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A DEMAND-SUPPLY ANALYSIS OF THE SPANISH EDUCATION WAGE PREMIUM*

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This paper estimates education demand for Spain in order to analyze whether variations of the education wage premium between the 1980s and 1990s can be explained within the framework of a supply and demand model. The evidence reveals a stable pattern of growth in the demand for education throughout the eighties and nineties, empirically showing that differences in the evolution of the education wage premium during the two decades can be explained by varying growth in education supply coupled with stable growth in education demand.

Key words: education wage premium, demand for education, labor supply.

JEL Classification: J24, J31, O33.

Recent decades have seen substantial heterogeneity in the evolution of the education wage premium, both across countries and over time [Katz, Loveman and Blanchflower (1995); Gottschalk and Smeeding (1997); Gottschalk and Joyce (1998); Acemoglu (2003)]. A natural starting point for the analysis of these differences is the demand-supply framework (D&S). The purpose of the D&S framework is to examine whether the evolution of the education wage premium can be approximated by supply-driven movements along a labor demand curve with a stable slope, plus shifts in labor demand. The results have been quite encouraging in a variety of contexts. Katz and Murphy (1992), for example, conclude that the education wage premium in the U.S. between 1963 and 1987 can be explained by steady, secular shifts in the demand for educated workers combined with observed changes in relative supply. Katz, Loveman, and Blanchflower (1995) show that the D&S framework is also useful for understanding the evolution of the wage premium in four OECD countries (the U.S., the U.K., Japan and France).

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Card, Kramarz, and Lemieux (1999) incorporate wage-setting institutions in a D&S framework and show that this helps to explain relative wage trends among less-skilled workers in the U.S., Canada and France in the 1980s. Acemoglu (2003) finds that the D&S with steady, secular shifts in the demand for educated workers can account for the differences in the evolution of wage inequality between Finland and Norway.

While the Spanish education wage premium has been studied quite intensively[1], the literature has not yet explored whether its evolution over time may fit within the D&S framework. The goal of this study is to ascertain whether the D&S framework can help explain the evolution of the education wage premium in Spain during the two decades between 1980 and 2000. The main finding is that the evolution of the premium during these two decades can be well approximated by combining the observed changes in labor supply with steady growth in the demand for education during the period 1980-2000. Interestingly, our estimates of the slope of the Spanish demand curve for education and education demand growth are quite similar to U.S. estimates.

One of the key elements of the D&S framework is the slope of the demand curve for education (which, in the standard D&S framework, is the inverse of the elasticity of substitution between more and less educated workers). The main difficulty faced when estimating this slope is that education supply and the education wage premium are determined simultaneously by demand and supply. Estimation therefore requires solving the standard identification problem (see Hamermesh, 1993, for a summary of this problem in the context of labor demand estimation). The empirical literature on the demand curve for education stretches back to the 1970s. Johnson (1970) estimates the elasticity of substitution between more and less educated workers to be 1.34 for a cross section of U.S. states in 1960. Ciccone and Peri (2005), using a panel of US states for the 1950-1990 period, and employing Acemoglu and Angrist's (2001) state-time-dependent child labor and compulsory school attendance laws as instruments for changes in the supply of education, find an elasticity of substitution of about 1.5. Angrist (1995) finds an elasticity of substitution of about 2 for data on Palestinian workers in the West Bank and the Gaza Strip during the 1980s; he uses the number of local higher-education institutions as an instrument for education supply. Fallon and Layard (1975), using cross-country data and employing income per capita as their instrument for education supply, obtain an estimate of 1.49 for the elasticity of substitution between more and less educated workers. Caselli and Coleman (2000) apply a D&S framework with endogenous technology to cross-country data and obtain an elasticity of substitution of 1.31. Katz and Murphy (1992) derive an elasticity of substitution of about 1.4 from U.S. time-series data for 1963-1987.

This paper uses the approach developed by Katz and Murphy to estimate the elasticity of substitution in a panel of Spanish regions, employing the beginning-of-period population structure as our instrument for education supply. The resulting estimate of the elasticity of substitution between more and less educated workers in Spain is close to the estimates reported by Katz and Murphy and Ciccone and Peri for the United States.

(1) See Abadíe (1997), Arellano, Bentolila and Bover (2001), Torres (2002) and Martinez-Ros (2001).

Our estimate of the slope of the Spanish demand curve for education for the 1980-2000 period allows us to examine the degree to which the D&S framework can be used to explain the evolution of the Spanish education wage premium during the period. Our chief empirical finding is that the evolution of the education wage premium as predicted by the framework fits quite closely with its actual evolution. For example, our estimates imply a fall in the relative wage of more educated workers during the 1980s of 1.45% and 0.93% in the 1990s, which comes close to the actual 1.52% drop in relative wages in the 1980s and the 0.92% decrease in the 1990s. Interestingly, we find similar annual growth rates for relative education demand in Spain (labor demand shifts) during the 1980s and the 1990s (2.7%). These estimates come close to estimates for the United States: for example, Katz and Murphy (1992) estimate the relative U.S. demand shifts to be about 3.3% per year, while Acemoglu (2002) reports an increase of about 2.5% annually.

One explanation for cross-country differences in the evolution of wage inequality, especially in Europe, is that wage-setting institutions differ by country [Acemoglu (2003); Card, Kramarz and Lemieux (1999); Abraham and Houseman (1993)]. Arguably, the most important institution for the Spanish case is collective wage bargaining; however, taking this into account does not affect our conclusion that the D&S framework is able to capture the evolution of the Spanish education wage premium.

The rest of the paper is structured as follows. Section 1 explains the Demand and Supply framework; Section 2 explains the data used and the measurement of relative wages and education supply; Section 3 presents the estimation and results; Section 4 evaluates the possible effects of collective bargaining on relative demand estimates; and Section 5 concludes.

1. The Demand and Supply Framework

According to the demand and supply framework, the wage of more relative to less educated workers (the education wage premium) is determined by education demand and supply. The simplest model of relative demand is based on the constant elasticity of substitution (CES) firm-level production function (see, for example, Katz and Murphy, 1992). The model assumes that firms *f* have access to the following production function:

$$Y = \left[A_f L^{\rho} + B_f H^{\rho} \right]^{\frac{1}{\rho}} \qquad [1]$$

where Y is output, H is the input of more educated (skilled) workers, and L the input of less educated (unskilled) workers. A_f and B_f denote the levels of factor-augmenting technology to which firms have access. It is straightforward to show that the production function parameter ρ determines the elasticity of substitution between factors σ. In particular, $\sigma = 1/(1 - \rho)$, which implies that $\rho \leq 1$ is necessary for the isoquants to be convex and the education demand curve to be well-defined. The case $\rho = 1$ corresponds to the case where the two types of labor are perfect substitutes, while $\rho \rightarrow -\infty$ implies that there is no substitutability at all between more and less educated workers).

Firms are assumed to take wages in the labor market as given when making their hiring decisions. Firms' demand for education, the demand for more relative to less educated workers *H/L*, can be obtained from their first-order conditions for profit-maximization as

$$\left(\frac{H}{L}\right)_D = \left(\frac{B_f}{A_f}\right)^{\sigma}\left(\frac{w^H}{w^L}\right)^{-\sigma}$$

where we have used that $\sigma = 1/(1 - \rho)$.

The D&S framework can be applied to the regional level by assuming that firms in region *i* have levels of factor-augmenting technology $A_f = A_i$ and $B_f = B_i$. A region's equilibrium education wage premium can now be determined by equating education demand with education supply H/L_{Si} in region *i* and solving for the relative wage for educated workers,

$$\left(\frac{w^H}{w^L}\right)_i = \left(\frac{B}{A}\right)_i\left(\frac{H}{L}\right)_{Si}^{-\frac{1}{\sigma}}$$

Taking logs on both sides yields

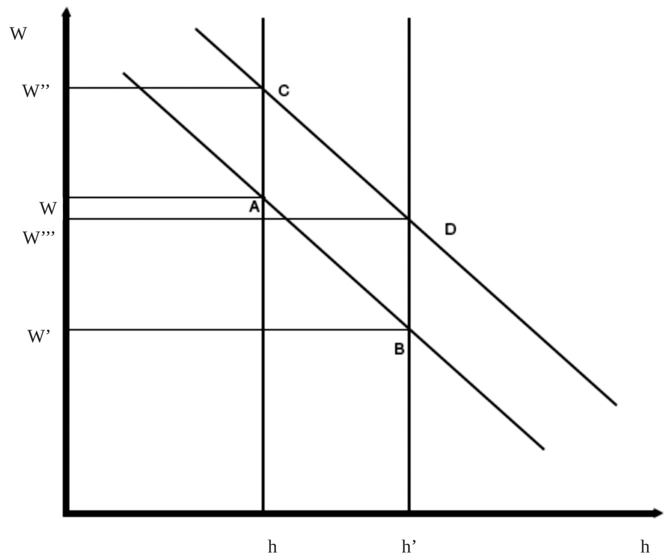
$$\omega_i = b_i - \frac{1}{\sigma}h_{Si}; \qquad [2]$$

where $\omega_i = Ln(w^H / w^L)$, $b = Ln\ (B/A)$ and $h = Ln\ (H/L)$. Taking differences over time (denoted by Δ) yields.

$$\Delta\omega_{it} = \Delta b_{it} - \frac{1}{\sigma}\Delta h_{iSt} \qquad [3]$$

Hence, log changes in the education wage premium, Δw_{it}, are equal to shifts in education demand, Δb_{it}, plus supply-driven movements along the education demand curve, $-\frac{1}{\sigma}\Delta h_{iSt}$. The strength of the effect of supply changes on the wage premium depends on the slope of the inverse education demand curve, $1/\sigma$, which is equal to the inverse of the elasticity of substitution between more and less educated workers. When the elasticity of substitution is high, supply changes will have small effects on the education wage premium (the inverse demand curve is flat). As the elasticity of substitution between more and less educated workers falls, the sensitivity of the education wage premium to changes in education supply increases. Figure 1 provides a graphic illustration of the relative wage effects of demand shifts and supply-driven movements along the demand curve. An increase in the relative supply, from *h* to *h'*, moves the equilibrium point along the downward-sloping inverse demand curve (A to B) and reduces the education wage premium. An increase in the relative demand for educated workers moves the equilibrium point to C and increases the education wage premium. When there are demand and supply shifts, the equilibrium rests at point D; in this case, the behavior of the education wage premium depends on which shift prevails.

Figure 1: THE RELATIVE DEMAND FOR EDUCATION



Source: Own elaboration.

The key feature of [3] from our point of view is that, once the elasticity of substitution between more and less educated workers has been estimated, it can be used to determine how supply and demand affect the evolution of the education wage premium. In order to resolve the standard simultaneous-equation identification problem, estimating the elasticity of substitution requires a valid instrument for shifts in the regional education supply.

2. DATA AND MEASUREMENT

2.1. Data

2.1.1. Individual Data

The wage data for this study comes from the Household Budget Survey (EPF) and Continuous Household Budget Survey (ECPF)[2], both of which cover a wide range of individual characteristics, such as education, age, region, annual earnings, type of employment contract, etc. The EPF is available for 1974, 1980-81 and 1990-1991; the ECPF is available for every quarter since 1985. Although there are some differences, the information in the EPF since 1980 is quite similar to that in the ECPF. As the focus here is on long-term trends, we will use data corresponding to 1980-1981, 1990-1991 and 2000-2001. The 1974 survey is not

(2) Dating from 1985, but with a change in methodology in 1997, and processed by the Spanish National Institute of Statistics (INE).

considered because it provides no usable education data. Other sources of wage data either lack needed individual data or cover only short periods. An additional advantage of the EPF is that the methodology used to compile this data has remained relatively stable since 1980.

Table 1: SCHOOLING GROUPS EQUIVALENCE. INDIVIDUAL AND AGGREGATE DATA

	EPF 81 survey groups	EPF 91 survey groups	ECPF 01 survey groups	SS07 survey groups	Years of schooling
Illiterate-Primary	1,2,3	1,2,3,4	1,2	1,2	0-6
Basic secondary and basic vocational	4	5,7	3	3,5	10,12
Advanced secondary and advanced vocational	5,6	6,7,8	4,5	4,6	12,14
College (short cycle)	7	9	6	7	15
College (long cycle)	8	10	7	8	17

Note: EPF survey: Encuesta de Presupuestos Familiares.

ECPF survey: Encuesta Continua de Presupuestos Familiares.

SS07 data and years of schooling from Serrano and Soler (2007).

Definition of the schooling groups:

Groups EPF 81: 1 - Illiterate, 2 - No certificate (of Primary studies), 3 - Primary, 4 - Basic secondary, 5 - Advanced secondary, 6 - Vocational training, 7 - College (short cycle), 8 - College (long cycle).

Groups EPF 91: 1 - Illiterate, 2 - No certificate, 3 - Basic Primary, 4 - Advanced primary, 5 - Basic secondary, 6 - Advanced secondary, 7 - Basic vocational training, 8 - Advanced vocational training, 9 - College (short cycle), 10 - College (long cycle).

Groups ACP F01: 1 - Illiterate, no certificate, 2 - Primary, 3 - Basic secondary and basic vocational training, 4 - Advanced secondary and advanced vocational training, 5 - Other advanced secondary studies, 6 - College (short cycle), 7 - College (long cycle).

Source: EPF Survey, ECPF Survey and Structural Wage Survey, Spanish National Statistics Institute.

We will focus on heads of households aged 20 to 65 who work full time and are not self-employed[3]. This provides us with a sample size of 11,402 workers for 1980/81; 8,838 for 1990/91 and 3,751 for 2000/01.

Our schooling data refers to the highest of the possible degrees attained upon the subject's completion of primary school, basic and advanced secondary and vocational school and college (short and long cycle). Table 1 gives the different education levels considered in the 1980-1981 and 1990-1991 EPF and the 2000-2001 ECPF and the same data aggregated into 5 homogeneous groups.

(3) Further details of the rationale for this selection, together with a description of the processing of the main data including wages, are given in Appendix A.

We use this information to assign years of schooling to each of the individuals in our sample[4].

Table 2: AVERAGE WAGES BY SCHOOLING GROUPS (IN PESETAS, 1980-2001)

		1980/81	1990/91	2000/01
Primary or less	size	8,104	5,872	2,135
	wage	617,374	1,252,468	1,934,716
	% women	5.2	7.5	8.6
Secondary	size	2,182	1,653	736
	wage	841,026	1,679,672	2,409,868
	% women	8.4	10.0	14.9
College	size	1,116	1,313	880
	wage	1,109,756	2,178,330	2,947,811
	% women	11.4	16.8	21.5
All	size	11,402	8,838	3,751
	wage	708,367	1,469,914	2,265,624
	% women	6.4	9.3	12.9

Source: Household Budget Survey (EPF) and Continuous Household Budget Survey (ECPF).

Table 2 shows some descriptive statistics for the sample taken for each year and level of education. For the sake of simplicity, the level of education was aggregated into three groups, primary or less, secondary and college, using the same aggregation criterion as that described in Table 1. The data shown are the sample size for each year and education group, average wages and the percentage of women in the sample.

Here, we must mention one very important final point that affects the adequacy of the data. The fact that the surveys covered only heads of households means that women are under-represented in the sample. As the table shows, the percentage of women among the total number of workers selected is 6.4% in 1980/81, 9.3% in 1990/91 and 12.9% in 2000/01. These figures are far below those yielded by labor-market surveys. For instance, according to the Economically Active Population Survey, published by INE (the Spanish National Institute of Statistics), women accounted for 27.9% of the total Spanish labor force between 1980 and 1981, 31.6% between 1990 and 1991 and 36.9% between 2000 and 2001, which, as will be explained later, will force us to seek a solution to the possible bias that may result from this under-representation.

(4) Serrano and Soler (2007) is used to assign these years.

2.1.2. Aggregate Supply Data

Our data on the aggregate supply of schooling comes from Instituto Valenciano de Investigaciones Económicas (IVIE) and Bancaja surveys (Serrano and Soler, 2007). The schooling data refers to employed workers in 17 Spanish regions from 1964 to 2007. The categories are:

1) Illiterate
2) No schooling or primary schooling only
3) Basic secondary
4) Advanced secondary
5) Basic vocational
6) Advanced vocational
7) Previous to college and short cycle college degree
8) Long cycle college degree

Less educated workers are defined as belonging to either the first, the second or the third group in Table 1. That is, workers with fewer than 14 years of schooling are considered to be less educated, while those with a college education are defined as more educated. These definitions are made for the purpose of comparison with other analyses for the U.S., U.K. and other countries[5].

2.2. Instruments

Our approach uses the beginning-of-period population structure as an instrument for analyzing regional changes in schooling supply. The required population data were drawn from the 1981 and 1991 Spanish Population Censuses provided by the National Statistics Institute.

3. Measurement and Descriptive Statistics

3.1. Education Wage Premium

Wages depend not only on schooling but on many other individual characteristics. To isolate the role of schooling, two approaches might be used. One could use all of the available individual characteristics to build a narrow definition of worker cohorts, then calculate the education wage premium as the wage of one narrowly-defined cohort relative to another, less-educated cohort that is very similar to the first in all other dimensions. However, this strategy requires many observations. We therefore focus on a second strategy based on Mincer wage regressions [Mincer (1974)]. Using *j* for individuals, *t* for years, and *i* for regions, we estimate

$$\ln(w_{it}^{j}) = \alpha_{it} + \beta_{it} S_t^j + \gamma_{it}^1 E_t^j + \gamma_{it}^2 (E_t^j)^2 + \mu_{it} X_t^j + \varepsilon_t^j . \qquad [4]$$

(5) Acemoglu (2002) defines this classification for the US. However, he considers this a simplification in a context in which there is a continuum of imperfectly substitutable skills.

The left-hand side is the log of individual wages and the right-hand side contains a list of explanatory variables: years of schooling (S_t^j), years of experience (E_t^j), and other k variables (represented by the $k \times 1$ vector (X_t^j) such as marital status, employment sector (agriculture worker) and gender. As usual, experience is calculated as age minus years of schooling minus six. The key parameter is β_{it}, that is, the percentage increase in wages (the return) from one year of schooling in any given region/year. Once we have estimated this return, we obtain the log education premium of workers with S^H years of schooling relative to workers with S^L years of schooling in region i for year t by multiplying the difference in years of schooling by the estimated return to schooling ($\hat{\beta}_{it}$).

The correct estimation of β_{it} raises two problems. Firstly, as advanced in the previous section, there is a sample selection problem due to the lack of data for some individuals. Since the survey focuses exclusively on heads of households, women are clearly under-represented in the estimation of [4]. The β_{it} estimate may therefore be biased. One solution would be estimate equation [4] for men and women separately and estimate the betas as the weighted average of the two. This would result in very low numbers of women for some regions, which would seriously affect our estimates. Another possibility would be to estimate [4] with weighted least squares (WLS), weighting by the percentage share of different groups defined by gender, region or level of education in the total number of employed in Spain. The 1991 and 2001 censuses provide sufficient data to perform the necessary calculations. Although the 1981 census data are not sufficiently detailed in this respect, the 1981 EPF data are adequate for calculation purposes. The latter strategy is the one chosen in the present study. The need to include women and find a means to avoid the above-mentioned bias arises from the fact that our supply variable captures both sexes jointly, thus justifying the need to resort to the process described in the preceding paragraph.

The second problem is that, with Mincerian wage regressions estimated using ordinary least squares, schooling can be correlated with unobservable characteristics (e.g., ability) that may also affect wages. While some Spanish studies have sought to address these concerns using instrumental variables, none of them use EPF or ECPF data, since these surveys do not provide suitable instruments. Nevertheless, there are two reasons to believe this concern should not affect our analysis. First, many studies have shown that the bias is quite small [Card (1999)]. Moreover, since our study focuses on the evolution of the education wage premium, our analysis will not be affected by the bias as long at the latter remains approximately constant in time. Another issue is that our estimating equation implies that the return to an additional year of schooling is independent of the level of schooling. In principle, we could relax this assumption by estimating the return to schooling only for those who have attained certain levels of education (degrees). But we do not have sufficient data to follow this approach for some of the smaller Spanish regions.

Table 3 contains our return-to-schooling estimates (or β) for Spain as a whole, obtained using the data on individuals available in our surveys. Here, it can be seen that return to education fell during the 1980s and 1990s. There are similar, previous findings for Spain; for example, Abadíe (1997) finds that Spanish wage inequality fell during 1980s, partly due to a decrease in the return to education[6]. We have only partial data for the 1990s. Barceinas, Oliver, Raymond and Roig (2000a) find an increase in return to education during the first half of the decade, while Izquierdo and Lacuesta (2007) find a notable decrease during the second half.

Table 3: SPANISH RETURNS TO EDUCATION. 1980/81, 1990/91 AND 2000/01

	1980-81	1990-91	2000-01
β_t	0.074	0.064	0.058
	(0.002)	(0.001)	(0.001)

Note: estimations for 1980/81 and 1990/91 are based on EPF and 2000/01 on ECPF. β_t represents the average return to education for Spain. Data in parenthesis represent standard deviations. The returns are estimated using Mincer equations and OLS. The sample is limited to heads of family and non self-employed workers.

Source: EPF and ECPF, Spanish National Statistics Institute.

3.2. Relative Supply

We first aggregate workers with only primary schooling and workers with lower-level secondary schooling using

$$L_{it} = L_{it}^2 + a_{it}^{L1} L_{it}^1 + a_{it}^{L3} L_{it}^3 + a_{it}^{L4} L_{it}^4 + a_{it}^{L5} L_{it}^5 + a_{it}^{L6} L_{it}^6$$

where a_{it}^{ln} is the efficiency of workers with n-level of schooling relative to workers with no schooling or primary schooling and n are the groups defined for the aggregate, Section 2.1.2. This efficiency parameter is obtained as the education premium in region *i* and year *t* of workers with n-level education relative workers with no more than a primary-school education. Here, the supply of more educated workers is obtained by aggregating workers with college (short and long cycle) educated workers using

$$H_{it} = H_{it}^7 + a_{it}^H H_{it}^8$$

where a_{it}^H is obtained as the education premium in region *i* and year *t* of workers with long-cycle college degrees with respect to short-cycle college workers. The log supply of education can now be obtained as

$$h_{i,t} = \ln \frac{H_{it}}{L_{it}}$$

(6) Other works point in a different direction however, maybe because of the use of different surveys to compare trends in the return to education. Barceinas, Oliver, Raymond, and Roig (2000b) estimate the return to education using EPF for 1980 and ECPF for 1985-1996. They found that return to education increased during this period, except between 1985 and 1991 when it fell. Their estimate of the return to education is 5.9% for 1980 and 7.0% for 1990. The estimates obtained in this paper are 6.4% for 1980, 6.0% for 1990, and 7.0% in 2000. The major differences in the 1990 figures are due to the use of ECPF data for that year.

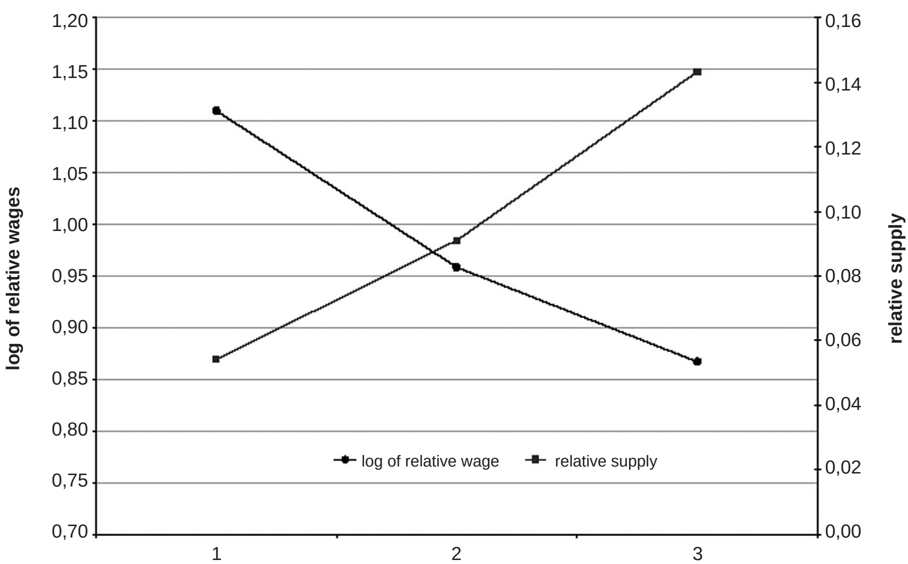
3.3. Descriptive Statistics

Figure 2 and Table 4 contain information on the education wage premium and the relative supply of schooling for the 1980s and 1990s. It can be seen that the education wage premium fell between 1980 and 1990 (from 1.11 to 0.95) and between 1990 and 2000 (from 0.95 to 0.86). The implied annual growth rates are equal to -1.52% during the 1980s and -0.92% during the 1990s. The (log) relative supply of schooling, on the other hand, increased from 0.054 to 0.091 during the 1980s and from 0.091 to 0.143 during the 1990s. The implied annual growth rates were 5.19% during the 1980s and 4.53% during the 1990s.

Whether the pattern in Figure 2 and Table 4 is sensitive to the way education groups are aggregated is an important issue. As a robustness check, therefore, we classify workers with advanced-secondary and vocational schooling in the higher education group and then repeat the analysis using this new classification. The results are shown in Figure 3 and Table 5.

Qualitatively, the evolution of the education wage premium and relative supply of schooling for this new classification is very similar to the one we obtained earlier.

Figure 2: Education Wage Premium and Relative Supply of Skills in Spain. 1980-2000 (More educated workers have previous to college or college education)



Source: Own elaboration with ECPF, EPF and Serrano and Soler (2007) data.

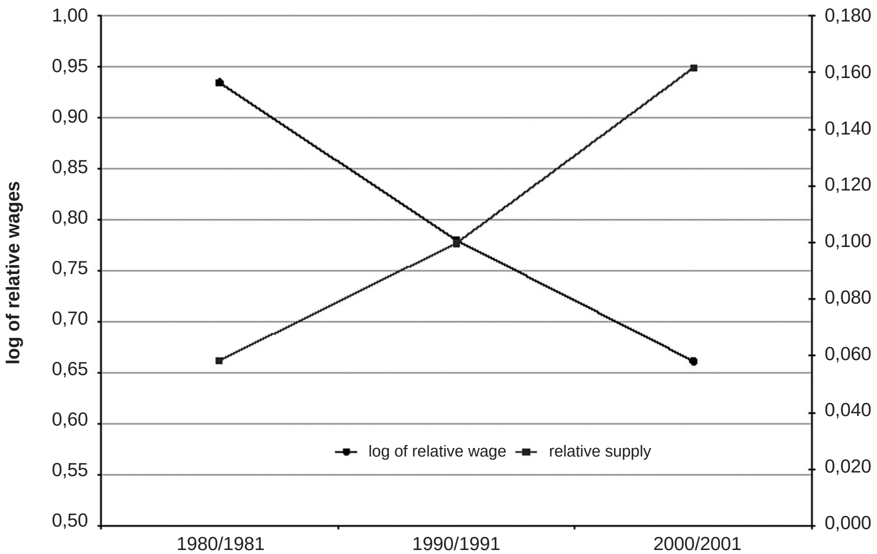
Table 4: RELATIVE SUPPLY AND LOGS RELATIVE WAGE IN SPAIN (I)

	1980-81	1990-91	2000-01
w_t	11.11	0.95	0.86
h_t	0.054	0.091	0.143

Note: w_t denotes the average Spanish education wage premium for year t while h_t represents relative supply for better-educated workers.

Source: EPF and ECPF, Spanish National Statistics Institute.

Figure 3: EDUCATION WAGE PREMIUM AND RELATIVE SUPPLY OF SKILLS IN SPAIN. 1980-2000 (MORE EDUCATED WORKERS HAVE ADVANCED SECONDARY, ADVANCE)



Source: Own elaboration with ECPF, EPF and Serrano and Soler (2007) data.

Table 5: RELATIVE SUPPLY AND LOGS RELATIVE WAGE IN SPAIN (II)

	1980-81	1990-91	2000-01
w_t	0.93	0.78	0.66
h_t	0.06	0.10	0.16

Note: w_t denotes the average Spanish education wage premium for year t while h_t represents relative supply for better-educated workers.

Source: EPF and ECPF, Spanish National Statistics Institute.

4. Estimation and Results

To gauge the extent to which the evolution of the education wage premium can be explained by the demand-supply framework, we estimate

$$\Delta\omega_{it} = \Delta b_t - \frac{1}{\sigma}\Delta h_{iSt} + (\Delta b_{it} - \Delta b_t) + \eta_{it} \qquad [5]$$

where Δw_{it} is the change in the log of education wage premium, Δb_t, the national shift in education demand, captures supply-driven movements along regional education demand curves, – (1/σ) Δh_{ist} are regional shocks to labor demand and Δ_{it} represents a time-region specific error.

Given that the data are for two decades of change, expression [5] can be rewritten as

$$\Delta\omega_{it} = \Delta b_{80} + (\Delta b_{90} - \Delta b_{80}) - \frac{1}{\sigma}\Delta h_{iSt} + (\Delta b_{i,80} - \Delta b_{80}) + \eta_{it} \qquad [6]$$

where Δb_{80} is average demand growth at national level for both decades, Δb_{90} – Δb_{80} is a dummy that takes a value of 0 for the eighties and 1 for the nineties in order to capture potential changes in relative demand growth for the second decade, and $b_{i,80}$ – Δb_{80} is the average value of change in demand for every region *i* relative to change in the nation as a whole. This last variable will be incorporated into estimation [6] as a regional effect.

Changes in the regional supply of educated workers are likely to be positively correlated to shifts in regional labor demand, which implies that the inverse elasticity of substitution between more- and less-educated workers cannot be estimated by applying ordinary least squares estimation to [6]. This positive correlation may be the result of worker migration to regions with rapidly rising wages, or it may reflect the fact that individuals living in regions where education is highly paid may decide to remain in school longer. Thus, it is necessary to find instruments for changes in the supply of education. Since the beginning-of-period population structure should be unaffected by shocks to regional labor demand, we will use the regional population structure in 1980 as an instrument for changes in the education supply during the 1980s and the population structure in 1990 as an instrument for changes in the education supply during the 1990s. Since changes in legal schooling requirements and better educational opportunities have increased the educational levels of younger people throughout Spain between 1980 and 2000, relative to that of the generation that retired during those two decades, changes in the regional supply of educated workers should also correlate to changes in the supply of education. Hence, the average level of schooling should have increased more rapidly in regions where there were more young people (under 20s) and more older people (between 60 and 64) in 1980 and 1990, relative to other regions. It suggests a positive link between these three variables. The beginning-of-period population share of 5-to-19-year-olds and 60-to-64-year-olds are the instruments used to estimate [6]. Hence, our identifying assumption is that this population share affects the change in the education premium over the course of the following decade only through its effects on the relative supply of education. Table 6, which presents the first stage regression, shows that both the 5-to-20-year-old and 60-to-64-year-old population share have a highly significant positive effect on the growth of education supply.

Table 6 contains the second stage results. The first column of results is the baseline specification. The remaining columns give the results of various robustness checks using other variables such as physical capital stock per worker for all sectors, information and communication technology (ICT) physical capital per worker, and employment level. Both the data on physical capital per worker and information and communication technology (ICT) physical capital per worker come from IVIE series. The ideas of these robustness checks using capital is that (some types of) capital may be complementary to educated workers and therefore affect the education wage premium. Total employment is included to test for aggregate scale effects.

Table 6: First Stage Regression

Dependent Variable: changes in log of relative supply

5-19 years old	4.60**
	(1.91)
60-64 years old	2.03*
	(1.11)
Constant	13.22**
	(6.05)
Fixed effects	yes
R^2	0.40
-statistic	5.18

Note: dependent variable is log changes in regional relative supply due to extra education

The regressors (the 5-19 and 60-64 age group variables) represent the regional share in total population of people aged between 5 and 19 and 60 and 64.

Data in parenthesis are standard errors.

** and * means significance at 5 and 10%.

Source: Own elaboration.

The results indicate that $-1/\sigma$ is between -0.775 and -0.865, meaning that the education demand curve is downward sloping. Moreover, our estimates are statistically different from zero at the 5% level. The implied elasticity of substitution between more- and less- educated workers is between 1.1 and 1.3. This value is very near to that found elsewhere. Johnson (1970), for example, estimates the elasticity of substitution between more- and less-educated workers to be 1.34 for a cross-section of U.S. states in 1960. Fallon and Layard (1975) find an elasticity of substitution between less- and more-educated workers of 1.49, using cross-country data. Angrist (1995) reports an elasticity of substitution of about 2 and

Caselli and Coleman (2000) estimate the elasticity of substitution between more- and less-educated workers to be approximately 1.3. Katz and Murphy (1992), using U.S. time-series data for the 1963-1987 period, report an inelasticity of about 1.4 for substitution between more- and less-educated workers. Using different estimation methods, Ciccone and Peri (2005) argue that the long-term elasticity of substitution in the U.S. between 1950 and the 1990s was between 1 and 2.

By how much did the demand for education increase in the 1980s and 1990s according to our estimates? The first column shows a demand shift of 0.267 during the 1980s, which represents an annual increase of about 2.7%. Since the difference between the pace of the education demand shifts during the 1990s relative to the 1980s is not statistically different from zero at any conventional level, we cannot reject the hypothesis that labor demand increased by the same amount during the 1990s as during the 1980s. Interestingly, our estimated increase in education demand for Spain is very similar to that estimated by Katz and Murphy (1992), who report a value of 3.3%.

Table 7: Relative Demand Estimation

	I	II	III	IV
Δb_{80}	0.267*	0.274*	0.268*	-1.841
	(0.160)	(0.117)	(0.160)	(4.002)
$\Delta b_{90} - \Delta_{80}$	0.033	0.049	0.058	0.063
	(0.094)	(0.103)	(0.091)	(0.137)
$-\frac{1}{\sigma}$	-0.847**	-0.865**	-0.796**	-0.775**
	(0.414)	(0.426)	(0.401)	(0.369)
Δk	–	0.917	–	–
	–	(0.280)	–	–
Δk_{ict}	–	–	-0.231	–
	–	–	(0.172)	–
Employment	–	–	–	0.326
	–	–	–	(0.636)
Adj. R^2	0.453	0.467	0.532	0.486
no.	34	34	34	34

Note: the dependent variable is changes in logs of regional relative wages due to extra education. Δb_{80} estimates changes in the (inverse) relative demand intercept from 1980 to 2000. $\Delta b_{90} - \Delta_{80}$ estimates differences in changes in the (inverse) relative demand intercept in the nineties relative to the whole period. $\frac{1}{\sigma}$ estimates the coefficient associated with changes in the log of the regional relative supply due to extra education or the (inverse) relative demand slope. Δk are changes in the logs of regional physical capital per workers and Δk_{ict} are changes in the logs of regional ICT physical capital per workers. Employment is the log of regional employment at the start of the period. All estimations are performed controlling for regional effects.

** implies significance at 5% and * at 10%.

Values in parenthesis are standard deviations.

Source: Own elaboration.

Summarizing, we find that education demand grew during the 1980s at a rate roughly similar to that for 1990s, and that this increase in education demand approximates that found for the United States during the same period. Moreover, the elasticity of substitution between more- and less-educated workers found by this paper is also quite similar to that estimated for the United States.

We are now ready to decompose the change in the education wage premium into the part attributable to changes in demand and the part attributable to changes in supply. Table 8 shows this exercise. Our results show that with no shift in education demand, the education wage premium would have fallen by -3.89% per annum between 1980 and 2000. By decade, this decrease would have been stronger during the 1980s (-4.15%) than during the 1990s (-3.63%). In the presence of education demand shifts only, the education wage premium would have increased by 2.7% during both the 1980s and the 1990s. The last column shows that the change in the education wage premium predicted by the demand-supply model comes close to the change actually observed in the premium.

Table 8: DECOMPOSITION OF RELATIVE WAGE CHANGES

	Wages[1]	Supply	Demand	Error
1980/81 – 1990/91	-1.52	-4.15	2.70	-0.06
1990/91 – 2000/01	-0.92	-3.63	2.70	0.01
Average	-1.22	-3.89	2.70	0.02

Note: *Supply* represents relative wage growth rates if the only changes are in relative supply, or changes in relative wages along the relative demand curve. Wages are the values for relative wages derived from Section 3.2. *Demand* represents the growth rates given by the common constant in [5]).

1. Observed.

Source: Own elaboration.

5. LABOR INSTITUTIONS

Let us now examine whether our conclusions above are robust to the influence of wage-setting institutions. These institutions have changed in almost all countries during recent decades. In Spain, a new system of labor regulation was introduced during the 1980s[7], the most important feature of which was centralized collective bargaining (CB)[8]. The latter involved wages being negotiated between unions and employer associations, as it did in other European Countries such as

(7) Ley del Estatuto de los Trabajadores (1980).

(8) Almost 50% of negotiations take place at sector-province level; 26.6% are sectoral negotiations at national level.

Germany[9]. CB agreements set a wage floor, and while only 18% of workers are paid at the negotiated rate, they tend to be the least-paid of all workers [see Dolado, Felgueroso and Jimeno (1997), hereafter DFJ][10]. Hence, education premia could, in principle, be affected by CB wage floors; more importantly from our perspective, the evolution of the education wage premium in Spain may have been affected by rising CB wage floors leading to wage compression (while there is a minimum wage in Spain, this wage is lower than the wage floor set by CB and is therefore not regarded as binding). One way to check whether trends in CB wage floors did indeed raise the wages of the less-educated workers, relative to the market-clearing wage level, is to examine whether the unemployment rate among less-educated workers increased more rapidly than it did among other education-classified worker groups. Table 9, which lists the percentage change in unemployment rates by worker education category, shows no marked differences between less- versus more-educated workers as far as unemployment trends are concerned. For example, between 1980 and 1985 (a time of rising overall unemployment), the rates for workers with higher versus lower levels of education exhibited similar trends. Between 1995 and 2000, a period characterized by falling unemployment, both the primary and upper-secondary education groups registered the same decline. In summary, over the course of the past twenty years, a similar trend has characterized unemployment among workers in all education categories.

Table 9: Growth in Unemployment Rate by Education Groups in Spain

	1980-85	1985-90	1990-95	1995-00
None or primary only	77.3	-20.1	51.6	-34.2
Lower-secondary	63.3	-35.4	28.8	-45.5
Upper-secondary	56.4	-27.5	47.2	-34.0
Tertiary	75.4	-27.4	31.1	-43.7

Source: Human Capital Series compiled by IVIE, Spain.

Another way to check whether our conclusions are driven by CB is to re-estimate our demand-supply model after excluding workers for whom the floors of CB agreements are likely to be binding. DFJ argue that bargained wages earned by the most CB-influenced workers fell below 125% of the minimum wage in the early 1980s and below 140% of the minimum wage in 1990. Because their study

(9) Despite very low union affiliation, almost 80% of workers are covered by some collective bargaining agreement as negotiations are binding for most non-union workers.

(10) Dolado, Felgueroso and Jimeno (1997) also explain that the real value of the minimum wage does not play any role in wage determination. They show that the lowest wages are determined by the CB wage floors, and that there is no link between the minimum wage and CB wage floors.

ends in 1996, we have to approximate their criterion for the year 2000. Our basic assumption is that the most CB-influenced workers earned less than the average wage reported for a Spanish "peon" (unskilled workers) in the 2002 Wage Structure Survey, if we take into account CPI inflation between 2000 and 2002. Table 10 shows the results of our re-estimated model, after eliminating workers whose wages fell below the specified cutoffs. Here, the slope and the intercept are almost identical to those obtained earlier. Our conclusions regarding the evolution of the Spanish education wage premium therefore continue to hold.

Table 10: Relative Demand Estimation exclusive of institutional effects

	I	II	III	IV
Δb_{80}	0.224*	0.234*	0.224*	-2.760
	(0.151)	(0.145)	(0.156)	(4.320)
$\Delta b_{90} - \Delta_{80}$	0.063	0.085	0.088	0.112
	(0.095)	(0.102)	(0.094)	(0.140)
$-\frac{1}{\sigma}$	-0.779*	-0.800*	-0.734*	-0.692*
	(0.481)	(0.490)	(0.476)	(0.424)
Δk	–	0.272	–	–
	–	(0.295)	–	–
Δk_{ict}	–	–	-0.213	–
	–	–	(0.195)	–
Employment	–	–	–	0.463
	–	–	–	(0.688)
Adj. R^2	0.23	0.13	0.17	0.22
no.	34	34	34	34

Note: the dependent variable is changes in logs of regional relative wages due to extra education. Δb_{80} estimates changes in the (inverse) relative demand intercept from 1980 to 2000. $\Delta b_{90} - \Delta_{80}$ estimates differences in changes in the (inverse) relative demand intercept in the nineties relative to the whole period. $\frac{1}{\sigma}$ estimates the coefficient associated with changes in the log of the regional relative supply due to extra education or the (inverse) relative demand slope. Δk are changes in the logs of regional physical capital per workers and Δk_{ict} are changes in the logs of regional ICT physical capital per workers. Employment is the log of regional employment at the start of the period. All estimations are performed controlling for regional effects.

** implies significance at 5% and * at 10%.

Values in parenthesis are standard deviations.

Source: Own elaboration.

6. Conclusions

The main aim of this paper was to examine the extent to which a demand-supply model may be used to explain the evolution of the education wage premium during the 1980s and the 1990s. Our key finding was that the evolution of relative supply was the main driving force behind the changes in this premium, since the demand for education rose at a similar pace during the two decades under study. We also found that the increase in the demand for education during our study period –about 2.7% per year– roughly approximated that estimated for other countries, including the United States. Another important finding was that the elasticity of substitution between more- and less-educated workers in Spain was around 1.1 and 1.3 respectively, and thus almost identical to that obtained in previous cross-country and cross-regional studies. Our study therefore suggests that the trend in the Spanish wage premium can be explained quite well by market forces.

Appendix A. The EPF 80-81/90-91 and ECPFs 1985-2005

A.1. Data

Household Budget Surveys and Continuous Household Budget Surveys supply data on consumption patterns, household and/or personal income (depending on the focus of the analysis) and other relevant data on household members. Although these sources do not provide homogeneous wage statistic series, they give important information that is relevant to this kind of analysis, for which they have become the main current source of statistics. However, a number of problems need to be considered.

First, despite the vast amount of information available, complete information is only available for the head of family[11]. Therefore, this paper, like other Spanish studies (Abadíe 1997), works explicitly with this selection. Nevertheless, as noted in Section 3.1, there are methods of extending the main findings to the working population as a whole.

Second, annual wages are not immediately determined by recorded earnings data. The main problem is that there is no information on hours worked or similar criteria unless the head of family has worked for more than thirteen hours during the reference week. The only solution to this problem is to use only those workers who reported working more than 13 hours and to assume that they worked full time.

Third, as in Abadíe (1997), groups of workers with unrepresentative characteristics were eliminated. For instance, education and experience cohorts were defined by five-year segments, and cohorts with fewer than fifteen records were deleted.

Fourth, the EPFs from 1980-81 and 1990-91[12] provide a broad spectrum of data for about 20,000 families for each year. However, the quarterly surveys use a smaller sample size, since their main objective is to offer a short-term analysis of consumption, rather than consumption structure. In any case, since 1997, the sample has doubled. This problem may be resolved by using two-year samples to improve

(11) For instance, the ECPFs records contain information about education only for this group

(12) These surveys' information include six quarters.

sample size, since the ECPFs poll the same family for six quarters, changing one sixth of this sample each quarter. It is therefore necessary to adjust household incomes, not only for the reasons described above but also for calendar effects when working with quarterly data. Further details of this will be given in subsection A.2.

The next problem stems from the heterogeneity of the definitions and the classifications of variables used. For example, different educational level classifications appear each year, partly because current surveys have modified their definitions over the years and partly because the Spanish legal definition of education changed during the period under study[13]. It is, nevertheless, possible to find a common denominator between the groups defined by each survey of individual data and for the aggregate supply data, as shown in subsection 3.1 and in Table 1. The years imputed are the same as those given in Serrano and Soler (2007).

To conclude, despite the limitations of working with these surveys, they possess many redeeming features which make them our best choice for data on the distribution of Spanish wage inequality between 1980 and 2000. Two main criteria may be cited in support of their use: first, the lack of any better alternatives for this period and, second, their usefulness as a basis for comparative analysis.

A.2. Refining wages from EPF and ECPFs

Once the surveys were selected for each year, they were then refined. As has been described, only data for heads-of-family working over thirteen hours per week were considered, after eliminating all self-employed wage-earners. Because ECPF surveys are restricted to household income data, this study focused exclusively on households where the head of family is the only worker. This is why women are under-represented and why this paper uses estimation [5].

Second, it was assumed that wages reported as below the legal minimum were either earned by part-time workers or individuals who had been unemployed for at least one year, or were erroneous replies. The records were modified when annual reported wages fell below the legal minimum for that year.

Third and last, calendar effects are taken into account in the ECPF information. The quarterly nature of these surveys implies that the wages reported might be influenced by the quarter in which they are given. To eliminate this effect, families with wages for all six quarters were taken first, and the calendar effect was analyzed in these cases. Then, a wage-level factor was obtained for each quarter. Once these factors were obtained, all the workers' wages were deflated. Finally, the quarterly average wage was taken and multiplied by four to give the yearly wage for all of the workers, regardless of which quarter they had worked. In this case, the six quarter samples were combined to increase (double) the ECPF sample size, although the same result could probably be obtained by using all four quarters in the calendar context.

(13) In the early nineties, the education law changed from an earlier one passed in 1970, which introduced compulsory schooling in Spain up to the age of fourteen. In 1990, the LOGSE extended compulsory schooling to the age of sixteen and changed the organization of the educational cycle, as the 1970 law had.

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RESUMEN

En este trabajo se estima la demanda de educación para España, con el objeto de analizar si la evolución de la prima a la educación durante los ochenta y noventa puede ser explicada en el marco de un análisis de oferta y demanda. Como principal resultado se encuentra que el crecimiento en la demanda fue muy similar entre los ochenta y noventa. Por ello, los resultados empíricos encontrados nos dicen que la diferente evolución de la prima a la educación entre ambas décadas puede ser explicada combinando los cambios en el crecimiento de la oferta con una demanda que creció de forma estable.

Palabras clave: premio a la educación, demanda de educación, oferta de trabajo.

Clasificación JEL: J24, J31, O33.