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HOW PROFITABLE IS TO STUDY IN SPAIN? AN EMPIRICAL INSIGHT USING A NEW SOURCE OF INFORMATION

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ABSTRACT

This paper presents empirical evidence on the returns to education in Spain using the *Survey on the Quality of Life in the Workplace*. Five waves (1999-2003) of the survey have been pooled to build a dataset for which Mincer-type earning functions are estimated. Unlike other analyses experience is computed as actual and not potential experience, and a variable capturing periods of unemployment is also included. We calculate the returns to education for male workers following the simplest Mincer's specification estimated by (a) OLS and (b) instrumental variables (IV) techniques as a means to deal with endogeneity concerns regarding schooling and find that returns to education for male salaried workers are 5.68 (OLS) and 7.37 (IV with a family background instrument) giving evidence of a slightly declining trend in the rate of return to education in Spain. Evidence against Mincer's underlying hypothesis of linearity of the returns to education in schooling is found when schooling attainment is taken as qualifications. Concerning the parallelism of log-earnings experience profiles across schooling levels, the inclusion of interaction terms between variables experience and education casts some doubts on the plausibility of this assumption in the private sector, although public sector's earning-experience profiles are more coherent with it. Moreover unlike previous international and Spanish studies the results provide evidence of larger returns among public employees. The empirical analysis is finally extended by focusing on regional differences, which are found to be large.

1. INTRODUCTION

Spain is one of the OECD countries having experienced a rapid economic growth in the last decades. This pace of economic growth together with changes in the structure of the economy has led to an enlarged demand for educated workers. Such rising demand has been encompassed on the supply side by a dramatic increase in the share of population holding formal qualifications. This process has been fostered, among other factors, by the high figures of unemployment among the younger cohorts, although a substantial reduction in those rates has been experienced in the last years. In any case Spain is still today one of the industrialised countries with the highest share of university students over population aged 18-22.

The fact that more education is rewarded by the labour market with higher incomes is one of the empirical regularities more widely observed for long, although the exact measure of these rewards and their evolution in time still generates a great deal of research. Based on a new data set referred to the immediate past years (1999-2003) this paper uses one of the most popular frameworks for approaching these issues, Mincer's original specification of his human capital earnings function. Mincer's specification and above all the interpretation of the estimated coefficient for schooling attainment as the private returns to education have been submitted to severe criticism. The paper deals with some of these issues and tries to give an answer to several questions: To what extent have changed what are conventionally called private returns to education as a result of the supply and demand relative performance? How acceptable are some of the assumptions embedded in Mincer's original specification in the Spanish case at the beginning of the 21st Century? How heterogeneous are rewards to education across Spanish regions? The paper is divided into the following sections: Section 2 discusses some factors regarding the estimation of rates of returns to schooling using Mincer's specification. Section 3 describes the data set used in the research and the estimation procedures. Section 4 details the empirical results and section 5 concludes.

2. ESTIMATING THE RATE OF RETURNS TO INVESTMENT IN SCHOOLING: SOME COMMENTS

As is well known, one of the methods conventionally used to measure the influence of education on earnings is the estimation of Mincer's functions in which the log of individual earnings is considered to be explained by a schooling attainment term and a quadratic working experience term. Mincer's (1974) work, where previous analyses by him and other authors were re-elaborated, leads to this empirical specification,

$$logY_i = \beta_o + \beta_1 S_i + \beta_2 X_i + \beta_3 X^2_i + u_i \quad [1]$$

where i is a subscript for individuals (i=1,2,...,n), $logY_i$ is log income, S_i is years of education, X_i is years of work experience after completing schooling and u_i is the random error term. As pointed out by Heckman *et al.* (2003), under special conditions –among which the assumption that direct costs are negligible- Mincer's framework captures two concepts: (a) a hedonic wage function that allows measuring how schooling and experience are rewarded in the labour market and (b) a rate of return to schooling (β_1) comparable with the return to alternative assets so allowing establishing the rationality of investment in education.

Two different theoretical motivations[1] led Mincer to equation [1]. Of these, the most frequently referred to is Becker's human capital theory according to which an individual chooses the length of his education so that the present value of the stream of future incomes is maximised, net of the direct costs of education. Assuming that these costs are negligible it follows from the equilibrium condition that the return to an additional year of schooling is approximately the difference in log wages between studying a given number of years and studying that number less one. Besides the triviality of direct costs of education Heckman *et al*. (2003) have listed some other embedded assumptions including stationarity of economic environment and perfect certainty about future earnings flows associated with different schooling levels, the absence of loss of work life from schooling, the linearity in schooling, the multiplicative separability between the schooling and experience components of earnings, and the absence of endogeneity of education[2]. Most frequently empirical exercises have incorporated other variables considered as relevant for the explanation of wage differences together with schooling and experience on the right side of equation [1]. Their inclusion in the estimations has the typical effect of increasing the R^2 but has been argued that this may lead to biased estimations of the coefficient associated with schooling given the frequent absence of independence between these new regressors and random perturbance element. The addition of these independent variables typically reduces the estimated returns to schooling since they in fact capture the mechanism through which the educated achieve better wage conditions[3].

Some of the concerns listed above, like that of linearity in schooling, separability between schooling and experience and the need of accounting for the endogeneity of schooling will be dealt with in the paper. Others, like that of tuition and fees, may not be so relevant in the Spanish case. Thus a recent exercise by Arrazola *et al.* (2001) estimates that opportunity costs account for more than 90 per cent of total private education cost. Public administration's strong

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[1] See Heckman *et al*. (2003) for the derivation of Mincer's earnings specification from his compensating differences model (1958) and the accounting-identity model of human capital formation (1974).

[2] Among the diverse strategies adopted to deal with this last concern are instrumental variable techniques, finding a paired comparison with similar ability (genetic twins, for example) and the use of proxies of the ability variable that are included as regressors in equation [1] (Heckman and Li, 2003). See Card (1999, 2000), Harmon *et al.* (2003) and Heckman *et al.* (2003) for comprehensive critical surveys of instruments used in recent international literature on the profitability of investment in education.

[3] This seems to be especially true in the case of choice variables linked with job characteristics and therefore endogenous. See Barceinas *et al*. (2001 and 2002), drawing on Mincer's discussion on the issue. In Barceinas *et al.* (2001) the inclusion of the highest number of control variables reduced returns to education from 8.2% to 6.5%.

subsidisation scheme (that assumes more than 80 per cent of total direct costs) leads to a large similarity in the private returns that they estimate through both internal rates of return formulation and human capital based standard Mincer specification. Some of the other assumptions, however, will not be considered here, although their potential impact in the interpretability of the results should not be neglected (particularly in what refers to the modelling of uncertainty about future returns at the time schooling decisions are made both in a static or dynamic setup, and in the impossibility for correcting for the potential selection bias).

3. MODELLING AND DATA DESCRIPTION

We use individual data obtained from a survey carried out by the Spanish Ministry of Economy and Social Affairs (*Ministerio de Trabajo y Asuntos Sociales; MTASS*), the so-called "Survey on the Quality of Life in the Workplace" ("*Encuesta de Calidad de Vida en el Trabajo"*, ECTV) which is yearly conducted since 1999 and contains information on a wide number of socio-economic and workplace variables for a sample of Spanish workers. Samples for the different years are not linked in any way, so it is impossible to match the evolving behaviour of individuals to obtain a panel of data and carry out longitudinal analysis. To increase the number of observations individual data for years 1999 to 2003 were pooled in a single database[4].

Relevant variables for this study include net revenues[5], schooling attainment and working experience. Variable 'schooling attainment' has been addressed in two ways. First, as in many studies, the original variable in the survey, categorical variable 'maximum qualification obtained by the individual' has been transformed by calculating the minimum number of years necessary to gain a qualification to allow the estimation of the internationally-comparable[6] return of an additional year of full-time education[7]. Alternatively, dummy variables were

[4] Unfortunately, the characteristics of the database do not allow controlling for the bias associated to the potential existence of a sample selection problem using Heckman's correction (something that is done in Alba-Ramírez and San Segundo, 1995; and Arrazola and Hevia, 2003, among others using different databases for Spain), since no information is available about the unemployed. However, as pointed out by Barceinas *et al.* (2002), some have recently expressed their concerns about the existence of potential problems associated with the usage of this procedure (Puhani, 2000). The problem seems no to be relevant for male workers in any case, given the international and also Spanish literature on the issue. Thus in the OLS estimation by Arrazola and De Hevia (2003) the selectivity correction reduced the returns to education of salaried men by 0.4 percentage points.
Year fixed effects were included in all estimations through dummy variables for the different years in the estimations with the aim of capturing both the impact of inflation and potential structural change.

[5] The actual value of earnings is unobserved in the survey where net monthly revenues are presented in the survey grouped in a high number of categories (13). The estimation of the different specifications involved allocating to all individuals in a given group the arithmetic midpoint between the threshold values of that group and using the information on average working hours per week to transform monthly revenues into wage per hour (a typical 4.1-week month was taken for these calculations). The alternative proposed by Stewart (1982) was also implemented, although no significant differences with OLS estimation were observed. The latter was preferred as it allowed robust estimations and accounting for potential endogeneity of schooling through IV techniques.

[6] The introduction of the reforms urged by the so called 'Bologna process' should improve comparability across Europe, at least at a University level, given the common definition of the measure units for academic burden. The introduction of the annex to the diploma, which includes a description of the formal education received in and after the official degree, during the professional career, will introduce significant changes in terms of accuracy of the calculation of the returns to education for the relevant group of workers.

[7] A measurement error may occur for individuals for which the completion of the degree took longer for which any amount of extra years of education could in fact be considered a sign of lesser productivity. Changes in the education laws in the last years have

introduced for the different levels of qualifications originally considered in the survey[8]. Education groups were previously clustered to allow correct estimation given the number of available observations[9], giving place to specification [2]:

$$lnw_i = \beta_o+\beta_1 E2_i+\beta_2 E3_i+\beta_3 E4_i+\beta_4 exp_i+\beta_5\ exp^2_i\ +u_i \qquad [2]$$

where, *lnw* is the Neperian logarithm of earnings, *Ei* is the maximum level of education reached by the individual considered, *exp* is the number of years of experience and *u* is as usual the error term.

Actual and not potential experience was used in the estimation. Variable experience has frequently (in a tradition inaugurated by the same Mincer, 1974) been defined as 'age' less 'years of schooling', less six, in the absence of information on actual experience. This overestimates working experience for employees that spent more than statutory number of years to obtain a qualification or for those for which finding a job took longer, and has raised some criticism since if variable schooling is potentially endogenous, the same concerns should apply to experience defined as a function of schooling attainment[10]. In this paper experience is defined as the age of the individual at the time of the survey less the age at which he obtained a tenure of at least three months[11]. An additional control variable was included, defined as the number of periods of involuntary unemployment experienced[12]. This is consistent with the interpretation of experience as a period of enrichment of individual's skills through informal training or on-the-job training. Experiencing periods of unemployment not only interrupts this preparation (except in those cases in which these are periods of education, which is not usually the case in Spain), but also raises some obsolescence concerns as they are longer.

As pointed out before, much research work has been concerned with potential endogeneity of schooling attainment, i.e. with the fact that some of the forces influencing education are also relevant when earnings are to be explained. Conventional procedures confirm

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increased the probability of error due to the alternative introduction and elimination of the possibility of devoting more than one natural year to complete what was originally planned for one academic year in the primary and secondary education.

[8] Something that has been considered more reasonable in countries with multiple education streams, which is the case of most EU countries. Card (1999) and Heckman *et al*. (2003) report results in which it is obvious that credentials may be more relevant than years of schooling even in the US, confirming the 'sheepskin effect' hypothesis.

[9] We first estimated Mincer's equation using 10 different levels that were then grouped into 4 clusters according to the similarity of the estimated coefficients to improve the estimations: E1 (Primary or less than primary and secondary education below compulsory threshold age), E2 (vocational education –first and advance levels- and post-compulsory secondary education), E3 (short duration university degree), E4 (long duration university degree and postgraduate degree –this last credential yielded significantly higher returns but was aggregated in this cluster due to the small number of relevant workers).

[10] As a reaction to these concerns variable experience has frequently been substituted by variable age, as has been recently done in the Spanish case by Barceinas et al. (2002) and Pons and Gonzalo (2002).

11 This alternative can also be subject to criticism, above all in the last years that have witnessed the concatenation of extremely short contracts sometimes for years at the beginning of many younger's working life. This effect should not be however largely relevant for total population, and is probably restricted to those cohorts.

[12] Original variable does not allow to control for length of these periods and also includes contracts whose rewards do not were sufficient to cope with.

the endogeneity of schooling attainment in our sample[13]. Family background is one of the most popular set of instruments[14] in this context. Previous analysis by Pons and Gonzalo (2002) has shown that parent's education and college availability are the instruments that fit the Spanish case better than those related with natural experiments like the season of birth and dummy variables capturing the effects of changes in the laws of education. In fact they conclude that the efficiency obtained using parent's education and university proximity is quite similar to that obtained using parent's education only. The instrument chosen here is then maximum qualification obtained by the head of household at the age of 18 (most frequently father's education), and the test conducted confirmed it as a suitable one[15] .

4. RESULTS

Table 1 displays the results obtained from the estimates based on Equations [1] and [2]. Both OLS[16] and IV techniques were used. Surprisingly in line with recent international comparative exercises (Trostel *et al*., 2002) IV results suggest that OLS estimates of the rate of return to schooling are biased downward by one third. The inclusion of the unemployment term, on its hand, marginally reduces the returns to education although being significant at 99% in all cases.

Table 1. Estimates for Mincer's specification (male salaried workers)

	(1)OLS	(2)IV	(3) OLS	(4) IV	(5)OLS	(6) IV	(7) OLS	(8) IV
Years of schooling	0.0587* (0.0009)	0.0751* (0.0022)	0.0568* (0.0010)	0.0737* (0.0023)				
E2					0.2058* (0.0078)	0.1547 (0.1507)	0.1902* (0.0079)	0.1318 (0.1423)
E3					0.4755* (0.0134)	0.7167** (0.3647)	0.4556* (0.0134)	0.7158** (0.3258)
E4					0.6648* (0.0141)	0.8178* (0.1382)	0.6461* (0.0142)	0.8003* (0.1250)
Experience	0.0260* (0.0010)	0.0262* (0.0010)	0.0273* (0.0010)	0.0273* (0.0010)	0.0271* (0.0010)	0.0275* (0.0010)	0.0285* (0.0010)	0.0288* (0.0010)
Experience squared	-0.0003* (0.0000)	-0.0003* (0.000)	-0.0003* (0.0000)	-0.0003* (0.0000)	-0.0004* (0.0000)	-0.0004* (0.0000)	-0.0004* (0.0000)	-0.0004* (0.0000)
Periods of unemployment			-0.0152* (0.0015)	-0.0120* (0.0015)			-0.0171* (0.0016)	-0.0154* (0.0019)
Constant	0.7123* (0.0161)	0.5313* (0.0273)	0.7415* (0.0164)	0.5517* (0.0286)	1.1382* (0.0128)	1.1553* (0.0453)	1.1584* (0.0131)	1.1304* (0.0462)
N	12169	11781	11603	11239	12346	12346	11770	11770
R-squared	0.3004	0.2839	0.3113	0.2937	0.2884	0.2521	0.3034	0.2612

*Significant at 1% level
**Significant at 5% level

[13] The Hausman (1978) test resulted in the rejection of the null hypothesis of exogeneity of schooling attainment, F(1, 11776)=63,27 prob >F=0.00. So $E(ui/Si) \neq 0$ and OLS gives biased and inconsistent estimates of the causal effect of schooling attainment on earnings.
[14] See Card (1999) for a critic discussion on family background instruments.
[15] *F statistic on excluded instruments test*, F(1, 13709)=3589.42, Prob>F=0.0000.
[16] All OLS estimations were carried out with standard errors robust in heterocedasticity, by means of the White's variance and covariance matrix after confirming the existence of heterocedasticity usually associated with cross-sectional data by the usual tests in which the null hypothesis of homocedascity was rejected (Cook-Weisberg chi2=19,594 Prob >chi2=0.0000; White General Test: 66.35169 p-value=3.7e-09; Breusch Pagan Lm statistic 19,594 p-value=0).

Note. The neperian logarithm of net hourly earnings (lnwh) is the dependent variable. IV estimations based on father's education (years of schooling and alternatively, levels of education). Robust standard errors are in parentheses.

4.1. Changing rates of return to education in Spain

The results in Table 1 can be compared to some of estimates in previous literature to assess the change in the returns to education in Spain in the last decades. Thus regarding table 1 (1), Barceinas *et al.* (2001)[17] found a stable pattern of returns to schooling during the 80's and first half of the 90's (around 7%), with small changes reflecting the GDP cycle and a rising trend in the end of that period (up to about 8%), this despite the significant increase in average number of years of schooling. According to the authors this was a consequence of changes in the structure of Spanish economy resulting in a demand for more educated workers that exceed corresponding supply. Estimated return to education in table 2(1) is significantly lower than those figures (5.87%). Given the progressive character of taxes on labour incomes (presumably a small) part of the difference may be due to the fact that dependent variable here is net incomes, while Barceinas *et al*. (2001) work with gross incomes. Other possible (although quite an unlikely one) explanation could be a dramatic change in the ability bias making more relevant the absence of control through IV. Divergences arising from the diversity in the sources of information could very likely explain another share of the difference. However, the possibility of a change in the imbalance between supply and demand for skilled workers cannot be neglected.

The comparison with the results by Pons and Gonzalo (2002), and Arrazola and De Hevia (2003) points in the same direction. In both cases the authors work with net hourly wages so previous discussion on this issue is not applicable here. Pons and Gonzalo's OLS estimates of the return to education for male salaried workers are 5.9% (based on 1994 Household Panel of the EU, PHOGUE) and 6.4%[18] (based on the 1991 Survey of Structure, Conscience and Biography of Class). Arrazola and De Hevia (2003) estimate a return of 6.4% using 1994 PHOGUE with different control variables. As discussed in section 2 the introduction of control variables has the effect of reducing the estimated coefficient of schooling attainment. The absence of such controls in Table 1 probably leads then to the overestimation of such returns when contrasted with other analyses. From the comparison of these figures a trend of diminishing returns to education in Spain seems to be unveiled. This is a pattern which seems to be shared by other EU countries while these returns are surprisingly rising in the US, something that has been argued to be the result of the inability of the European economy to create jobs with higher skills requirements at a rate sufficiently high to equal or exceed that of supply.

[17] Based on small samples containing information about head of household; gross incomes on annual basis (no information on hours of work was available).

[18] The IV estimate of the return to education is 10.7 using parents' education as instrument.

The analysis of the returns to qualifications points out in the same direction. Vila and Mora (1998) provide useful benchmark results for comparison between estimates in equation (5) in table 1 and corresponding figures in 1981 and 1991. Bearing in mind that they are based on a different source, estimated coefficients for education levels in equation (5) are on average 10 percentage points lower than those of 1991 which in their turn were (a) very similar to those of 1981 for short and long cycles university degrees, and (b) already showed a decline trend in returns to education for lower qualifications. Although it is necessary to be cautious given the different nature of the data analysed in both papers, this trend may be a reflection of the law of diminishing returns to the formation of human capital at the margin as the level of per capita income increases, as suggested by Pscharopoulos (1994). Moreover the less than proportional reduction in the returns for more educated workers could reflect the structural changes having taking place in the Spanish economy with an increase in the share of activities demanding more skilled workers. In any case, however, the increase in the supply for skilled workers seems to have exceeded the corresponding demand resulting in a declining trend for the returns to education that was not evenly distributed among qualifications.

4.2. Linearity in schooling

The estimation of the returns to schooling in table 1 as both years of education and qualifications allows assessing the linearity of those returns[19], that is, the underlying assumption in Mincer (1974) that 'each additional year of schooling has the same proportional effect on earnings, holding constant years in the labour market' (Card, 1999). As shown in table 1, after correcting for endogeneity in schooling, coefficient for category E2 (vocational and upper secondary education) is not significant, what indicates that this group's behaviour does not significantly differ to that of E1 (Primary or less than primary and secondary education below compulsory threshold age) which is group of reference, this despite the additional amount of years of schooling involved. Moreover, nor the difference between coefficients for categories E3 (short cycle university degree) and E4 (long cycle university degree) is significant, showing a similar behaviour for all workers with university degrees independent the average number of years needed to achieve them[20], a result that could be related to signalling hypothesis. These results are very clearly observed in table 2, where earning premiums associated with successive levels of education are displayed. According to OLS estimates getting a short university degree

[19] An alternative approach applied to US data by Heckman *et al*. (2003) by including indicator variables for each year of schooling allowed rejecting linearity in line with previous studies thus giving evidence of 'sheepskin effects' according to which exceptionally large rates of returns were observable in schooling levels associated with degree completion years.

[20] These results may be misleading. As shown by Heckman *et al* (2003) based on US data, the relaxation of some of the assumptions in Mincer specification shows that this type of specification dramatically understates the return to finishing high school although the differences in the estimates are less pronounced for college completion.

diploma, which on average implies three additional years of schooling when compared with the immediate lower level of education, yields a reward which is close to 31%. A figure that annualised is rather similar to the premium from long-cycle university degrees. Both figures rise dramatically in IV estimation that as pointed out significantly reduces the differences between both university qualifications. Although this is a dubious calculation given changes in Spanish educative system (a change in primary and secondary qualifications that increased minimum school leaving age) 'jumping' from E2 to reference group E1 implied 4-5 years of additional schooling for most workers in the sample. An outstanding difference is then observed in the rates of return of an additional year of schooling depending on the educative period involved. The influence of schooling attainment as a continuous variable on earnings seems to be underestimated in the case of university graduates and overestimated for lower levels of education. In both cases this pattern is underlined when potential endogeneity of schooling is considered. The assumption of linearity of schooling has proved then to be quite strong at least in the Spanish case, where simultaneously working with the two alternative specifications of education seems to be a reasonable option.

Table 2. Percentage earning premiums associated with clusters of qualifications

Differences in earnings by education level	OLS	IV	Differences in earnings by education level	OLS	IV
E2/E1	22.85	16.74	E3/E2	30.96	75.41
E3/E1	60.89	104.77	E4/E2	58.25	94.08
E4/E1	94.41	126.57	E4/E3	20.83	10.64

4.3. An insight in the multiplicative separability between schooling and experience

One of the implicit assumptions in Mincer's original specification is that of multiplicative separability between education and experience. Table 3 presents the results of an estimation where schooling is allowed to vary by experience. As observable the simple interaction term between years of schooling and experience is significant, and this is also the case when the full range of interaction terms is included. Figure 1 plots experience-earnings profiles by qualification groups as estimated in the specifications in table 3 (4) and (6). According to the predicted log earnings profiles for different qualifications a convergence trend with experience is observable. The assumption of parallelism in log-earnings experience profiles across schooling levels seems then only reasonably acceptable in the 14-28 yrs. of exp. range while convergence is clearly apparent in upper experience levels. Two significantly divergent patterns are embedded in this result: Figure 1.b shows that the assumption seems less unrealistic in the case of public sector employees, except for the group of less qualified workers, an effect that may be linked to the difficulties of entrance in that positions that preserve workers from competition independent considerations about obsolescence that may be applicable in general terms. The different job structure in both sectors can also contribute to these results.

Table 3. Education-experience interactions and employment sector

	Total		Private		Public	
	(1)	(2)	(3)	(4)	(5)	(6)
Education (years of schooling)	0.0484* (0.0019)	0.0374* (0.0027)	0.0395* (0.0021)	0.0304* (0.0030)	0.0626* (0.0044)	0.0624* (0.0073)
Experience=exp	0.0191* (0.0015)	0.0412* (0.0041)	0.0154*	0.0336* (0.0047)	0.0291* (0.0038)	0.0688* (0.0096)
Experience squared=exp^2	-0.0002* (0.0000)	-0.0008* (0.0001)	-0.0002* (0.0000)	-0.0007* (0.0001)	-0.0004* (0.0000)	-0.0012* (0.0002)
Education×exp	0.0005* (0.0000)	-0.0056* (0.0007)	0.0007* (0.0001)	-0.0045* (0.0009)	-2.24e-06 (0.0002)	-0.0072* (0.0015)
Education×exp^2		0.0001* (0.0000)		0.0001* (0.0000)		0.0001* (0.0000)
Education ^{2}×exp		0.0003* (0.0000)		0.0003* (0.0000)		0.0003* (0.0000)
Education ^{2}×exp^2		-8.32e-06* (1.10e-06)		-8.08e-06* (1.45e-06)		-5.51e-06* (1.98e-06)
Constant term	0.8993* (0.0233)	1.020* (0.0302)	0.9859* (0.0259)	1.081* (0.0324)	0.8081* (0.0658)	0.8177* (0.0932)
R2	0.2922	0.3005	0.2191	0.2241	0.3625	0.3741
n	12,169	12,169	9,867	9,867	2,302	2,302
Ho: b4=b5=b6=b7=0		F(4, 12161) = 41.94 Prob > F = 0.0000		F(4, 9859) = 23.88 Prob > F = 0.0000		F(4, 2294) = 9.74 Prob > F = 0.0000

*Coef significant at 1% level
Robust standard errors are in parentheses
Note. The neperian logarithm of net hourly earnings (lnwh) is the dependent variable.

Figure 1.a. Experience-earnings profiles (private sector)

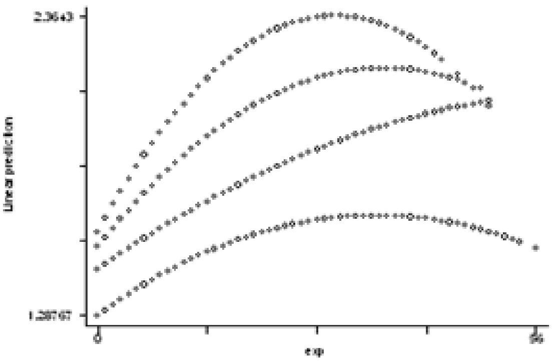
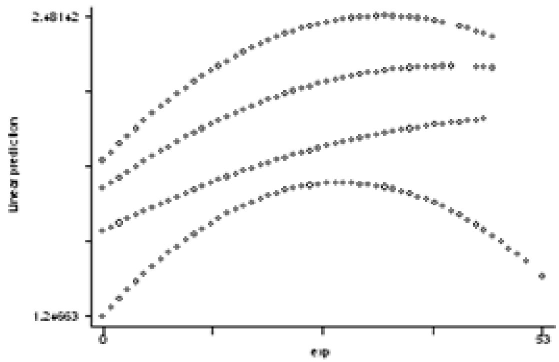


Figure 1.b. Experience-earnings profiles (public sector)



In terms of rewards for education and contrary to international evidence (see Psacharopoulos, 1994) and some previous Spanish evidence (Barceinas *et al.*, 2001), table 3 shows that when the simplest specification [1] is separately estimated[21] for both sectors the returns to education are much larger in the public sector than in the private one (6.25% and 5.18% respectively). The difference is significantly more accentuated when the whole set of interaction terms is considered (6.24% and 3.04%, respectively). When only the simplest interaction term is included this variable is not significant for public sector, thus estimated coefficient for schooling is altered only marginally, whilst it is in the private sector case, where it results in a reduction of more than one percentage point, from 5.18 to 3.95. Finally overall

[21] Detailed results for this model are not reported here but are available under request.

average returns to education fall from 5.89 (Table 1) to 4.84% (table 3-1) when the interaction between education and experience is included, and further to 3.74% (Table 3-2) when all the interaction terms are considered. We next turn to estimating the return to education by regions.

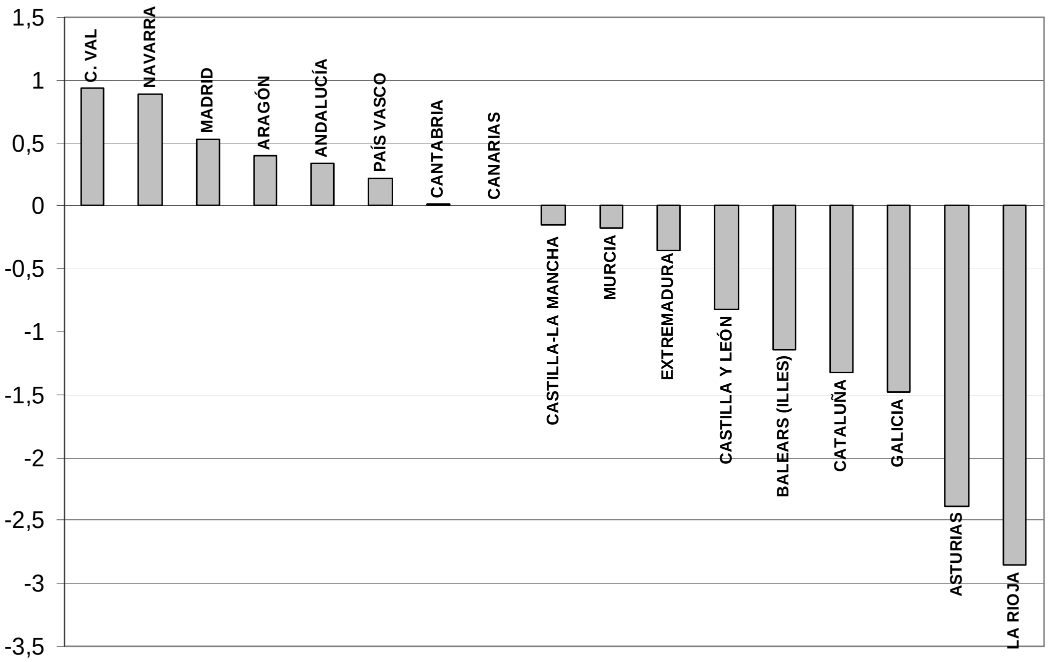
4.4. Regional differences

Education is one of the competences that have been transferred to regional governments in Spain, in a process that started soon after the Regional Fundamental Laws begun to be approved after the political transition to democracy. Not only education but also other related policies, like social and active labour market policies experienced the same process. This obviously opens a room for manoeuvre for regional governments that may be interested in promoting education (that can be fostered by increasing individual returns) as a mean for reaching higher levels of productivity and so of per capita income. Despite the fact that (a) the existence of a national common legislation on issues like the basic national curriculum and organisation of qualifications and (b) the relative short period of time elapsed in some cases since the responsibilities were effectively transferred put a doubt on the possibility that regional policies have already influenced the relative rewards to education[22] it may be pertinent to gain some insight on the territorial aspects of this rewards and to explore some of the likely explanations for these differences which are not restricted to education rewards.

Separate equations have been estimated for the 17 Spanish autonomous regions[23] to test the hypothesis of the existence of heterogeneous returns to education linked to territorial variables which seems largely confirmed by the results. Estimates of the rates of return to education in Figure 2 depict a situation of great diversity[24]. Overall IV estimates of the Mincerian returns to education from table 2 are 7.51% (7.37% when controlling by periods of unemployment). The extremes are Navarra and Comunidad Valenciana, with estimated returns over 8%, and La Rioja, with an estimated return of 4.51% to each year of additional education.

[22] The existence of interregional migration flows adds complexity to the interpretation of these coefficients.
[23] The cities of Ceuta and Melilla, in North Africa, were excluded from the estimation due to sample size restrictions.
[24] The estimated coefficient for the variable capturing the number of involuntary unemployment periods shows the expected sign in all cases except in the Comunidad Valenciana, where it is not significant (in Galicia and the Basque Country despite not being significant the sign is as expected). Sign for the square of experience is consistently negative, although not significant in Extremadura.

Figure 2. Returns to education (region-Spain)



Comparative international analyses have underlined the relevance of factors like (a) per capita income, (b) average educational attainment and (c) the percentage of GNP spent on education[25] when explaining cross-country differences in the returns to education. However, when regional returns to education are plotted against the corresponding GDP per capita and the expenditure on education, both variables fail to provide a convincing argument about the origin of the observed regional differences in the returns. This is quite in line with recent international evidence based on comparable data (Trostel *et al*., 2002). Regarding the relative supply of skilled workers as measured through the share of actives holding a university degree, a slightly positive relationship is observed (thus indicating that in general terms skilled supply is corresponded by a high demand for educated workers, although an extremely diverse pattern emerge. Thus according to *Working Force Survey* (2003) the share of university graduates over actives was 25.1 for Spanish male actives and reaches its highest level in the Basque Country (39%) and Madrid (33.7%), among other regions, with returns to education higher than Spanish average. However, other regions with equally relatively large returns to education exhibit low shares of university graduates (i.e. Valencia, 20.5%). Diverse combinations of supply and demand for skilled workers, probably related whith demography and job structure underlie these results which need deeper analysis.

[25] Although a recent paper based on comparable data (Trostel *et al*., 2002) found only slight relationships between these explicative factors and returns to education by country.

5. CONCLUSIONS

Answering to the title of the paper, to study in Spain seems still to be a profitable investment. Despite the evidence of a trend of reduction of returns to education both before and after accounting for endogenity in schooling, rates of return which have been estimated to be in the environment of 6-7% do not seem to be in risk of being seriously challenged by alternative investments. However, education may not be equally rewarding across Spanish regions for which a fork of 4.4-8.4% has been estimated. Tentative explanations have been informally tested in the paper, although a definitive more convincing explanation needs deeper research.

Regarding the plausibility of assumptions underlying Mincer's specification, the paper provides evidence against the multiplicative separability of education and experience and explores the differences between public and private sector underlying the results. Unlike international and some Spanish literature returns to education seem to be largely higher for public sector employees than for their private sector counterparts. Finally linearity of earnings in schooling is explored through the calculation of the earnings premium associated with consecutive levels of education. The polarisation of the rewards in two groups: university graduates versus the rest of salaried gives some evidence of nonlinearity which is confirmed when relative returns to diverse qualifications are calculated, which implies the implicit assumption that every additional year of education is equally rewarded independent the education level involved.

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