OBTAINING LIFETIME EARNINGS PATTERNS FOR SPAIN

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P. T. N.º 22/09

Paper presented to the Workshop "Modelling the Effects of Pensions and Other Welfare State Transfers in an Ageing World". Institutional support from the Spanish Science and Technology System (Project Nos. SEJ2006-4444 and ECO2008-04997/ECON) and the Catalan Government Science Network (Project Nos. SGR2005-177 and SGR2005-460 as well as XREPP-Xarxa de Referència en Economia i Política Públiques) is acknowledged.

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Edita: Instituto de Estudios Fiscales
N.I.P.O.: 602-09-006-9
I.S.S.N.: 1578-0252
Depósito Legal: M-23772-2001

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ABSTRACT

In this paper we estimate lifetime earnings profiles for Spain using a new dataset based on administrative Spanish Social Security records from 1980 to 2005. In order to exploit the panel data dimension of this dataset, we use panel data estimates specially derived for censored data as is the case. First, the Kalwij (2003) estimator derived for censored panel data is used in order to estimate the censored earnings, controlling for some identifiable variables. Second, to overcome the main weakness of the above-mentioned method, i.e. the impossibility of identifying variables that remain constant over time, the Wooldridge (1995) and Zabel (1992) estimators are used. Finally, given the difficulties in isolating the period, age and cohort effects, linked by a linear relationship, a test of uniform wage growth is performed following Fitzenberger *et al*. (2001). As the results of the test indicate that the age and period (or cohort) effects can be considered as separable, a tentative age profile can be derived for each gender and contribution group.

Keywords: Age-earning profiles, censoring, cohort analysis, panel data.
JEL classification: J31.

1. INTRODUCTION

There is a broad literature about the life-cycle earnings patterns of workers in developed economies, starting with Mincer's (1974) famous work. In fact, the so-called 'Mincerian human capital earnings function' –in which the natural logarithm of earnings is expressed as a linear function of the number of years of school completed and as a quadratic function of potential work experience– is the standard specification of age-earning profiles empirical studies.

Most empirical works initially carried out cross-sectional estimations, mainly due to the non-availability of datasets needed for longitudinal analysis. Nevertheless, some doubts have arisen about the traditional literature based on Mincerian specification. On the one hand, the quadratic empirical approximation could not be the best, as shown in Murphy and Welch (1990) for USA and Robinson (2003) for UK. Both studies found that alternative specifications –such as a cubic or quartic– fit the data better. On the other hand, some studies found that cross-section profiles fail to portray earnings growth over the working life, so they should not be used as longitudinal earnings profiles. In particular, Gohmann et al (1998) found evidence that age-earnings profiles are not stable over time and, consequently, care must be taken in using cross-sectional data. Thornton et al (1997) found that the U-inverted shape characteristic of cross-sectional profiles is not always observed in longitudinal profiles.

As a result of an increasing availability of micro datasets, the possibility of performing longitudinal analyses on earning profiles has been increasing. Nevertheless, in this context an identification problem arises. Clearly three types of factors are associated with earnings evolution over time, namely, age, cohort and period effects. Age effects refer to how age and accumulation of work experience by individuals affect their life cycle earnings pattern, over or under the rate of general productivity growth of the economy. The cohort (birth year) effect refers to special events or characteristics that can affect a specific cohort. For example, Freeman (1979) and Welch (1979) analysed the impact of the cohort size on the earnings evolution, and found that the entry into the labour market of baby boom generations had a sizeable negative effect on the relative wage of young workers. Finally, period effects comprise cyclical factors with similar impact on all ages and cohorts. A cross-sectional age earning profile obtained for individuals of different ages in a given year shows differences among cohorts due to age and cohort factors. Meanwhile, a longitudinal earning profile obtained for an individual shows age and period effects. In a panel data set, combining cross-section and time series data, the three effects are confounded. In fact, it is not possible to separate them as there is a simple linear combination between them –birth year is the difference between current period and age.

In this context, Hanoch and Honig (1985) use panel records of US Social Security Administration for 1951-1974 period to estimate age-profiles earnings of white married males and unmarried females covering ages 40-68. They find that the age-earnings path shows a similar shape but a later decline than in the cross-section estimations. Beaudry and Green (2000) use the data of Canadian Surveys of Consumer Finance from 1971-1993 to estimate longitudinal profiles for cohorts –defined as individuals born in two consecutive years– of both sexes divided in two education groups. They find that cohort earnings profiles have shifted down for successive cohorts, especially for high-educated male. Beach and Finnie (2004) replicate the Beaudry and Green analysis using tax-based longitudinal microdata for 1982-99 period obtaining similar results. Nevertheless, they find that profile earnings of young cohorts (entering the nineties) have steepened. Gosling, Machin and Meghir (2000) analyse the wage differential between young and old cohorts in UK, and find that new cohorts experience a permanent loss in their earnings pattern due to a lower entry wage. Fitzenberger *et al.* (2001) compare the wage trends across cohorts of German male employees using a random sample of social security data for 1976-1984 period. In order to try to identify the period, cohort and age effects, they develop a parsimonious approach assuming separability among them and testing the existence of a uniform wage trend. Using quantile regression techniques they obtain earnings profiles for different education groups as functions of age, cohort and year. Defining a 'macroeconomic wage trend' they conclude that German wage structure was quite stable in that period.

In this paper we analyse the evolution of career profiles in the Spanish labour market over the 1981-2005 period. The structure of the paper is as follows: Section 2 is devoted to describing the database. We use a longitudinal dataset built from the continuous working life sample (MCVL)[1], a random sample of administrative records of the Spanish Social Security administration containing censored individual earnings histories. In Section 3 empirical framework and estimation procedure employed are described. The results of this estimation are presented in Section 4 which also discusses the past evolution of wages and the possibility of extrapolating those tendencies into the future. In particular, the existence of a separable age and period effect is tested following Fitzenberger *et al.* (2001). Finally, Section 5 summarises main conclusions.

2. THE DATA

The MCVL is a sample extracted from Spanish Social Security administrative data. 4% of individuals registered with the Social Security administration –either

[1] *Muestra Continua de Vidas Laborales*.

contributing or receiving benefits– over the sampling year are selected and their entire life time history is joined to the data set. Hence, although it is not a pure panel, this dataset is rich in longitudinal data.

Among the difficulties of dealing with such a large administrative dataset –the sample size is around a million people in year 2005– the most challenging for our purpose is dealing with empty contribution bases and relating contribution, affiliation and benefit data from the same individual all defined with different time periods/units. In particular, in order to extract reliable contribution data in a specific time unit, it is necessary to follow up all the contracts in which individuals have been involved computing time and contribution separately, in order to avoid a wrong correspondence between working time and contribution per unit of time.

To avoid those difficulties we obtain, for each year and individual, the average contribution per hour. For that purpose we use all the contracts that this individual has over the year, using their initial and final date and the share of time worked –a variable accounting for part-time work. Nevertheless, as this variable is not completely reliable, we decide to focus on full-time work in order to avoid possible biases, keeping a considerable data set[2]. In particular, once we eliminate part-time contracts and individuals contributing to special regimes, we are left with more than 800,000 individuals, with 25 annual contribution bases at the most, the average being 12.5 per individual. This will be the base of our estimation. Hence we do not take into account changes in time worked. This would indeed complement the present analysis but, given that the database only registers unemployment periods when there is a benefit associated, for the moment we opt to leave aside this analysis.

With respect to the empty contribution bases, we opt for not filling in the missing data, though we eliminate the contracts more affected by this problem (15%). Given that there seems to be no systematic reason for these missing values, any criteria chosen to fill the gaps would bias results in any direction.

3. ESTIMATION PROCEDURE

In this paper we estimate longitudinal earning profiles for Spanish workers combining cross-section and longitudinal data directly extracted from Social Security administrative records. Unfortunately, for a worker "i" the available variable is the "covered earnings" (c_{it}) in year "t" instead of "earnings" in each year (w_{it}). The relationship between the earnings (unobserved variable) and covered earnings (observed variable) is given by a double censored function, that is:

[2] Only 10% of the registered contracts are part time.

$$c_t = \begin{cases} l_t \text{ if } w_t < l_t \\ w_t \text{ if } l_t \le w_t < u_t \\ u_t \text{ if } w_t \ge u_t \end{cases} \quad (1)$$

where l_t and u_t are the lower and upper thresholds mandated annually[3]. For each person "i" in our sample who works in the period of analysis, we observe covered earnings (c_t) each year. The goal in this step is the generation of the earnings of the workers in our sample. In order to estimate the earnings we assume a panel data model. The specific model is:

$$w_{it} = \beta_{0t} + X_{it}\beta_t + \alpha_i + \varepsilon_{it} \quad t = 1,....,T \quad (2)$$

where X_{it} is a vector of exogenous explanatory variables, β_{0t} is a time-specific intercept parameter to control for calendar time effects, including cycle and trend, β_t is the vector of the parameters of interest and α_i is an unobserved individual specific effect. The error term ε_{it} is assumed to be normally distributed with zero mean and variance σ_t^2, $\varepsilon_{it} \left| x_{it}, \alpha_i \approx N\left(0, \sigma_{\varepsilon,t}^2\right)\right.$.

The model outlined in equations (1) and (2) is usually referred to as a doubly censored regression panel data model, and two econometric characteristics need to be taken into account. Firstly, the fact that the covered earnings are observed only in the case that they are located between the thresholds. Bearing in mind that very few observations of our database are bottom censoring, we can neglect left censoring[4]. The equation (1) becomes:

$$c_t = \begin{cases} w_t \text{ if } w_t < u_t \\ u_t \text{ if } w_t \ge u_t \end{cases} \quad (3)$$

Secondly, the possibility that the unobserved job-individual specific characteristics (α_i) are correlated with the explanatory variables, due, for instance, to the omission of educational variables or unobserved individual-specific preferences over work or leisure. Ignoring the censoring structure and the correlation between α_i and x_{it} leads to inconsistent parameter estimates. The way we deal with these problems is described below. First, in (a) and (b) two alternative estimation methods are outlined. Second, the approach taken in order to try to separate the cohort, period and age effect is explained in (c).

a) *Estimators for censored panel data*

This paper adopts a panel data tobit model as formulated in Kalwij (2003). Kalwij's estimator does not allow for an arbitrary relationship between the

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3 We apply the log transformation to wages for statistical reasons.

4 Another reason is that some of the few cases were below the minimum threshold indicating non-reliable information.

unobserved individual effects and the covariates, but rather requires a pre-specified relationship. Following Chamberlain (1984) and Wooldridge (1995), Kalwij (2003) parameterises this relationship as a time-invariant function of the covariates and a random individual specific error term, which can be formalised as a linear function of the across-time averages of the covariates, that is:

$$\alpha_i = \gamma_0 + \overline{X}_i \gamma + v_i \tag{4}$$

where $\overline{X}_i = \frac{1}{T}\sum_{t=1}^{T} X_{it}$, $v_i | x_i \approx N(0, \sigma_v^2)$. Substituting equation (4) into (2) yields:

$$w_{it} = \pi + X_{it}\beta + \overline{X}_i \gamma + \xi_{it} \tag{5}$$

where $\pi = \beta_0 + \gamma_0$, $\xi_{it} = v_i + \varepsilon_{it}$ with $\xi_{it} \approx N(0, \sigma_t^2)$ and $\sigma_t^2 = \sigma_v^2 + \sigma_{\varepsilon,t}^2$. Following Kalwij (2003)[5] to overcome the sensitivity of the parameter estimates with respect to this parameterisation, the estimation method proposed is based on taking first differences in equation (5), which eliminates the random effects –or equivalently from equation (2)– to yield the following model:

$$\nabla w_i = \nabla X_i \beta + \eta_i \tag{6}$$

and the observed variable:

$$\nabla c_i = \begin{cases} \nabla w_i & \text{if } w_{it-1} < u \text{ and } w_{it} < u \\ 0 & \text{if } w_{it-1} > u \text{ and } w_{it} > u \\ u - w_{it-1} & \text{if } w_{it-1} < u \text{ and } w_{it} > u \\ w_{it} - u & \text{if } w_{it-1} > u \text{ and } w_{it} < u \end{cases}$$

where $\nabla w_i = w_{it} - w_{it-1}$, $\nabla x_i = x_{it} - x_{it-1}$, $\eta_i = \xi_{it} - \xi_{it-1} = \varepsilon_{it} - \varepsilon_{it-1}$. Although the job-individual specific effects no longer appear in equation (6) they are still present in the selection part of the model. The cost of applying this estimator is that in the estimation one can use only those individuals with observed earnings in two consecutive periods.

The estimation procedure is done in two steps in order to take into account arbitrary serial correlation. The correlation coefficient between ε_{it} and ε_{it-1} is

[5] Another approach is the semiparametric estimator proposed by Honoré (1992), which does not require the assumption that the error term is normally distributed. As a result, it is impossible to derive the marginal effect of the covariate on the observed dependent variable. Generally one cannot use parameter estimates from nonlinear fixed effects estimators to predict levels of the outcome variable. This is because it is not possible to impute the fixed effect from such models. However, nonlinear estimators can be used to infer signs of parameters or they can be thought of as a specification check of more efficient, but also more restrictive, random effects models. If there is no correlated heterogeneity then estimates from these models will be similar. The asymptotic distribution and estimator performance is discussed in Honoré (1992).

denoted by ρ. In the first step the maximum likelihood procedure is employed to estimate equation (6). The probability of observing non-top coded values for the dependent variable in periods "t" and "t-1" is given by:

$$\Pr\left(w_{it-1} < u, w_{it} < u\right) = 1 - \Phi_2\left(\frac{u - \pi_{t-1} - X_{i,t-1}\beta - \overline{X}_i\gamma}{\sigma_{t-1}}, \frac{u - \pi_t - x_{i,t}\beta - \overline{x}_i\gamma}{\sigma_t}, \rho\right) = \tag{7}$$
$$= 1 - F(c_i, X_i, \theta)$$

where $\theta_t = \left(\pi_{01}, \pi_{0T}, \beta, \gamma, \sigma_1, \ldots, \sigma_T, \rho_2, \ldots, \rho_T\right)$ and $\Phi_2(\ )$ denotes the cumulative bivariate standard normal distribution function. The density function of the observed values is:

$$f\left(\eta_i \middle| w_{it-1} < u, w_{it} < u\right) = \frac{1}{\sigma_\eta}\phi\left(\frac{\eta_i}{\sigma_\eta}\right) \times$$
$$\times \Phi\left(\min\left\{\frac{u - \pi_{0,t-1} - X_{it-1}\beta - \overline{X}_i\gamma - \frac{\sigma_{\eta t-1}}{\sigma^2_{\eta,t}}\eta_{it}}{\sqrt{\sigma^2_{t-1} - \sigma^2_{\eta t-1}\sigma^{-2}_{\eta,t}}}, \frac{u - \pi_{0,t} - X_{it}\beta - \overline{X}_i\gamma - \frac{\sigma_{\eta t}}{\sigma^2_{\eta,t}}\eta_{it}}{\sqrt{\sigma^2_t - \sigma^2_{\eta t}\sigma^{-2}_{\eta,t}}}\right\}\right) \tag{8}$$

The maximum likelihood estimate of $\theta_t = \left(\pi_{01}, \pi_{0T}, \beta, \gamma, \sigma_1, \ldots, \sigma_T, \rho_2, \ldots, \rho_T\right)$ is given by:

$$\hat{\theta}_t = \underset{\theta}{\arg\max} \sum_{i=1}^{N} \begin{bmatrix} \left(1 - I\left(w_{it-1} < u, w_{it} < u\right)\right) \ln\left(1 - F\left(c_i, x_i, \theta\right)\right) + \\ + I\left(w_{it-1} < u, w_{it} < u\right) \ln f\left(\eta_i \middle| w_{it-1} < u, w_{it} < u\right) \end{bmatrix} \tag{9}$$

Maximum likelihood is applied to each period (t=2,…,T) and next, as a second step, a Minimum Distance Estimator is employed to impose the restriction $\left(\beta_t = \beta, \gamma_t = \gamma, \forall t\right)$ using the optimal weighting matrix. Only the β parameters corresponding to the time varying regressors are identified. For the time constant regressors only $(\beta_t + \gamma_t)$ is identified. Once the parameters of interest are estimated, we obtain fitted values of the endogenous variable. The goal is the generation of the earnings of the workers in the sample. Two situations are considered:

a) The earnings are fully observed. In this case, the covered earnings are located between the thresholds $(w_t < u_t)$. The "observed earnings" are equal to the covered earnings, i.e. $w_t = c_t$.

b) w_t is not observed whenever it exceeds u_t. The imputation is exclusively needed for those observations such that $c_t = u_t$.

We focus on the imputation of w_{it}. After the estimation of the parameters we can impute the earnings in the case of $c_t = u_t$ or $c_t = l_t$. An easy way for the imputation of w_{it} is given by $\hat{w}_{it} = \hat{\pi} + \hat{\beta}X_{it} + \hat{\gamma}\overline{X}_i + \hat{\sigma}\lambda\left(\frac{u - \hat{\pi} - \hat{\beta}X_{it} - \hat{\gamma}\overline{X}_i}{\hat{\sigma}}\right) + \hat{\varepsilon}_{it}$, where $\lambda(.) = \phi(.) \big/ 1-\Phi(.)$, $\phi(.)$ and $\Phi(.)$ denote, respectively, the density and the distribution function of a standard normal; $\hat{\varepsilon}$ are random values from a normal distribution with zero mean and $\hat{\sigma}$ is the variance.

b) *Including constant regressors*

The main limitation of the procedure described above is that it uses the first difference of observations. If one of the explanatory variables is constant over time for each worker and period we have $\nabla x_i = x_{it} - x_{it-1} = 0$. For this methodological reason the previous approach discards the variable "cohort" in the analysis. If we want to evaluate the cohort effects on earnings patterns we need to work with the estimation procedure developed in Wooldridge (1995) and Zabel (1992), based on similar assumptions such as the Kalwij (2003) estimator, including the parameterisation of the covariates-random effects relationship as expressed in (4). The estimation of equation (2) is done in two steps, the first of which gives the maximum likelihood estimates for each period (Tobin, 1958):

$$\hat{\theta}_t = \underset{\theta}{\operatorname{argmax}} \sum_{i=1}^{N} \left[\begin{array}{l} (1 - I(w_{it} < u)) \ln\left(\left[1 - \Phi\left(\dfrac{u_{it} - \pi_i - X_{it}\beta - \overline{X}_i\gamma}{\sigma_t}\right)\right]\right) + \\ + I(w_{it-1} < u)\left[\dfrac{1}{\sigma_t}\phi\left(\dfrac{w_{it} - \pi_i - X_{it}\beta - \overline{X}_i\gamma}{\sigma_t}\right)\right] \end{array} \right] \tag{9.1}$$

Maximum likelihood is applied to each period (t=2,…,T) and next a Minimum Distance Estimator is employed to impose the restriction $(\beta_t = \beta, \gamma_t = \gamma, \forall t)$ using the optimal weighting matrix.

c) *Testing the uniform wage growth hypothesis*

In order to obtain an age-earning profile we need to elaborate further. First, we aim to check the assumption that pure age effects have a stable life-cycle pattern in different periods and for different cohorts. To that end, we explore whether wage trends are uniform across cohorts, by introducing testable separability conditions implying uniformity of wage trends.

Following Fitzenberger *et al*. (2001), we assume that wage growth can be characterised as the sum of a pure ageing effect and a pure time effect in the following way:

$$w_{it} = \alpha_i + \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 Age + \beta_4 Age^2 + u_{it} \tag{10}$$

Then the life-cycle wage growth is independent of the calendar year t. This is the so-called uniform wage growth hypothesis. It implies that each cohort faces the same wage growth over the life-cycle. Economy-wide shifts are common to all cohorts in the same year but they occur at different points during the life-cycle of each cohort. If the separability condition (10) holds, we can construct a "life-cycle wage profile" independently of the calendar year and a macroeconomic time trend independently of age. Finally, condition (10) is violated if "interaction terms" of period and age enter the equation of relevance. In order to test the uniform wage growth hypothesis, we consider the following interaction terms of age and time in the equation, that is:

$$\begin{aligned} w_{it} &= \alpha_i + \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 Age + \beta_4 Age^2 + \beta_5 (t \times Age) + \\ &+ \beta_6 \left(t \times Age^2\right) + \beta_7 \left(t^2 \times Age\right) + \beta_8 \left(t^2 \times Age^2\right) + \mu_{it} \end{aligned} \qquad (11)$$

Fitzenberger *et al.* (2001) proposed a formal test of the uniform growth. The null hypothesis requires that all coefficients of the interaction terms be jointly zero:

$$H_0 : \beta_5 = \beta_6 = \beta_7 = \beta_8 = 0 \qquad (12)$$

According to (12), only if the separability condition holds is it meaningful to construct a life-cycle wage profile as a function of pure ageing and a macroeconomic trend variable.

Second, we want to evaluate the effect of the birth cohort on earnings. In order to do so we incorporate two new variables in the regression equation:

$$w_{it} = \alpha_i + \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 Age + \beta_4 Age^2 + \beta_5 coho_{1i} + \beta_6 coho_{2i} + u_{it} \qquad (13)$$

where "coho" is a dummy variable with values: $coho_{1i} = \begin{cases} 1 \text{ if } 1958 \le \text{birth} < 1968 \\ 0 \text{ otherwise} \end{cases}$

and $coho_{2i} = \begin{cases} 1 \text{ if birth} \ge 1968 \\ 0 \text{ otherwise} \end{cases}$, so that the variables are equal to one when the individual belongs to the baby boom cohorts –those born between 1957 and 1977.

$$\begin{aligned} w_{it} &= \alpha_i + \beta_0 + \beta_1 t + \beta_2 t^2 + \beta_3 Age + \beta_4 Age^2 + \beta_5 coho_1 + \beta_6 coho_2 + \\ &+ \beta_7 (t \times Age) + \beta_8 \left(t \times Age^2\right) + \beta_9 \left(t^2 \times Age\right) + \beta_{10} \left(t^2 \times Age^2\right) + u_{it} \end{aligned} \qquad (14)$$

As before, the test of the uniform wage growth hypothesis requires that all coefficients of the interaction terms be jointly zero:

$$H_0 : \beta_7 = \beta_8 = \beta_9 = \beta_{10} = 0 \qquad (15)$$

Besides cohort, the estimation considers some other variables that hold constant over the time period covered by data. Hence the data set is split in order to be able to apply the second method according to gender and 12 contribution groups –the latter is also needed in order to consider differential contribution thresholds.

4. MAIN RESULTS

Results are presented as follows. First, Table 1 presents the results of estimating equation (11) according to (8) for each gender and contribution group (CG). In Table 2 we present the results of estimating equation (14) according to (9.1). Wages are expressed in hourly wages at 1983 constant prices. We focus on national workers, because immigration in Spain is too recent to capture differential age effects.

In Tables 1 and 2 one observes the following. First, the coefficients associated with age and its square are significant, showing that the effect of age is positive and decreasing. The latter is not so for females. The coefficient for age square is only significant for some contribution groups. As we will see later, the low level of female labour participation over the period analysed does not allow clear patterns to be observed for female lifetime earnings. The period effect is almost always significant but small, while its second order effects tend to be non-significant. In any case it is important to bear in mind that period, age and cohort are mixed due to the linear relationship linking them. The goal of the estimation is to predict wage –provided that the uniform wage growth hypothesis holds– more that fully identifying relationships among variables. The interaction terms combining age and period are not significant, showing that the separable specification holds and hence we can accept the uniform wage growth hypothesis. Together with the coefficients the result for the chi test clearly accepts the null hypothesis.

In Table 2 we can also observe the cohort effect of being a baby boomer having a negative and significant effect in some contribution groups. The sign tends to be negative, being higher for early baby boomers –those born between 1957 and 1967. The significant cases are concentrated in this cohort and in males. The failure to be significant might be due to the fact that baby boomers are a substantial share of the individuals observed in the sample and for some ages there are scarce empirical counterparts. Baby boomers were born between 1957 and 1977, they are from 4 to 24 years old in 1981 and from 24 to 48 in 2005.

Table 1

ESTIMATION RESULTS OF EQUATION (12) BY GENDER AND CONTRIBUTION GROUP (CG)

MALE												
	CG1	**CG2**	**CG3**	**CG4**	**CG5**	**CG6**	**CG7**	**CG8**	**CG9**	**CG10**	**CG11**	**CG12**
CONST	0.0161	0.9819	1.5068	1.1661	0.0161	0.9593	0.1020	0.0048	0.4407	0.2391	0.5020	1.0214
	(3.0184)	(2.5798)	(2.3526)	(2.0119)	(1.2319)	(2.8271)	(0.9303)	(2.9277)	(1.1221)	(1.5688)	(0.9391)	(1.7636)
Year	0.0026	0.0032	0.0133	0.0114	0.0217	0.0462	0.0122	0.0682	0.0082	0.0234	0.0117	0.0105
	(13.2146)	(17.0895)	(22.6869)	(18.3931)	(12.279)	(4.2117)	(10.4171)	(8.2115)	(6.1943)	(8.8670)	(3.6751)	(3.2662)
Year2	0.0050	-0.0014	0.0051	-0.0006	0.0027	0.1703	0.0038	-0.0011	-0.0040	-0.0013	0.0006	-0.0003
	(0.7957)	(-0.4534)	(4.7009)	(-4.7153)	(1.6738)	(0.5598)	(1.9831)	(-0.086)	(-2.266)	(-3.042)	(9.1773)	(-0.009)
Age	0.0021	0.1226	0.0019	0.0074	0.0137	0.0262	0.0003	0.0001	0.0042	0.0266	0.0077	0.0066
	(12.1378)	(7.6827)	(13.0326)	(14.4091)	(10.721)	(3.3238)	(9.5715)	(9.4921)	(7.2462)	(7.0110)	(2.1130)	(1.6424)
Age2	-0.0045	-0.0836	-0.0009	0.0035	0.0025	0.2779	-0.0002	0.0887	0.0041	-0.0022	-0.0007	-0.0029
	(-2.8701)	(-2.0053)	(-6.7784)	(6.3367)	(1.9282)	(0.3430)	(-3.3559)	(0.1123)	(2.0129)	(-2.805)	(-1.489)	(-0.549)
Year·Age	0.0006	0.0068	-0.0018	-0.0018	-0.0015	-0.0040	-0.0029	0.0002	-0.0020	0.0025	0.0050	0.0002
	(1.2057)	(1.0610)	(-2.1016)	(-1.4959)	(-1.745)	(-1.356)	(-1.0279)	(0.5158)	(-0.871)	(1.8382)	(0.8894)	(0.4957)
Year2·Age	0.0046	0.0010	0.0003	-0.0044	0.0041	0.0021	0.0046	0.0001	0.0020	-0.0033	0.0013	0.0031
	(0.2471)	(0.2481)	(0.4862)	(-0.3179)	(0.6403)	(0.0694)	(0.3361)	(0.1917)	(0.2199)	(-0.503)	(0.4274)	(0.0344)
Year·Age2	0.0002	-0.0048	-0.0007	0.0008	0.0066	-0.0641	0.0054	0.0004	0.0003	0.0013	0.0025	0.0051
	(0.3214)	(-0.2535)	(-0.3972)	(0.7017)	(0.3937)	(-0.185)	(0.1656)	(0.1295)	(0.1659)	(0.1234)	(0.7282)	(0.2550)
Year2·Age2	0.0080	0.0021	0.0091	0.0064	0.0162	-0.0006	0.0009	-0.0037	0.0099	0.0229	0.0165	-0.0401
	(0.0684)	(0.0709)	(0.0934)	(0.1689)	(0.1687)	(-0.192)	(0.2188)	(-0.173)	(0.2115)	(0.2064)	(0.2154)	(-0.075)
chi test	2.3160	4.1314	5.4831	3.6751	5.2605	3.9480	4.0801	2.7523	2.8555	5.3777	4.1618	2.7229

(Follow)

(Continuation)

	FEMALE											
	CG1	**CG2**	**CG3**	**CG4**	**CG5**	**CG6**	**CG7**	**CG8**	**CG9**	**CG10**	**CG11**	**CG12**
CONST	0.0166	0.3568	0.2256	0.0152	0.8353	0.0096	0.0716	0.0210	0.0057	0.1813	0.7490	0.9305
	(2.0374)	(1.2307)	(1.9744)	(1.3661)	(2.5379)	(1.6637)	(1.3078)	(4.9748)	(1.5146)	(1.4574)	(1.0580)	(1.9783)
Year	0.0069	0.0044	0.0053	0.0546	0.0899	0.0049	0.0160	0.0015	0.0005	0.0109	0.0021	0.0079
	(2.1581)	(2.1838)	(2.8205)	(3.7992)	(2.6016)	(1.6387)	(1.3200)	(2.6657)	(4.3353)	(7.0918)	(4.7319)	(1.0632)
Year2	-0.0027	-0.0042	-0.0019	0.0012	-0.0056	-0.0008	-0.0018	-0.0028	0.0005	0.0333	-0.0016	0.0051
	(-2.4333)	(-1.6873)	(-0.9228)	(0.6683)	(-0.344)	(-1.663)	(-0.3162)	(-0.372)	(0.0288)	(1.8352)	(-0.825)	(0.0013)
Age	0.0066	0.0106	0.0059	0.0063	0.0555	0.0047	0.0076	0.0023	0.0036	0.0029	0.0060	0.0141
	(4.1894)	(2.0494)	(4.0763)	(1.9865)	(2.8822)	(1.7242)	(2.3838)	(1.9583)	(4.2724)	(6.8982)	(3.6661)	(1.1118)
Age2	-0.0030	0.0024	-0.0030	-0.0005	-0.0032	0.0028	-0.0019	0.0039	-0.0053	-0.0010	-0.0054	-0.0013
	(-0.3078)	(0.1604)	(-0.3983)	(-2.1047)	(-0.043)	(0.0584)	(-0.3636)	(0.518)	(-0.117)	(-0.171)	(-1.152)	(-0.405)
Year·Age	0.0025	0.0019	0.0001	0.0004	0.0037	0.0011	0.0015	0.0049	-0.0587	0.0116	0.0031	0.0007
	(0.7497)	(1.4459)	(1.4277)	(1.3820)	(0.7536)	(0.1226)	(0.9871)	(1.0902)	(-0.043)	(0.0909)	(3.2957)	(0.6592)
Year2·Age	-0.0030	0.0034	-0.0025	0.0029	-0.0060	-0.0032	-0.0007	0.0040	0.0239	-0.0040	-0.0045	-0.0024
	(-0.1137)	(0.2043)	(-0.1639)	(0.1311)	(-0.094)	(-0.015)	(-0.1135)	(0.0412)	(0.1829)	(-0.186)	(-0.026)	(-0.248)
Year·Age2	-0.0004	-0.0022	0.0058	-0.0002	0.0095	0.0039	0.0035	0.0055	0.0168	-0.0037	-0.0028	-0.0029
	(-0.2660)	(-0.0659)	(0.0888)	(-0.2178)	(0.1929)	(0.1954)	(0.0504)	(0.1066)	(0.1607)	(-0.018)	(-0.064)	(-0.015)
Year2·Age2	0.0029	0.0003	0.0055	0.0322	-0.0003	0.0040	-0.0086	-0.0081	0.0050	0.0034	0.0134	0.0159
	(0.1496)	(0.1102)	(0.0181)	(0.2279)	(-0.109)	(0.1053)	(-0.0488)	(-0.079)	(0.0255)	(0.0864)	(0.1356)	(0.0353)
chi test	3.2865	3.7463	2.3299	3.0350	3.0210	2.7660	2.6393	2.9665	3.4674	3.4591	4.9649	1.9816

Source: Own elaboration.

Note: t-statistics in brackets. For the sake of clarity we omit results regarding γ in equation (5).

Table 2

ESTIMATION RESULTS OF EQUATION (15) BY GENDER AND CONTRIBUTION GROUP (CG)

MALE												
	CG1	**CG2**	**CG3**	**CG4**	**CG5**	**CG6**	**CG7**	**CG8**	**CG9**	**CG10**	**CG11**	**CG12**
CONST	0.0151	0.9860	1.5074	1.1673	0.0164	0.9617	0.0960	0.0095	0.4453	0.2419	0.4998	1.0208
	(2.9488)	(2.2745)	(2.7899)	(1.8024)	(1.1481)	(3.0613)	(1.0137)	(2.6222)	(0.8809)	(1.7410)	(1.3088)	(2.0692)
$Coho_1$	-0.0630	-0.0637	-0.0005	-0.0266	-0.0033	-0.0157	0.0442	-0.0071	0.0031	-0.0160	0.0455	0.0216
	(-5.225)	(-3.333)	(-1.654)	(-2.337)	(-1.887)	(2.0054)	(1.2299)	(-6.333)	(1.5569)	(-2.665)	(1.6696)	(1.8895)
$Coho_2$	0.0004	-0.0201	0.0902	-0.0039	-0.0023	-0.0033	0.0020	-0.1024	-0.0014	-0.0158	0.0042	0.0197
	(0.0025)	(-1.926)	(0.8746)	(-1.966)	(-0.459)	(1.1126)	(0.9987)	(-1.888)	(-1.453)	(-2.000)	(0.2254)	(1.4885)
Year	0.0013	0.0022	0.0168	0.0108	0.0169	0.0424	0.0127	0.0702	0.0005	0.0228	0.0138	-0.0079
	(12.8600)	(17.3623)	(23.0726)	(18.1892)	(11.8742)	(3.7770)	(10.7420)	(8.4933)	(6.4511)	(9.1631)	(3.5811)	(-3.274)
$Year^2$	0.0005	-0.0039	0.0033	-0.0028	0.0044	0.1626	0.0016	-0.0002	-0.0032	-0.0010	0.0011	-0.0049
	(0.7876)	(-0.2751)	(4.7497)	(-4.7747)	(1.7120)	(0.5618)	(1.7936)	(-0.023)	(-2.210)	(-2.913)	(9.4337)	(-0.018)
Age	0.0036	0.1208	0.0007	0.0007	0.0130	0.0265	0.0018	0.0073	0.0018	0.0285	0.0049	0.0122
	(12.2084)	(7.8700)	(12.7125)	(14.4839)	(10.5188)	(3.3242)	(9.5888)	(9.5733)	(7.4538)	(7.0246)	(1.9250)	(1.7229)
Age^2	-0.0019	-0.0875	-0.0028	0.0032	0.0000	0.2773	-0.0012	0.0876	0.0024	-0.0049	-0.0008	-0.0006
	(-2.8134)	(-2.3348)	(-7.0868)	(6.1923)	(1.9358)	(0.4589)	(-3.2116)	(0.0204)	(2.1515)	(-2.639)	(-1.567)	(-0.246)
$Year \cdot Age$	0.0074	0.0092	-0.0030	-0.0029	-0.0007	-0.0008	-0.0022	0.0028	-0.0004	0.0048	0.0020	0.0039
	(1.1386)	(1.2841)	(-2.0776)	(-1.4075)	(-1.7448)	(-1.668)	(-1.2567)	(0.6199)	(-1.088)	(1.9603)	(0.8627)	(0.5482)
$Year^2 \cdot Age$	0.0000	0.0040	0.0020	-0.0043	0.0033	0.0009	0.0073	0.0045	0.0010	-0.0028	0.0014	0.0010
	(0.0545)	(0.0621)	(0.5681)	(-0.1760)	(0.2871)	(0.1377)	(0.1506)	(0.2126)	(0.1286)	(-0.166)	(0.4380)	(0.1361)
$Year \cdot Age^2$	-0.0033	-0.0028	-0.0026	-0.0011	0.0002	-0.0580	0.0017	0.0025	0.0019	0.0007	0.0014	0.0058
	(-0.1280)	(-0.0097)	(-0.7766)	(0.6326)	(0.0402)	(-0.093)	(0.4594)	(0.1974)	(0.3507)	(0.4463)	(1.0932)	(0.1869)
$Year^2 \cdot Age^2$	0.0007	0.0043	0.0065	0.0100	0.0136	-0.0004	0.0055	-0.0044	0.0090	0.0150	0.0233	-0.0416
	(0.0747)	(0.1661)	(0.2461)	(-0.0859)	(0.1636)	(-0.010)	(0.1588)	(-0.107)	(0.0943)	(-0.105)	(0.0571)	(-0.237)
chi test	2.6016	3.8823	5.5372	3.9872	5.5210	3.4841	3.7230	2.6406	3.1560	5.1692	3.7835	3.0171

(Follow)

(Continuation)

	FEMALE											
	CG1	**CG2**	**CG3**	**CG4**	**CG5**	**CG6**	**CG7**	**CG8**	**CG9**	**CG10**	**CG11**	**CG12**
CONST	0.0194	0.3480	0.2218	0.0167	0.8420	0.0057	0.0789	0.0232	0.0036	0.1833	0.7498	0.9285
	(1.7182)	(0.9325)	(2.3338)	(1.3479)	(2.5709)	(2.0424)	(1.2765)	(5.4082)	(1.4297)	(1.5693)	(1.2684)	(1.9615)
$Coho_1$	-0.0016	0.0020	0.0001	-0.0076	-0.0020	-0.0014	-0.0001	0.0093	-0.0036	-0.0012	0.0008	0.0005
	(-1.0300)	(0.3366)	(0.7862)	(-1.3566)	(-1.6659)	(-2.646)	(-0.4613)	(0.7895)	(-1.988)	(-1.232)	(0.4896)	(0.9876)
$Coho_2$	0.0062	-0.0038	-0.0023	-0.0137	-0.0031	-0.0081	-0.0020	-0.0001	0.0059	0.0005	-0.0023	0.0008
	(0.3699)	(-0.8745)	(-1.8556)	(-1.9322)	(-1.9879)	(-1.468)	(-1.2565)	(-0.565)	(0.9568)	(0.3547)	(0.8975)	(1.0235)
Year	0.0092	0.0039	0.0074	0.0535	0.0908	0.0047	0.0165	0.0019	0.0081	0.0123	0.0040	0.0100
	(1.8008)	(1.9666)	(2.9308)	(3.9678)	(2.8740)	(1.3534)	(1.3334)	(2.7686)	(3.9937)	(6.8515)	(4.6626)	(0.7643)
$Year^2$	-0.0047	-0.0001	-0.0035	0.0031	-0.0032	-0.0028	-0.0055	-0.0020	0.0044	0.0325	-0.0004	0.0030
	(-2.1136)	(-1.7956)	(-0.4640)	(0.9149)	(-0.2747)	(-1.591)	(-0.1689)	(-0.435)	(-0.088)	(2.0398)	(-0.951)	(0.3225)
Age	0.0032	0.0189	0.0034	0.0038	0.0535	0.0045	0.0080	0.0024	0.0009	0.0014	0.0078	0.0118
	(4.4833)	(2.0025)	(3.8076)	(2.1471)	(2.7996)	(1.2504)	(2.4159)	(2.3269)	(4.2434)	(6.9916)	(3.9755)	(1.4941)
Age^2	-0.0005	0.0054	-0.0031	-0.0031	-0.0008	0.0048	-0.0042	0.0002	-0.0022	-0.0002	-0.0037	-0.0017
	(0.2044)	(0.1753)	(-0.0786)	(-2.3882)	(-0.0987)	(0.1120)	(-0.4736)	(0.3030)	(-0.235)	(-0.126)	(-1.489)	(-0.723)
$Year \cdot Age$	0.0050	0.0029	0.0029	0.0038	0.0032	0.0035	0.0044	0.0040	-0.0647	0.0082	0.0024	0.0029
	(1.1410)	(1.0179)	(1.6425)	(1.4471)	(0.7220)	(0.1708)	(0.7135)	(1.1915)	(-0.113)	(0.1072)	(3.7319)	(0.2958)
$Year^2 \cdot Age$	-0.0030	0.0005	-0.0052	0.0027	-0.0051	-0.0011	0.0021	0.0050	0.0238	-0.0013	-0.0037	-0.0046
	(-0.1236)	(0.0270)	(-0.1634)	(0.2264)	(-0.0622)	(-0.224)	(-0.1255)	(0.0504)	(0.1204)	(-0.1091	(-0.006)	(-0.083)
$Year \cdot Age^2$	-0.0037	-0.0021	0.0059	-0.0018	0.0042	0.0038	0.0059	0.0057	0.0188	-0.0012	-0.0009	-0.0043
	(-0.1610)	(-0.1682)	(0.1024)	(-0.2959)	(0.1820)	(0.0308)	(0.1413)	(0.2471)	(0.1538)	(-0.058)	(-0.059)	(-0.086)
$Year^2 \cdot Age^2$	0.0020	0.0049	0.0016	0.0324	-0.0010	0.0026	-0.0118	-0.0087	0.0038	0.0047	0.0160	0.0246
	(-0.0213)	(0.1426)	(0.1092)	(0.0112)	(-0.0355)	(0.0825)	(-0.0133)	(-0.241)	(0.2019)	(0.1509)	(0.0561)	(0.1567)
chi test	2.8838	2.8303	2.2178	3.1986	2.7847	2.7986	2.3803	3.0793	3.9046	2.8615	5.1714	1.8361

Source: Own elaboration.

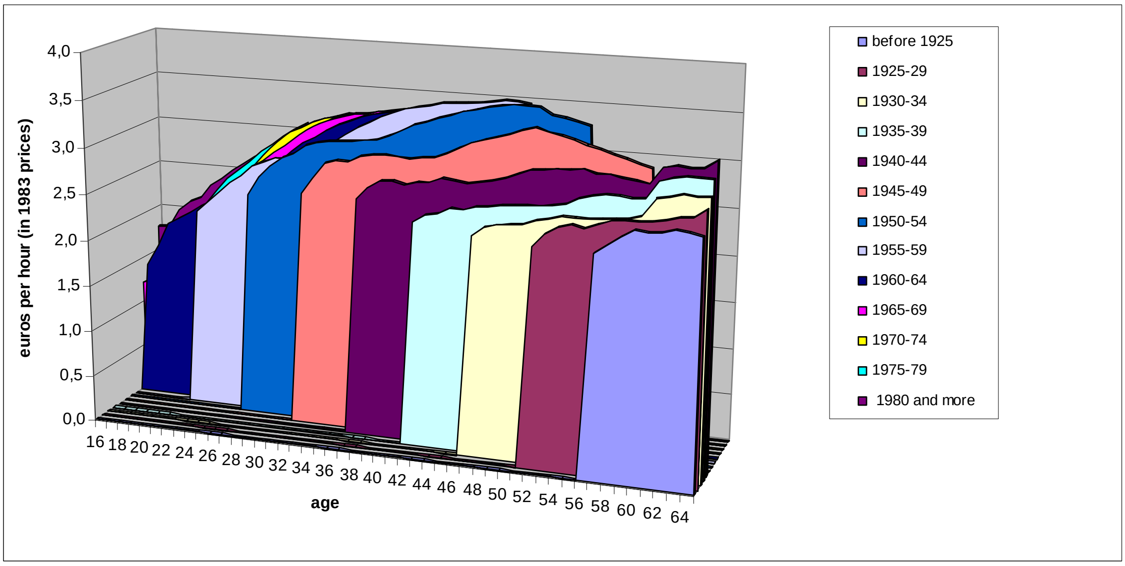
Note: t-statistics in brackets. For the sake of clarity we omit results regarding γ in equation (5).

We show the estimated earnings profiles below. First, the longitudinal cohort profiles for the cohorts observed in our data sample are shown in Figure 1, which plots the resulting average real lifetime earnings profiles for the observed cohorts (defined as five-group birth years). We can observe some expected trends. First, the lifetime profiles show the expected shape: wage increases with age though the rate of increase tends to stabilize or decrease at some point in middle age. Second, the successive cohorts improve their performance, showing increases in productivity. Interestingly, females experience greater and constant improvements, while for males those are concentrated for some cohorts. As a consequence, the gender gap tends to decrease in successive cohorts.

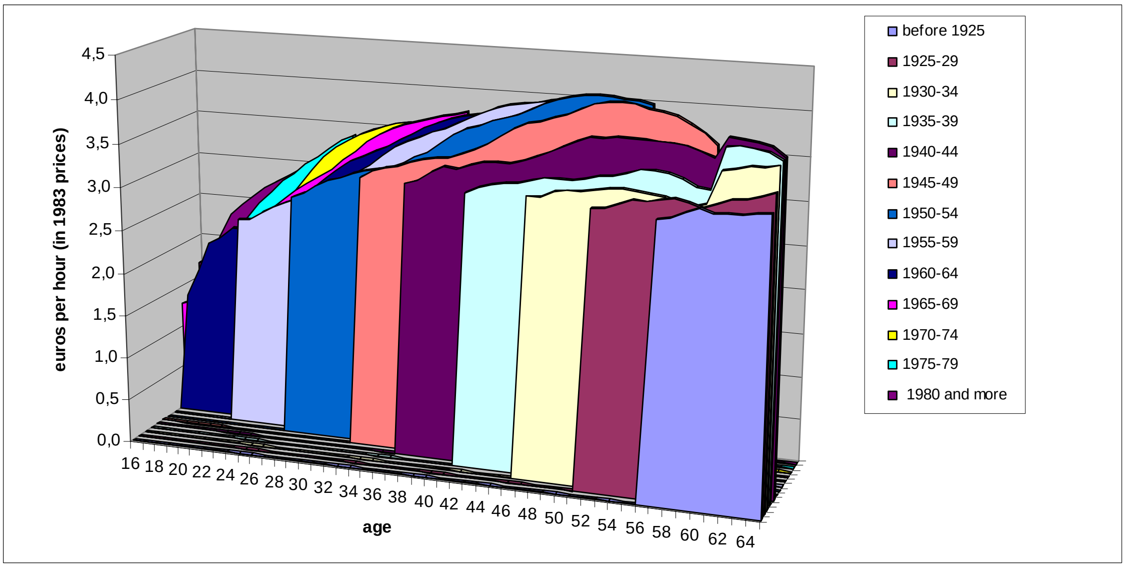
Figure 1

LONGITUDINAL EARNING PROFILES BY COHORT

a) Male



b) Female



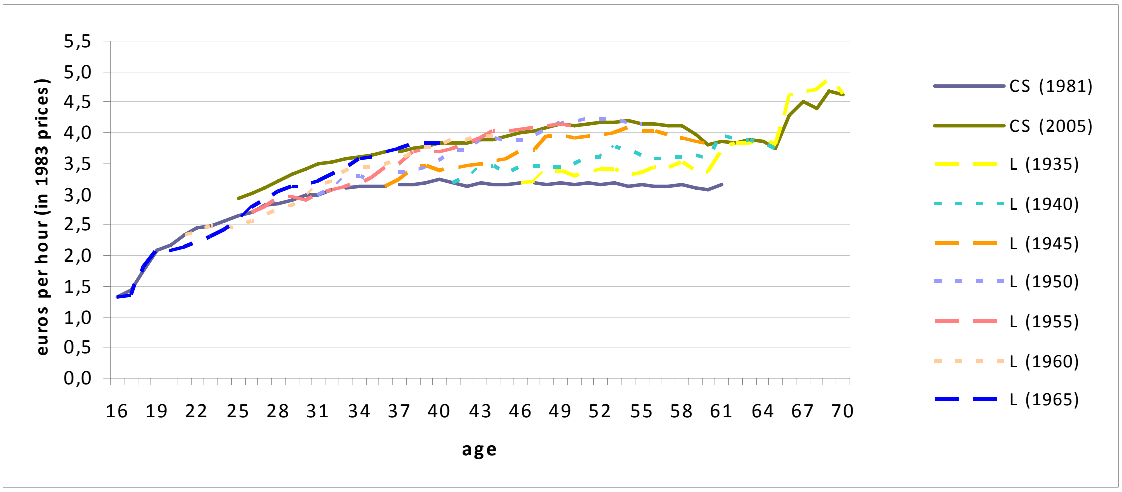
Source: Own elaboration.

Second, the longitudinal profiles of succeeding cohorts are compared to the cross-sectional earning profiles in Figure 2. Clearly, longitudinal earning profiles show a steeper shape and latter decline than in the cross section[6]. Furthermore, some cohort effects are apparent as the longitudinal profiles overcome the cross-section that should be enveloping them. As in Figure 1, one can observe that females experience stronger cohort differences than males.

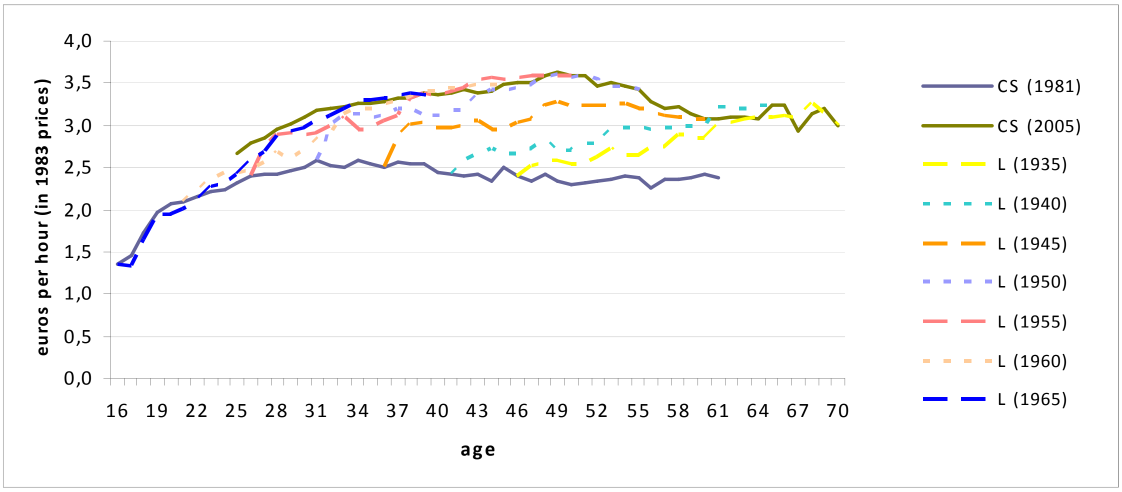
Figure 2

LONGITUDINAL (L) EARNING PROFILES VS CROSS-SECTION (CS) EARNING PROFILES

a) Male



b) Female



Source: Own elaboration.

Finally, Figure 3 tries to isolate in a more detailed way the pure age effects by plotting the deviation of the growth rate of wage along the life cycle with respect to the average. First, by subtracting the annual average growth, the

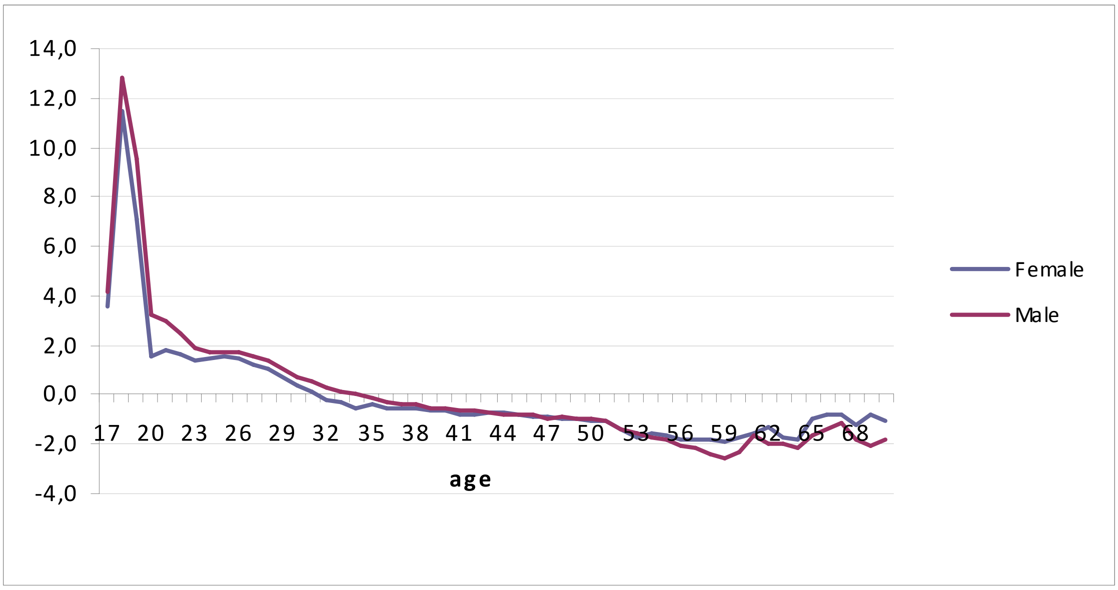
6 The irregular path from age 65 to 70 might be reflecting a selection bias.

period effect is somehow corrected. Figure 4 shows the value of these series of average annual growth rate for females and males, detailing the latter by contribution group. Clearly those annual growth rates show a strong cyclical variation. Second, by taking the average, the cohort effects tend to be smoothed[7]. The resulting age effects might give interesting insights with respect to future long-term prospects of productivity growth in the face of demographic ageing.

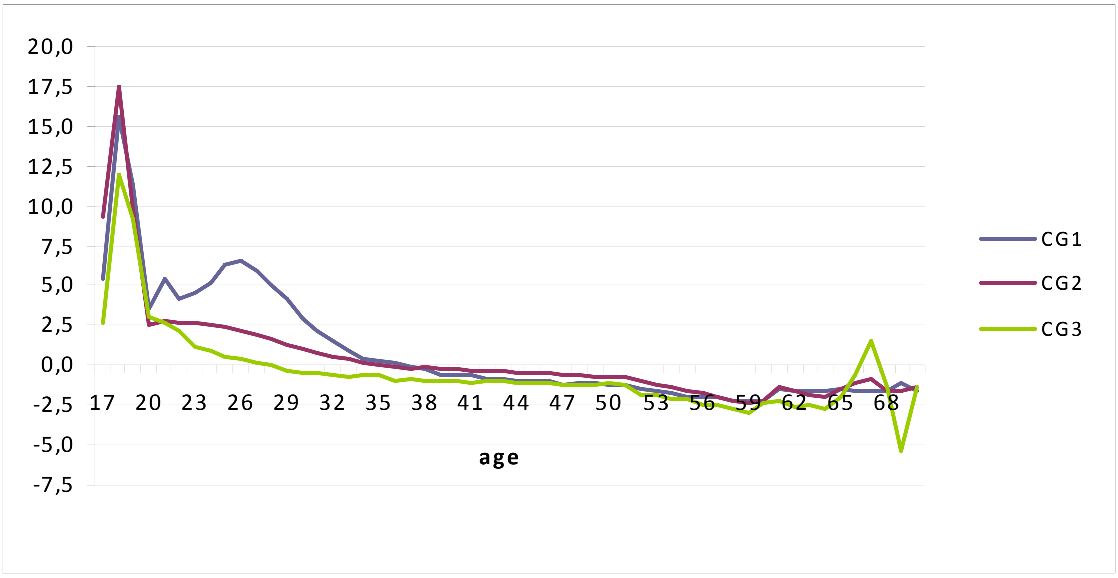
Figure 3

DEVIATION OF EARNINGS GROWTH RATE FROM AVERAGE

a) By gender



b) Males by contribution group (CG)



Source: Own elaboration.

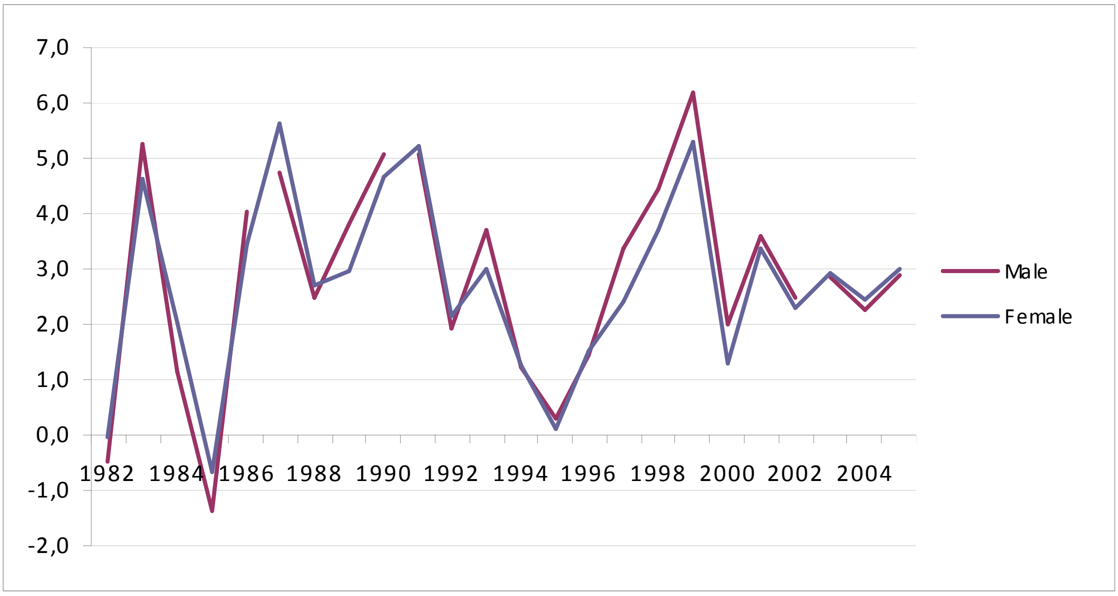
Note: CG1 from 1 to 3; CG2 from 4 to 8 (plus 13 with similar contribution bases); CG3 from 9 to 11.

[7] Note that the average is the average of the subgroup defined by age and gender (see Figure 3).

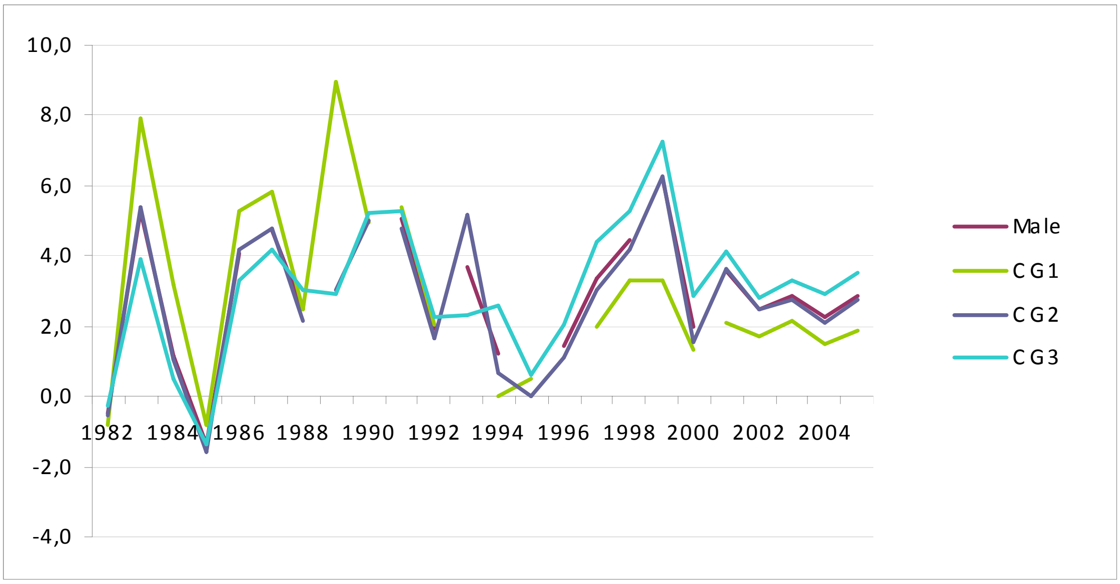
Figure 4

ANNUAL GROWTH RATES OF WAGES BY GENDER AND CONTRIBUTION GROUP

a) By gender



b) Males by contribution group (CG)



Source: Own elaboration.

Note: CG1 (from 1 to 3); CG2 from 4 to 8 (plus 13 with similar contribution bases); CG3 from 9 to 11.

Panel *a*) in Figure 3 compares the average growth rates of wage for males to those for females. This "net" wage growth shows the expected pattern, as long as it tends to be positive and high at the beginning and decrease along the life cycle, becoming negative at the end. The increase at the end of the working life –showing a considerably erratic behaviour– might reflect a selection bias. Females show a similar pattern, though the growth rate is lower except after the childrearing period –from age 50 on[8]. Clearly, the low female participation rates and the changes occurring in the last decades do not permit clear conclusions to be established on their future wage trends without further investigation.

Panel *b*) in Figure 3 focuses on males, comparing the net growth rate over the life cycle of the different contribution groups. The expected life cycle evolution is again observed. Regarding contribution group we find, to some extent, the expected patterns as the growth rate starts being higher, the higher the qualification level. The life cycle wage growth is more stable for the medium contribution groups and always lower for the low contribution group. The growth rate for more educated workers is higher once they enter the labour market –later than the rest of groups– and grows for longer as could be expected. Nevertheless, their wage growth falls below that of the second group for those older than 38. In order to interpret this result one can observe the average growth rate by year and contribution group in Figure 3.b). Quite surprisingly, along the second half of the observed period the growth rate of highly educated male's wages falls below the rest of groups, while the opposite happens for the low educated. As said above Beaudry and Green (2000) find a similar cohort effect in Canada. This effect is probably not totally smoothed in Figure 3, by taking the deviation from the average wage growth of wages, due to the lack of empirical counterpart of the baby boom generation.

5. MAIN CONCLUSIONS

In this paper we analysed the extent to which a new longitudinal data set allows us to obtain lifetime earnings profiles for Spanish workers over the period 1981-2005. In order to mitigate the problem of censored data, we have obtained uncensored wages using two alternative panel data estimator approaches suitable for this purpose.

Although it is not possible to disentangle the age, period and cohort effects, the hypothesis that period and age effects are separable is not rejected in the data set and some evidence of cohort effects is also obtained. As a result, the age effects can be investigated by computing the average growth rate of wage over the lifecycle, smoothing the period and cohort effects. First, the observation of the growth rate by age shows the expected pattern: the growth rate of wages tends to be higher at the beginning of the working life and lower

[8] The relatively high growth rate in the last years before retirement might be reflecting a relatively stronger selection bias. A special feature of the Spanish labour market is that females are at the moment retiring later than men, probably trying to reach the minimum eligibility requirements.

afterwards. The estimated age effects might be useful in deriving future expected trends in labour productivity growth in the face of population ageing. This is especially true for males, as the evolution of female wages is more subject to uncertainty. In fact, males experience higher growth rates than females from the beginning of the working life until the end of the childrearing period. Nevertheless, changing patterns of female participation worsen the possibility of deriving clear conclusions about future female wage trends.

Second, cohort effects seem to be important and to affect some contribution groups differently, but the extension of the panel does not yet allow complete testing of the wage losses experienced by the baby boom generations with a scarce empirical counterpart.

Finally, the more educated males show a relatively higher growth rate with age for the first two decades, once they enter the labour market with some delay. Their growth rate falls below the average group later on. Interestingly, the growth rate of the lower contribution group is always below the rest, while the growth rate of the medium and top groups cross around age 40, probably due to a negative cohort effect experienced by highly educated baby boomers from 1990 on.

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SÍNTESIS

PRINCIPALES IMPLICACIONES DE POLÍTICA ECONÓMICA

En este trabajo se utiliza la Muestra Continua de Vidas Laborales (MCVL) para obtener perfiles longitudinales de salarios para España. Existe abundante literatura, tanto teórica como empírica, acerca de cómo evoluciona el salario de los individuos a lo largo del ciclo vital. Los primeros trabajos obtenían una forma de U invertida, mostrando una caída del salario al final de la vida laboral. Sin embargo pronto se observó que esa caída podía deberse al empleo de datos de corte transversal –es decir referidos a un mismo año–, que recogían, además de las diferencias de edad los llamados efectos cohorte –diferencias entre grupos de edad debidas a otros factores como la educación, cultura, etc. Si esas diferencias tienen un efecto positivo sobre el salario de las sucesivas cohortes cabe esperar que el perfil de salarios de ciclo vital se despegue del de corte transversal, creciendo más y no llegando a decrecer, quizá, al final de la vida laboral. Pero en realidad los datos adecuados para el análisis de los perfiles de salarios son los datos de panel que recogen a los mismos individuos –panel puro– o individuos similares –pseudo panel– a lo largo de varios períodos.

El trabajo empírico con datos de carácter longitudinal es relativamente reciente, especialmente en España, debido a la escasez de datos de este tipo. Hasta la publicación de la MCVL las únicas fuentes de datos longitudinales que contenían información de salarios eran el panel de IRPF y el panel de hogares de la Unión Europea. La MCVL ofrece nuevas posibilidades al respecto ya que puede incluir datos de base de cotización –en la mayoría de casos igual al salario– de hasta 25 años atrás hasta la fecha de extracción. En concreto, para este estudio se han utilizado los datos de bases de cotización correspondientes al período 1981-2005.

Los resultados de este trabajo permiten extraer interesantes conclusiones. En primer lugar se comparan los perfiles de corte transversal correspondientes a los años inicial y final (1981 y 2005) con los perfiles longitudinales de algunas cohortes. Así puede observarse cómo los perfiles de corte transversal presentan la tradicional forma de U invertida, si bien la caída final es leve. En cambio, el perfil longitudinal se despega del de corte transversal, creciendo por encima, y hay indicios de un crecimiento del salario real con la edad mayor en las primeras décadas y menor al final de la vida laboral. En segundo lugar, el análisis de los resultados por sexo revela un mayor crecimiento salarial para las mujeres –es decir, una reducción de gap salarial–, probablemente debido a efectos cohorte más marcados por los fuertes cambios que ha experimentado la participación laboral femenina en el período de análisis. Finalmente, se realiza también un análisis según el nivel de cualificación de los trabajadores, limitado en este caso a los varones ya que para las mujeres se dispone de un número insuficiente de observaciones. Como cabe esperar, se encuentra que a todas las edades la tasa de crecimiento del salario es mayor cuanto mayor es también el nivel de cualificación. Sin embargo esa pauta se invierte para los más cualificados a partir de los 38 años, ya que su tasa de crecimiento pasa a ocupar el segundo puesto. La explicación de este fenómeno parece ser un efecto cohorte negativo que afecta a los trabajadores más cualificados de la generación del baby boom. El hecho de ser una generación relativamente numerosa, combinado con un cierto nivel de sobre-educación parece dar lugar a un efecto negativo en los salarios, observado también en otros países, como Canadá.

El análisis realizado tiene interés en sí mismo, para la teoría del capital humano, pero tiene también implicaciones variadas en otras disciplinas. En concreto la forma del perfil salarial de ciclo vital es relevante de cara al análisis de la sostenibilidad del estado del bienestar ante el envejecimiento de la población. Es conocido el hecho de que el futuro aumento esperado de la tasa de dependencia demográfica se traducirá, de cara a la política pública, en una aumento de magnitud similar del ratio entre perceptores de transferencias (en especie o en metálico) y pagadores de impuestos. Indiscutiblemente, en este contexto cobra relevancia el análisis de la relación entre edad y salario –la base imponible de una parte considerable de los impuestos. De hecho, los estudios que evalúan la sostenibilidad del estado del bienestar ante el envejecimiento demográfico requieren la adopción de supuestos respecto a la evolución futura de la productividad con la edad.

Asimismo, la forma del perfil salarial tiene consecuencias sobre la magnitud del gasto en pensiones, el gasto social más voluminoso y más afectado por el envejecimiento. En concreto, la fórmula que determina la pensión de entrada al sistema depende de dos factores relacionados con el nivel de cotizaciones y del número de años cotizados. El primer factor se concreta en la base reguladora que, en su forma actual, promedia las bases de cotización de los últimos quince años. El segundo factor define el porcentaje de base reguladora que se recibirá como pensión en función de los años cotizados. Claramente si el salario real crece con la edad, emplear bases de cotización pasadas en el cálculo de la base reguladora tiene un efecto negativo sobre el derecho a pensión. Si bien la fórmula de la pensión incluye la actualizando al IPC correspondiente de todas las bases con más de dos años de antigüedad, esa corrección no permite la recuperación del crecimiento real. Sólo aquellos trabajadores que experimentan shocks negativos –reducción del salario o períodos de desempleo– al final de su carrera laboral ven su derecho a pensión mejorado por recurrir a bases pasadas.

Entre las recomendaciones del Pacto de Toledo se incluye el empleo de las bases de cotización correspondientes a toda la vida laboral para calcular la base reguladora. Ciertamente esa medida refuerza el carácter contributivo del sistema –la correspondencia entre cotizaciones y prestaciones– pero produce también un efecto reductor del derecho a pensiones. En el momento de creación del sistema español de Seguridad Social, en la base reguladora se consideraban únicamente las bases de cotización de los dos últimos años. Posteriormente se introdujo el cambio de 2 a 8 años –Ley 26/1985– y, finalmente, a instancias del Pacto de Toledo, la Ley 24/1997 elevó el número de años considerados hasta 15. Algunos estudios han tratado de analizar el efecto de esa medida. Dado el mayor aumento del salario en los primeros años de vida laboral, cabe esperar que la extensión de esta medida más allá de los 15 años –en la línea de las recomendaciones del Pacto de Toledo– produzca recortes mayores de la pensión de entrada.

Finalmente otro resultado a resaltar es que esta medida tiene un efecto redistributivo menos obvio y probablemente inesperado, en relación al tema que no ocupa –la forma del perfil de ciclo vital. Dado que los trabajadores más cualificados suelen tener tasas de crecimiento salariales mayores, el hecho de recurrir a bases de cotización pasadas, les produce mayores recortes en la pensión inicial.

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