ON THE REWARDS TO EDUCATION IN SPAIN: ENDOGENEITY AND REGIONAL DIFFERENCES.

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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 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 to better approach the underlying concept, this being specially relevant given high unemployment rates in Spain and average length of these periods among certain groups. 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. The consideration of schooling attainment as qualifications allow relaxing Mincer's underlying hypothesis of linearity of the returns to education in schooling. Evidence against this assumption is displayed. We also test the parallelism of log-earnings experience profiles across schooling levels. The empirical analysis is finally extended by focusing on regional differences.

JEL Classification: J24, I21
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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 increase in the demand for educated workers. A rising demand that has been encompassed on the supply side by a dramatic increase in the share of population holding formal qualifications in a process fostered, among other factors, by the high figures of unemployment among the younger cohorts.

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 large number of contributions to the specialised literature

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This paper examines these issues through one of the most popular approaches, Mincer's original specification of his human capital earnings function, where the log of the hourly net wage is the dependent variable and regressors include (a) schooling attainment -that has been considered here both as years of schooling and as qualifications- and (b) a quadratic term of experience which is estimated in a way that is argued to fit the underlying concept in Mincer (1974). Unlike much of previous literature in the article actual and not potential experience is computed and controlled by the existence of interruptions due to unemployment. The constant and an error term complete the estimated model.

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 criticisms. These refer, among other aspects, to Mincer's underlying assumptions of absence of direct costs for education, linearity in schooling, multiplicative separability between schooling and experience components of earnings, and the absence of endogeneity of education. After arguing that direct costs for education may be negligible in the Spanish case according to previous analyses, the paper addresses the rest of these issues. In short, several questions are to be answered in this paper. To what extent have private returns to education changed as a result of 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? The focus of the paper is then extended by considering regional differences through separate estimations for each of the 17 Spanish regions, and some alternative explanations for the substantial diversity which is observed are briefly examined.

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, and surveys some recent developments in the Spanish literature on the issue. Section 3 describes the data set used and discusses 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 and a reference to the Spanish case

As is well known, one of the methods conventionally used to measure the influence of education on earnings is the estimation of so-called Mincerian functions in which the log of individual earnings is considered to be explained by both schooling attainment and a quadratic working experience term. This approach (Mincer, 1974) where previous research by Mincer and other authors was re-elaborated, leads to the following empirical specification,

$$logY_i = \beta_o + \beta_1 S_i + \beta_2 X_i + \beta_3 X^2_i + u_i \qquad (1)$$

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 in Heckman, Lochner &Todd (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.

Since its formulation *Mincer's earnings function* is one of the empirical specifications more widely estimated in different times and places. Most text books describe Mincer's equation as the culmination of the analysis of the implications of the human capital model for the age-earnings profile and state that it still provides a reasonably accurate insight in this relationship not only in the United States, but also in the labour markets of many other countries despite the diversity observable in their respective labour market institutions, and despite the various critiques to which it has been subject, some of which will be addressed later in this paper. Two different theoretical motivations led Mincer to equation 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 in 1974). 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. As pointed out before this rate of return to schooling can be then compared with the interest rate to test the optimality of investment in education from the individual's point of view.

The already mentioned triviality of direct costs of education is one of the assumptions embedded in Mincer's earnings function. Heckman et al. (2003) have listed some others 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.

The plausibility of this last assumption has been challenged by numerous articles, to the point that it has become the object of interest of a large deal of analyses where Mincer's specification is estimated. The basic argument here is that it may exist some unobserved forces may simultaneously exert an influence on (a) workers' schooling attainment and (b) their earnings. As pointed out in Bound, Jaeger & Baker (1995) 'ability' may be one of these forces, since it is quite likely correlated with earnings (those with higher ability earn more) and also with schooling attainment (those with higher ability gain higher qualifications). Different strategies have been adopted to deal with the bias arising from endogeneity. Instrumental variables techniques (IV) are one of the most widely used. This requires finding an instrument which is uncorrelated with the true measure of schooling and uncorrelated with the measurement error (see Card, 1999 and 2000; Harmon, Oosterbeek & Walker, 2003, and Heckman et al., 2003, for comprehensive critical surveys of instruments used in recent international literature). Other approaches include 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 (see Heckman and Li, 2003).

Some of the latest contributions to the Spanish literature regarding the estimation of Mincer's equation already introduce controls for the endogeneity of education through IV techniques that resulted in a new 'generation' of analyses. Previous pieces of work are surveyed in Oliver, Raymond, Roig & Barceinas (1999), Arrazola, De Hevia, Risueño & Sanz (2003). More recent contributions offer estimations that have been obtained by the consideration of both supply (like proximity to college or Education Laws reforms) and demand-side instruments like the more traditionally used instruments such as family background including parent's education, for example. Results are difficult to compare given the diversity of sources and the inclusion of different restrictions in the sample finally used, and by the inclusion of different sets of control variables. This is something also very common in international literature. In fact most frequently empirical exercises have included these other variables considered as relevant for the explanation of wage differences together with schooling and experience quadratic term on the right side of equation (1). Their inclusion in the estimations has the typical effect of increasing the coefficient of determination. However it 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 inclusion of these independent variables have the practical effect of reducing the estimated returns to schooling due since they in fact capture the indirect mechanisms through which the educated achieve better wage conditions. This seems to be especially true in the case of choice variables linked with job characteristics and therefore endogenous (see Barceinas, Oliver, Raymond & Roig, 2001 and 2002, drawing on Mincer's discussion on the issue). In Barceinas et al. (2001) for example the inclusion of the highest number of control variables reduced returns to education from 8.2% to 6.5%.

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 Arrazola et al. (2001) estimate that opportunity costs account for more than 90 per cent of total private education cost. Public administration's strong subsidisation scheme (that assumes more than 80 per cent of total direct costs) leads to largely similar private returns when they are estimated 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

Our data are drawn 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). This is an annual study conducted since 1999 which contains information on a wide number of socio-economic, satisfaction 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. 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 & San Segundo, 1995 and Arrazola & De Hevia, 2003 and 2006, among others using different databases for Spain), since no information is available about the unemployed. However, as pointed out in 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 in Arrazola and De Hevia (2003) for example 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 with the aim of capturing both the impact of inflation and potential structural change.

Relevant variables for this study include net revenues, schooling attainment and working experience. Variable 'schooling attainment' has been addressed in two ways. First, as in many studies, categorical variable 'maximum qualification obtained by the individual' has been transformed to fit the original equation proposed by Mincer. Thus, minimum number of years necessary to gain a qualification was calculated to generate a new variable using information provided by the survey to allow the calculation of the internationally-comparable return of an additional year of full-time education. Alternatively, dummy variables were introduced for the different levels of qualifications originally considered in the survey. Education groups were previously clustered to allow correct estimation given the number of available observations, 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. 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 with the aim of improving 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.)

Actual and not potential experience is used in the estimation. Variable experience has frequently (in a tradition inaugurated by Mincer, 1974), been defined as age less years of schooling (in the Spanish case this is usually the minimum number of years necessary to complete the stated level of education), 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. As a reaction to these concerns variable experience has frequently been substituted by variable age, as recently done in the Spanish case in Barceinas et al. (2002) and Pons & Gonzalo (2002). In this paper variable experience is defined as the age of the individual at the time of the survey less the age at which he obtained tenure of at least three months. This alternative can also be subject to criticism, above all in the last years when many young workers experience a concatenation of extremely short contracts sometimes for years at the beginning of their working life. This effect should not be however largely relevant for total population, and is probably restricted to those cohorts. To more accurately approximate actual experience an additional control variable was considered, including the number of periods of involuntary unemployment experienced by the individual. 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 formation (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, a vast corpus of literature 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 the endogeneity of schooling attainment in the sample. Thus Hausman test (Hausman, 1978) resulted in the rejection of the null hypothesis of exogeneity of schooling attainment and therefore OLS gives biased and inconsistent estimates of the causal effect of schooling attainment on earnings. Regarding IV estimations, as pointed out in section 2, family background is one of the most popular set of instruments in this context (Card, 1999). There is a strong correlation between an individual’s education and that of his parents. The latter variable can then be considered as a potential instrument. According to Mora (1996) the access to post-compulsory education, both secondary and university degrees, is strongly linked in Spain to economic and education characteristics of family environment. Moreover previous analysis by Pons and Gonzalo (2002) has shown that parent’s education and college availability (existence of a university in the province where the individual lived at the age of 16) are 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 is then maximum qualification obtained by the head of household at the age of 16 (most frequently father's education). As pointed out in Bound et al. (1995) if the instrument is only weakly correlated with schooling attainment then besides obtaining estimations with large standard errors, even a weak correlation between father's education and the error in the original equation could lead to a large inconsistency in IV estimates likely to even surpass that of OLS estimates. The F statistic on excluded instruments test confirmed head of household education as a satisfactory instrument.

4. Results

Table 1 displays the results obtained from the estimates based on Equations (1) and (2). Both OLS and IV techniques were used. Surprisingly in line with recent international comparative exercises (Trostel, Walker & Woolley, 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. All OLS estimations were carried out with standard errors robust in heteroscedasticity, by means of the White's variance and covariance matrix after confirming the existence of heteroscedasticity usually associated with cross-sectional data by the usual tests (Cook-Weisberg chi2; White General Test; Breusch Pagan LM statistic) in which the null hypothesis of homoscedascity was rejected.

4.1 Changing rates of return to education in Spain

The results in Table 1 can be compared to some of those in previous literature to assess the change in the returns to education in Spain in the last decades. Among others, two papers provide estimations that can be used as a reference. In terms of equation 1 in table 1, comparable estimations Barceinas et al. (2001) depicted a stable pattern of returns to schooling during the 80's and the 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 per capita years of schooling of the labour force. This is explained by the authors in relation with changes in the structure of Spanish economy so that demand for more educated workers overtook supply. Estimated return to education in table 1(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 from Pons & Gonzalo (2002) and Arrazola & 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, ECHP) and 6.4% (based on the 1991 Survey of Structure, Conscience and Biography of Class). The IV estimate of the return to education is 10.7 using parents' education as instrument. The estimation in Arrazola and De Hevia (2003) results in 6.4 per cent return using 1994 ECHP with different control variables. As discussed in section 2 the introduction of control variables has always the effect of reducing the estimated coefficient of schooling attainment. The absence of such controls in Table 1 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 apparently shared by other EU countries while these returns are surprisingly rising in the US, what has been interpreted as the result of the inability of the European economy to create jobs with higher skills requirements at a rate sufficiently high to equals or overpass that of supply.

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 education	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
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 in parentheses.

The analysis of the returns to qualifications points out in the same direction. Thus Vila & Mora (1998) provide useful benchmark results for comparison between estimates in equation 5 in table 1 and corresponding figures in 1981 and 1991. Although it has to be taken into consideration 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 declining 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 Psachapoulos (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, 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, a result that could be related to the 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 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 groups of qualifications

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

4.3 An insight in the multiplicative separability between schooling and experience

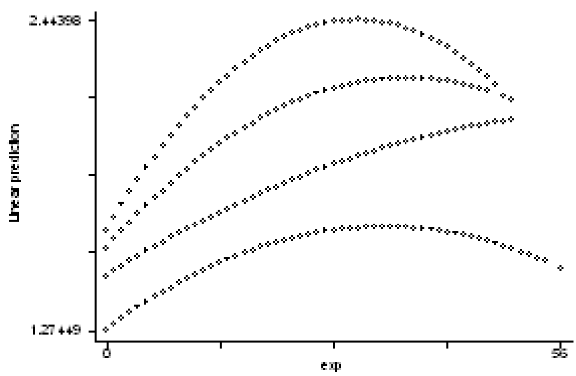
As pointed out in the introduction of this paper, one of the implicit assumptions in Mincer's original specification is that of multiplicative separability between education and experience. Table 3 presents the results of a model where schooling is allowed to vary by experience. A set of interaction terms between both variables (as linear and quadratic terms) was included as a regressor in the estimation of equation (1) with the aim of testing this condition. 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–column 2. 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 reasonable in the 14-28 yrs. of exp. range while convergence is clearly apparent in upper experience levels. Overall 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. It is necessary to point out that these results are not only compatible with the existence of different return to experience profiles between educational groups. Since there is a correlation between experience and age, another interpretation would be that different age cohorts who made educational choices under different educational regimes experience different returns to the respective levels of education. As long as cohorts are not followed over time these two concepts cannot be separated empirically. We next turn to estimating the return to education by regions.

Table 3. Education-experience interactions

	Total	
	(1)	(2)
Education (years of schooling)	0.0484* (0.0019)	0.0374* (0.0027)
Experience=exp	0.0191* (0.0015)	0.0412* (0.0041)
Experience squared=exp2	-0.0002* (0.0000)	-0.0008* (0.0001)
Education*exp	0.0005* (0.0000)	-0.0056* (0.0007)
Education*exp2	-	0.0001* (0.0000)
Education 2*exp	-	0.0003* (0.0000)
Education 2*exp2	-	-8.32e-06* (1.10e-06)
Constant term	0.8993* (0.0233)	1.020* (0.0302)
R2	0.2922	0.3005
N	12,169	12,169
Ho:b4=b5=b6=b7=0	-	F(4, 12161) = 41.94 Prob > F = 0.0000

*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. Experience-earnings profiles



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. This obviously opens a room for manoeuvre for regional governments that may be interested in promoting education (which can be fostered by increasing individual returns) as a mean for reaching higher levels of productivity and so of per capita income. Despite (a) the existence of a national common legislation on issues like the basic national curriculum and organisation of qualifications (b) and the relative short period of time elapsed in some cases since the responsibilities were effectively transferred puts a doubt on the possibility that regional policies have already influenced the relative rewards to education[1] 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[2] to test the hypothesis of the existence of heterogeneous returns to education, something which seems largely confirmed by the results. Estimates of the rates of return to education in Table 4 offer a picture of great diversity[3]. Overall IV estimates of the Mincerian returns to education from table 2 are 7.51% (7.37% when controlling by periods of unemployment). Table 4 shows that the extremes are Navarra and Comunidad Valenciana, with estimated returns over 8%, and La Rioja, with an estimated return of 4.51% per year of additional education.

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 in explaining cross-country differences in the returns to education. Although a deeper analysis of this issue exceeds the scope of this paper, figure 2 and 3 plot regional returns to education against the corresponding GDP per capita and the expenditure on education, both failing to provide a convincing argument about the origin of the observed regional differences in the returns. This is quite in line with international evidence based on comparable data (Trostel *et al*., 2003).

Regarding the relative supply of skilled workers as measured through the proportion of active population holding a university degree, a slightly positive relationship seems to be unveiled by the tendency curve in Figure 4 (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 *Labour Force Survey* (2003) the share of university graduates over actives was 25.1 for Spanish male actives and reached its highest level in the Basque Country (39%), Navarre (33.8%) and Madrid (33.7%), Aragon (29.6%) and Cantabria (29.4%), in all cases regions where the returns to

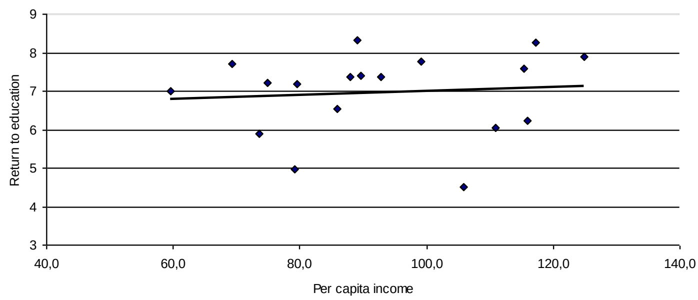
[1] The existence of interregional migration flows adds complexity to the interpretation of these coefficients.

[2] The cities of Ceuta and Melilla, in North Africa, were excluded from the estimation due to size restrictions.

[3] 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 the sign is as expected despite not being significant). Sign for the square of experience is consistently negative, although not significant in Extremadura.

education are higher than Spanish average. However, other regions with equally relatively large returns to education exhibit low shares of university graduates. This is the case of Valencia (20.5%) or Andalusia (20.8%). In some cases a small share of university graduates is associated with low returns (this is the case of the Balearic Islands, 16.1%, Castilla-La Mancha, 16.9%, Extremadura, 18%, the Canary Islands, 20.7%, Murcia, 21.8%, Galicia, 22.7% and La Rioja, 24.9%). Asturias, Catalonia and Castilla y León are on their hand examples of regions where a higher-than-average presence of active holding a university degree (27.8%, 26.2% and 25.5%) coexists with returns lower than Spanish average, especially in the former case. Diverse combinations of supply and demand for skilled workers underlie these results which need deeper analysis.

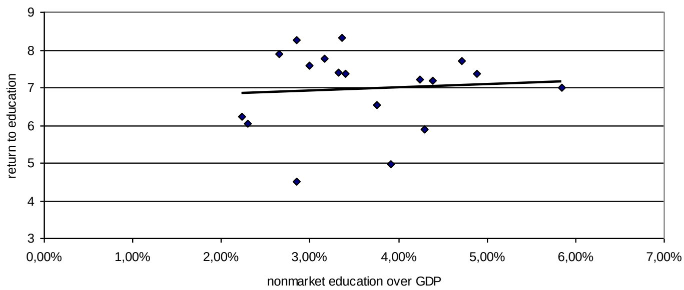
Figure 2. GDP per capita (PPS, UE-25 2000-2003 = 100) and return to education across regions



Source: Table 4 and Spanish Regional Accounts, INE (National Institute of Statistics)

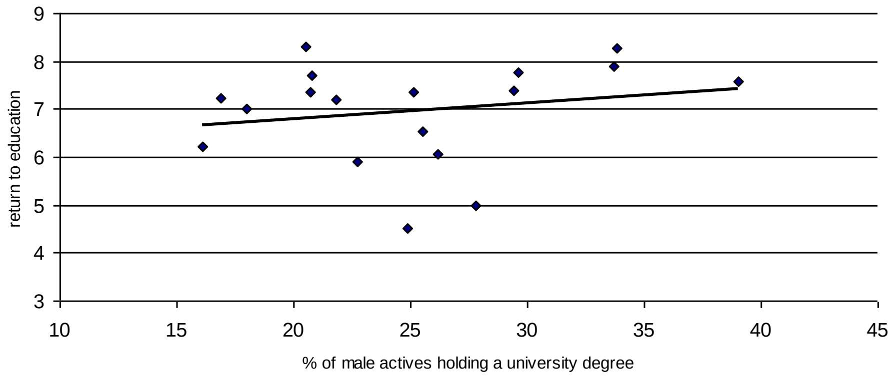
Two additional explanations can be added: First, in their analysis Pastor et al. (2008) show that the economic sector can be a significant factor in explaining the regional component of returns to education (the election of more rewarding economic sectors of activity as an indirect way for making private educative investment profitable). Second, Simón et al. (2006) underline the very significant role that institutional aspects, and specifically the collective bargaining system, play in explaining regional wage differences in Spain, which are found to be larger than in many other countries. Thus in their analysis it is shown that inter-regional differences in both actual and bargained wages are extremely similar. This fact points at the predominately sub-national (provincial and regional) structure of collective bargaining as a key factor in explaining regional wage differences. A non-competitive factor that could equally explain at least part of the differences observed in Table 4.

Figure 3. Expenditure on education (nonmarket education as % GDP) and return to education across regions



Source: Table 4 and Spanish Regional Accounts, INE (National Institute of Statistics)

Figure 4. Share of male actives holding a university degree and returns to education across regions



Source: Table 4 and Labour Force Survey 2003, INE (National Institute of Statistics)

Table 4. Wage equation. Salaried men by region of residence (IV estimation)

	Andalucía		Aragón		Asturias		Islas Baleares	
	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff
Schooling	0.0807* (0.0051)	0.0771* (0.0055)	0.0765* (0.0117)	0.0777* (0.0117)	0.0531* (0.0138)	0.0498* (0.0142)	0.0614* (0.0132)	0.0623* (0.0134)
Exp	0.0238* (0.0033)	0.0251* (0.0033)	0.0340* (0.0047)	0.0332* (0.0047)	0.0308* (0.0053)	0.0331* (0.0054)	0.0199* (0.0047)	0.0241* (0.0049)
Exp^2	-0.0002* (0.0000)	-0.0002* (0.0000)	-0.0004* (0.0001)	-0.0004* (0.0001)	-0.0004* (0.0001)	-0.0005* (0.0001)	-0.0002*** (0.0001)	-0.0003** (0.0001)
Unempl.[1]		-0.0096* (0.0037)		-0.0229* (0.0047)		-0.0194** (0.0092)		-0.0152* (0.0052)
Constant	0.4965* (0.0647)	0.5406* (0.0695)	0.4209* (0.1449)	0.4540* (0.1458)	0.7275* (0.1741)	0.7727* (0.1816)	0.7532* (0.1440)	0.7150* (0.1455)
N	1362	1222	502	490	556	548	448	428
R^2	0.2950	0.2997	0.2128	0.2235	0.2104	0.2310	0.2627	0.2867

	Canarias		Cantabria		Castilla-La Mancha		Castilla y León	
	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff
Schooling	0.0794* (0.0094)	0.0737* (0.0099)	0.0798* (0.0175)	0.0696* (0.0097)	0.0655* (0.0101)	0.0604* (0.0082)	0.0605* (0.0082)	0.0739* (0.0170)
Exp	0.0231* (0.0048)	0.0222* (0.0049)	0.0334* (0.0057)	0.0243* (0.0038)	0.0257* (0.0004)	0.0245* (0.0030)	0.0262* (0.0030)	0.0360* (0.0055)
Exp^2	-0.0003* (0.0001)	-0.0002** (0.0001)	-0.0004* (0.0001)	-0.0003* (0.0000)	-0.0003* (0.0000)	-0.0003* (0.0000)	-0.0003* (0.0000)	-0.0005* (0.0001)
Unempl.[1]		-0.0213* (0.0070)			-0.0147* (0.0048)		-0.0169** (0.0075)	-0.0550* (0.0153)
Constant	0.4806* (0.1228)	0.5720* (0.1314)	0.5130** (0.2136)	0.6701* (0.1162)	0.7246* (0.1221)	0.7369* (0.0964)	0.7411* (0.0968)	0.6189* (0.2114)
N	592	555	382	697	673	1241	1203	367
R^2	0.2503	0.2757	0.2685	0.3760	0.3857	0.2896	0.2991	0.3193

	Catalonia		C.Valenciana		Extremadura		Galicia	
	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff
Schooling	0.0722* (0.0071)	0.0830* (0.0085)	0.0831* (0.0091)	0.0649* (0.0075)	0.0701* (0.0083)	0.0531* (0.0124)	0.0589* (0.0121)	0.0722* (0.0073)
Exp	0.0337* (0.0045)	0.0289* (0.0034)	0.0281* (0.0035)	0.0081** (0.0040)	0.0106* (0.0041)	0.0300* (0.0043)	0.0306* (0.0044)	0.0335* (0.0046)
Exp^2	-0.0004* (0.0001)	-0.0004* (0.0001)	-0.0003* (0.0001)	-0.0000 (0.0001)	-0.0000 (0.0001)	-0.0004* (0.0001)	-0.0004* (0.0001)	-0.0004* (0.0001)
Unempl.[1]			0.0012 (0.0045)		-0.0088** (0.0042)		-0.0124 (0.0092)	-0.0024 (0.0049)
Constant	0.5049* (0.0876)	0.4118* (0.0998)	0.4265* (0.1046)	0.7478* (0.0889)	0.6567* (0.0972)	0.5445* (0.1479)	0.5093* (0.1460)	0.5047* (0.0916)
N	603	913	876	557	483	683	653	590
R^2	0.2650	0.2155	0.2211	0.2887	0.3271	0.2484	0.2537	0.2642

	Madrid		Murcia		Navarra		Basque Country		La Rioja	
	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff	Coeff
Schooling	0.0792* (0.0075)	0.0790* (0.0077)	0.0841* (0.0119)	0.0762* (0.0128)	0.0442* (0.0124)	0.0451* (0.0120)	0.0759* (0.0129)	0.0826* (0.0120)	0.0728* (0.0087)	0.0719* (0.0091)
Exp	0.0300* (0.0037)	0.0309* (0.0036)	0.0275* (0.0076)	0.0311* (0.0048)	0.0242* (0.0049)	0.0257* (0.0050)	0.0319* (0.0048)	0.0300* (0.0079)	0.0203* (0.0038)	0.0206* (0.0038)
Exp^2	-0.0003* (0.0001)	-0.0003* (0.0001)	-0.0002 (0.0002)	-0.0003* (0.0001)	-0.0003* (0.0001)	-0.0003* (0.0001)	-0.0004* (0.0001)	-0.0003*** (0.0002)	-0.0002* (0.0001)	-0.0002* (0.0001)
Unempl.[1]		-0.0105*** (0.0061)				-0.0218* (0.0083)	-0.0046 (0.0063)	-0.0111* (0.0035)		-0.0070*** (0.0041)
Constant	0.4286* (0.0996)	0.4242* (0.1042)	0.5258* (0.1619)	0.5198* (0.1704)	0.9004* (0.1512)	0.8943* (0.1468)	0.5220* (0.1730)	0.5651* (0.1661)	0.5767* (0.0991)	0.5918* (0.1037)
N	1234	1209	412	580	373	363	551	398	645	629
R^2	0.3226	0.3296	0.3226	0.2888	0.2472	0.2662	0.2973	0.3408	0.3550	0.3536

5. Conclusions

Education seems still to be a profitable investment in Spain. 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 for such differentials 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. 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 annualised returns to education are compared to the estimated coefficient for the continuous variable 'years of schooling', something that challenges the implicit assumption that every additional year of education is equally rewarded independent the education level involved.

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