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Sheepskin Effects in the Spanish Labour Market: A Public–Private Sector Analysis

EMPAR PONS and JUAN M. BLANCO

University of Valencia, Spain

ABSTRACT *The aim of this paper is to contrast the nature of the effect of education, Human Capital or Screening, in the Spanish labour market. We use the Hungerford and Solon methodology to distinguish between the returns to schooling from mere years of schooling as a reflection of their productive–enhancing contribution (human capital) and the returns to schooling from academic certificates as signals of the individual's ability (sheepskin effects). We separate our data into public and private sector workers. In the public sector the institutional restriction in the access and in the wage settings might force certificate rewards. Those not necessarily should be interpreted as sheepskin effects but as a consequence of educational requirements at initial hiring. Higher flexibility in the private sector would allow employers to use certificates as signals of greater productivity. Our estimations show little evidence of sheepskin effects in the private sector concentrated in upper secondary education while we observe certificate rewards for all diplomas in the public sector.*

KEY WORDS: Returns to schooling; human capital; screening; public–private sector

Introduction

The relationship between schooling, productivity and earnings has long interested economists. In an attempt to explain this relationship, two main theories have emerged: Human Capital, and Screening. According to Human Capital theory, education increases individual productivity and, therefore, earnings (for example, Becker, 1964). According to the Screening hypothesis, individuals with more schooling tend to earn more, not because (or, at least, not only because) schooling makes them more productive, but rather because it signals their pre-existing productive abilities. Since these abilities are non-observable, schooling is used as a proxy and, as a result, higher levels of education are correlated with higher earnings (for example, Spence, 1973).

Correspondence Address: Empar Pons, Department of Economic Analysis, University of Valencia, Campus dels Tarongers, s/n, Edificio Departamental Oriental, 46022 Valencia, Spain. Email: Amparo.Pons@uv.es

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The model used in most studies to investigate the relation between education and earnings is that developed by Mincer (1974), who proposes a linear relation between the logarithm of earnings and years of schooling, and interprets the coefficient of this variable as the return of one year of schooling. The Mincer equation, however, does not allow us to determine the nature of these returns. One of the more frequently used contrasts of the Screening model versus Human Capital theory is based on the so-called 'sheepskin argument'. This argues that workers are rewarded not for the productive–enhancing contribution of schooling, but rather for obtaining the certificate that comes with completing a particular level of schooling. Based on a sheepskin prediction, 'wages will rise faster with extra years of schooling when the extra years also convey a certificate'. Hungerford and Solon (1987) generalize the Mincer wage equation by treating the relationship between log wages and years of schooling as a discontinuous *spline function* with discontinuities at every certificate year. This allows us to distinguish between the returns to schooling from mere years of schooling as a reflection of their productive–enhancing contribution (human capital) and the returns to schooling from certificates as signals of the individual's ability. Since Hungerford and Solon presented evidence of significantly larger returns to certificate years in the United States, many further attempts have been made, including those by Belman and Heywood (1991), Heywood (1994) or Arkes (1999), for instance, in the USA, by Ziderman (1990) in Israel, and by Patrinos (1996) in Guatemala. In Spain, San Martín (2001) studies the linearity of returns to schooling using the ECBC (Social Biography and Class Structure Survey) and identifies a positive but no significant sheepskin effect for those individuals that have undergone at least 11 years of education.

The aim of this paper is to delve more deeply into the contrast between human capital and screening theories applying the Hungerford and Solon method to the Spanish case separating into public and private sectors. We think that it is convenient to separate these subsamples due to the great differences in the characteristics of public and private sectors.[1] Public sector wages are established annually by the Government's Budget and have five components. Three of them are guaranteed—the *base wage*, an *extra posting payment* and a *specific complement*—and two components are variable—a *productivity complement* and *other components*. These variable components (representing around 5% of total wages) are used more to remunerate ordinary or extraordinary tasks during overtime hours than a proper productivity mechanism. Wage setting in the private sector allows for better adaptation to worker's productivity. This flexibility is conditioned by the role of unions in the wage setting.[2] Although the unionization rate is especially low in Spain (about 10%), unions play a role in private sector wage setting. In fact, about 90% of employees are covered by collective bargaining.[3] Unions and employers bargain first at national level. This agreement constitutes a benchmark for further bargaining at provincial level, sector level or firm level. It is this final level in which workers and employers are able to obtain an agreement as to take into account productivity, and this fact allows them, for example, to pay extra for a certificate, giving rise to a signalling mechanism and reward years of schooling as mere human capital. Finally, as in other countries, earning dispersion in Spain is usually smaller in the public sector than in the private sector. This fact could merely reflect more flexibility in the private sector as well as the concern of public administrators for pay equality.[4]

These differences between the public and private sector condition the sheepskin effect analysis. In fact, we think that we could only appropriately talk about sheepskin effects in the private sector. Since it is believed the public sector is not competitive and less concerned about profits and efficiency, why should this sector pay extra for credentials? And, in case diploma effects exist, would they really be proper sheepskins? Wage fixing in the non-competitive public sector could reward diplomas by *Credentialism.* This theory is based on a sociological argument proposed by Berg (1970) and Dore (1976) and argues that firms offer higher wages and better jobs to educational certificate holders even if these individuals are not more productive. However, it is difficult to explain why a private sector firm would want to act in this way since by doing so it places itself at a competitive disadvantage. Moreover, we know that educational certificates are requirements to access to a public job, and so public sector can reward certificates merely as consequence of this mechanism. In the non-competitive Spanish public sector, wages are closely linked to jobs instead of to individual productivity: the employer is not able to pay more to a single individual for having a certificate. Instead, an educational requirement is set for each job. It is believed that payment and job conditions are better off in the public sector in comparison with the private one, and consequently a queue to obtain these jobs does actually exist. Educational requirement is an easy-to-follow and hard-to-corrupt selection rule. This requirement system is a phenomenon that is not in essence equivalent to screening, although it is observationally equivalent to sheepskin since we would observe in both a positive correlation between certificates and wages. Additionally to the specific educational requirement, in order to get a public sector post an examination must be passed. The topics to be examined include specific knowledge and skills needed to match each job. So it supposes an extra selection rule and, also, allows for acquiring specific human capital. As a consequence of these public sector peculiarities, we would expect certificates to be more important than years of schooling for civil servants (the last as reflection of human capital), but we can distinguish these certificate effects (forced by educational requirement or credentialism) from sheepskins (certificates as signal of higher productivity). On the other hand, the private sector presents more flexibility in the wage setting and private employers might pay for certificates if they adequately signal worker productivity; that is, if they act as proper sheepskins. Additionally, years of schooling might also be rewarded as a reflection of their productive–enhancing contribution.

Previous reflections on public and private sectors differences lead us to concentrate our contrast of sheepskin effects in the private sector—although we cannot omit the public sector workers of the analysis (e.g., estimating only a wage equation for private sector workers) as it could arise a self-selection bias. This problem emerges from the fact that the employment status of a given individual and, thus, the probability that this individual is in the sample to be used for estimation depends on variables that also affect wages. In order to take this into account, when we analyse the public–private wage differentials we estimate the influence of the wage differential on sector choice simultaneously with the public–private wage equations through an Endogenous Switching Regression model using a maximum likelihood procedure.

This paper is organized as follows. In the next section the methodology of estimation is presented. It includes the Hungerford and Solon specification and the Endogenous Switching Regression model for the public–private sector estima-

tion. The third section presents the data-set and the variable definitions. The results are presented in the subsequent section. Our overall conclusions are explained in the final section.

Methodology

Hungerford and Solon's Equation

Hungerford and Solon's specification states the relationship between the natural wage logarithm, log *W*, and years of schooling, *S*, as a discontinuous spline function with discontinuities at certificate years. In our case, EGB (General Basic Education that includes Primary School and Lower Secondary School), Upper Secondary and University Certificates.[5] After 1970, eight years of schooling were made compulsory for Spanish students by the *Ley General de Educación y Financiamiento de la Reforma Educativa*. Although, given that our sample contains individuals who are not affected by this law and the percentage of individuals that do not obtain the EGB Certificate, we still consider the possibility that this certificate may also have a sheepskin effect. Certificates corresponding to voluntary education, Upper Secondary and University certificates, are also tested. Different from, for example, Hungerford and Solon (1987) or San Martín (2001), we have preferred to define the sheepskin effects by certificates than by a specific number of years for certain certificates. We think there are some problems in identifying degrees with years of schooling in our sample due to the different possible options of the Spanish education system, especially in upper secondary school and university.

The Hungerford and Solon method involves regressing log *W* on total years of schooling *S*, on the three certificate dummy variables (D_{EGB}, D_{USEC} and D_{UNI}) and on the interacted variables of these dummies with the additional years of schooling over the corresponding certificate. We can summarize our specification in the following equation:

$$\ln W = \alpha + \beta S + \lambda_1 D_{EGB} + \gamma_1 (S - S_{EGB}) D_{EGB+} + \lambda_2 D_{USEC} + \gamma_2 (S - S_{EGB} - S_{USEC}) D_{USEC} + \lambda_3 D_{UNIV} + \delta F + u \quad (1)$$

where S_{EGB} and S_{USEC} represent the years of schooling necessary to obtain EGB and Upper Secondary Certificates, respectively, *F* is a set of other variables that influence wage setting and *u* is an error term. The significance of the parameters of the certificate dummy variables λ_1, λ_2, and λ_3 allows us to test for sheepskin effects and the interacted variables for the returns to schooling in the different education segments. So, β would reflect the return to the years necessary to attain the first segment, $\beta + \gamma_1$ would be the estimated return for the second segment, and $\beta + \gamma_1 + \gamma_2$ the returns for the third.

The Switching Regression Model for the Public–Private Sector Estimate

Pay differentials between the public and the private sectors have been traditionally measured by estimating separate wage equations for the workers of each sector. Denoting the public and private sectors as sectors 1 and 2, we can express the corresponding wage function as:

$$\ln W_1 = X\beta_1 + u_1 \tag{2}$$

$$\ln W_2 = X\beta_2 + u_2 \tag{3}$$

where ln W_j is the natural log of wages in sector j, β_j are the vector coefficients associated with wage-determining attributes X, and u_j is a disturbance term. Ordinary least squares (OLS) estimation of wage equations may be biased since individuals sort themselves into the sector that pays higher, and therefore their decision depends on the wage differential, ln W_1 – ln W_2. That is, as an endogenous selection process determines the assignment of workers to both sectors, they are not randomly distributed. In addition to this process of self-selection, there is an employer selection conditioned by individual characteristics, Z. This can be summarized in the following switching equation:

$$I = (\ln W_1 - \ln W_2)\gamma + Z\alpha - \varepsilon \tag{4}$$

where I is a latent variable that determines the sector in which the individual is working, γ and α are vectors of parameters to be estimated, and ε is a disturbance term. If $I > 0$, the individual is selected into sector 1 (the public sector); otherwise sector 2 is the outcome. Substitution of equations (2) and (3) into equation (4) gives the reduced form of the switching equation:

$$I = X(\beta_1 - \beta_2)\gamma + Z\alpha - e = D\delta - e \tag{5}$$

where $D = [X, Z]$, $\delta = [(\beta_1 - \beta_2)\gamma, \alpha]$ and $e = (u_1 - u_2)\gamma - \varepsilon$. For the error terms, u_1, u_2 and e, a normal distribution is assumed with mean zero and covariance matrix:[6]

$$\Sigma = \begin{bmatrix} \sigma^2_1 & \sigma_{12} & \sigma_{1e} \\ \sigma_{12} & \sigma^2_2 & \sigma_{2e} \\ \sigma_{1e} & \sigma_{2e} & 1 \end{bmatrix}$$

The latent variable I is not observed; its dichotomous realization I^*, however, is observed. So:

$$I^* = 1 \quad \text{If } I > 0$$
$$I^* = 0 \quad \text{If } I \leq 0$$

The OLS estimation of equations (2) and (3) would be biased if $E[u_1 / I > 0] \neq 0$ and $E[u_2 / I \leq 0] \neq 0$, as happens in this case:

$$E[u_1 / I > 0] = -\sigma_{1e} \frac{\phi(D\delta)}{\Phi(D\delta)} \tag{6}$$

$$E[u_2 / I \leq 0] = \sigma_{2e} \frac{\phi(D\delta)}{1 - \Phi(D\delta)}, \tag{7}$$

where $\sigma_{1e} = cov(u_1, e) = \rho_{1\varepsilon}\sigma_1, \sigma_{2e} = cov(u_2, e) = \rho_{2\varepsilon}\sigma_2$, and ϕ and Φ are the density function and the distribution function of a normal distribution $N(0, 1)$, respectively. Taking account of these expressions, equations (2) and (3) become:

$$\ln W_1 = X\beta_1 - \sigma_{1e}\frac{\phi(D\delta)}{\Phi(D\delta)} + v_1 \tag{8}$$

$$\ln W_2 = X\beta_2 + \sigma_{2e}\frac{\phi(D\delta)}{1-\Phi(D\delta)} + v_2 \tag{9}$$

where $E[v_1 / I^* = 1] = 0$ and $E[v_2 / I^* = 0] = 0$. The estimation of equations (8) and (9) by the two-step method gives us consistency (Heckman, 1979). To obtain efficiency, we need to estimate equations (2), (3) and (5) simultaneously by maximizing the logarithm of the likelihood function, taking two-step estimations as initial values.[7]

Data-set and variable description

Data

Our data are delivered from the Social Biography and Class Structure Survey (ECBC) carried out in 1991. ECBC data are the only Spanish data-set that provides two measures of education: the certificates obtained, and the total number of years of schooling. We need both data in order to apply the Hungerford and Solon method. Additionally, this data-set offers wide-reaching information on family background, personal details and labour market characteristics.

The ECBC sample size is 6632 and presents a lack of equiprobability: some individuals are more likely to be interviewed than others. For example, young individuals and individuals living in Madrid are over-represented. Additionally, given that the upper-social classes are usually of a small size in reduced samples like this, individuals with upper secondary school and university studies are over-represented. Weights to solve these factors are provided and allow us to obtain a random sample. After removing inconsistencies and wage missing values, we are left with a sample of 2195 salaried workers aged 19–65.[8] Of these, 1344 are working in the private sector and 851 in the public sector.[9]

One final point, our subsample of public workers is over-represented (38% of the salaried workers). Because higher-educated individuals are over-represented and, as we will see later, these individuals have a higher probability of being allocated in the public sector, this subsample has the same problem. We solve this weighting our data to obtain both the main statistics and the estimates.

Variable Description

Mainly, we need to explain the education variables built to apply the Hungerford and Solon method. First, as we have already mentioned, with ECBC data it is possible to compute the number of years of schooling as well as the level of the highest certificate obtained by the individual. Knowing the total years of schooling (those years that end in the attainment of a certificate plus those years that do not end with a complete certificate) avoids the problem of a shortage of variability

in the years of schooling that arises when a specific number of years is assigned. Second, as we have also already pointed out, we have encountered some difficulties in assigning a specific number of years for certain certificates in our sample as is usual in other studies. In upper secondary education and university studies there are different options that require a different number of years to attain the certificate, and these are not necessarily accumulative. There are also difficulties in our sample with respect to the coexistence of individuals who study under different education systems (before or after 1970), which affects the number of academic upper secondary education years (one additional year for individuals who studied after 1970). We assign the certificates taking all of these particularities into account. Specifically, D_{EGB} takes the value one if the individual attains eight years of schooling, D_{USEC} is one if the BUP or Upper Vocational certificate is obtained, and D_{UNI} is one if the University (Long Cycle) certificate is obtained.[10]

The dependent variables are the sector in which the individual is working (public or private) and the logarithm of the net hourly wages. The regressors in the wage equations are the education variables specified in equation (1), potential experience (calculated as present age minus the age the individuals say they had when they started to work), experience squared, tenure (measured in years) and region and municipal size dummies to control for territorial differences.

In the wage equation estimates, in order to correct the sample selection problems that arise from the previous choice of participating in the labour market, we have used a model with two equations, the proper wage equation and a reduced-form participation equation:

$$P = K\alpha + \mu \tag{10}$$

where P is a dichotomy variable that takes value one if the individual works and zero if he/she does not work, K is a vector including both the variables of the wage equation and additional variables that affect the labour market participation decision, α is a vector of parameters and μ an error term. We estimate this participation equation by a probit model and calculate the Inverse Mills Ratio (IMR):

$$IMR = \frac{\phi(K\hat{\alpha})}{\Phi(K\hat{\alpha})} \tag{11}$$

where ϕ and Φ are the density and the distribution functions of a normal distribution $N(0,1)$, respectively and α the vector of estimated coefficients of the labour market participation equation. Next we include the IMR in the wage equation as an additional regressor.[11] Table A1 in appendix A provides the descriptive statistics of the labour market participation equation and Table A2 presents the Probit model results.

The regressors in the public–private sector selection function are gender, age, marital status (married), the number of dependent children and a set of background characteristics such as the father's and the mother's education, if the father is/was a civil servant and the mother a working woman.[12] We add to the switching function all the variables included in the wage equation except experience and experience squared (which are substituted for dummy variables of age) and tenure.[13] Moreover, schooling is considered in levels instead of years given the requirements of access to the public sector. The summary statistics of the wage equation for total, public and private subsamples are presented in Table 1.

Table 1. Descriptive statistics: wage equations

	Public sector mean	Private sector mean
Wage[a]	813.69 (370.55)	577.98 (332.67)
Years of schooling, *S*	11.360 (3.914)	8.945 (3.586)
Experience	20.412 (5.781)	17.575 (4.422)
Tenure	13.099 (10.236)	11.239 (10.326)
Woman	0.414	0.322
Number of observations by year of education		
$S = 0$	1	6
$S = 1$	2	2
$S = 2$	2	11
$S = 3$	6	59
$S = 4$	2	18
$S = 5$	4	38
$S = 6$	9	44
$S = 7$	18	53
$S = 8$	96	395
$S = 9$	34	77
$S = 10$	24	85
$S = 11$	59	67
$S = 12$	40	80
$S = 13$	24	98
$S = 14$	187	81
$S = 15$	108	79
$S = 16$	87	59
$S = 17$	93	73
$S = 18$	23	12
$S = 19$	24	4
$S > 20$	8	3
Total observations	851	1344

We omit the statistics for region and municipal size dummy variables. Standard deviation in parentheses.
[a]Variable in pesetas/hour.

We can observe that, overall, public sector workers earn more than private sector workers. Although in absolute terms (standard deviation) wages dispersion is slightly higher in the public sector, in relative terms (coefficient of variation calculated as the standard deviation on the mean) the dispersion is higher in the private sector (0.455 in the public sector and 0.575 in the private sector). Public sector workers take longer careers than private sector workers do, and experience and tenure are higher for public sector workers. It is possible to explain this year of schooling difference by the fact that, even in the lowest public jobs, a considerable amount of education is a requirement to occupy hose jobs.[14] Furthermore, the Spanish public sector hires a higher proportion of professionals than the private sector does. Albert *et al.* (1999) show that the rapid rise of the public sector labour market has also been accompanied by an intense change in the composition of public employment by educational levels increasing the proportion of public employees who have university studies (45% in the public sector and 12% in the private sector in 2002; Spanish Labour Force Survey). Table 1 also provides the number of observations by year of education. It shows sufficient observations of

Table 2. Descriptive statistics: sector selection function

	Total mean	Public sector mean	Private sector mean
Public sector worker	0.387		
Woman	0.349	0.414	0.3252
Married	0.602	0.676	0.572
Number of dependent children	0.745 (1.006)	0.839 (1.071)	0.707 (0.976)
Age 26–35 years	0.301	0.341	0.284
Age 36–45 years	0.267	0.280	0.262
Age 46–55 years	0.146	0.182	0.131
Age 56–65 years	0.077	0.091	0.071
EGB	0.509	0.402	0.552
Upper secondary education	0.210	0.205	0.212
University (Short Cycle)	0.092	0.203	0.047
University (Long Cycle)	0.078	0.151	0.048
Less than EGB (father)	0.338	0.249	0.374
Upper secondary education (father)	0.076	0.116	0.059
University (father)	0.066	0.127	0.042
Father's education unknown	0.006	0.004	0.007
Less than EGB (mother)	0.346	0.259	0.381
Upper secondary education (mother)	0.033	0.046	0.027
University (mother)	0.015	0.026	0.010
Mother's education unknown	0.142	0.155	0.137
Public sector father	0.176	0.303	0.124
Working mother	0.202	0.163	0.218
Observations	2195	851	1344

We omit the statistics for region and municipal size dummy variables. Standard deviation in parentheses.

non-degree years of schooling to estimate a Hungerford and Solon discontinuous spline function. With regard to the sector selection equation (Table 2), the more significant facts are that public sector workers have higher levels of education and have more educated parents. In this sector, there is also a higher proportion of women, of married individuals, of workers with dependent children and a higher proportion of workers whose father is/was also public sector workers.

Estimation Results

Wage equations are presented in Table 3. Mincer specification results are presented in the second and third columns to compare with our Hungerford and Solon results. We observe that the returns to schooling in the public and the private sectors are the same: 3.9%. We were surprised by both the low levels of returns to schooling obtained and by the fact that returns to schooling were similar. These phenomena emerge when we take account of selectivity bias and estimate a switching estimation model by maximum likelihood. When we estimated these coefficients with OLS, returns to schooling of 6.3% in the public sector and

Table 3. Switching model: wage equations

	Mincer function		Spline function (a)		Spline function (b)	
	Public sector	Private sector	Public sector	Private sector	Public sector	Private sector
Constant	6.343 (55.55)	5.785 (97.47)	6.619 (39.76)	5.917 (66.44)	6.639 (39.97)	5.918 (66.34)
Total years of schooling, S	0.039 (8.29)	0.039 (11.50)	−0.017 (−0.86)	0.027 (2.26)	−0.017 (−0.86)	0.027 (2.25)
$(S - S_{EGB})^{*}D_{EGB}$			−0.001 (−0.03)	−0.018 (−0.90)	−0.002 (−0.08)	−0.018 (−0.89)
$(S - S_{EGB} - S_{USEC})^{*}D_{USEC}$			0.049 (2.23)	0.046 (2.99)	0.057 (2.33)	0.059 (3.16)
D_{EGB}			0.146 (2.01)	−0.024 (−0.59)	0.148 (2.06)	−0.024 (−0.60)
D_{USEC}			0.177 (2.65)	0.096 (1.78)	0.184 (2.75)	0.095 (1.76)
D_{UNI} (Short Cycle)					−0.047 (−0.89)	−0.062 (−1.04)
D_{UNI} (Long Cycle)			0.140 (2.22)	0.093 (1.57)	0.125 (1.88)	0.073 (1.21)
Experience	0.019 (4.19)	0.019 (6.58)	0.016 (3.52)	0.019 (6.65)	0.016 (3.37)	0.019 (6.60)
Experience^2	−0.0003 (−3.12)	−0.0003 (−5.97)	−0.0002 (−2.65)	−0.0004 (−6.48)	−0.0002 (−2.52)	−0.0004 (−6.43)
Tenure	0.003 (1.38)	0.079 (5.82)	0.003 (1.64)	0.008 (6.08)	0.003 (1.67)	0.008 (6.08)
IMR	−0.074 (−2.07)	−0.020 (−0.74)	−0.082 (−2.28)	−0.016 (−0.61)	−0.083 (−2.31)	−0.017 (−0.66)
Woman	−0.089 (−2.35)	−0.272 (−10.14)	−0.074 (−1.95)	−0.265 (−10.04)	−0.071 (−1.86)	−0.264 (−9.97)
σ public sector	0.381 (40.96)		0.358 (41.60)		0.359 (40.45)	
σ private sector		0.385 (22.21)		0.361 (21.20)		0.366 (21.38)
ρ public sector	−0.616 (−9.58)		−0.430 (−4.05)		−0.4370 (4.05)	
ρ private sector		0.760 (17.37)		0.704 (11.77)		0.725 (13.30)
Observations	851	1344	851	1344	851	1344

We omit the results for region and municipal size dummy variables. t–value in parentheses.

5.2% in the private sector were obtained (these results are provided in Table A3 in appendix A).[15]

The fourth column of Table 3 shows the estimation results of a spline function corresponding to equation (1). We observe that the coefficient of EGB years is positive and significant only in the private sector (2.7%). However, in this sector there is no reward other than the accumulated years of schooling. In the public sector, by contrast, there are no returns to EGB years but there is a positive diploma effect for this certificate, D_{EGB} (in fact, we find certificate rewards for all certificates in the public sector). In the upper secondary education segment, we do not observe any additional returns to schooling in either sector but we do observe a positive diploma effect in the public sector (in the private sector this is only significant at the 90% level of confidence). In the university segment, there are positive returns to these years of schooling in both the private and the public sectors (4.9% and 4.6%, respectively). This is to a certain extent surprising as we expected no returns to years of schooling in the public sector. Usually, the public sector cannot adapt wages to individual productivity (which we calculate through years of schooling). However, it is within this higher educated group that the majority of responsibility jobs are concentrated and where there are more possibilities for promotion depending on personal productivity. Moreover, we once again observe a diploma effect in the public sector and none in the private sector. The total returns for this education segment are 4.9% for the public sector and 7.3% for the private sector (the sum of returns to EGB and university years of schooling).

Given that in the public sector the Short Cycle University certificate or the attainment of three years of the Long Cycle certificate facilitates the access of individuals to certain occupations, we consider an alternative specification that includes a dummy variable D_{UNIS} if the individual has attained the Short Cycle certificate or three years of the Long Cycle certificate. As the fifth column of Table 3 shows, no diploma effect is found for this certificate in either the public sector or the private sector.[16]

With respect to the others variables of the wage equations, in all specifications of the wage equations experience, experience squared and tenure coefficients show the expected sign; women hold a disadvantaged position although the coefficient of this variable is lower in the public sector and IMR is only significant in the public sector. Of some special interest are the signs of the correlation coefficient in the public sector $\rho_{1\varepsilon}$ and in the private sector, $\rho_{2\varepsilon}$. These correlations appear in the conditional expectation of wages:

$$E[\mathrm{In}W_1 / I > 0] = X\beta_1 - \rho_{1\varepsilon}\sigma_1 \frac{\phi(D\delta)}{\Phi(D\delta)} \tag{14}$$

$$E[\mathrm{In}W_2 / I < 0] = X\beta_2 + \rho_{2\varepsilon}\sigma_2 \frac{\phi(D\delta)}{1-\Phi(D\delta)} \tag{15}$$

Our estimate of $\rho_{1\varepsilon}$ is negative whereas $\rho_{2\varepsilon}$ is positive, and both coefficients are significant. These results imply that those who actually work in the public sector and those who actually work in the private sector perform better than any random individual would have. This fact suggests that separate OLS estimations of the public–private wage equations would not be consistent as the mean of the error term would not be zero. This is especially important since we can observe

differences in the OLS and Endogenous Switching estimates in terms of magnitude and differences between sectors (these being favourable to the public sector when using OLS estimates and similar when using the Endogenous Switching model).

Table 4 presents estimates of the sector choice equation. We can observe in the second column (corresponds to the Mincer specification of the wage equations) that more schooling makes a person more likely to obtain a public sector job. Women have a higher probability of becoming public sector workers; the number of dependent children has a positive effect; older individuals are more likely to be found public sector workers than younger people. As regards the family background variables, an individual whose father works in the public sector or has a secondary education has a positive probability of public employment. An individual whose mother has less than a primary education, on the other hand, has a lower probability and having a working mother is not significant. Similar results are obtained with the spline specifications (third and fourth columns in Table 4).

To summarize, our results show that diploma effects are concentrated in the public sector. In the private sector we only observe them for upper secondary education. This finding are in line with Van der Gaag and Vijverberg (1988) who, using a similar methodology, find for males in the Republic of the Ivory Coast

Table 4. Switching model: switching equation (public = 1, private = 0)

	Mincer function	Spline function (a)	Spline function (b)
Constant	–1.600 (–9.67)	–1.812 (–10.99)	–1.792 (–10.88)
EGB	0.291 (2.40)	0.492 (3.72)	0.483 (3.66)
Upper secondary education	0.464 (3.55)	0.748 (5.23)	0.735 (5.15)
University (three years)	1.251 (8.34)	1.475 (9.28)	1.484 (9.32)
University (five years)	1.342 (8.42)	1.355 (8.20)	1.351 (8.20)
Woman	0.236 (3.35)	0.249 (3.48)	0.245 (3.43)
Married	0.129 (1.55)	0.142 (1.59)	0.142 (1.60)
Number of dependent children	0.074 (1.95)	0.068 (1.64)	0.066 (1.62)
Age 26–35 years	0.371 (3.95)	0.377 (3.81)	0.370 (3.77)
Age 36–45 years	0.349 (3.11)	0.365 (3.08)	0.360 (3.05)
Age 46–55 years	0.627 (5.03)	0.643 (4.90)	0.634 (4.85)
Age 56–65 years	0.671 (4.25)	0.715 (4.27)	0.701 (4.21)
Less than EGB (father)	–0.023 (–0.27)	–0.005 (–0.06)	–0.004 (–0.05)
Upper secondary education (father)	0.207 (1.90)	0.167 (1.47)	0.164 (1.45)
University (father)	0.113 (0.99)	0.120 (1.00)	0.114 (0.95)
Father's education unknown	–0.606 (–1.86)	–0.607 (–1.68)	–0.611 (–1.72)
Less than EGB education (mother)	–0.272 (–3.22)	–0.290 (–3.25)	–0.289 (–3.26)
Upper secondary education (mother)	–0.202 (–1.19)	–0.226 (–1.28)	–0.230 (–1.31)
University (mother)	–0.021 (–0.08)	–0.087 (–0.35)	–0.086 (–0.35)
Mother's education unknown	–0.061 (–0.64)	–0.083 (–0.83)	–0.084 (–0.85)
Public sector father	0.340 (4.51)	0.395 (5.01)	0.386 (4.91)
Working mother	–0.077 (–1.08)	–0.078 (–1.02)	–0.076 (–0.99)
Log-likelihood	–1805.872	–1772.823	–1772.104
Observations	2195		

We omit the results for region and municipal size dummy variables. *t*-value in parentheses.

additional returns for the Secondary and Post-Secondary certificates in the public sector whereas in the private sector only the Secondary Education certificate is rewarded. Stelcner *et al.* (1989) for Peru, also estimating an endogenous switching model, find that Secondary and Post-Secondary certificates are premiumed in both sectors. Therefore, our results contrast with Heywood (1994) who observes for the USA strong evidence of diploma effects in the private sector but little evidence of them in the public sector. In this paper we use different methodology to Heywood in the estimation of the sector wage equations. Heywood does not take into account the idea that sector choice could be endogenous. However, this fact is not the cause of our different results since we have estimated separate equation for public and private sectors by OLS (Heywood's method) and the results, presented in Table A3 in appendix A, show the same pattern that was found when we used an endogenous switching model estimation (but with different magnitude in the returns to years of schooling). Hence, differences in the results presumably arise from the differences in the wage settings in Spain and the United States.

Conclusions

In this paper we have applied the Hungerford and Solon (1987) method to examine alternative explanations for the relation between education and earnings in the Spanish labour market, separating into public–private sectors. We differentiate, specifically, between Human Capital theory and the 'sheepskin' argument. The latter argues that workers are rewarded not for the productive–enhancing contribution of schooling (the case put by Human Capital theory), but rather for the signal sent by the certificates individuals obtain with the completion of different levels of schooling (sheepskin argument). We have extended our analysis to distinguish between public and private sectors as we expected different behaviours within these groups. It is often assumed that the application process and the wage settings in the public sector are more formal and less competitive than the public sector and, therefore, makes more use of diplomas. Although this mechanism not should lead us to interpret them as a sheepskin effect since wages are closely linked to jobs instead of to individual characteristics and the public employer is not able to pay more to a single individual for having a certificate. We think that diploma effects in the public sector are due to educational requirement and they are used more as an strict selection rule of applicant than a signal mechanism. However, higher flexibility of the private sector would allow us to observe, if any, proper sheepskin effects.

When we examined the public and private wage differences using an Endogenous Switching Regression model, we observed differences between sectors in both the retribution of the years of schooling and in the diploma effects. While in the public sector there is only a positive return to university years, in the private sector we obtain a positive return to EGB and university years of schooling. In the public sector all certificates are rewarded, confirming our initial hypothesis that the diploma effects would be substantial in this sector. In the private sector, however, only the Upper Secondary Education certificate has a positive sheepskin effect (only significant at 90%).

The global conclusion is that there is little evidence of sheepskin effects in the Spanish labour market. Certificate rewards were found basically in the public sector where we think the institutional restriction in the access and wages setting

would force a diploma effect that should be differentiated from a proper sheepskin effect; that is, certificates as signal of individual productivity. This would seem to confirm the validity of Human Capital theory; that is, that education enhances individual productivity.

Notes

1. See López (1993) and Albert *et al.* (1999) for an analysis of the wage setting and employment policy in the Spanish public sector.
2. For more information on collective bargaining structure in Spain, see Jimeno and Toharia (1994).
3. In fact, in Spain we cannot distinguish between union and non-union sectors since the unionization rate is so low everywhere and unions' influence is so high in all economic private sectors.
4. Albert *et al.* (1999) and García *et al.* (1997) analyse in detail wage differentials between the public and private sectors in Spain.
5. Using OECD and EUROSTAT classification, the Spanish education system until 1990 was divided in three segments. First, EGB (eight years) included primary school (five years) and lower secondary school (three years). In the second segment, upper secondary school, a choice was made between vocational studies comprised of two cycles, lower (two years) and upper (three years), and academic studies that included BUP (three years) and an additional pre-university course (COU) necessary to move on to the following segment. Finally, university includes Short Cycle (three years) and Long Cycle (five years) university studies.
6. The variance of the error term of the reduced-form switching equation has been set equal to one in order to identify the whole parameters (see Maddala, 1983, pp. 223–224).
7. See Maddala (1983, p. 224) for a derivation of the likelihood function.
8. There is a quite high percentage (20%) of workers that no report information about their wage.
9. Our samples are similar to previous studies using ECBC. See for example, San Martín (2001) for total sample analysis or Albert and Moreno (1998) for public–private sector analysis.
10. In order to test for non-linearity in the returns to schooling, there is a method consisting of regressing log W on a set of dummy variables for $S = 1, 2, 3 \ldots 17$. This specification imposes no restriction on the shape of the earning/schooling profile. It would treat log W as a step function of years of schooling with a separate step for each year and allow sheepskin effects to be identified. Even though we are able to know the exact number of years of schooling, we think the small size of the sample due to the fact that we split it by sector could cause estimation problems in case of including a dummy variable for each year of schooling. Additionally to this fact, the mentioned difficulties in assigning a specific number of years for certain certificates do not allow us to interpret accurately the results for each year.
11. Taking this into account, in equations (2) and (3) the X vector includes the IMR and β includes the associated parameter to IMR; that is, the covariance between the unobservable factors that affect the labour participation and those that affect wages. Then, we estimate the equations (2), (3) and (5) model by the maximum likelihood method.
12. In order to avoid loosing observations, we include a dummy variable that takes the value one when the father or mother's level of education is unknown.
13. We omit tenure because, in order to identify the γ parameter in equation (5), at least one of the variables entering the wage equations must not appear in the switching function (see Maddala, 1986, p. 1639).
14. In Lassibille (1998), with the Spanish Household Budget Survey (1990–1991), a similar difference in years of schooling was observed.
15. A similar pattern emerges in Van der Gaag and Vijverberg (1988) and Hartog and Oosterbeek (1993).
16. We also considered the possibility that D_{UNIS} only takes the value of one if the individual obtained the Short Cycle certificate and the results were very similar.

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Appendix A

Table A1. Descriptive statistics: labour market participation equation

	Mean
Participate	0.409
Years of schooling	8.510 (4.119)
EGB	0.509
Upper secondary education	0.173
University (three years)	0.053
University (five years)	0.058
Age 26–35 years	0.244
Age 36–45 years	0.220
Age 46–55 years	0.168
Age 56–65 years	0.173
Married	0.647
Woman	0.497
Householder	0.450
Number of dependent children	0.672 (0.984)
Number of income receivers	1.672 (0.819)
Partner upper secondary	0.121
Partner university	0.071
Observations	6194

We omit the results for region and municipal size dummy variables. Standard deviation in parentheses.

Table A2. Probit model: labour market participation equation

Constant	1.211 (10.12)
Years of schooling	–0.008 (–0.62)
EGB	0.280 (1.37)
Upper secondary education	0.632 (3.59)
University (three years)	0.164 (1.22)
University (five years)	0.173 (2.18)
Age 26–35 years	0.071 (1.18)
Age 36–45 years	–0.044 (–0.60)
Age 46–55 years	–0.342 (–4.31)
Age 56–65 years	–0.830 (–9.76)
Woman	–0.516 (–11.61)
Married	–0.337 (–6.13)
Household head	0.306 (5.95)
Number of dependent children	–0.068 (–2.75)
Number of income receivers	–0.704 (–27.10)
Partner upper secondary	–0.019 (–0.32)
Partner university	0.251 (3.24)
Log likelihood	–3224.614
Observations	6194

We omit the results for region and municipal size dummy variables. *t*-value in parentheses.

Table A3. OLS wage equation estimations

	Mincer function		Spine function	
	Public sector	Private sector	Public sector	Private sector
Constant	5.635 (94.81)	5.708 (85.60)	6.090 (58.63)	5.893 (80.12)
Years of schooling, total	0.063 (19.54)	0.052 (13.03)	–0.008 (–0.62)	0.329 (3.59)
$(S - S_{EGB})^*D_{EGB}$			0.010 (0.47)	–0.017 (–1.04)
$(S - S_{EGB} - S_{USEC})^*D_{USEC}$			0.052 (3.20)	0.056 (3.34)
D_{EGB}			0.149 (3.43)	–0.026 (–0.72)
D_{USEC}			0.182 (3.31)	0.092 (1.86)
D_{UNI} (Long Cycle)			0.101 (1.83)	0.073 (0.84)
Experience	0.028 (6.69)	0.023 (7.73)	0.023 (5.60)	0.021 (7.49)
Experience^2	–0.0004 (–5.13)	–0.0004 (–6.75)	–0.0003 (–4.45)	–0.0004 (–7.11)
Tenure	0.004 (2.34)	0.008 (5.52)	0.005 (2.73)	0.008 (5.91)
IMR	–0.105 (–3.20)	–0.031 (–1.09)	–0.105 (–3.03)	–0.020 (–0.78)
Gender (woman)	–0.010 (–0.42)	–0.236 (–9.34)	–0.007 (–0.28)	–0.245 (–9.68)
Adjusted R^2	0.470	0.369	0.495	0.389
Observations	851	1,344	851	1,334

We omit the results for region and municipal size dummy variables. *t*-value in parentheses.