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Private rates of return to human capital in Spain: new evidence[1]

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I. Introduction

Education is the most direct form of investment in human capital that individuals may undertake. The estimation of the private returns on that investment is one of the most widely studied topics in labour economics, giving rise to a huge empirical literature. Our objective is contributing to this literature from the perspective of the Spanish case.

The Spanish economy has undergone a rapid process of structural change during the last two decades. One of the factors underlying this change is the improvement in the quality of labour. Although the educational level of the Spanish population is lower than the European average, during the last two decades an important increase in the stock of human capital of the labour force has taken place. While the share of the labour force holding post-compulsory qualifications in 1980 was 14.5%, this share rose to 35.3% in 1998. It was in the eighties when a huge public financial effort made available post-compulsory education to larger sections of the schooling age population through public supply or publicly subsidised private supply. This increase has affected both male and female population, but in the last case the effect has been stronger. This has had as a result a growing female participation in all the cohorts.

The fact that the Spanish economy has been suffering since the beginning of the eighties the highest unemployment rate in Europe is probably not neutral to the explanation of the increasing demand for education. A reduction in private opportunity costs of further education and the differentials both in wages and job probability in favour of those with higher education has, probably, led to a rising in the decisions to continue into higher education for a growing proportion of youth.

A large stream of Spanish evidence has been published for the last ten years. The availability of different micro data sets during those years has made progress on the field possible. Several conclusions can be drawn from this literature. Firstly, returns to education, when years of schooling is used as independent variable, tend to show a range of values between 5 and 7%.

These Figures reflect different weights from gender and sector. Also, higher returns for women than for men and higher returns for private than for public sector is a common finding. Secondly, when using qualifications as independent variable, results tend to show linearity. Thirdly, return differentials between compulsory and higher education tended to increase during the eighties. Fourthly, those authors trying to test for signalling effects have only found evidence of weak signalling.

Our results tend to confirm those findings for different samples. Moreover, they show robustness to different specifications departing from the basic mincerian model. However, we find a difference in the level of returns. In our case, returns take values around 8%. Most of the previous Spanish evidence is drawn from data of 1990 or 1991. Our results for 1994 and 1995 would imply a slight increase in the returns to education during the first half of the nineties. This is a remarkable result given the growth in the supply of educated workers during the period.

The datasets used were the Household Budget Survey 1990/91 (HBS 90/91) that offers information about all members of 20.000 households, especially in those aspects related to qualifications attained, annual net income, as well as their employment status. Unfortunately, information about hours worked is not available. The Continuous Household Budget Survey 1985-1996 (CHBS 1985-1996) is a quarterly survey with a sample based on 3.000 households. Variables are the same than those in the HBS 90/91, but are provided only for the head of household. The Household Budget Survey 1980 (HBS 80) has the same structure than CHBS but with a sample of 24.000 households. The Wage Structure Survey 1995 (WSS-95) is a employer survey of 175.000 wage earners, which contains an important amount of characteristics related to each worker (qualification, tenure, type of contract, type of job, sector, firm size, and so on). Wages are gross and net and they are provided on hourly, monthly and annual basis. Finally, the European Household Panel 1994 (ECHP 94) offers information about 8.000 surveyed households. Basic personal characteristics are provided for each individual as well as labour market status. For the employed, information is given on gross and net wages, and worked hours.

This survey provides information about educational level, and also on age leaving education, which allow us to approximate 'real' years of schooling.

All surveys were purged dropping those observations with wages below minimum wage, younger than 18 years and older than 65 years and, in ECHP 94, individuals whose approximated 'real years of schooling' were evidently atypical. In the case of male observations, only full-time workers were included in the used samples. For female workers both full and part-time workers made up the samples but a control dummy is included.

In order to make easier the comparability of surveys, all the considered wages were re-converted to gross wages before taxes and social security payments. This process was carried out with a specific program, and according to the information from the taxable units. Our work makes new contributions to the Spanish evidence. In the first place, we have compiled a homogeneous database in terms of wage definition. In this sense all estimations are run using gross wages as dependent variable. Most of the Spanish literature is based on a definition of net wage close to the concept of 'take home income'. So our results are more clearly interpretable in terms of the effect of human capital on productivity. In the second place, new contributions are made on three topics. Firstly, we control for endogeneity of schooling by using instrumental variable estimators. The results show stability of the results under conditions specified later in the text. Secondly, we introduce the effect of unemployment on returns to education under different hypothesis by using a internal rate of return approach whose results are consistent with those obtained from a standard mincerian specification using qualification dummies. Finally, we test sheepskin effects by estimating the effects on returns to education of repeated years to get a degree, and we present several tests in order to demonstrate the validity of the human capital theory versus the signalling theory.

The chapter is organised in six additional sections. The second section establishes the cross-section and time series results of returns to education from a basic mincerian model. After that, we test in the third section the robustness of the basic model results by introducing different controls on that model. Section four introduces the effect of unemployment on returns. In section five, the possibility of signalling and sheepskin effects is tested with different procedures. The sixth section addresses the problems of schooling endogeneity and instrumental variable estimations. The main conclusions drawn from the previous work are given in the final section.

II. Private returns to human capital in Spain in a basic mincerian framework

In this section we deal with the main results obtained from a parsimonious mincerian model estimated by OLS, with years of education and quadratic on potential experience as independent variables and logarithm of hourly gross wages (including all Social Security payments)[2] as dependent variable. We use a sample of full-time workers from ECHP-1994 and WSS-1995.[3][4] Table 1 shows the results obtained using years of schooling[5] in 1994 and 1995, with values ranging between 7.5-10%. From these results, three main simple facts can be pointed out. Firstly, full-time male and female workers have returns to schooling around 8% per annum. Secondly, full-time female returns are slightly higher than men's. And, finally, when in the WWS-1995 part-time female workers are controlled for, female returns are still higher than in the previous case.

The fact that we are working with an employer survey (WWS-1995) and a household survey (ECHP-1994) may produce some distortions in the comparability of working hours. The latter may be behind the dissimilarities in the results obtained for men.[6] Because we might grant a higher reliability in wages and working time data to employer surveys, returns to education for full-time male workers in Spain in the mid Nineties must be slightly higher than 8%. For women, rates of return estimates depend on the inclusion of a dummy

[2] Because employer costs include all Social Security payments, we consider that the wage definition has to include them.
[3] In annex II it can be found the principal characteristics of different datasets used, and also the sample selection criteria. This applies to the whole work.
[4] Workers with earnings below minimum wage in annual terms were excluded from all samples.
[5] ECHP-1994 provides information on age leaving education. Therefore, in this case it is possible to aproximate the real years of schooling for each degree. But in order to compare the different datasets, the results obtained are allocated to the rest of sources.
[6] Returns to each year of schooling range from 7.46% in 1994 to 8.20% in 1995.

variable that controls for the part-time.[7] While full-time female workers show almost the same return to schooling in both years and surveys (8.29% in 1994 and 8.27% in 1995), when part-time is controlled for a quite important difference appears: from 8.28% in 1994 to 10.02% in 1995. Again, since this 10% comes from a sample that was addressed to employers this result seems to be more reliable.

Table 1. Returns to additional year of schooling (%). Gross hourly wages

	Men	Women[1]	Women[2]
ECHP-1994	7.46	8.29	8.28
WSS-1995	8.20	8.27	10.02

1. Only full time. 2. All the sample controlling for part time.

Generally speaking, it seems that the similarity between full-time male and female rates and the differences between returns to schooling when female part-time is controlled for, could be explained by the following reasons. Firstly, dissimilarity between sample sizes used in both surveys: while the WSS-1995 includes 118.027 observations for males and 30.769 for females, in the ECHP-1994 we have only 2.181 for males and 848 for females. Secondly, sample selection bias can affect women estimates (see later on).

The only Spanish data set that allows estimation of our mincerian equations for a series of years to track the evolution of returns through time is the CHBS 1985-1996. This source presents two limitations. First, the necessary data on earnings is only given for head of households and on annual basis (no information on hours of work), and, secondly, sample size is rather small.

We tested whether results coming from that source were consistent with those obtained from larger samples using also heads of households for some of the years for which CHBS and that other larger samples were available. These larger samples of heads of households are available in two data sets: HBS-1990/91 and ECHP-1994. The returns obtained from samples of male head of households for annual earnings were 6,9% in 1990 and 7,2% in 1994. These

[7] In the Spanish case and in those years, the proportion of female working in part-time was near 15%.

returns are almost identical to those found using the CHBS samples for the same years. Therefore, one should be reasonably confident on the consistency of the results obtained using the CHBS series. That is, it is possible to claim that there are not big differences between the return for heads of households and the return for the whole population. Additionally, HBS-1980/81, which has exactly the same characteristics than CHBS, except for a larger sample, was included in the time series.

Estimated equations are shown in Table 2 and graphed in Figure 1. The time profile of returns shows three phases: an increase from 1980 to 1985, a slight decreasing trend from 1985 to 1993 and, finally, a rapid increase from that year until 1996, last year of the sample. Altogether, the range of returns runs from 5.9 in 1980 to 8.1 in 1996. In order to test the equality of coefficients over time, temporal dummies with schooling interactions were included in the equation. Only in one case the coefficient was not significant and the test about structural change showed statistical significance.[8] It should be mentioned that variations in returns over time seem to be related to the GDP cycle. Therefore, these small changes in rates of return may reflect differences in demand for human capital depending on cyclical sensitivity.

[8] First, results from a restricted estimation (with all years) where compared with the unrestricted yearly models. Wald test, likelihood ratio and Lagrange multipliers were 182, 181 and 179 respectively. That implies the rejection of the null of equality of return rates.

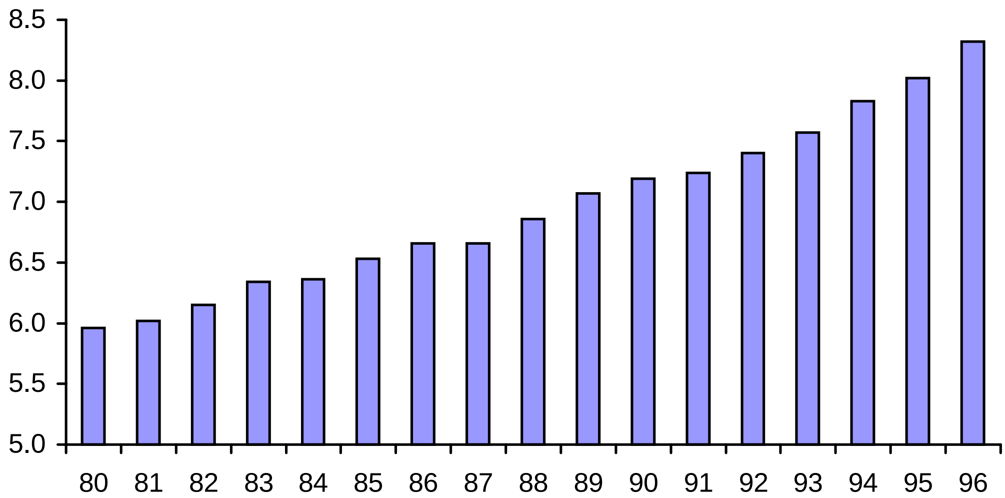
Table 2. Rates of returns to education 1980, 1985-1996. Heads of household.
Annual gross wages.

	Constant	Schooling	Experience	Experience^2	n	R^2
1980	13.6347	0.0594	0.0277	-0.0004	6948	0.33
	463.2	*55.2*	*17.2*	*15.8*		
1985	12.9910	0.0756	0.0380	-0.0006	901	0.31
	111.9	*18.7*	*5.3*	*-4.8*		
1986	13.0360	0.0728	0.0429	-0.0007	782	0.29
	99.9	*13.3*	*5.8*	*-5.5*		
1987	13.2124	0.0695	0.0355	-0.0005	874	0.26
	121.3	*15.0*	*5.6*	*-5.0*		
1988	13.1537	0.0713	0.0433	-0.0006	842	0.24
	107.2	*15.4*	*5.9*	*-4.9*		
1989	13.2014	0.0707	0.0492	-0.0007	818	0.26
	103.6	*15.9*	*6.3*	*-5.7*		
1990	13.2016	0.0705	0.0545	-0.0008	767	0.27
	111.5	*13.9*	*8.3*	*-7.6*		
1991	13.2732	0.0686	0.0550	-0.0008	728	0.26
	97.4	*13.2*	*6.7*	*-5.7*		
1992	13.5442	0.0671	0.0417	-0.0006	748	0.27
	125.8	*15.3*	*6.3*	*-5.2*		
1993	13.6436	0.0670	0.0361	-0.0005	700	0.26
	118.3	*13.0*	*5.5*	*-4.7*		
1994	13.6836	0.0719	0.0317	-0.0004	687	0.25
	104.1	*12.7*	*4.5*	*-3.2*		
1995	13.5391	0.0730	0.0459	-0.0006	644	0.29
	106.3	*14.9*	*5.8*	*-4.8*		
1996	13.5509	0.0808	0.0401	-0.0005	618	0.25
	93.0	*13.8*	*4.7*	*-3.5*		

White robust t-statistic below the coefficient
Source: HBS 80/81 and CHBS 85-96.

Figure 1. Rates of return to education and 70% confidence interval. 1980-1996.[9]

Figure 2. Per capita years of schooling of the labour force (head of household) in Spain 1980-1996.



Source: Labour Force Survey

The rather stable pattern of returns to schooling takes place in a period with a steady increase in the average level of qualification of the labour force, as shown in Figure 2. Certainly, many factors could help to explain these changes[10], but the most important one is the increase in the supply of education by the public sector. In any case, the most important feature that

[9] The intermediate years between 1980 and 1985 had been obtained by a simple linear interpolation.

[10] For instance, the important decrease in the labour force without any formal study is related to demographic trends, specially in the agricultural sector.

appears in this increasing average level of schooling and the stability (and even a certain increase) in returns to education is that demand for more educated people has even overtaken supply. Probably, both the technological change and the tertiaritation that the Spanish economy has undergone in the last twenty years could explain this match between supply and demand.

III. Variations on the basic model

III.1. Qualifications as independent variable

The use of qualifications as independent variable allows us to test the linearity hypothesis, which underlines the years of schooling approach. Marginal rates of return for each level of education are shown in Table 3. Patterns from both data sets look very similar. We find growing returns as we move up the educational ladder, specially from primary to upper secondary. On the other hand, the returns show a relative stability in the university cycles. Furthermore, a common feature to all four estimates is a larger jump of returns in two levels, namely upper secondary and and upper vocational. [11]

Female returns tend to be higher than male returns. This is clearer for the sample WSS-95. Moreover, in this sample differentials increase with qualification. This is not the case of ECHP-94 where differentials tend to decrease as the level of education increases.

[11] In the case of the vocational path, the only level considered was upper vocational due to comparability problems between both data sets. Therefore, the rate of return to upper vocational was calculated with respect to lower secondary.

Table 3. Marginal rates of return (%)
Gross hourly wages. In percentage

	ECHP-1994		WSS-1995	
	Men	Women	Men	Women
Primary	1.0	-2.8	1.2	3.7
Lower secondary	3.8	4.4	3.5	5.2
Upper secondary	9.6	9.6	10.3	12.4
Short University cycle*	10.2	11.0	9.3	8.0
Long University cycle*	10.0	9.1	11.2	14.6
Primary	1.0	-2.8	1.2	3.7
Lower secondary	3.8	4.4	3.5	5.2
Upper vocational	7.0	6.9	8.6	10.5
Years schooling	7.4	8.3	8.2	10.0

*In this case, the 6 real years of schooling required to obtain the university degree was divided in 3,5 years to short University cycle and 2,5 to long University cycle.

Table 4 shows the estimated wage premiums over no schooling for the case of the WSS-95 sample.[12] Results are depicted in Figures 3 and 4 relating wage premiums to number of years required for the completion of each degree, in general and vocational paths. These Figures show that rewards increase with higher qualification, while after 8 years of schooling the linear approximation could be adequate. Female returns are higher than male's.

Table 4. Wage premium (reference: no schooling)
WSS-1995. Gross hourly wages.

	Men	Women
Primary	6.1	18.6
Lower secondary	16.6	34.1
Upper secondary	57.8	83.7
Short University cycle	90.3	111.7
Long University cycle	118.3	148.3
Primary	6.1	18.6
Lower secondary	16.6	34.1
Lower vocational	42.3	64.4
Upper vocational	59.4	86.7

[12] In using WSS-95 only, vocational education can be further disaggregated.

Figure 3. Wage premium (reference: no schooling). General path.
Gross hourly wages. WSS-1995

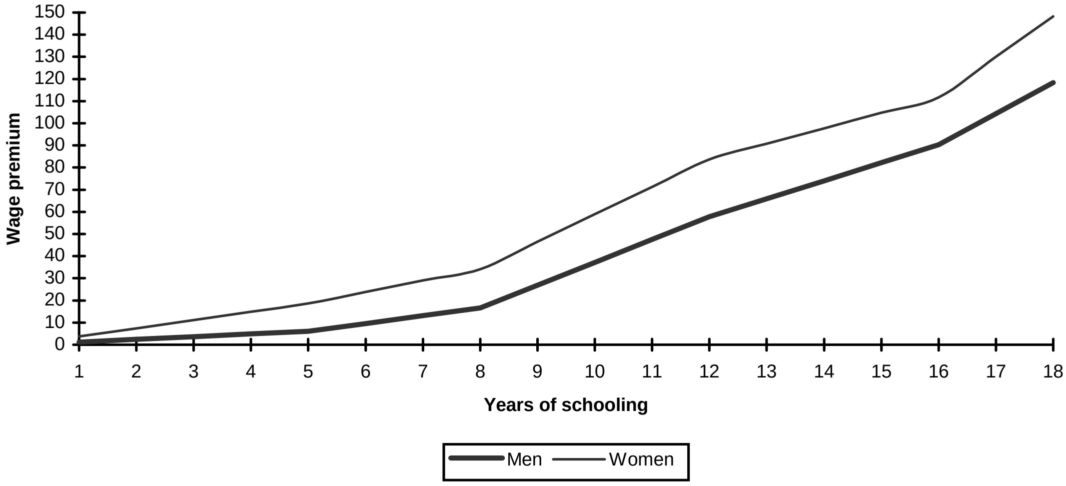
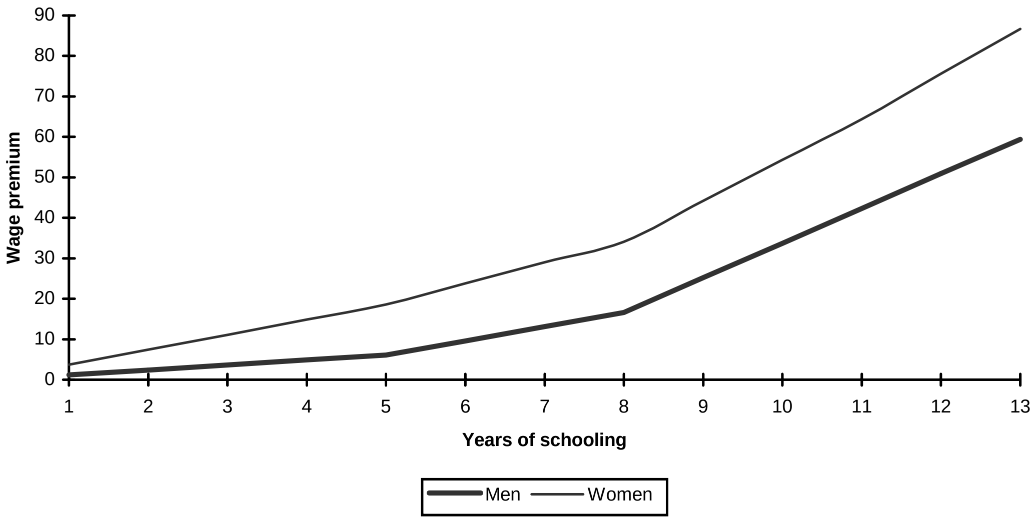


Figure 4. Wage premium (reference: no schooling). Vocational path.
Gross hourly wages. WSS-1995

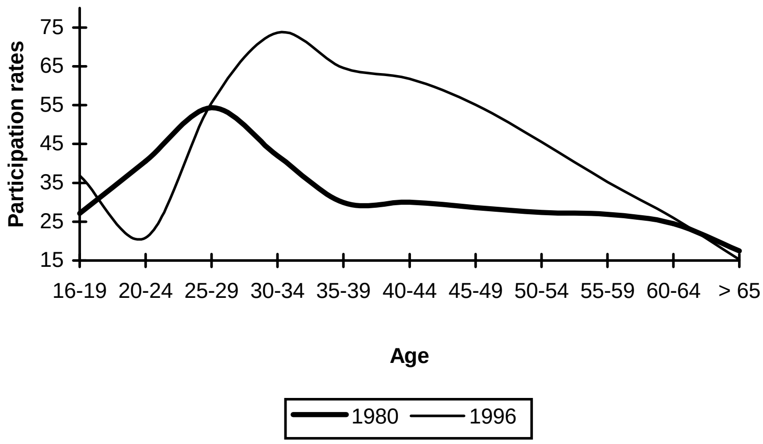


III.2. Other forms of human capital

Human capital theory considers other forms of human capital investment, than schooling. Experience and tenure increase human capital accumulation along the life cycle. It should be taken into account that the most common measure of experience is potential (as a difference between age and years of schooling). This definition of experience should not be a serious problem in the case of men. However, for women the approach is more ambiguous because of their discontinuities in the labour market.[13]

Moreover, two points should be added to those discontinuities in relation to the use of a cross-section. Firstly, we should expect female attachment to labour force to increase with education. Secondly, we are dealing with different generations in the sample. In the Spanish case, a clear generational break took place in the last twenty years in terms of labour market attachment of female working age population.[14]

Figure 5. Female participation rates. 1980-1996



Source: EPA (2nd quarter of each year).

WSS-1995 provides information on tenure. This information has allowed us to qualify potential experience as the 'experience up to current job'. That is, the number of years that go from leaving school until the current job. It should be reminded that this definition of experience is also potential. But, on the other hand, this approach allows us to consider experience and tenure separately. In the specification that includes both previous experience and tenure the rates of return to education show a slight decrease in relation to the standard mincerian specification. The rate for men decreases from 8,2% to 7,4% and the rate for women from 10,0% to 8,3%.

[13] In the Spanish case, only a few previous works have used tenure. Vid., Cañada (1993), De la Rica and Ugidos (1995), Ullibarri (1996), Garcia et al. (1997) and Salabarria and Ullibarri (1997). Other works use age instead of experience (Lassibille (1994), Garcia et al. (1997), Oliver et al. (1998) and Lassibille and Navarro (1998)).

[14] Obviously, the second point is partially a result of the first given the general increase of female education stock during those twenty years.

Table 5. Returns to other forms of human capital in Spain. 1995
Marginal returns to tenure and experience.
Gross hourly wage. WSS-1995

	Men	Women
Returns to previous experience		
Years of previous experience		
1	1.8	2.6
5	1.6	2.2
10	1.4	1.7
20	1.0	0.8
30	0.5	-0.1
Returns to tenure		
Years of tenure		
1	4.1	7.9
5	3.7	6.7
10	3.2	5.2
20	2.1	2.2
30	1.1	-0.7

As the specification of both previous experience and tenure is quadratic to calculate their returns we need to do it for a certain number of years. Table 5 shows returns to experience and tenure for 1, 5, 10, 20 and 30 years of previous experience and tenure. Three points are worth mentioning from these results. Firstly, returns to education are far higher than returns to experience and tenure. Secondly, returns to both experience and tenure are higher for women but decrease at a faster rate probably reflecting a different life cycle in the labour market. In addition, it should be noticed that differentials are much larger in tenure than experience. Finally, returns to tenure are more important than returns to experience.

III.3. Additional control variables and selection bias

Earnings functions with additional control variables coming from three different surveys have been estimated. From the HBS-1990/91 sector and regional dummies were taken, while from the ECHP-1994 regional dummies were used. Finally, from the WSS-1995 we used sector, job contract (fixed term vs. non-fixed term), company ownership (private or public) and plant size. Table 6 summarises our results. As expected, a clear and common feature appears: rates of return slightly go down. Using the highest number of control variables a to 6.5%). These results seem to show that larger reductions are related to the introduction of choice variables. In this sense, the larger reduction observed in the case of WSS-1995 would be due to the fact that all the variables included refer to job characteristics, and therefore endogenous choice variables. In fact, the change in job characteristics is one of the mechanisms through which the more educated achieve higher wages, so some of these estimates tend to underestimate the real return to education. (Mincer, 1974).

An interaction between regions and schooling has been used, searching for different rates of return among regions. Table 7 shows results obtained from different regions, with a sample from ECHP-1994 and HBS 1990-91. Overall, results are relatively stable to such inclusion[15], with a Figures ranging from 6,3% to 7,3% in 1990-91 and from 6,1% to 8,6% in 1994. Despite that, the null hypothesis that all regions share a common return to education is statistically rejected.

Table 6. Rates of return to education including additional variables. Men
In percentage

HBS 1990-91. Annual gross wage						
Rate of return	7,0	6,8	7,0	6,8	7,0	6,7
Sectors		**X**		**X**		**X**
Regions			**X**	**X**		
Regions*Schooling					**X**	**X**
Adjusted R^2	0,39	0,40	0,40	0,41	0,40	0,41
ECHP 1994. Hourly gross wage						
Rate of return	7,5	6,6	7,3	6,6	8,6	7,6
Sectors		**X**		**X**		**X**
Regions			**X**	**X**		
Regions*Schooling					**X**	**X**
Adjusted R^2	0,34	0,38	0,36	0,40	0,36	0,40
WSS-1995. Hourly gross wage						
Rate of return	8,2	7,6	8,1	7,4	7,5	6,5
Contract		**X**				**X**
Ownership			**X**			**X**
Size of firm				**X**		**X**
Sector					**X**	**X**
Adjusted R^2	0,38	0,42	0,39	0,47	0,43	0,53

X: variables included in each specification.

Table 7. Rates of return by regions (%).
Men. Gross hourly wages.

[15] Nevertheless regional disaggregation is not neutral in relation to the dispersion obtained.

	HBS-91/91	ECHP-94
Madrid	7,0	8,6
Northwest	6,7	6,1
Northeast	7,3	7,8
Centre	6,8	7,7
East	7,3	6,7
South	6,9	6,9
Canary Islands	6,3	6,9

Another modification in our parsimonious mincerian model has been the correction of female sample selection bias, using Heckman's two-step approach, but with some modifications, because we try to separate the decision to participate from the probability of being employed. For women, employment probability is related to some household and personal characteristics, but it is also a situation conditioned by the decision to enter the labour market. To tackle these two different process, we estimated two different probits, the first one aimed at obtaining the probability to enter the labour market, while the second one allow us to know the female employment probability. Both models yielded expected results with respect to behaviour variables. (Full results are included in Annex III)

When controlling for selection bias returns are slightly lower than when no selection bias control was used. That is, for ECHP-1994 (with hourly wages as independent variable), returns vary from 7,45% to 7,36% depending on whether or not part-time is controlled for, whereas when selection bias is not controlled for returns were 8,29% and 8,28% respectively.

Table 8. Return to schooling for women with selection bias correction (%). ECHP-1994
Gross hourly wages.

	Returns	Sample
All women	7,45	848
Part-time controlled	7,36	906

These results allow us to reconsider the female returns obtained using the more simple mincerian framework. The sample selection bias, that clearly affects both the rate of female participation in the labour market and their probability to be employed, has an increasing impact on female returns.

III.4. Private and public sector samples

As it has been suggested in the first section there are clear differences between male and female returns when sector is considered[16], with a higher return in private than in public. For instance, taken gross yearly wages, differences in men are higher than 1 percentage point (from 6.06% in public sector to 7.12% in private one), while in female difference reach 1.5 points (from 5,71 in private sector to 7,26 to public sector). The same pattern arises when gross hourly wages are taken into account, but with lower differentials between both sectors: from 6.43% to 6.89% for males and no differences for women.

Table 9. Returns to additional year of schooling (%). Private and public sectors
Gross hourly and yearly wages.

	HBS-1990/91[1]		ECHP-1994[2]	
	Values	n	Values	n
Men				
All	7,00	9.743	7,46	2.181
Private	7,12	7.192	6,89	1.612
Public	6,06	2.551	6,43	569
Women[3]				
All	7,53	3.133	8,29	848
Private	5,71	1.689	6,81	477
Public	7,26	1.444	6,92	371

1. Gross yearly wages. 2. Gross hourly wages. 3. Only full time female.

For males, both phenomena (higher returns in private sector and a decreasing pattern between yearly and hourly wages) can be explained by differences in the process of wage determination between sectors. Whereas in the private sector men show higher dispersion in hourly wages and in the number of hours worked, in the public sector these differences tend to be lower. In Table 10 gross hourly wages differentials have been reproduced for private and public sector by educational level, taking long university cycle as reference. According to these Figures, wage differentials by level of education are higher in private than in public sector.

[16] We define the public sector in terms of national accounts, that is the entrepreneurial public activity is not considered as public but as private.

Table 10. Male wages differences between levels and sector. 1994
Percentage difference between each gross hourly wages level and long University cycle

	Private	Public
Without studies	39.2	45.3
Primary	39.9	48.2
Compulsory	40.4	52.4
Upper vocational	48.9	58.7
Upper secondary	58.3	60.3
Short University cycle	76.0	80.0
Long University cycle	100.0	100.0
Total	47.4	67.1

Source: ECHP 94.

Our results suggest that the process underlying male's rates is even stronger in the female's case. Table 11 shows the percentage of employment by educational level in both private and public sector. A strong concentration of higher educated women in public sector is found as compared to male employment. However, it should be pointed out that concentration has not been increasing but rather decreasing.

	1.980			1.996		
	Private	Public	All	Private	Public	All
Without studies						

Male	93,3	6,7	100,0	93,1	6,9	100,0
Female	95,4	4,6	100,0	90,9	9,1	100,0
Primary						
Male	89,4	10,6	100,0	90,2	9,8	100,0
Female	93,7	6,3	100,0	92,2	7,8	100,0
Secondary						
Male	87,5	12,5	100,0	89,1	10,9	100,0
Female	86,2	13,8	100,0	87,1	12,9	100,0
Upper secondary						
Male	79,5	20,5	100,0	79,7	20,3	100,0
Female	73,4	26,6	100,0	76,4	23,6	100,0
University						
Male	54,6	45,4	100,0	59,0	41,0	100,0
Female	36,9	63,1	100,0	45,3	54,7	100,0

Source: LFS 1980, 1996

III.5. Hourly and yearly wages

Returns to schooling from hourly wages can be only estimated using the 1994 and 1995 surveys. Since these data sets are rather new, most empirical work on returns for Spain uses yearly wages. This element leads us to estimate returns with the same data purging process but with different wage definitions, yearly and hourly. Additionally, a second aspect related to the definition of number of hours worked suggests the necessity to discuss differences between hourly and yearly returns. First of all, it should be mentioned that hours worked variable used in our simple mincerian model was defined as number of hours in collective bargaining agreements. However, that number of hours may not necessarily be the actual working time, because sickness and absenteeism provoke a non-negligible impact on them. Table 12 shows the number of hours worked by educational level and from it two important features appear. First, actual hours are lower than bargained ones and, second, the difference is higher the lower the educational level. In any case, only when actual hours of work are considered[17] the Card (1999) decomposition of returns to education appears. As Card has shown for the US, approximately two thirds of returns to education of annual wages observed in the Nineties can be explained by the effect of hourly-wages, while the rest is accounted for hours per week and weeks per year of work. In the Spanish case the estimation of the actual number of hours worked for the higher level of education is not reliable and this can explain that the Card decomposition does not follow the expected pattern. Results with both definitions of wages (hourly bargained and yearly) are shown in Table 13. Results suggest that the definition does not matter in relation to returns. This conclusion is clear for men in the ECHP-1994 as well as in the WSS-1995 for both men and women. Only in the female sample from ECHP-1994 a slight difference appears in the University cycles (long and short) in favour of hourly wages, but this result probably has to do with the small sample used.

Table 12. Yearly male hours worked, by educational level. 1995
Average number and difference in percentage

Level	Real	Bargained	Difference(%)
1	1.664	1.757	-5,3
2	1.675	1.764	-5,0
3	1.624	1.754	-7,4
4	1.671	1.733	-3,6
5	1.651	1.747	-5,5
6	1.670	1.745	-4,3
7	1.671	1.727	-3,2
8	1.655	1.718	-3,6
All	1.655	1.750	-5,4

Table 13. Rates of return to education by levels and years in relation to no schooling[1]. 1994-1995
Gross yearly and hourly wages. In percentage

	ECHP-1994				WSS-1995			
	Men		Women		Men		Women	
	Yearly	Hourly	Yearly	Hourly	Yearly	Hourly	Yearly	Hourly
Levels								
Primary	0.82	0.98	-2.82	-2.79	1.20	1.22	3.66	3.71
Lower secondary	1.81	2.05	-0.07	-0.08	2.05	2.08	4.47	4.27
Upper secondary	4.30	4.58	2.69	2.87	4.69	4.81	7.07	6.98
Upper vocational	3.54	3.94	2.20	2.59	4.43	4.57	6.78	6.67
Short University cycle	5.29	5.66	3.92	4.55	5.53	5.64	7.01	6.98
Long University cycle	6.23	6.43	4.84	5.30	6.44	6.57	8.26	8.24
Years	7.46	7.46	7.37	8.29	8.02	8.21	10.04	10.02

1. Female part-time controlled for.

IV. Unemployment and returns to education

In the previous sections the estimation of returns to education was carried out

[17] Nevertheless, it must be said that in the Spanish case actual hours of work from higher levels of education are totally unknown.

assuming that individuals have the same unemployment probabilities. In addition, it seems quite clear that unemployment could have affected both the demand for higher levels of education and rates of return to education. Table 14 shows unemployment rates by level of education and age[18], obtained from the LFS samples. As it can be seen we have had a high, persistent and not equally distributed unemployment. This fact probably affects previous rates of return estimates. The effect of unemployment can appear through two ways: by modifying the opportunity cost of education and by affecting future earnings since differently qualified people have different probabilities to enter and to leave unemployment. For these reasons, when unemployment is taken into account returns may be subject to a non-negligible impact.

[18] Individuals younger than 30 years have not been included due to comparability problems in terms of job search for different education levels.

Table 14. Unemployment rates by educational qualifications. Men. (%)

	All	30-44 years	45-65 years
LFS-1990			
Illiterates	15.9	17.0	14.1
Primary	10.5	8.8	6.5
Lower secondary	14.4	7.1	4.9
Upper secondary	11.7	6.2	4.5
Lower vocational	17.0	7.8	6.5
Upper vocational	10.0	5.2	2.3
University (SC)	5.6	2.6	3.2
University (LC)	6.9	4.3	1.1
Total	11.8	8.0	7.6
LFS-1994			
Illiterates	25.8	32.5	21.7
Primary	17.8	17.2	11.6
Lower secondary	24.0	16.3	8.4
Upper secondary	17.2	9.6	6.5
Lower vocational	24.1	14.5	10.7
Upper vocational	17.4	9.7	8.4
University (SC)	9.0	5.5	2.4
University (LC)	10.8	5.6	2.9
Total	19.5	14.8	12.0
LFS-1995			
Illiterates	24.87	29.0	21.5
Primary	16.26	16.7	10.8
Lower secondary	21.51	14.8	9.2
Upper secondary	15.49	8.7	6.5
Lower vocational	20.22	10.6	6.9
Upper vocational	14.46	7.7	7.6
University (SC)	9.70	6.2	2.7
University (LC)	10.40	6.4	2.0
Total	17.63	13.4	11.1

The effect of unemployment on returns to education has not been almost treated in the literature. Rather the focus has tended to be on the effect of education on unemployment. However, it is obvious that education affects the probabilities to be unemployed and the duration of unemployment spells and, as a result, average life cycle incomes and, therefore, their respective rates of return to education. In this sense, the number of contributions, which analyse the first approach, is not very extensive. We should mention specially Ashenfelter and Ham (1979), Nickell (1979) and Groot and Oosterbeek (1992). In this literature, the forgone earnings are included in the estimation of returns to schooling, but some dissimilarities in the treatment appear. A fundamental difference is introduced in Ashenfelter and Ham with respect to the other authors. Ashenfelter and Ham aim to disentangle the importance of voluntary and involuntary unemployment on unemployed hours, whereas in the case of Nickell and Groot and Oosterbeek unemployment is basically seen as involuntary. This means that non-pecuniary benefits may arise from a situation of unemployment that should be taken into account. Even though a part of unemployment should be considered voluntary, in the Spanish case we can assume that most of it is involuntary. Therefore we can exclude non-pecuniary benefits from the calculation without much cost.

From this point of view, our approach tries to introduce the effect of unemployment on both, costs and benefits, by using the Internal Rate of Return (IRR). This system has three stages. First, a probit model estimates the probability of being unemployed, taking into account different qualifications and ages. Second, an earnings equation is run with the sample of wage earners, controlling for selection bias.[19] Finally, the age-earnings profiles are derived taking into account both the probabilities of being employed and the unemployment benefits, and the IRR is calculated.[20] From a conceptual point of view, the effects of considering the unemployment probabilities in the evaluation of the rates of return to education can be seen in Figure 6. The age-earnings profiles without unemployment effect are shown in panel a), and the unemployment effect on the age-earnings profiles is introduced in panel b). The figure shows that when unemployment probabilities are taken into account, and comparing two educational levels (upper secondary and university), two different effects appear. The opportunity cost of remaining in the educational system becomes smaller and, on the other hand, the extra income of the more educated population increases because unemployment tends to affect with

[19] For 1990/91 and 1994, a Probit model was estimated with the data obtained from the HBS-1990/91 and ECHP-1994, respectively. This estimation was used to control for selection bias, but the forecasted probabilities to be employed were estimated from the Spanish LFS samples due to the greater reliability of this source.

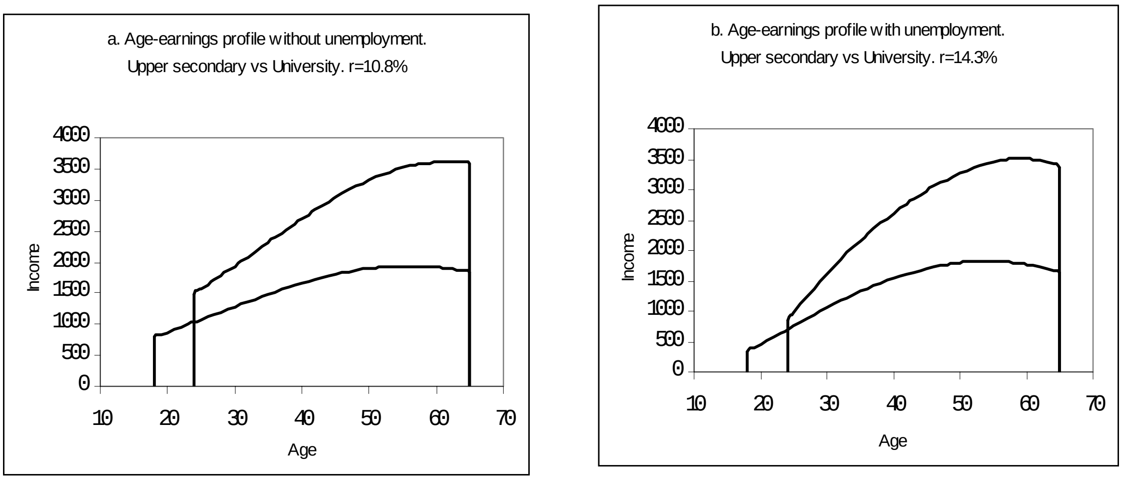
[20] The age-earnings profiles are calculated according to this formula:

$$\tilde{Y} = (\exp\{\hat{Y} + \tfrac{1}{2}\hat{\sigma}^2\}) * f(x) + b * [1 - f(x)]$$

where $\hat{Y}$ is the fitted earning value, f(x) is the employment probability, *b* is the unemployment benefit and σ is the standard error of the regression. In addition, we have assumed that unemployment benefits are 60% of the last wage that the unemployed population that receive unemployment benefits were 42,9% in 1990, 57,8% in 1994 and 62,9% in 1995. Finally, wage has been estimated using the forecasted value according to age and qualification. Because we are interested in private returns, unemployment benefits are considered. If, on the contrary, we would try to estimate the social returns, unemployment benefits would not have been considered.

more intensity less educated people (see Table 14). As a consequence, in our example the internal rate of return increases from 10,8% to 14,3%.

Figure 6. Unemployment effects on return to education.



Results with this approach are shown in Table 15 in which we compare returns under three different assumptions: firstly, no unemployment effect (by dummy coefficients and by IRR), secondly, taking unemployment effect into account but not including benefits and, finally, including benefits. Results show that if unemployment effect is not included, differences between the IRR approach and the standard econometric procedure through coefficients are negligible. However, as it was expected, rates of return increase considerably when employment probability is included in the calculation. For example, the rate of return to university long cycle increases by 21,0% in 1990/91, 36,2% in 1994 and 31,4% in 1995. When additionally benefits are included returns change by 13,6%, 15,2% and 14,3%, respectively. Obviously, the largest increases are obtained in 1994 and 1995 because the unemployment rates for these years are higher than for 1990. On the other hand, it is important to note that the most important modification occurs in upper vocational returns. However, despite that change in the upper vocational returns, its rate continues being the lowest one, except for 1990/91 where the upper vocational return with probability to be employed and unemployment benefits is higher that the lower vocational and upper secondary rates. To sum up, when we consider the probability to be employed the rates of return increase on average 16% in 1990/91, 46% in 1994 and 43% in 1995. But, when unemployment benefits are included these increases go down to 8%, 17% and 20% respectively.

Table 15. Returns to education (IRR) obtained in relation to compulsory level. HBS-1990, ECHP-1994 and WSS-1995. Men. Gross hourly or annual wages
In percentage

	Standard	IRR	Unemployment	Unemp.+ benefit
HBS-1990 (annual wages)				
Upper-secondary	7.0	6.9	7.3	6.9
Lower-vocational	6.9	6.8	6.9	6.6
Upper-vocational	6.9	6.7	8.6	7.8
University short cycle	7.7	7.6	9.2	8.5
University long cycle	8.1	8.1	9.8	9.2
ECHP-1994 (hourly wages)				
Upper-secondary	9.7	10.1	14.8	11.6
Upper-vocational	7.0	6.9	10.5	8.1
University short cycle	9.4	9.6	14.2	11.4
University long cycle	10.1	10.5	14.3	12.1
WWS-1995 (hourly wages)[21]				
Upper-secondary	10.3	10.7	15.2	12.8
Lower-vocational	12.8	13.7	20.3	16.7
Upper-vocational	8.6	8.7	13.2	10.9
University short cycle	9.2	9.4	13.1	11.1
University long cycle	10.2	10.5	13.8	12.0

V. Signalling and sheepskin

The screening hypothesis (Arrow, 1973; Spence, 1973; Stiglitz, 1975) states that education, and by extension qualifications, primarily acts as a screening device for workers as opposed to enhancing their productivity. In any case, it is possible to claim that education could have two effects on returns: a direct one by increasing productivity and an indirect one in signalling innate or pre-existing ability. Therefore, if we are interested in rates of return to educational

[21] It is important to note that due the WSS-1995 is an employer sample does not contain unemployment information, so it was not possible controlling by selection bias. Nevertheless, we applied the probabilities of being employed derived from EPA-95 in order to calculate the age-earnings profile and obtain the IRR.

investment it will be important to know the signalling weight.[22] In the following paragraphs, a set of procedures is employed to test the signalling hypothesis: comparisons of rate of return to schooling between screened *versus* unscreened groups and life cycle wages profile of differently qualified people.

V.1 Screened vs. unscreened rate of return method

First we compare rates of return of a particular sub-sample of the population as a theoretically unscreened group with other theoretically screened sub-samples. This approach is carried out in two ways.

The first one consists of comparing the rates of return between self-employed (unscreened group) and wage earners (screened group). The hypothesis is that since the self-employed have no need to signal ability, any return to education must represent a true return to human capital investment. Thus, the difference between the wage earners rate of return and the self-employed rate of return is a measure of signalling effect of education.

The second one is based on the comparison of rates of return between private (competitive and, therefore, unscreened group) and public sector (non-competitive and, therefore, screened group). If signalling theory is correct, then non-competitive sectors should show higher rates of return than competitive do, because in the first situation productivity is less important and a signalling process is more likely to take place.

Table 16 shows the results obtained with the self-employed approach[23] using a sample from HBS-1990/91 and ECHP-1994 controlling, in the last case, for selection bias because the choice between self-employed and wage earners

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[22] Remember that social rates of returns to education would be overvalued if signalling were present, while private returns are not affected.

[23] A statistical approach to average earnings by educational level between wage earners and self-employed has two different problems. Firstly, it is well known that self-employed understate their income, and this problem is more severe when we consider a higher level of education. Secondly, on average (overall and by educational level) self-employed are older that wage earners. This problem does not allow comparing directly differences between income and educational qualifications. To tackle with those problems, a set of filters has been used aimed to work with a more homogeneous sample.

may not be random. It seems quite evident that returns to education for self-employed are dramatically lower than those for wage earners do (7,0 % vs. 3,1 % in HBS 90/91 and 8,2% vs. 3,9 in ECHP 94). However, the explanatory capacity of the self-employed model is very low. Nevertheless, because of the effect of income underreporting of self-employed these results need to be interpreted with a lot of care. Another potential source of problems with the results related to self-employed arise from the heterogeneity of individuals that could belong to this group: temporal wage earners 'unemployed', wage earners transformed to self-employed, anticipated pensioner, and so on. In this sense, it is very difficult to derive a common behaviour through an earning equation to self-employed.

Table 16. Rates of return to wage earners vs self-employed

	HBS 1990-91		ECHP 1994.	
	Wage earners	Self-employed	Wage earners	Self-employed
Constant	13.0851	13.7989	13.4891	13.0701
	(711.8)	(235.2)	(292.9)	(36.9)
Schooling	0.0700	0.0316	0.0821	0.0359
	(67.1)	(10.6)	(34.6)	(4.3)
Experience	0.0456	0.0304	0.0517	0.0419
	(41.7)	(9.6)	(17.9)	(4.4)
$Experience^2$	-0.0006	-0.0004	-0.0006	-0.0006
	(-29.4)	(-9.0)	(-13.2)	(-4.1)
lambda			-0.5939	0.4058
			(-5.6)	(2.5)
Adjusted R^2	0.39	0.08	0.40	0.07
n	9743	2459	2181	473

White robust t-statistic in parentheses.

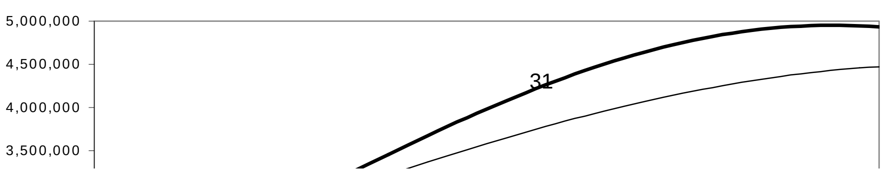
Secondly, Table 17 shows results from earnings equations with a sample from public and private sectors, based on HBS-1990/91 and ECHP-1994. Some authors consider that sector choice is not random. Thus, results from separated estimations by sector could have a specification bias (Arabsheibani and Rees, 1998). To tackle this problem, Heckman's procedure has been applied in two stages. However, in the ECHP-1994 case the *lambda* coefficients were not significant, so we preserve the model without selection bias correction. The rates of returns to education based on HBS-1990/91 are higher in private than in public sector, while rates of return based on ECHP-1994 are very similar, suggesting that a strong version of signalling theory must be rejected. Figure 7 displays university long cycle profiles based on these estimations. It can be emphasised that at the beginning of working life wages are higher in the public than in the private sector (13% higher), which accords to a weak degree of signalling. Nevertheless, when the life cycle is considered, wages in private sector grow faster than in public sector, ending up with higher income. This result is clearly against the main forecast of the signalling approach. This finding suggests that a weak version of signalling theory is present in the moment in which employers hire their work force. But, when 'true' productivity appears, employers pay higher wages to the most productive people.

Table 17. Earnings function by sector. Men.

	HBS 1990-91. Annual wages		ECHP 1994. Hourly wages.	
	Private	Public	Private	Public
Constant	12.9428	13.2171	5.9551	6.3079
	(478.6)	(106.6)	(104.1)	(64.1)
Schooling	0.0834	0.0666	0.0689	0.0643
	(27.9)	(15.0)	(19.4)	(15.5)
Experience	0.0527	0.0351	0.0391	0.0321
	(35.5)	(15.7)	(11.8)	(5.5)
Experience^2	-0.0007	-0.0004	-0.0005	-0.0004
	(-27.9)	(-11.7)	(-7.5)	(-3.5)
lambda	-0.2245	0.0726		
	(-4.4)	(1.5)		
Adjusted R^2	0.35	0.39	0.26	0.34
n	7192	2551	1612	569

White robust t-statistic in parentheses.

Figure 7 Earnings functions. Public and private sectors. HBS-1990/91. Men. (18 Years of schooling= Long University Cycle).



V.2 Returns to tenure and qualifications groupings method.

A second procedure is based on the comparisons of life cycle return profile of differently qualified people. That is, signal theory suggests that the educational process acts only as a ‘filter’, separating individuals with higher innate ability (with higher educational level) from the rest. But, when employers increase their knowledge about the ‘true’ productivity of their employees (as a result of their experience) these differences must decrease. Alternatively if those differences increase signalling theory must be rejected.

To test the signalling theory we use three different methods. First, we compare coefficients of the independent variable ‘years of education’ for different samples, which differ in tenure in their actual job. If signalling is present, the explained capacity of those coefficients (measured through t-statistics) must decline. In Table 18 results obtained with different samples for different tenure degrees based on WSS-1995 data are presented. These values do not seem to support the signalling theory. Rates of return increase with a higher level of tenure: from 6,3% (first year of tenure) to 8,3% (with a tenure between 1 to 5 years) and to 7,9% (between 5 to 20 years). Otherwise, a slight decrease of rates of return appears, but only after more than 20 years of tenure (6,5%).

Table 18. Earning functions by level of tenure in the current job. WSS-1995 Men.
Hourly gross wage.

	Years of tenure in current job			
	(0,1]	(1,5]	(5,20]	(20,50]
Constant	6.2582	6.1305	6.2648	6.3694
	(297.7)	(422.6)	(453.6)	(162.1)
Schooling	0.0634	0.0828	0.0794	0.0651
	(48.1)	(108.0)	(144.3)	(100.4)
Experience	0.0336	0.0402	0.0424	0.0482
	(27.5)	(42.9)	(51.8)	(23.1)
Experience^2	-0.0004	-0.0005	-0.0005	-0.0006
	(-18.9)	(-28.6)	(-37.9)	(-20.6)
Adjusted R^2	0.25	0.38	0.34	0.30
n	12561	25654	48711	31101

White robust t-statistics in parentheses.

A second method is the P-test, proposed by Psacharopoulos (1979). According to the weak version of this test, employers pay initially higher wages to more educated workers. In the strong version, these differences persist throughout the life cycle. If screening hypothesis is correct, then we should find a convergent profile once employers have adjusted wages of higher educated people to their *true* productivity. Otherwise, if a divergent profile of wages between different level of education is found, then signalling theory must be rejected. Table 19 shows estimations of earning functions by qualifications. First of all, it should be noticed the strong significance of the 'tenure' coefficients, and their positive sign. This is *prima-facie* evidence against the strong hypothesis. As can be observed, the "compulsory-upper secondary-university long cycle" pattern shows a divergent process in the profile income-tenure. That is, workers with higher degree of education have higher wages and similar returns to tenure than the less educated ones even when employers have had the possibility to see their 'true' productivity.

Table 19. Earning functions by educational qualifications. WSS-1995. Men.

Dependent variable: log (Gross hourly wages)

	Compulsory	Upper sec	Lower voc.	Upper voc.	Short cycle	Long cycle
Constant	6.8936	7.1712	7.0153	7.1929	7.5026	7.6811
	(1000.0)	(599.2)	(462.8)	(750.5)	(542.7)	(582.2)
Tenure	0.0402	0.0535	0.0512	0.0519	0.0493	0.0607
	(61.0)	(43.3)	(33.2)	(41.5)	(27.5)	(27.8)
Tenure^2	-0.0004	-0.0008	-0.0007	-0.0008	-0.0007	-0.0011
	(-19.6)	(-21.4)	(-14.6)	(-20.8)	(-13.6)	(-14.9)
Previous exp	0.0193	0.0185	0.0197	0.0189	0.0214	0.0268
	(24.8)	(13.2)	(10.5)	(14.5)	(13.3)	(15.3)
Previous exp^2	-0.0003	-0.0001	-0.0002	-0.0002	-0.0002	-0.0002
	(-13.4)	(-2.4)	(-2.5)	(-4.4)	(-2.8)	(-2.4)
Adjusted R^2	0.35	0.28	0.40	0.37	0.28	0.27
n	33208	12709	5797	9961	6329	7058

White robust t-statistics in parentheses.

Finally, an additional test, also suggested by Psacharopoulos (1979), is based on comparing the mid-to-early career earnings ratio for different sectors as years of schooling increase. A compatible behaviour with signalling theory should show a steady decrease and higher ratios in non-competitive than in competitive sectors.[24] The result of this method is presented in Table 20. Results, based on samples from HBS-1990/91 and WSS-1995, do not support the signalling hypothesis. In the HBS-1990/91 sample, the hypothesis that the non-competitive sector (public sector) has higher ratios than the competitive sector has not been proved. Moreover, the required decreasing pattern of this ratio is not clear, either.

Table 20. Income ratios at the middle and at the beginning of professional careers by years of schooling and by sectors.

	WSS-1995									HBS 1990/91	
Years	Extrac.	Manuf.	Utilit	Cons..	Trade	Hotels	Trans	Finan	Buss	Public	Privat
8	1.6	1.6	2.1	1.6	1.7	1.5	1.6	1.2	1.7	1.4	1.3
10	1.7	1.7	2.1	1.8	1.9	1.4	1.7	1.7	1.8	1.3	1.2
11.5	1.6	1.6	2.0	1.5	1.8	1.5	1.6	1.4	2.1	1.3	1.4
13	1.8	1.7	1.9	1.7	1.9	1.2	1.8	1.7	1.7	1.2	1.2
16	1.4	1.7	1.7	1.7	1.9	1.8	1.5	1.6	1.9	1.3	1.3
18	1.5	1.7	2.1	1.6	1.7	1.4	1.5	1.5	1.9	1.2	1.7

For the WSS-1995, income in the middle and at the beginning of their career belongs to tenures greater than 8 years and lower than 3 years, respectively. For the HBS-1990/91, for ages higher than 35 and 45 years old and lower than 25 years, respectively.

In summary, it seems quite clear that a strong version of signalling theory must

[24] See Cohn *et al*. (1987) , for a formal presentation.

be rejected in the Spanish case. Our results confirm those reached by other Spanish studies (Lassibille, 1994; Corugedo, 1998; Blanco and Pons, 1999; Pons, 1999). Notwithstanding that, our work suggests that a weak impact of signalling should be considered. Then, it is not possible to assert that the signalling hypothesis does not contain some elements of truth. However our empirical evidence shows that signalling does not explain return differentials, we can also claim, based on the showed empirical evidence, that signalling essentially does not explain return differentials.

V.3. Sheepskin effect

Another form of signalling is the so-called sheepskin-effect (Layard and Psacharopoulos, 1974; Park, 1999), which can be tested by comparing the effect on wages coming from 'actual' schooling years needed to attain a certain qualification. Several authors have argued that some degree of signalling could appear if people with qualifications earn more that those without them, but with the same amount of years of schooling. Note though that higher wages are likely to reflect higher productivity if individuals with higher abilities are more likely to obtain higher qualifications.[25] According to this interpretation, qualification acts as a proxy of ability (as a 'credential'), or as a proxy for benefit accruing from educational investment. Nevertheless, it is important to note that we have no information to determine the precise reason to explain why some individuals require more years to obtain a qualification (for instance, whether it is due to less ability or because the individual studied part-time while working). In this sense, it is not possible to asset whether the sheepskin effects imply signalling. So, in this part of the work we are only concerned with establishing if sheepskin effects exist.

Following Park (1999), the specification used to estimate this effect is

$$\ln W = \alpha_0 + \alpha_1 X + \sum_{i \in I} \sum_{j \in J} \beta_{ij} D(level = i) * D(S = j) + \eta$$

[25] Layard and Psacharopoulos (1974), p. 989.

Where X is a set of control variables, S is the number of years of schooling, I = {Compulsory. (3), Upper secondary (4), Lower vocational (5), Short University cycle (6) and Long University cycle (7)} and J = {8,9,21}. D = (level=i) is a dummy variable that takes value 1 when individual has the same educational level *i*. In addition, D(S = j) is also another dummy variable which takes value 1 if S = j and η is the random disturbance.[26] Rates of return to education had been calculated according to:

$$\frac{\widehat{W}_{i,j} - \widehat{W}_{3,8}}{n * \widehat{W}_{3,8}}$$

where

$$\widehat{W}_{i,j} = \exp \{\alpha_0 + \alpha_1 X + \sum_{i \in I} \sum_{j \in J} \beta_{ij} D(level = i) * D(S = j) + \frac{1}{2}\sigma^2\}$$

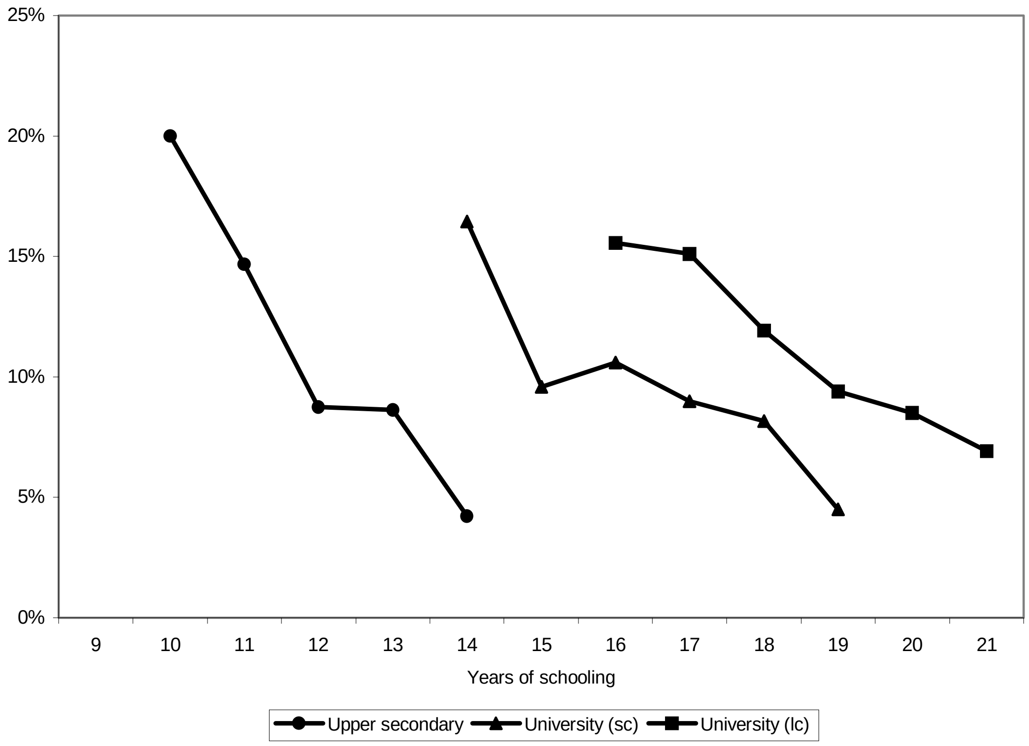
is the predicted value, and *n* is the number of years of education between the minimum compulsory level (3,8) and (i,j).

Data used for this analysis comes from ECHP-1994 and results from this approach are shown in Table 21. The experience variable was calculated using average men values obtained from the sample. Then, this average experience was assigned to individuals with 8 years of schooling (compulsory level). From this experience, each additional year of education was subtracted to build different predicted values. Figure 8 shows rates of return to education for Upper Secondary, Short University and Long University cycles. Results point to a negative influence of returns when more years of schooling are needed. For instance, those individuals who finish their university long cycle studies after 16 years of schooling have a rate of return of 15,6%. This rate decreases to 15,1% if they need 17 years and it goes down even more, to 11,9%, if the total amount

[26] For instance, D413 represent a dummy variable with 1 for individuals with upper secondary level and 13 years of schooling. Official years of schooling from 1970 on to reach each qualification are the following (between brackets number of years of previous system): BUP and COU is the sum of three different levels, that is, Lower Vocational with 10 years (and without previous existence), BUP with 11 years (10 years) and COU with 12 years (11 years). Short University Cycle, on the other hand, needs 15 years (14 years), while Long University Cycle demands 17 years, (16 years).

of years needed is 18. The profile is clearly decreasing until a low rate of return of 6,9% with 21 years of schooling.[27] This decreasing pattern of returns appears because additional schooling years imply higher opportunity costs and, at the same time, a shorter payback period. In addition, it is also possible that individuals who finalise their studies in a shorter period are more skilful and have a higher productivity and, therefore, obtain higher rates of return.[28]

Figure 8. Sheepskin effects. Hourly gross wages. Men



[27]These results are similar than those obtained by Corugedo (1998) using a different approach.
[28] It should be noticed, again, that extra years of schooling could reflect interrupted process connected with work, illness or familiar responsibilities, as well as less ability to achieve qualification. Nevertheless, we have no information about these different causes.

Table 21. Sheepskin-effects. Men. ECHP-1994

	Coefficient	t-student
Constant	6.1555	122.4
Schooling	0.0410	8.1
Experience	0.0441	15.1
Experience^2	-0.0006	-10.4
D38	-0.0495	-1.6
D39	-0.0872	-1.9
D310	-0.0379	-0.8
D410	0.2434	2.6
D411	0.2516	3.6
D412	0.1681	2.9
D413	0.2091	2.8
D414	0.0597	0.4
D512	0.0629	1.0
D513	0.1955	2.9
D514	0.0804	1.0
D515	0.1466	1.6
D614	0.5209	7.1
D615	0.3315	4.0
D616	0.4184	4.4
D617	0.3835	3.4
D618	0.3760	3.8
D619	0.1708	1.6
D716	0.6129	7.2
D717	0.6488	8.4
D718	0.5635	5.8
D719	0.4776	3.7
D720	0.4613	3.1
D721	0.3906	2.7
Adjusted R^2	0.40	
n	2181	

An initial signal effect clearly appears when sheepskin is considered. But this effect may only reflect a weak signal if individuals with the same qualification tend to earn similar wages along the life cycle.

VI. Endogeneity and Instrumental Variables

As it is well known, OLS estimates of returns to education are biased if explanatory variables in earnings equations are not exogenous. The most common problem that arises from endogeneity is the 'ability bias'. It appears because the error term from the mincerian earnings equation reflects, among other factors, innate ability of individuals. If the more skilful people are those who obtain higher qualifications, the random disturbance and the regressor (that is, years of schooling) will be correlated, and estimates will be inconsistent. Instrumental Variables is an ordinary procedure to deal with this problem. However, adequate instruments must be found, that is, instruments related to years of schooling but not correlated with the random disturbance.

We have chosen ‘age’ (E) as the most adequate instrument for the Spanish case, because the important change of our educational system, had a higher impact on younger than on older generations. Several factors have raised the average level of education of new generations: increasing public educational supply, its growing extension to young population and increasing compulsory age. Hence our instruments are independent of individuals innate ability. The functional form adopted includes a quadratic in age and some spline terms. That is

$$S_i = \varphi_0 + \varphi_1 E + \varphi_2 E^2 + \delta_1 [D_1 (E–E_1)] + \delta_2 [D_2 (E–E_2)] +....+ \delta_m [D_m (E–E_m)] + v_i$$

Where

$D_j = 0$ if $E \leq E_j$
$D_j = 1$ if $E > E_j$ to $j = 1,2,...m.$

This approach demands for a correct selection of E_j. To do so, a stepwise process was followed. In the first step, all possible values of E_1 were used and was selected the best-adjusted model. In a second step, with a fixed E_1, identical processes were made with all possible values of E_2 ($E_2 > E_1$), and so on.[29]

This procedure shows that the instrument defined is adequate, because growing supply of schooling is related to increasing years of schooling. That is, instruments are correlated with the regressor. At the same time, it seems reasonable to expect that innate ability has not changed in the last fifty years.

[29] This process leads us to the following instruments HBS-1990/91: constant, E, E2, E28, E36 and E58; ECHP 94: constant, E, E2, E27, E38 and E48 and WWS-95: constant, E, E2, E28, E30, E40, E49 and E59. where Ei=E-i. For instance, E28 = Age-28.

This assumption implies that instruments chosen are independent from the random disturbance.

Results from OLS vs IV are shown in Table 22. The IV estimates are presented in three different forms. The first one is the standard IV estimate, the second one consists in correcting for differences in individual specific coefficient and self-selection bias by employing a method proposed by (Garen, 1984), which is applicable when the choice variable is continuous. Finally, the third one is based on a split-sample instrumental variable (SSIV) (Angrist and Krueger, 1995), which uses one-half of a sample to estimate parameters of the first-stage equation. First-stage parameters estimates are then used to construct fitted values and second-stage parameters estimates in the other half a sample. This approach was applied only to the WSS-1995 sample, because it contains a large number of observations.

Reliability of IV vs. OLS was tested through Sargan, Hausman and Bound (1995) tests. Sargan test rejects the null hypotheses only with WSS-1995 estimations, as a result of the large number of observations (118.027).[30] In Hausman test, high values of χ^2 statistics implied the rejection, in all cases, of the null hypothesis of exogeneity of the education variable. Finally, F-statistic Bound test of excluded instruments suggests that the instrument used is correct.

Regarding the values of rates of return found with IV, there is no substantial change in comparison with those obtained with OLS, but slightly higher values for HBS-1990/91 and ECHP-1994, and slightly lower values for WSS-1995. These results provide evidence that in the Spanish case the ability bias seems to be not very important and that differences between OLS and IV estimates do not follow a clearly identifiable pattern.

Table 22. Earnings equations. OLS vs Instrumental Variables. Men

[30] As it is well known, if the test is consistent, the standard hypothesis test approach gives an asymmetrical treatment to the Error Type I (erroneous rejection of a true null hypothesis) vs the Error Type II (erroneous acceptance of false null hypothesis). When the sample size increase, the probability of Error Type II tends to zero, meanwhile the probability of Error Type I is maintained constant at the significance level selected. As a consequence, if the significance level is not reduced when the sample size increases, a major sample size would mean a major probability of rejection of the null hypothesis. In the limit, all the models and constrains would be rejected to the extent that they constituted a simplified representation of the reality.

HBS 1990/91. Annual gross wage
(nº of obs. = 9743)

	OLS		IV		IV Garen	
	Coef.	t-statistic	Coef.	t-statistic	Coef	t-statistic.
Constant	13.0851	711.8	12.9008	201.7	12.9493	207.2
Schooling	**0.0700**	67.1	**0.0823**	12.7	**0.0745**	11.7
Experience	0.0456	41.7	0.0520	40.1	0.0505	39.5
Experience^2	-0.0006	-29.4	-0.0007	-24.3	-0.0006	-23.8
Adjusted R^2	0.39		0.38		0.40	
Sargan			2.50	5.99*		
Hausman			82.4	7.82*		
Bound (F)			151.8			

ECHP 1994. Hourly gross wage
(nº of obs. = 2181)

	OLS		IV		IV (Garen)	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
Constante	13.5449	296.2	5.6651	24.0	5.6443	25.2
Schooling	**0.0721**	29.3	**0.0913**	3.5	**0.0954**	3.9
Experience	0.0369	13.0	0.0476	7.6	0.0445	7.6
Experience^2	-0.0004	-8.0	-0.0006	-3.8	-0.0005	-3.9
Adjusted R^2	0.34		0.32		0.36	
Sargan			0.2	5.99*		
Hausman			26.3	7.82*		
Bound (F)			35.4			

WSS 1995. Hourly gross wage.
(nº of obs. = 118027)

	OLS		IV		IV (Garen)		SSIV (n=59975)	
	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic	Coef.	t-statistic
Constant	6.0425	932.9	6.0324	255.5	6.1119	258.3	6.0141	167.8
Schooling	**0.0821**	242.6	**0.0747**	30.6	**0.0649**	26.3	**0.0771**	20.2
Experience	0.0500	124.5	0.0576	123.0	0.0555	117.5	0.0571	75.2
Experience^2	-0.0006	-76.9	-0.0007	-72.3	-0.0007	-70.4	-0.0007	-42.7
Adjusted R^2	0.38		0.38		0.40		0.15	
Sargan			26.08	9.49*				
Hausman			846.3	7.82*				
Bound (F)			972.5					

* Indicate chi squared critic values at the 5% level.

Our result contrast with a common finding in international empirical literature that shows greater changes between rates of return obtained by OLS and by IV. In some cases IV returns doubled those from OLS. In our opinion, as it can be drawn from Annex I inspired in Card (1998), in some cases IV estimator do not represent sample average returns to education but those from particular groups, far away from the sample average, and strongly correlated with used instruments.

VII. Conclusions

The main conclusions of this work could be grouped in the following items:

I) Global returns to education

1) Regarding private returns to education in terms of hourly wages, our findings suggest estimates of around 8% per year. This result is fairly robust to different samples used in this work and to different estimation methods.

2) The historical evolution of returns to education from 1980 until 1996 shows a certain increasing trend in spite of the increase in the human capital stock of the Spanish population. In fact, in 1980, the estimated rate of return was 6% per year and the average years of schooling of the labour force was 6. On the other hand, in 1996, the estimated rate of return to education was 8% and per capita years of schooling of the labour force increased to 8.3. So, the rate of return to education has increased by 33% and this figure has been accompanied by a 45% increase in the human capital stock of the Spanish labour force. In other words, the race between human capital demand and supply has shown a certain advantage for the demand side, which translates in an increase of the returns to education, in spite of the big push experienced by the educational supply.

3) When evaluating the returns to education it is possible and convenient to take the different employment probabilities associated to the different

educational levels into account. We find that as the educational level increase, the unemployment probabilities become smaller. This relationship may represent an important link, which permits the more educated individuals take full advantage of their high educational level. In this case, the returns to education tend to surpass by around two points the previously mentioned range of 7%-8% that is implicitly calculated considering the hypothesis that the unemployment rate is similar among different educational levels.

4) The returns to education by sex are relatively similar when both samples (males and females) are considered in a homogeneous way; in particular, considering full time workers and dropping anomalous observations.

5) Regarding private and public sector, the returns to education appear a bit higher in the private sector, result that could be a consequence of wage-setting mechanisms for civil servants in the public sector.

6) Considering other forms of human capital different from schooling, the obtained evidence shows returns to previous experience between 2% and 0.5%, whereas returns to tenure are in the range of 4% to 1%. Clearly, both types of human capital have a lower rate of return than schooling.

II) Returns to education and educational level

1) Education in Spain seems to show increasing returns, in the sense that the returns to an extra year of education manifest a certain trend to grow when the number of years of education increases. However, after 8 years of schooling, corresponding to lower secondary, the increase in the returns to education becomes moderate. In fact, an upper limit must exist, because otherwise, from an individual point of view, there would be no incentives to stop the accumulation of human capital.

2) University education becomes more profitable than vocational education. This could be a consequence of the way in which the vocational path is

selected (in some cases the vocational path is selected as a second option) and can reflect differentials in innate ability of the individuals. It could also be considered as a by-product of the relative failure of the vocational path in Spain.

III) Is education only a filter device?

If education is a filter, social returns to education would be much lower than private returns and the whole educational system should be considered inefficient and a wasting screening method that absorbs a disproportionate amount of resources. We have dedicated important efforts to discriminate between the human capital hypothesis and the screening hypothesis. Our results indicate certain signalling when the individual enter the labour market (weak signalling) but there is no evidence that this is so over time. So, high private educational return could be considered as an indication of a high social return. In other words, if there is no screening, or if the screening hypothesis does not explain the lion part of the wage differentials, investing in human capital could be a profitable activity from an individual is as well as public or social point of view.

IV) Methodological questions.

Traditionally, the estimation of the educational returns by fitting Ordinary Least Squares (OLS) to Mincer type equations has been submitted to the criticisms of ability and endogeneity bias, since unobservable individual ability would be correlated with the random term. In our case, the estimates by IV, when the sample is adequately purged, become similar to those obtained by OLS. An explanation of why in some cases both methods could produce important differences is suggested taking the Spanish case as a reference.

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Annex I. An interpretation of Instrumental Variables estimator in the presence of between groups heterogeneity

The aim of this section is to show that using IV, instead of OLS, estimates may increase considerably. This possibility is related to the heterogeneity of rates of return to education in a process in which IV are correlated with regressors of those individuals with higher returns. To simplify, consider deviations from average, only one explanatory variable and the following Data Generating Process (DGP):

$\begin{pmatrix} Y_1 \\ Y_2 \end{pmatrix} = \begin{pmatrix} X_1 & 0 \\ 0 & X_2 \end{pmatrix} \begin{pmatrix} \beta_1 \\ \beta_2 \end{pmatrix} + \begin{pmatrix} \varepsilon_1 \\ \varepsilon_2 \end{pmatrix}$ Where β_1 and β_2 are the effects of X_1 and X_2 for observation groups 1 and 2. It is a normal mistake to take a common β and to estimate it by IV using “Z” as an instrument. The estimation of the common β is

$\tilde{\beta} = \dfrac{\sum Y_1 Z_1 + \sum Y_2 Z_2}{\sum X_1 Z_1 + \sum X_2 Z_2}$ Additionally, estimators for each group of IV are:

$$\tilde{\beta}_1 = \frac{\sum Y_1 Z_1}{\sum X_1 Z_1} \quad ; \quad \sum Y_1 Z_1 = \tilde{\beta}_1 \cdot \sum X_1 Z_1$$

$$\tilde{\beta}_2 = \frac{\sum Y_2 Z_2}{\sum X_2 Z_2} \quad ; \quad \sum Y_2 Z_2 = \tilde{\beta}_2 \cdot \sum X_2 Z_2$$

Substituting: $\tilde{\beta} = \dfrac{\tilde{\beta}_1 \cdot \sum X_1 Z_1 + \tilde{\beta}_2 \sum X_2 Z_2}{\sum X_1 Z_1 + \sum X_2 Z_2}$

$$\tilde{\beta} = \tilde{\beta}_1 \frac{\sum X_1 Z_1}{\sum X_1 Z_1 + \sum X_2 Z_2} + \tilde{\beta}_2 \frac{\sum X_2 Z_2}{\sum X_1 Z_1 + \sum X_2 Z_2}$$

if $\tilde{w} = \dfrac{\sum X_1 Z_1}{\sum X_1 Z_1 + \sum X_2 Z_2} = \dfrac{1}{1 + \dfrac{\sum X_2 Z_2}{\sum X_1 Z_1}}$

we can define $\tilde{\beta} = \tilde{w} \cdot \tilde{\beta}_1 + (1 - \tilde{w}) \cdot \tilde{\beta}_2$. Then,

$limp(\tilde{\beta}) = \beta_1 \cdot limp(\tilde{w}) + \beta_2 \cdot (1 - limp(\tilde{w})) = \beta_1 \cdot w + \beta_2 (1 - w)$. Because

$$\hat{\rho}_{Z_1X_1} = \frac{\sum X_1Z_1}{\left(\sum X_1^2\right)^{1/2}\left(\sum Z_1^2\right)^{1/2}} = \frac{\hat{\sigma}_{Z_1X_1}}{\hat{\sigma}_{Z_1}\cdot\hat{\sigma}_{X_1}}$$

$$\sum X_1Z_1 = \hat{\rho}_{X_1Z_1}\cdot\hat{\sigma}_{X_1}\hat{\sigma}_{Z_1}$$

$\sum X_2Z_2 = \hat{\rho}_{X_2Z_2}\cdot\hat{\sigma}_{X_2}\hat{\sigma}_{Z_2}$ Then, $\tilde{w} = \dfrac{1}{1+\left(\dfrac{\hat{\rho}_{X_2Z_2}}{\hat{\rho}_{X_1Z_1}}\right)\cdot\left(\dfrac{\hat{\sigma}_{X_2}\cdot\hat{\sigma}_{Z_2}}{\hat{\sigma}_{X_1}\cdot\hat{\sigma}_{Z_1}}\right)}$ As a result:

$$limp(\tilde{\beta}) = \beta_1\cdot w + \beta_2\cdot(1-w)$$

$$w = \frac{1}{1+\left(\dfrac{\rho_{X_2Z_2}}{\rho_{X_1Z_1}}\right)\left(\dfrac{\sigma_{X_2}\cdot\sigma_{Z_2}}{\sigma_{X_1}\cdot\sigma_{Z_1}}\right)}$$

When the correlation between instrument and regressors is considered, the following extreme situations may appear:

If $\rho_{X_1Z_1} \to 0 \Rightarrow w \to 0 \Rightarrow limp(\tilde{\beta}) = \beta_2$ or if $\rho_{X_2Z_2} \to 0 \Rightarrow w \to 1 \Rightarrow limp(\tilde{\beta}) = \beta_1$

This phenomenon could be illustrated in IV as Table 1 shows. For instance, if in HBS-1990/91 we use the whole sample, the return to education that is obtained from OLS is 10,2% whereas for IV the result is 18,0%. Nevertheless, and in order to homogenise this sample, the standard filter was applied to it. This filter consists, basically, of excluding those individuals with earnings below the minimum wage. From the application of the filter three different samples can be derived: one with all observations (entire sample), a second one only including those individuals whose wage is above minimum wage (filtered sample), and finally a sample with the individuals with wages below minimum wage (excluded individuals). It is quite reasonable to think that these two different groups of individuals, because of their different educational and wage characteristics, probably have different rates of return to education (β_1 and β_2). In this case, using IV, the return to education is 8,2% for the filtered sample, 18,9% for the excluded individuals and 18,0% for the entire sample. What is remarkable is that 25% of the sample in the IV estimation determines the estimated coefficient when the whole sample is used. That is, according to the previous formulation the IV estimators basically reflect the return to education of the sub-sample that has the higher rate of return. The same conclusion is obtained when we consider the WSS-1995 sample, where 12% of observations, that form the excluded individuals, almost determine the rate of return obtained for the entire sample.

Table 1. OLS vs. IV estimation to different sub-samples.

HBS-1990/91(Gross annual wage)

Sample		Rate	n	R^2	R^2 (Step 1)
entire	OLS	**10.2%**	13035	0.34	
	IV	**18.0%**	13035	0.19	0.13
filtered	OLS	**7.0%**	9743	0.39	
	IV	**8.2%**	9743	0.38	0.07
excluded	OLS	**8.1%**	3292	0.14	
	IV	**18.9%**	3292	-0.02	0.16

WSS-1995 (Gross annual wages)

		Rate	n	R^2	R^2 (Step 1)
entire	OLS	**9.1%**	134880	0.35	
	IV	**17.6%**	134880	0.10	0.05
filtered	OLS	**8.0%**	118027	0.38	
	IV	**7.5%**	118027	0.38	0.05
excluded	OLS	**6.7%**	16853	0.18	
	IV	**21.7%**	16853	-0.37	0.05

Annex II. Datasets

The main characteristics of the datasets used are described in the following paragraphs.

Household Budget Survey 1990/91 (HBS-1990/91).

This survey offers information about income and expenditure of about 20.000 households. It has information for all members of each household, especially in those aspects related to qualifications obtained, their annual net income, as well as their position in the labour market (employed/unemployed, sector, occupation, self-employer/wage earner). Nevertheless, while information about the total number of hours worked is not disposable, provides information on whether individuals worked less than 30 hours during the week of reference. This information has allowed us to discriminate between full and part-time job. The criteria to include observations in the sample were the following: wage earners, living in national territory, full time, wages above minimum wage. The application of the criteria left a disposable sample of 10.470 observations from the original 23.381.

Continuous Household Budget Survey 1985-1996 (CHBS-1985-1996).

From 1985 onwards, this survey is made each quarter with a sample based on 3.000 households. This survey has the structure of a rotating panel, that is, an eighth of the sample is changed each quarter. Variables are the same than those that appears in the HBS 1990/91, but are disposable only for the head of the household.

Wage Structure Survey 1995 (WSS-1995).

Offers information about structure and distribution of wages for Spain and individually**,** besides an important amount of characteristics related to each worker (qualification obtained, tenure, kind of contract, type of job, sector, firm size, market for the company products, bargaining process and ownership). This survey is built with all the companies with more than 10 wage earners at October the 31st of 1995. Wages are gross and they are presented as hourly, monthly or annual amount. The sample used included individuals with wages above the minimum wage, with tenure of at least one year, full-time and younger than 65 years. This delimitation of the sample leads us to an useful sample of 114.773 observations for men and 34.467 for women.

European Household Panel 1994 (ECHP-1994).

This survey offers information about all people that lives in one of the 8.000 surveyed households. Basic personal characteristics are provided for each individual as well as labour market status. For the employed, information is given on industry, firm's type of ownership, gross and net wages, and worked hours. It provides information about educational level, and also on age leaving education. Since we know also the qualification attained then actual years of schooling, for each qualification, can be known. We purged the survey dropping those individuals with wages below minimum wage. Observations for which the difference between the implied number of years to complete a degree by the individual and the officially required years was larger than five were dropped. The application of these criteria resulted in a sample with 2.193 men and 1.050 women.

Annex III. Estimated equations

Table 1. Standard Mincerian equations.

	Men	Women[1]	Women[2]
HBS-1990/91. Annual gross wage			
Constant	13.0851	12.9897	11.0961
	711.8	*543.7*	*139.3*
Schooling	0.0700	0.0753	0.1286
	67.1	51.1	53.8
Experience	0.0456	0.0313	0.0591
	41.7	20.2	20.1
Experience^2	-0.0006	-0.0003	-0.0008
	-29.4	-10.7	-12.8
Part time			-0.6839
			-9.6
Adjusted R^2	0.39	0.47	0.38
n	9743	3133	5054
ECHP-1994. Hourly gross wage			
Constant	5.9341	5.7855	5.8043
	124.7	*77.5*	*80.8*
Schooling	0.0746	0.0829	0.0828
	29.0	*20.7*	*21.4*
Experience	0.0393	0.0396	0.0378
	13.2	*8.2*	*8.3*
Experience^2	-0.0004	-0.0005	-0.0005
	-8.0	*-4.8*	*-4.8*
Part time			0.3918
			7.3
Adjusted R^2	0.34	0.40	0.43
n	2181	848	906
WSS-1995. Hourly gross wage			
Constant	6.0425	5.8823	5.3685
	897.7	*524.1*	*348.6*
Schooling	0.0821	0.0828	0.1002
	255.9	*126.3*	*121.2*
Experience	0.0500	0.0472	0.0649
	121.3	*66.8*	*69.3*
Experience^2	-0.0006	-0.0005	-0.0008
	-75.9	*-37.7*	*-42.9*
Part time			-0.1397
			-14.2
Adjusted R^2	0.38	0.37	0.30
n	118027	30769	40912

1 Only full time, 2 All the sample controlling part time.
White robust t-statistic below the coefficient

Table 2. Standard Mincerian equations by educational levels

	HBS-1990/91 Annual gross wage		ECHP-1994 Hourly gross wage		WSS-1995 Hourly gross wage	
	Men	Women	Men	Women	Men	Women
Constant	13.3268	11.3677	6.2690	6.3804	6.4491	5.7711
	687.4	*128.8*	*128.6*	*70.0*	*753.1*	*192.2*
Primary	0.1648	0.4058	0.0490	-0.1394	0.0608	0.1857
	12.6	*7.5*	*1.2*	*-1.6*	*8.7*	*6.6*
Lower sec.	0.2879	0.7390	0.1640	-0.0066	0.1661	0.3414
	19.1	*12.6*	*4.4*	*-0.1*	*22.7*	*12.0*
Upper sec.	0.5685	1.2630	0.5492	0.3445	0.5778	0.8371
	32.8	*20.8*	*12.0*	*3.9*	*72.4*	*29.1*
Lower voc.	0.4255	1.0755			0.4230	0.6439
	20.4	*16.0*			*49.5*	*21.4*
Upper voc.	0.6327	1.3396	0.5126	0.3367	0.5944	0.8673
	28.9	*19.5*	*10.8*	*3.6*	*75.0*	*29.5*
University (sc)	0.9041	1.8155	0.9060	0.7279	0.9031	1.1170
	48.4	*31.1*	*19.1*	*8.2*	*104.1*	*36.6*
University (lc)	1.1000	2.0150	1.1568	0.9545	1.1828	1.4830
	50.1	*31.5*	*23.7*	*10.6*	*132.1*	*48.9*
Experience	0.0452	0.0580	0.0463	0.0456	0.0517	0.0650
	40.4	*19.2*	*16.2*	*11.5*	*128.3*	*68.5*
Experience^2	-0.0006	-0.0007	-0.0006	-0.0007	-0.0006	-0.0009
	-29.7	*-12.1*	*-11.6*	*-8.4*	*-84.8*	*-43.9*
Part time		0.6845		0.3387		-0.1416
		9.6		*6.5*		*-14.4*
Adjusted R^2	0.40	0.38	0.40	0.51	0.40	0.31
n	9743	5054	2181	906	118027	40912

White robust t-statistic below the coefficient

Table 3. Rate of return to other forms to human capital. WSS-1995. Hourly gross wage

	Men	Women
Constant	6.3631	5.6863
	1025.8	*407.3*
Schooling	0.0741	0.0834
	195.3	*102.0*
Previous exp.	0.0188	0.0268
	50.8	*30.7*
Previous exp^2	-0.0002	-0.0005
	-25.1	*-19.6*
Tenure	0.0426	0.0821
	118.7	*85.6*
Tenure^2	-0.0005	-0.0015
	-48.7	*-46.6*
Part time		-0.0117
		-1.2
Adjusted R^2	0.42	0.43
n	118027	40912

White robust t-statistic below the coefficient

Table 4 a. Rate of return including additional variables. Men. HBS-1990/91. Annual gross wage

Constant	12.9394	13.0978	12.9472	12.9359	13.0807
	570.5	*521.5*	*451.5*	*569.3*	*710.5*
Schooling	0.0681	0.0696	0.0676	0.0675	0.0699
	61.1	*66.5*	*60.4*	*31.4*	*33.5*
Experience	0.0445	0.0459	0.0450	0.0450	0.0460
	40.6	*42.2*	*41.1*	*41.0*	*42.1*
$Experience^2$	-0.0006	-0.0006	-0.0006	-0.0006	-0.0006
	-28.5	*-30.0*	*-29.0*	*-28.9*	*-29.8*
Industry	0.2060		0.1958	0.2023	
	12.5		*11.7*	*12.2*	
Building	0.1259		0.1259	0.1265	
	7.0		*6.9*	*7.0*	
Services	0.1795		0.1807	0.1821	
	11.0		*11.0*	*11.2*	
Northwest		-0.0514	-0.0451		
		-2.4	*-2.1*		
Northeast		0.0433	0.0439		
		2.2	*2.2*		
Centre		-0.0396	-0.0286		
		-2.1	*-1.5*		
East		0.0203	0.0276		
		1.0	*1.4*		
South		-0.0281	-0.0138		
		-1.4	*-0.7*		
Canary I.		-0.0993	-0.0930		
		-4.0	*-3.8*		
Northwest*s[1]				-0.0029	-0.0034
				-1.3	*-1.5*
Northeast*s				0.0030	0.0029
				1.4	*1.4*
Centre*s				-0.0008	-0.0014
				-0.4	*-0.7*
East*s				0.0040	0.0035
				1.9	*1.7*
South*s				-0.0005	-0.0013
				-0.2	*-0.6*
Canary I.*s				-0.0062	-0.0068
				-2.3	*-2.6*
Adjusted R^2	0.40	0.40	0.41	0.41	0.40
n	9743	9743	9743	9743	9743

1. Indicate the interaction between schooling and region.
White robust t-statistic below the coefficient

Table 4b. Rate of return including additional variables. Men. ECHP-1994. Hourly gross wage

Constant	5.9287	6.0885	6.0732	5.9374	5.9327
	126.2	*109.3*	*110.5*	*125.7*	*126.8*
Schooling	0.0648	0.0733	0.0638	0.0856	0.0752
	22.8	*28.5*	*22.7*	*27.5*	*22.2*
Experience	0.0377	0.0399	0.0383	0.0401	0.0385
	13.0	*13.5*	*13.3*	*13.6*	*13.4*
Experience^2	-0.0004	-0.0005	-0.0005	-0.0005	-0.0005
	-8.3	*-8.3*	*-8.6*	*-8.4*	*-8.7*
Industry	0.1430		0.1337		0.1359
	6.2		*5.8*		*5.9*
Building	0.0099		0.0022		0.0044
	0.2		*0.1*		*0.1*
Services	0.2718		0.2614		0.2618
	10.8		*10.6*		*10.7*
Northwest		-0.3034	-0.2807		
		-8.3	*-7.9*		
Northeast		-0.0896	-0.0786		
		-2.8	*-2.5*		
East		-0.1255	-0.1107		
		-4.0	*-3.6*		
South		-0.1718	-0.1646		
		-5.0	*-4.9*		
Centre		-0.1947	-0.1856		
		-5.3	*-5.3*		
Canary I.		-0.2212	-0.1998		
		-4.1	*-3.9*		
Northwest*s[1]				-0.0251	-0.0229
				-7.7	*-7.1*
Northeast*s				-0.0076	-0.0066
				-2.7	*-2.3*
Centre*s				-0.0091	-0.0078
				-3.2	*-2.7*
East*s				-0.0161	-0.0155
				-5.2	*-5.1*
South*s				-0.0183	-0.0173
				-5.4	*-5.3*
Canary I.*s				-0.0168	-0.0149
				-2.9	*-2.7*
Adjusted R^2	0.38	0.36	0.40	0.36	0.40
n	2181	2181	2181	2181	2181

1. Indicate the interaction between schooling and region.
White robust t-statistic below the coefficient

Table 4c. Rate of return including additional variables. Men. WSS-1995. Hourly gross wage

Constant	6.2923	6.0564	5.9809	6.0800	6.2711
	878.8	*935.6*	*932.4*	*829.0*	*787.1*
Schooling	0.0759	0.0805	0.0745	0.0754	0.0648
	221.8	*236.1*	*222.3*	*211.4*	*187.3*
Experience	0.0408	0.0491	0.0462	0.0465	0.0360
	99.9	*122.8*	*120.5*	*116.7*	*93.9*
Experience^2	-0.0004	-0.0005	-0.0005	-0.0005	-0.0004
	-62.7	*-76.1*	*-76.4*	*-72.3*	*-60.2*
Contract	-0.2708				-0.2277
	-93.3				*-81.2*
Ownership		0.1873			0.0363
		49.1			*9.9*
Size of firm 2			0.0789		0.0780
			24.1		*25.1*
Size of firm 3			0.1890		0.1805
			52.6		*53.0*
Size of firm 4			0.2821		0.2619
			79.3		*77.3*
Size of firm 5			0.3873		0.3473
			121.5		*113.8*
Extrac				0.1652	0.1235
				19.0	*15.8*
Manuf				0.0528	-0.0142
				13.1	*-3.7*
Utilities				0.3118	0.2060
				48.3	*35.3*
Trade				-0.0038	-0.0664
				-0.7	*-12.5*
Hotels				-0.0947	-0.1394
				-14.7	*-23.4*
Transcom				0.1017	0.0130
				19.0	*2.6*
Finance				0.3419	0.2123
				63.2	*40.1*
Busser				0.0147	-0.0144
				2.2	*-2.2*
Adjusted R^2	0.42	0.39	0.47	0.43	0.53
n	118027	118027	118027	118027	118027

White robust t-statistic below the coefficient

Table 5. Return to schooling in women with selection bias correction (%). ECHP-1994

Constant	6.0721	6.1250
	42.0	*43.5*
Schooling	0.0745	0.0736
	13.3	*13.6*
Experience	0.0331	0.0306
	6.0	*5.8*
Experience^2	-0.0005	-0.0004
	-4.5	*-4.5*
Part time		0.3854
		7.2
Lambda 1	0.0660	0.0629
	1.2	*1.1*
Lambda 2	-0.2148	-0.2353
	-2.8	*-3.1*
Adjusted R^2	0.41	0.43
n	906	848

White robust t-statistic below the coefficient

Table 6. Rate of return. Public and private sector

	HBS-1990/91				ECHP-1994			
	Men		Women		Men		Women	
	Public	Private	Public	Private	Public	Private	Public	Private
Constant	13.3903	13.0179	13.2458	13.0923	6.3079	5.9551	6.2305	5.8627
	370.0	*562.5*	*347.6*	*360.2*	*64.1*	*104.1*	*47.9*	*63.1*
Schooling	0.0606	0.0712	0.0726	0.0571	0.0643	0.0689	0.0692	0.0681
	33.9	*47.9*	*35.3*	*22.2*	*15.5*	*19.4*	*10.6*	*12.3*
Experience	0.0337	0.0491	0.0177	0.0350	0.0321	0.0391	0.0296	0.0399
	17.1	*37.6*	*8.5*	*16.7*	*5.5*	*11.8*	*4.4*	*6.6*
Experience^2	-0.0004	-0.0006	-0.0001	-0.0005	-0.0004	-0.0005	-0.0004	-0.0005
	-11.6	*-27.2*	*-2.9*	*-11.0*	*-3.5*	*-7.5*	*-2.8*	*-4.3*
Adjusted R^2	0.39	0.35	0.46	0.29	0.34	0.26	0.34	0.28
n	2551	8192	1444	1689	569	1612	371	477

White robust t-statistic below the coefficient

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