and re-enter poverty experiencing a row of one to two year spells.

Table 5

NUMBER OF SPELLS OF POVERTY AND NON-POVERTY IN TOTAL SAMPLE.

Sample restricted to individuals who are poor in 1994 and consecutive observation in panel ECHP 1994-2000

Number of occurrences	Poverty		Non-poverty	
	Freq.	%	Freq.	%
1	2,388	65.17	1,670	45.58
2	1,003	27.37	678	18.5
3	268	7.31	92	2.51
4	5	0.14		
Total individuals	3,664	100	2,440	66.59

Table 6

FREQUENCY DISTRIBUTIONS OF ELAPSED DURATIONS BY ORDER OF OCCURRENCE.

Sample restricted to individuals who are poor in 1994 and consecutive observation in panel. ECHP 1994-2000

Elapsed duration	First poverty spell		First non-poverty spell		Second poverty spell		Second non-poverty spell		Third poverty spell		Third non-poverty spell	
	Freq	%	Freq	%	Freq	%	Freq	%	Freq	%	Freq	%
1	1,718	46.89	1085	44.47	691	54.15	413	53.64	201	73.63	81	88.04
2	705	19.24	461	18.89	332	26.02	225	29.22	59	21.61	11	11.96
3	389	10.62	276	11.31	146	11.44	85	11.04	13	4.76	—	—
4	205	5.59	185	7.58	72	5.64	47	6.1	—	—	—	—
5	179	4.89	145	5.94	35	2.74	—	—	—	—	—	—
6	123	3.36	288	11.8	—	—	—	—	—	—	—	—
7	345	9.42	—	—	—	—	—	—	—	—	—	—
Total individuals	3,664	100	2440	100	1276	100	770	100	273	100	92	100
Mean (Std. Dev.)	2.50 (1.97)		2.47 (1.75)		1.77 (1.04)		1.70 (0.89)		1.31 (0.56)		1.12 (0.33)	

4. ECONOMETRIC APPROACH: A MULTI-STATE MULTI-SPELL HAZARD MODEL

The main aim of this paper is the estimation of the individual probability of leaving poverty taking account the effect across multiple spells of the current poverty and non-poverty duration, the occurrence of multiple-spells, the lagged poverty and non-poverty duration and individual and household characteristics. For that purpose, we use life-table analysis and estimate a multivariate multi-state multi-spell model for the exit and re-entry hazards.

Our strategy consists in simultaneously estimating up to four hazard rates at once, mirroring the individuals' complete poverty history[28]. We first choose to examine the persistence of poverty for individuals who are poor in 1994 and have consecutive interviews in the panel, rather than examining the incidence of poverty for the entire ECHP sample. Later, we will restrict our analysis to those individuals who become poor in 1995 (inflow sample or sample of new entrants) in order to understand the relevance of left-truncation on results.

In the hazard methodology, the probability of leaving poverty h_{pi} (or re-entering poverty, h_{ri})[29] may be defined as the limit of the conditional probability of a transition taking place in a small interval dt after time t if no transition occurs until t, when that interval approaches to zero. The exit rate is modelled by means of a multiplicative separable function of three terms: the baseline exit hazard, the covariates and unobserved heterogeneity.

Formally:

$$h_{pi}(T_{pi}, X_{pi}(t), \theta_i) = \lim_{dt \to 0} \frac{\Pr(t + dt > T_{pi} \geq t \mid T_{pi} \geq t)}{dt} = \lambda_0(T_{pi}) \exp\{X_{pi}(T_{pi})'\beta\}\theta_i \quad (1)$$

In this equation, subscripts i indicate the individual and p the period in poverty. The term T_{pi} is the latent current duration of individual *i´s p*'th poverty spell, $\lambda_0(t)$ is the interval-specific baseline hazard rate, which is unknown; X_{pi} is a vector of time-invariant and time-varying covariates for individual i, β is the vector of unknown parameters to be estimated, and finally θ_i captures unobserved individual characteristics that could affect the hazard but are unobservable in the data, such as social exclusion problems together with attitudes towards claiming benefits of finding a job, motivation, inherent ability, and so on.

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[28] We have performed our estimations in Stata 9.2.

[29] We here present the econometrics for the analysis for the probability of leaving poverty in order to simplify notation, however our analysis includes the estimation of both the probability of leaving and re-entering poverty in an analogous way.

Defining the probability of surviving through any interval dt after having survived the preceding j intervals as $(1-h_{pi})$, the probability of ending a spell of poverty in the p^{th} interval is given by the hazard function[30]:

$$h_{pi} = \Pr\left[T_{pi} = t\right] = h_{pi} \prod_{s=1}^{t_p-1}(1-h_{pi}^{s}) \qquad (2)$$

where t_p represents poverty duration. However, given that there are some poverty spells that continue to proceed at the end of the sample period, right censored spells also contribute to the likelihood. Their contribution can be expressed as[31]:

$$\Pr\left[T_{pi} > t\right] = \prod_{s=1}^{tp}(1-h_{pi}^{s}) \qquad (3)$$

Given that we are interested in incorporating multiple spells of poverty and non-poverty to our analysis, our likelihood function contains several components that capture multiple exits from poverty to non-poverty and viceversa. In particular the likelihood for any observed individual i can be expressed as:

$$Li = \left\{\prod_{s=1}^{tp1}(1-h_{p1i}^{s})\right\}^{(1-d1i)} \times \left\{h_{p1i}\prod_{s=1}^{tp1-1}(1-h_{p1i}^{s})\prod_{s=1}^{tr1}(1-h_{r1i}^{s})\right\}^{d1i(1-d2i))} \times$$
$$\left\{h_{p1i}\prod_{s=1}^{tp1-1}(1-h_{p1i}^{s}) \times h_{r1i}\prod_{s=1}^{tr1-1}(1-h_{r1i}^{s})\right\}^{d1id2i} \times \left\{h_{r1i}\prod_{s=1}^{tr1-1}(1-h_{r1i}^{s})\prod_{s=1}^{tp2}(1-h_{p21i}^{s})\right\}^{d2i(1-d3i)} \times$$
$$\left\{h_{r1i}\prod_{s=1}^{tr1-1}(1-h_{r1i}^{s}) \times h_{p2i}\prod_{s=1}^{tp2-1}(1-h_{p2i}^{s})\right\}^{d2id3i} \times \left\{h_{p2i}\prod_{s=1}^{tp2-1}(1-h_{p2i}^{s})\prod_{s=1}^{tr2}(1-h_{r2i}^{s})\right\}^{(1-d4i)d3i} \times$$
$$\left\{h_{p2i}\prod_{s=1}^{tp2-1}(1-h_{p2i}^{s}) \times h_{r2i}\prod_{s=1}^{tr2-1}(1-h_{r2i}^{s})\right\}^{d4id3i} \qquad (4)$$

Where h_{p1} and h_{p2} are discrete hazard rates for leaving poverty during her first and second poverty spell and h_{r1} and h_{r2} are discrete hazard rates for re-entering poverty during the corresponding periods out of poverty in between. Furthermore, t_{p1} and t_{p2} represent poverty spell durations while t_{r1} and t_{r2} are the non-poverty spell durations once the first poverty spell ends (t_{r1} takes place

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[30] We omit t, X and θ to simplify notation.

[31] Similarly to the exit rate, the hazard rate for re-entry is given by an analogous expression (where "p" changes to "r"): $h_{ri}(T_{ri}, X_{ri}(t), \theta_i) = \lambda_0(T_{ri})\exp\{X_{ri}(T_{ri})'\beta\}\theta_i$. Thus the probability of ending a spell of non-poverty in the r^h interval is given by: $h_{ri} = \Pr[T_{ri} = t] = h_{ri}\prod_{s=1}^{tr-1}(1-h_{ri}^{s})$ and the contribution to the likelihood of non-poverty spells that continue to proceed at the end of the sample is $\Pr[T_{ri} > t] = \prod_{s=1}^{tr}(1-h_{ri}^{s})$.

between the t_{p1} period and just before t_{p2}) or the second one finishes (t_{r2} takes places after t_{p2}). Finally, d_{1i}, d_{2i}, d_{3i}, d_{4i} are dummy variables that allow us to distinguish between censored and completed poverty and non-poverty spells.

The first component in (4) captures the likelihood that the individual during her first poverty period remains in poverty all the period under study (a). The second and third component account for the likelihood of individuals who exit during their first poverty period to their first non poverty period, remaining in this state the rest of years (b) or re-entry again to their second poverty experience (c). Within these some will remain in their second poverty experience the rest of years (d) or they will exit to their second non-poverty experience (e). Finally, the last two components capture the likelihood that the individuals who enter a second non-poverty period remain in this state the rest of years (f) or exit to a new poverty experience (h)[32].

In our estimations we use a quadratic form for the baseline hazard rate as in Biewen (2006) given that our results from life-tables confirm the adequateness of this particular form of duration dependence (see also figures 2 and 3). In order to take unobserved heterogeneity into account, a finite-mixture unobserved heterogeneity distribution with unknown support points is also considered[33]. Therefore, the likelihood function for individual i is obtained by integrating the following conditional likelihood distribution:

$$L_i(\beta,\theta,\gamma,\pi)=\prod_{s=1}^{S} L(\beta,\gamma \mid \theta=s)\pi(s) \qquad (5)$$

where θ are the location points, π the probability associated to them, and *s* the number of support points.

Note here that the use of current duration as an explanatory variable for leaving a poverty or non-poverty spell stems from the idea that it appears reasonable to think that there is something about the length of the period of time spent either in poverty or out of poverty that affects the probability of a

[32] In (4) d_{li} allows for making a separation between censored and uncensored durations during the first poverty period, taking value 1 when the individuals exit to their first non-poverty period and 0 in the rest of cases; $d_{2i\,i}$ equals 1 when the individuals during their first non-poverty period exit to a second poverty period (0 in the rest of cases); d_{i3} allows a separation between censored and uncensored durations in the second poverty period, it takes value 1 when the individuals exit to non-poverty during the second poverty period (0 in the rest of cases); finally, d_{i4} distinguishes between censored and completed duration in the second non-poverty period, it takes value 1 when the individuals exit to poverty during the second non-poverty period (0 in the rest of cases).

[33] Heckman and Singer (1984) demonstrate that standard parametric form assumptions for unobserved heterogeneity might be biased when the chosen distribution for the unobservable term is incorrect. For this reason, they solve this problem by assuming that unobserved heterogeneity is discretely distributed with unknown support points.

household leaving or returning to this situation. This reasoning appears straightforward in clear-cut definitions of other possible individual states like unemployment, where the loss of human capital or the end of benefit reception while unemployed makes it reasonable to expect a different escape rate from unemployment as unemployment duration increases. Why would this be the case for the state of poverty?.

In the case of poverty, the definition of *state of poverty* is not so clear-cut. The division between being poor or not is a thin line in the income distribution. Is it reasonable then to expect that the opportunities to move up in the income distribution for households in its lowest tail will be affected by the time they remain in low income? Theoretically, when a household enters poverty, household members would start to use up their savings in order to maintain their previous level of welfare. The longer the household is poor, the more likely the household's savings will have ended and the more likely the household is to suffer a welfare loss. This welfare loss may imply a loss of household members' opportunities (due to the costs of undertaking them) that may bring the household out of poverty. These opportunities include the members' search for a job if unemployed, the members' investment in education that will help them enter the labour market in an advantageous position or the departure of members from the household to create a new one. Other effects on the exit hazard rate could be imposed by the means-testing and receipt duration schemes of state benefits paid to the lowest tail of the income distribution. Hence, it would be reasonable to think that the probability that a household jumps out of the lowest tail of the income distribution could be affected by poverty duration.

A similar reasoning would apply to the probability of returning to poverty. As Gardiner and Hills (1999) point out, the income mobility process is not random and low-income escapers are more likely to drop back into the poorest than those who never suffered low-income. Clearly, duration out of poverty in this case is expected to play a similar role: the longer the time the individual is out of poverty, the lower the probability of returning to it.

The study of the relationship between the duration of a poverty spell and the escape and re-entry rate will test this correlation in order to find out if it is constant in time or it changes after some duration of a poverty or non poverty spell. Obviously, one should note that, in the case of poverty or non-poverty, the difficulties in detecting this correlation and disentangling it from unobserved heterogeneity may be larger than for other definitions of individual or household states. The reason is the larger amount of events that affect the value of the household income and the time span needed in order to detect this correlation due to both the time it takes a household to use up its savings and the long-term nature of the effects of a household's low income period on most household members' labour market opportunities and correlated decisions.

The covariates included in our estimations will try to capture the differences in individual characteristics but also those related to the composition of the household they belong to (number of dependent children, age and education of the household head, number of earners in the household, etc.) and household members labour market attachment (whether the household head or other adults are in paid work, etc.). The only individual variable included is gender while age will probably be captured by the age of the household head.

5. RESULTS

In a first approach to measuring the relevance of spell duration on the probability of leaving poverty we report life-table estimates of the probability of leaving and re-entering poverty. These results assume that the population is homogeneous in characteristics. We begin by analysing the whole sample of spells irrespective of their order and follow by distinguishing the order of each spell occurrence. In a second step we report results on estimations of transition rates using multivariate hazard regression models. This second approach to measuring transition risk provides a generalization of life tables estimations when transition rates are allowed to vary not only with the elapsed time at risk but also with observed and unobserved individual characteristics.

5.1. Life-table estimates of transition rates

Tables 7 and 8 display the life-table estimates of hazard rates, survival probability and cumulative failure. Table 7 illustrates that both types of spells show a decline of the transition hazard as duration evolves, thus supporting the idea of *negative duration dependence* for both situations. However, some differences are already observable between the exit and re-entry hazards. First, the probability of returning to poverty is significantly lower than the probability of exiting from poverty. Thus, non-poverty spells, in general are of a longer duration than poverty spells. Secondly, the re-entry hazard continues to decline after three years of spell evolution while the exit hazard rate experiences a rapid decline during the first three years but is fairly constant from then onwards.

Distinguishing the order of spells and thus analysing the effects of spell accumulation is one our main objectives. Therefore, in table 8 we include results on transition rates for each spell type by their order of occurrence. We can see that the results in table 7 are similar to those obtained for the first spell of poverty or non-poverty in table 8 but are clearly different from those obtained for the second poverty spell. This result underlines the importance of taking multiple spells into account and in considering the differential hazard rate implied by accumulation of multiple experiences in and out of poverty.

Regarding the results for the first poverty spell, we can see that hazard rates decline rapidly during the first two years of observed poverty spell duration, thus supporting the idea of *negative duration dependence*. However, the hazard stays fairly constant from two up to seven years duration. Indeed, some 47 per cent of individuals in their first poverty spell left after one year of observation in the panel (note here that the real spell could be much longer) while out of those that remain poor, just over a third 32 per cent left poverty in the following year. In contrast, from the third to the sixth year of observation, the exit hazard rate fell only by two percentage points, from 23.5 to 21.4. Combining this relative high hazard rates for the first poverty spell with the results on the first spell survival function suggests that the majority of individuals in our sample experience relatively short poverty spells while some minority (a fifth of the sample) experience relatively long spells: 62 per cent of individuals remain poor only during one year, 44 per cent two years, 35 per cent at least 3 years and just about 19 per cent seven or more years.

Table 7

LIFE TABLES ESTIMATES OF HAZARD RATES, SURVIVAL PROBABILITY AND CUMULATIVE FAILURE FOR ALL POVERTY EXITS AND RE-ENTRIES.

Based on all poverty spells observed from ECHP waves 1994-2000 for individuals who are poor since 1994

Interval (years)		Total number of individuals at risk Total (individuals)	Deaths	Lost	Survival (%)	Cum. Failure (%)	Std. Error	Hazard (%)	Std. Error
All exits									
1	2	5218	1900	715	60.91	39.09	0.7	48.59	1.08
2	3	2603	718	378	42.79	57.21	0.75	34.94	1.28
3	4	1507	323	225	32.88	67.12	0.75	26.2	1.45
4	5	959	163	114	26.94	73.06	0.75	19.87	1.55
5	6	682	111	103	22.2	77.8	0.74	19.3	1.82
6	7	468	87	36	17.91	82.09	0.73	21.4	2.28
7	8	345	0	345	17.91	82.09	0.73	0	-
All re-entries									
1	2	3302	947	632	68.29	31.71	0.85	37.69	1.2
2	3	1723	351	346	52.82	47.18	0.98	25.54	1.35
3	4	1026	155	206	43.95	56.05	1.04	18.33	1.47
4	5	665	63	169	39.18	60.82	1.09	11.48	1.44
5	6	433	38	107	35.26	64.74	1.15	10.54	1.71
6	7	288	0	288	35.26	64.74	1.15		

Table 8

LIFE TABLES ESTIMATES OF HAZARD RATES, SURVIVAL PROBABILITY AND CUMULATIVE FAILURE BY ORDER OF OCCURRENCE.

Based on all poverty spells observed from ECHP waves 1994-2000 for individuals who are poor since 1994

Interval (years)		Total number of individuals at risk	Deaths	Lost	Survival (%)	Cum. Failure (%)	Std. Error	Hazard (%)	Std. Error
First poverty spell (1)									
1	2	3664	1326	392	61.76	38.24	0.83	47.27	1.26
2	3	1946	519	186	44.47	55.53	0.88	32.57	1.41
3	4	1241	246	143	35.11	64.89	0.87	23.51	1.49
4	5	852	151	54	28.69	71.31	0.85	20.15	1.63
5	6	647	111	68	23.49	76.51	0.83	19.91	1.88
6	7	468	87	36	18.95	81.05	0.8	21.4	2.28
7	8	345	0	345	18.95	81.05	0.8	0	—
First non-poverty spell (2)									
1	2	2440	736	349	67.51	32.49	0.98	38.79	1.4
2	3	1355	295	166	51.86	48.14	1.1	26.23	1.51
3	4	894	144	132	42.84	57.16	1.14	19.05	1.58
4	5	618	63	122	37.99	62.01	1.16	11.99	1.51
5	6	433	38	107	34.19	65.81	1.2	10.54	1.71
6	7	288	0	288	34.19	65.81	1.2		
Second poverty spell (3)									
1	2	1276	491	200	58.25	41.75	1.44	52.77	2.3
2	3	585	190	142	36.72	63.28	1.54	45.35	3.2
3	4	253	77	69	23.78	76.22	1.55	42.78	4.76
4	5	107	12	60	20.07	79.93	1.64	16.9	4.86
5	6	35	0	35	20.07	79.93	1.64		
Second non-poverty spell (4)									
1	2	770	206	207	69.09	30.91	1.79	36.56	2.5
2	3	357	56	169	54.89	45.11	2.21	22.9	3.04
3	4	132	11	74	48.54	51.46	2.66	12.29	3.7
4	5	47	0	47	48.54	51.46	2.66		

(Sigue)

(Continuación)

Interval (years)		Total number of individuals at risk	Deaths	Lost	Survival (%)	Cum. Failure (%)	Std. Error	Hazard (%)	Std. Error
Third poverty spell (5)									
1	2	273	83	118	61.21	38.79	3.33	48.12	5.13
2	3	72	9	50	49.49	50.51	4.43	21.18	7.02
3	4	13	0	13	49.49	50.51	4.43		
Third non-poverty spell (6)									
1	2	92	5	76	90.74	9.26	3.94	9.71	4.34
2	3	11	0	11	90.74	9.26	3.94		
Fourth poverty spell (7)									
1	2	5	0	5	100	0	0		

The interesting question we pose is: Do these conclusions differ for those individuals that experience a second occurrence in poverty (after having experienced a period of non-poverty)? We observe that the probability that an individual leaves poverty when experiencing a second occurrence is significantly higher than it was during his/her first poverty spell. Indeed, during the first year the hazard rate in the second poverty period is 5.5 percentage points higher than in the first one. Interestingly this difference increases up to a 14 and 19 per cent more during the second and third year. Therefore, we find evidence that individuals remain a relatively shorter time in poverty if they have managed to leave deprivation for some time most recently.

Turning to results on non-poverty spells, we observe that the shape of the first re-entry hazard is also consistent with negative duration dependence up to the third year, remaining constant thereafter. Interestingly we find little differences in the annual hazard rates of returning to poverty depending on the order of the non-poverty spell. The largest difference is observable after durations of three years or more and, in contrast with the impact of spell order in poverty experiences, re-entry hazard rates in the second non-poverty spell are lower than in the first one. This implies that once you have managed to step out of poverty once, the accumulation of non-poverty spells plays in your favour by reducing the probability of coming back to poverty.

Figures 2 and 3 plot exit and re-entry hazard rates by spell order using the previous life-table results[34]. The common pattern of these estimated hazard functions is that they all show some negative duration dependence and all exits (and all re-entries) show a similar pattern to that of the first one. However,

[34] This figures plot estimations of the poverty hazard rates, which may be interpreted as a type of sample-average hazard function without controlling for individual characteristics.

thanks to the separate estimations by spell order, we already observe that the size of the second spell poverty hazard is significantly higher than that of the first spell, at least during the first three years of spell duration. In addition, the size of the first re-entry poverty hazard is higher than that of the second one.

Figure 2

LIFE-TABLE HAZARD RATES AS DURATION EVOLVES. ECHP 1994-2000

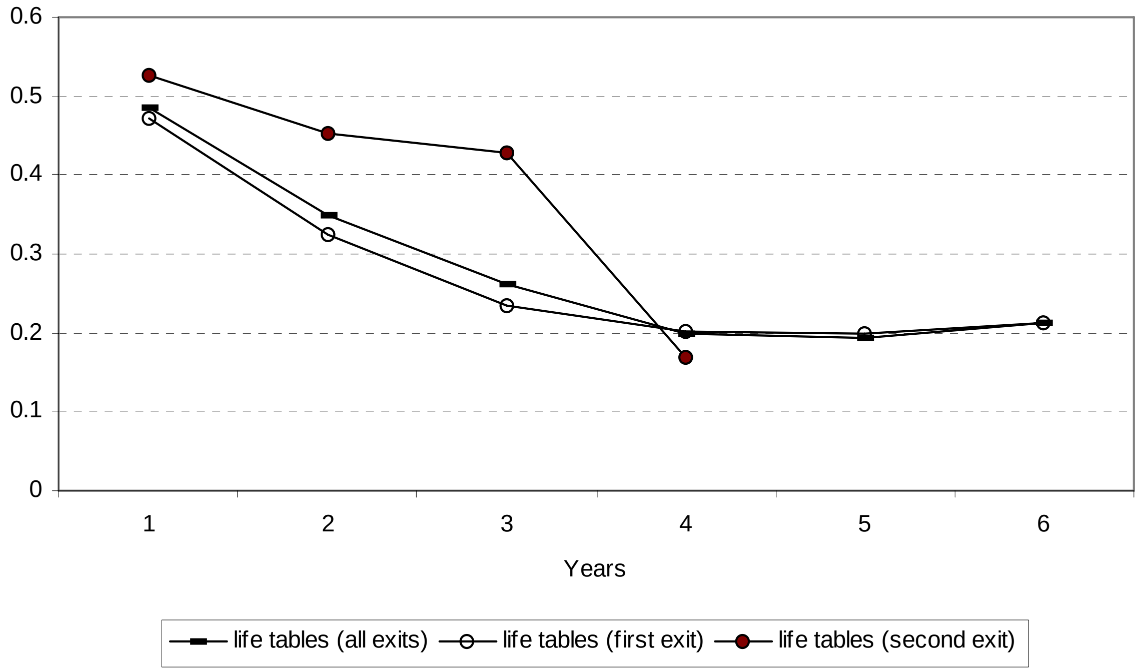
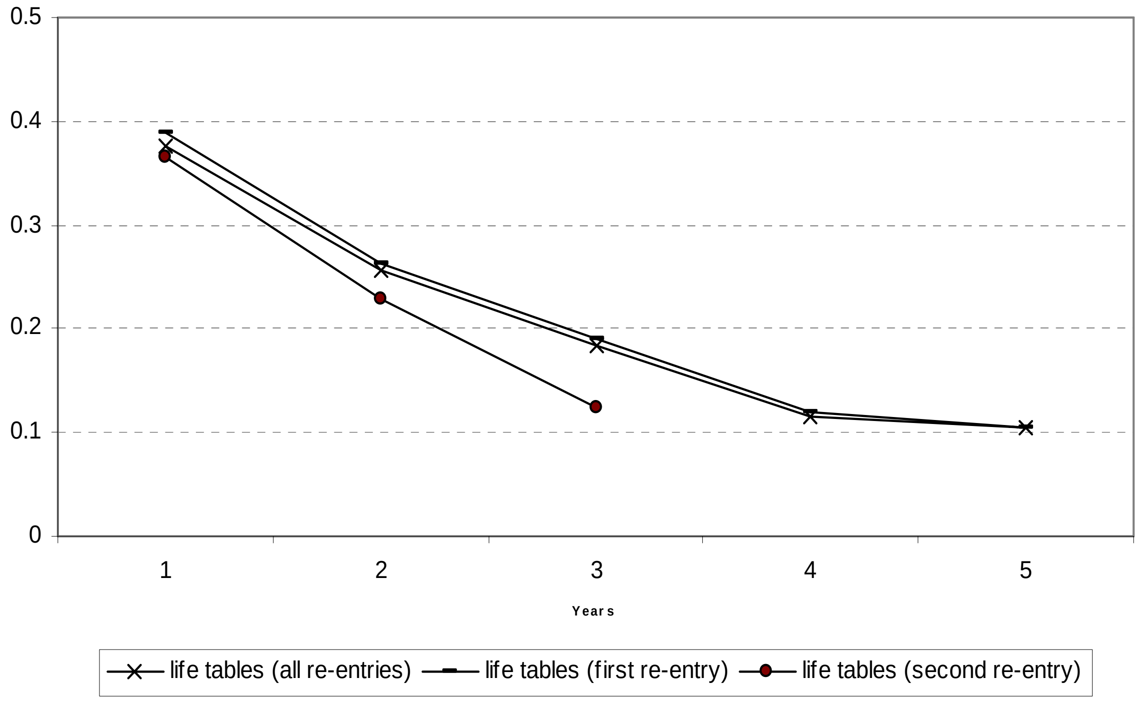


Figure 3

LIFE-TABLE HAZARD RATES AS DURATION EVOLVES, BY SPELL ORDER. ECHP 1994-2000



5.2. The main characteristics of individuals under analysis

In any case, before going into a more detailed multivariate analysis of spells we must focus our discussion on the comparative characteristics of the samples of spells to be used in regressions. In order to do this we have constructed table 9 where one can compare the characteristics of the sample of individuals who experience a left-censored spell (*first spells*, second pair of columns and *second spells*, third pair of columns) with those of the sample of individuals whose first transition into poverty is observed (*inflow sample of spells*, fourth pair of columns)[35].

Results show that there are some differences in the characteristics of individuals who suffer some left-censored poverty spell and those who are observed to enter poverty within the observation window. In particular, these differences are related to the household and household's head socioeconomic and demographic characteristics more than to individual characteristics, even if the individual's age and labour status is somewhat different too. In fact, if we were to use a sample of new entrants to poverty instead of using one of individuals with left-censored poverty spells, our sample would contain significantly younger individuals, more often active, living in households whose head is relatively young (often below 49 years of age), more educated, more often employed full time or unemployed but rarely retired, with more adults in the household (more often active in the labour market) and fewer children, whose main income source is wages and whose total household income is nearer to the poverty line and, in some cases, it is declared to be temporarily zero. As it could be expected, the characteristics of the sample of new entrants turn to be most similar to those of the individuals who, having experienced a first left-censored spell, register a second poverty spell within the observation window.

Focusing on poverty spell duration, we can see that the elapsed duration of poverty spells for individuals with left-censored spells is significantly longer: 2.4 years compared to 1.7 years (more than eight months longer). This result clearly reflects the duration bias of choosing to discard left-censored spells completely when analysing poverty dynamics. Including the first and second spell in the analysis reduces duration to 2.2 years and includes non-poverty spells of a mean duration of 1.7 years in between poverty spells.

In general, after these descriptive analyses, we can assert that the inclusion of left-censored spells in the regressions will influence multivariate results on first-spell hazard rates for the case of Spain by including individuals who have experienced poverty more persistently and, in general, due to their household composition, are

[35] Note that an extended version of this table appears in table A8. Note also that right censoring may imply also that the complete duration of a spell is not observed. We here refer to spell duration in terms of the observation of the spell since it begun until it finishes or suffers from attrition.

less likely to transit out of poverty. These effects, however, are less observable for those individuals who suffer poverty recall given that their characteristics are much more similar to those of a sample of new-entrants to poverty.

Table 9

CHARACTERISTICS OF SPELLS SAMPLES, INDIVIDUALS: MEANS

Characteristics	POVERTY SPELLS							
	All spells (5,218 indiv.) (5,113 weight)		**First spell (3,664 indiv.)** (3,387weight)		**Second spell (1,276 indiv.)** (1,416 weight)		**Inflow sample of spells – new entrants (1,593 indiv.)** (1,632 weight)	
Individual Characteristics	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**
Age	36.2	22.2	36.6	22.8	35.4	21.2	35.2	21.2
Aged 16-29	0.23	0.42	0.22	0.42	0.24	0.43	0.25	0.43
Aged 60+	0.18	0.39	0.20	0.40	0.16	0.37	0.16	0.37
Child, below 16 years old	0.23	0.42	0.24	0.43	0.22	0.42	0.22	0.42
Labour status								
Working (+15 hours/week)	0.22	0.41	0.20	0.40	0.24	0.43	0.26	0.44
Working (less 15 hours/week)	0.02	0.15	0.03	0.16	0.02	0.13	0.01	0.10
Unemployed	0.14	0.35	0.15	0.35	0.13	0.34	0.14	0.34
Discouraged worker	0.00	0.06	0.00	0.05	0.01	0.09	0.02	0.13
Economically inactive	0.38	0.49	0.39	0.49	0.38	0.49	0.36	0.48
Unemployment experience								
Had unemployment spell last 5 years	0.29	0.45	0.29	0.45	0.30	0.46	0.31	0.46
Main income source								
No income from any source	0.28	0.45	0.28	0.45	0.29	0.46	0.27	0.44
Wages and salaries	0.13	0.34	0.12	0.33	0.14	0.35	0.19	0.39
Self-employment or farming	0.06	0.24	0.05	0.23	0.06	0.24	0.06	0.23
Pensions	0.11	0.31	0.11	0.32	0.09	0.28	0.09	0.29
Unemployment benefits	0.06	0.23	0.06	0.24	0.05	0.22	0.06	0.25
Any other social benefits	0.08	0.27	0.08	0.27	0.08	0.27	0.06	0.24
Private income	0.05	0.22	0.04	0.20	0.06	0.24	0.04	0.21

(Sigue)

(Continuación)

Characteristics	POVERTY SPELLS							
	All spells (5,218 indiv.) (5,113 weight)		**First spell (3,664 indiv.)** (3,387weight)		**Second spell (1,276 indiv.)** (1,416 weight)		**Inflow sample of spells – new entrants (1,593 indiv.)** (1,632 weight)	
Individual Characteristics	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**
Household Characteristics								
Household structure								
Total household members	4.16	1.68	4.09	1.69	4.27	1.63	4.17	1.59
Number of adults in household	3.06	1.35	2.99	1.32	3.15	1.43	3.10	1.40
Number of 0-5 children	0.23	0.42	0.24	0.43	0.21	0.41	0.21	0.41
Main income source								
Wages and salaries	0.34	0.47	0.32	0.47	0.38	0.49	0.42	0.49
Self-employment income	0.15	0.36	0.13	0.33	0.18	0.38	0.14	0.34
Pensions income	0.23	0.42	0.25	0.43	0.20	0.40	0.21	0.41
Unemployment income	0.12	0.33	0.14	0.35	0.10	0.30	0.13	0.34
Transfers income	0.10	0.30	0.11	0.31	0.10	0.30	0.08	0.27
Private income	0.05	0.21	0.05	0.22	0.04	0.19	0.03	0.16
Poverty Gap (as % of poverty line)								
0-10%	0.23	0.42	0.22	0.42	0.23	0.42	0.28	0.45
Zero income	0.02	0.14	0.00	0.00	0.06	0.25	0.07	0.25
Household head characteristics								
Household head aged 30-39	0.20	0.40	0.20	0.40	0.23	0.42	0.27	0.44
Household head aged 40-49	0.26	0.44	0.24	0.43	0.29	0.45	0.28	0.45
Household head aged 50-59	0.23	0.42	0.23	0.42	0.23	0.42	0.19	0.39
Household head aged 60+	0.24	0.43	0.29	0.45	0.16	0.36	0.12	0.33
Female household head	0.14	0.35	0.12	0.33	0.18	0.38	0.15	0.35
Separated, Divorced or Widowed	0.10	0.30	0.12	0.32	0.06	0.23	0.07	0.25
Head is in paid work, more than 15 hours	0.53	0.50	0.48	0.50	0.60	0.49	0.64	0.48
Head is working part-time	0.07	0.25	0.07	0.25	0.08	0.27	0.05	0.21

(Sigue)

(Continuación)

Characteristics	POVERTY SPELLS							
	All spells (5,218 indiv.) (5,113 weight)		**First spell (3,664 indiv.)** (3,387weight)		**Second spell (1,276 indiv.)** (1,416 weight)		**Inflow sample of spells – new entrants (1,593 indiv.)** (1,632 weight)	
Individual Characteristics	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**	**Means**	**S.E.**
Head retired	0.15	0.36	0.19	0.39	0.09	0.28	0.06	0.24
Head unemployed	0.17	0.38	0.15	0.35	0.23	0.42	0.23	0.42
Number of earners in household (active)	1.61	1.14	1.58	1.13	1.63	1.15	1.61	1.12
Head university education	0.04	0.20	0.03	0.16	0.07	0.25	0.06	0.24
Head secondary education	0.07	0.25	0.06	0.24	0.08	0.28	0.13	0.34
Characteristics of Spells								
Non-censored observations	0.63	0.48	0.68	0.47	0.59	0.49	0.68	0.47
Elapsed duration (years)	2.17	1.68	2.40	1.90	1.82	1.06	1.67	1.19
Lagged poverty duration (years)			—	—	1.78	1.02		
Lagged accum. pov. duration (years)			—	—	3.60	1.39		
Lagged non-poverty duration (years)			—	—	1.73	1.02		
Lagged accum. non pov. duration (years)			—	—	1.73	1.02		

Note: These results omit the percentage of missings in variables for which children have no information available.

5.3. Estimation results on poverty exits and re-entries

In this sub-section, we estimate several discrete hazard models taking into account the individual's complete poverty history in order to analyse the determinants of leaving or re-entering poverty in Spain. We are interested in obtaining the evolution of the hazard rates as poverty and non-poverty spells evolve after controlling for demographic and socio-economic individual and household characteristics as well as for lagged poverty and non poverty durations. Table 10 presents the estimated hazard regressions for all exits and re-entries controlling for unobserved heterogeneity and distinguishing the spell order, therefore considering multiple exit and multiple re-entry periods are not independent.

Table 10

DISCRETE HAZARD MODELS FOR ALL POVERTY EXITS AND ALL POVERTY RE-ENTRIES, BY SPELL ORDER CONTROLLING FOR UNOBSERVED HETEROGENEITY

Variables	First poverty exit			First poverty re-entry (after first poverty exit)			Second poverty exit (after first poverty re-entry)			Second poverty re-entry (after second poverty exit)		
	Coef.	Std. Err.	P>z	Coef.	Std. Err.	P>z	Coef.	Std. Err.	P>z	Coef.	Std. Err.	P>z
Intercept	-1.636	0.292	***	0.754	0.326	**	-0.790	0.505		1.822	0.893	**
Spell characteristics:												
Duration	-0.053	0.102		0.173	0.168		1.157	0.326	***	1.002	0.716	
Duration square	-0.024	0.013	*	-0.094	0.027	***	-0.331	0.073	***	-0.412	0.187	**
lagged poverty duration (years)							-0.600	0.068	***			
lagged non poverty duration (years)										-0.825	0.165	***
Individual characteristics:												
Gender (Male)	0.109	0.054	**	-0.051	0.073		-0.026	0.105		0.007	0.172	
Household characteristics												
(Number of earners in the household)/10	0.276	0.028	***	-0.323	0.040	***	0.114	0.054	**	-0.301	0.107	***
Head of Household implyment+education												
Head employed or other situation	ref			ref			ref			ref		
Head retired	-0.053	0.124		-0.025	0.189		-0.330	0.317		0.689	0.510	
Head unemployed	-0.779	0.077	***	0.573	0.100	***	-0.683	0.137	***	1.360	0.246	***
Head university education	0.440	0.138	***	-0.712	0.183	***	0.455	0.200	**	-1.126	0.388	***
Head secondary education	0.628	0.110	***	-0.649	0.129	***	0.931	0.192	***	-0.253	0.324	
Head less than secondary education	ref			ref			ref			ref		
Head of Household Age:												
Less than 30	ref			ref			ref			ref		
30-39	-0.325	0.118	***	-0.309	0.144	**	-0.161	0.211		0.403	0.336	

(Sigue)

(Continuación)

Variables	First poverty exit			First poverty re-entry (after first poverty exit)			Second poverty exit (after first poverty re-entry)			Second poverty re-entry (after second poverty exit)		
	Coef.	**Std. Err.**	**P>z**	**Coef.**	**Std. Err.**	**P>z**	**Coef.**	**Std. Err.**	**P>z**	**Coef.**	**Std. Err.**	**P>z**
40-49	-0.339	0.115	***	0.185	0.132		0.110	0.190		0.413	0.288	
50-59	-0.100	0.123		0.260	0.139	*	-0.082	0.203		0.017	0.311	
60+	0.251	0.141	*	-0.005	0.183		-0.374	0.240		-0.359	0.455	
Household composition:												
One person household	0.984	0.244	***	-1.902	0.313	***	0.043	0.455		-2.959	1.117	***
Single parent with one or more children	1.051	0.199	***	-1.256	0.246	***	0.400	0.364		-1.844	0.733	***
Couple, no children	0.372	0.211	*	-1.236	0.267	***	0.409	0.393		-2.158	0.751	***
Couple one child (child aged < 16)	1.940	0.217	***	-1.030	0.254	***	0.686	0.398	*	-2.146	0.727	***
Couple two child (child aged < 16)	1.058	0.187	***	-0.351	0.224		0.156	0.347		-2.216	0.700	***
Couple three or more(child aged < 16)	ref	-	-	ref	-	-	ref	-	-	ref	-	-
Couple, one or more children (at least one child aged ≥16)	0.908	0.183	***	-1.035	0.216	***	0.943	0.323	***	-2.127	0.683	***
Other households	1.275	0.188	***	-1.287	0.225	***	0.494	0.339		-1.471	0.699	**
Mass points and probability												
θ_1	5.025***											
θ_2	-0.365***											
$Pr(\theta_1)$	0.07											
$Pr(\theta_2)$	0.93											
Sample-individual observations)	9,163			6,028			2,256			1,306		
Log-likelihood	-9537.8403											

*** Indicates significance at 1 per cent; ** indicates significance at 5 per cent; * indicates significance at 1 per cent.

Note: The reference individual is a female living in a household whose head is less than 30 years of age, has less than secondary education and is employed. The household includes three or more children.

The models considered include covariates such as gender, head of household age and education, head job positions, head qualifications, household composition distinguishing single parents and households with dependent children, number of earners in the household, the length of the current poverty and non-poverty spells and the duration of the lagged poverty spells[36].

As one would expect, table 10 confirms that effects of a covariate on poverty exit and re-entry has the opposite sign in most of the cases. Thus, characteristics that help in leaving poverty also help in avoiding recurrence. Exit rates from poverty are higher if the household's head has a high educational attainment or there is a large number of earners in the household.

However, there are variables that present interesting differences in their effects depending on the number of accumulated spells experienced by individuals. Figure 4 plots the baseline hazard rate (duration and duration squared) for poverty and non-poverty exits after controlling for observed and unobserved heterogeneity. For individuals in their first observed poverty spell (many of them left-censored), we find some negative duration dependence from the first year of spell duration[37], while for those experiencing a second spell, the effect of duration is very different. Indeed, the probability of leaving a second poverty spell increases (even if less and less each year) at the beginning of the spell (approximately during three years), and from then after, the probability of leaving poverty starts to decrease as duration evolves. This pattern could be explained by the fact that first spells may have lasted more than three years already at their first year of observation. We also observe here that, for the case of poverty recall the hazard shifts down when spells accumulate (i.e. the second non-poverty spell observed). Thus, the probability of re-entering poverty in a second period does not grow when the non-poverty spell is still short (as it happens with poverty spells) it is always lower as the spell duration increases. Finally, we observe that the probability of experiencing a second re-entry poverty is lower for those that had one already and, in contrast, having a second exit from poverty is relatively more probable when one has had one.

[36] We fitted a variety of other alternative specifications. For example, we considered including unemployment rates and GDP growth rate but they were not statistically significant and the distribution of the estimated parameters was very imprecise. Therefore, these covariates we not kept in the specifications reported here.

[37] Biewen (2003) finds similar negative duration effects for German individuals.

Figure 4

THE SHAPE OF THE PREDICTED HAZARD RATE FOR POVERTY AND NON POVERTY EXITS

(after controlling for observed and unobserved heterogeneity) at the mean of covariates. ECHP 1994-2000.

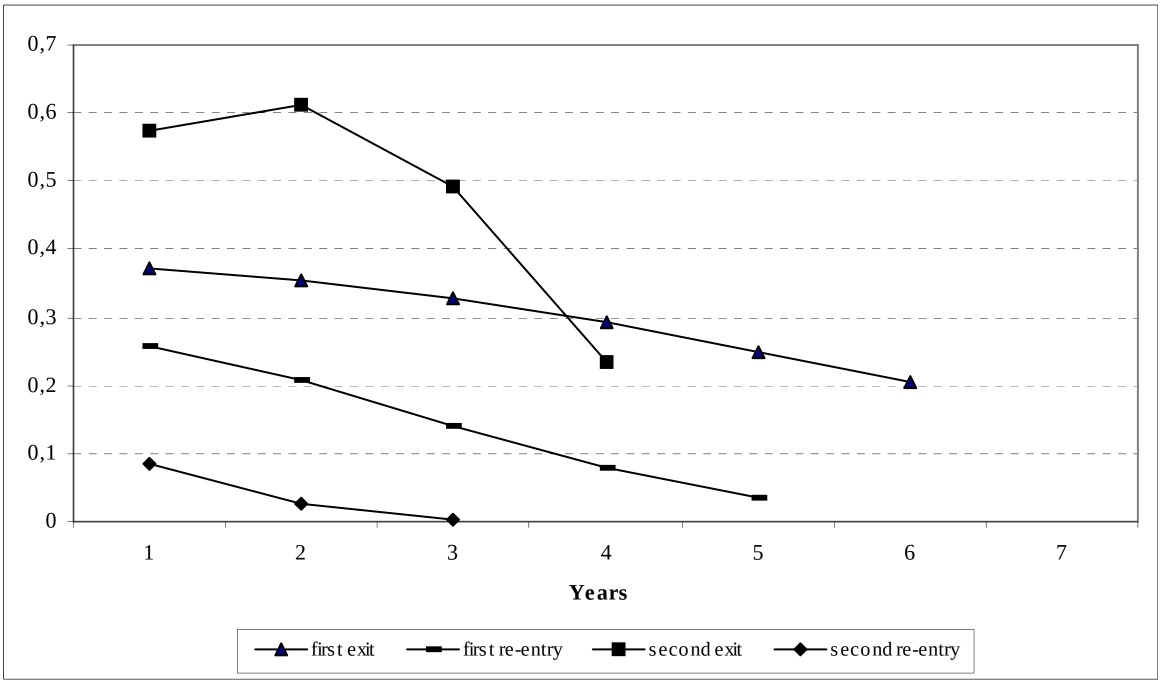


Table 10 also shows that individuals who experience poverty are more likely to experience poverty in future because the longer the previous poverty duration the less likely individuals will leave a second poverty spell. Therefore, there is some state dependence poverty effect[38] in individuals' poverty histories. Alternatively, the time spent out of poverty plays the opposite role: the longer the time the individual is out of poverty, the lower the probability of poverty recall. Also, the effects of lagged durations have the expected sign: lagged poverty duration reduces the exit hazard and lagged non-poverty duration reduces the recall hazard.

Another observable difference in the effect of covariates when spells accumulate is that being a male significantly increases the individual's probability of leaving poverty (a 11.5 percent) only for first spells while for a second poverty spell, gender does not affect the individual's chances to leave poverty[39].

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[38] This effect is also detected by Biewen (2004) in Germany. The explanations are twofold. On one hand, individuals who are poor in one period have observed characteristics such as human capital decay, unemployment, health problems or difficult living arrangements that make poverty prone. On the other hand, they present unobserved heterogeneity terms such as lack of intelligence or ability, low levels of motivation or unfavourable attitudes.

[39] Other papers with a variety of data sources for Spain in different periods since 1985 find that couples with three or more children are the type of households who are more prone to be in poverty - see García-Serrano *et al.* (2001), Cantó (2002 and 2003), Cantó and Mercader (2002),

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