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XARXA DE REFERÈNCIA EN ECONOMIA APLICADA

How market access shapes human capital investment in a peripheral country

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Abstract

Human capital endowment is one of the main factors influencing the level of development of a region. This paper analyses whether remoteness from economic activity has a negative effect on human capital accumulation and, consequently, on economic development. Making use of microdata this research proves that remoteness from economic activity has contributed to explain the divergences in the level of education observed across Spanish provinces over the last 50 years. The effect is significant even when controlling for the improvement of education supply. Nonetheless, the accessibility effect has been petering out since the 1960s due to the decreasing barriers to mobility.

Keywords: regional development, human capital, market access

JEL classifications: O10; R11; R40

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1. Introduction

Differences in development and income inequality across regions are one of the most intensely studied issues in the economics literature. Despite the efforts spent on promoting development and reducing inequality through an active and continuous regional policy, these inequalities persist over time. In the case of the Spanish regions, the process of convergence of income per capita that started in the 1950s has been strongly attenuated since the early eighties 1980s (De la Fuente, 2008). In 2007 the richest region, Madrid, had an income per capita 60% higher than Extremadura, the poorest (Cuadrado-Roura, 2010).

The growth literature has emphasised the importance of human capital accumulation for growth (Barro and Sala-i-Martin, 2004). Barro (1991) and Mankiw et al. (1992) showed that the educational level across countries is a significant variable in explaining differences in growth rates across countries. At the regional level, De la Fuente (2008) argues that education policy has been an important factor in reducing regional differences in income per capita in Spain, especially during the period 1985-2005. He finds that after TFP, human capital is the most important factor contributing to the convergence of productivity, explaining from 20% to 30% of convergence during the period. The spatial differences in human capital endowment translate into spatial productivity differentials. De la Fuente and Domenech (2006) estimate an average contribution of human capital to regional productivity differentials for the period 1960-2000 of 39.8%. Pablo-Romero and Gómez-Calero (2008) find that differences in human capital endowments were able to explain 29.3% of provincial (NUTS 3) differences in productivity for the period 1990-1999.

Given the importance of human capital endowments for growth and the differences across regions, it is worthwhile to analyse the determinants of educational incentives from a spatial perspective. The objective of this paper is to investigate how the individuals' birthplace may affect their educational attainment. Specifically, we aim to test the impact that access to economic activity may have on the educational attainment of individuals born in different Spanish provinces.

Essentially, the individual's decision to invest in education depends on the expected wage premium and the cost of education. Moreover, the cost of education will be affected by each individual's ability so as that the higher the ability the lower the cost. Hence, for a given cost of education and ability level, incentives to schooling will depend on the wage premium. There is extensive literature that relates the spatial variation of economic activity to a measure of market potential since the work of Harris (1954). More recently, the New Economic Geography (NEG) has provided a theoretical basis for this relationship. In this context, the development of the NEG model by Fujita, Krugman and Venables (1999) predicts that nominal factor prices vary across locations depending on their market access. In particular, they predict higher nominal wages paid by those firms close to large markets, the so-called "wage equation". The underlying hypothesis is that firms located in remoter markets will incur in higher trade costs on the sales to markets as well as on the purchases of intermediate inputs. As a consequence, the value-added attributable to the production factors will be higher. Several papers have provided evidence of the impact of market access on nominal wages, namely Redding and Venables (2004) for 101 countries; Hanson (2005) for North American counties; Mion (2004) for Italian provinces; Brakman, Garretsen and Scramm (2004) for German districts; and both Head and Mayer (2006) and Breinlich (2006) for European Union regions.

But market access can have an additional impact on factor accumulation. The theoretical and empirical research that looks at the relationship between human capital accumulation and access to markets is mostly based on the work by Redding and Schott (2003). These authors extend the standard two-sector Fujita et al. (1999) economic geography model by allowing for unskilled individuals to endogenously choose whether to invest in education. They show that countries located on the economic periphery that face higher trade costs have a lower skill premium and, consequently, the incentive for education falls. By reducing human capital accumulation, remoteness lowers gross domestic product. Based on their theoretical work, Redding and Schott (2003) provide evidence that countries with lower market access have lower levels of educational attainment. Using trade flow data, they construct a theoretically consistent measure of market access, following the methodology of Redding and Venables (2004), in order to approximate countries' remoteness from world economic activity. They regress educational attainment –defined as the proportion of the population which has completed secondary and tertiary education- on market access for a cross-section of 105 developed and developing countries. The results of their regression show that the estimated coefficient on market access is positive and statistically significant.

Following Redding and Schott (2003), several papers have analysed whether a similar relationship between educational attainment and market access can also be found across regions. Breinlich (2006) adapts the approach developed by Redding and Venables (2004) to the regional level to test whether the market access variable is as significant on a regional level as it is in the international context to explain the spatial structure of income. Specifically, he applies the NEG framework to examine the role that proximity to markets plays in explaining regional per capita income levels in the EU and finds a positive and significant effect of market access on Gross Value Added per capita. More interestingly for our purposes, he disentangles the different channels through which market access affects income levels, finding that physical and human capital accumulation plays a more important role than direct trade costs in explaining income differences.

López-Rodriguez, Faíña and López-Rodriguez (2007), once again within the NEG framework, analyse the impact of distance from markets on the level of human capital accumulation. These authors use aggregate data on educational attainment from the EU Labour Force Survey for a sample of 203 EU-15 regions defined at NUTS 2 level. The authors use the percentage of population aged 25-64 with low, medium or high levels of education as the dependent variables. Market access is computed as a distance weighted sum of regional GDPs for all EU-27 NUTS 2 regions. By regressing education attainment levels on market access, the authors conclude that there is a positive correlation between medium and high levels of regional educational attainment and access to economic mass. These results are confirmed when using years of schooling as the dependent variable instead of educational attainment.

Adopting the same methodological approach as the previous papers, Karahasan and López-Bazo (2013) test the impact of market access on regional variability in human capital across Spanish provinces. They posit that the market access variable is capturing the effects of variables that can affect the spatial distribution of human capital unless they are included in the estimation. In particular, they hypothesize that the non-inclusion of regional differences in industrial structure as well as the non-control for spatial dependence can bias estimates of the market access variable. The reason is that in the Redding and Schott (2003) model, individuals' education decision is affected by more variables than just market access, and these variables can be proxied by the regional industrial mix and the control of spatial dependence. Karahasan and López-Bazo (2013) use aggregate data for the 47 Spanish mainland provinces for the period between 1995 and 2007. They use a market potential à la Harris calculated with gross value added and road distance and average years of education as the dependent variables. A significant effect of market potential is found when no additional variables are included in the estimated equation. Moreover, the coefficient on market potential is decreasing over time, reflecting a progressively lower effect of trade. However, when the sectoral composition of employment and spatial dependence are controlled for, the role of market access in explaining regional differences in human capital accumulation becomes almost negligible.

However, the Redding and Schott theoretical model assumes the immobility of human capital. Given that labour mobility is significantly lower between countries than within countries, this model is applicable to cross-country data, but it is more dubious with cross-regional data in the same country. In this sense, we should consider the possibility that the spatial distribution of human capital within a country might be the result of educated workers migrating to regions where their educational investment can obtain higher returns. Crozet (2004) tests the hypothesis that workers within countries move taking into account the market potential of the potential destination regions as a result of the forward effect of Krugman's model. Crozet finds support for this hypothesis by using data for five European countries.[1] However, the simulations of agglomeration dynamics with his estimated parameters would explain a very limited centripetal force due to migration. Paluzie et al. (2009) estimate a model with Spanish data on the lines of Crozet to explain migration flows in the 1920s, the 1960s and 2000-2004. The authors find a positive but decreasing effect of market potential on migration in all three periods. Unlike the two first periods, when manufacturing activity potential was the driving force, in the third period migration was attracted by the market potential of service activities while manufacturing showed a negative effect on migration. Additionally, the estimates of the parameters controlling for the effect of trade costs and labour mobility costs witnessed a continuous decline through the three periods. It should be taken into account that the authors do not control for the education level of migrants.

The objective of our paper is to empirically test the impact that an individual's place of birth has on individual education decisions in a country located at the periphery of Europe which in recent decades has witnessed a process of international integration. In particular, our hypothesis is that as integration both at national and European level has been increasing, the local conditions in terms of market potential in the education decision have become less relevant due to a drop in mobility costs. We use micro-data to look at the relationship between human capital and access to market at regional level. The use of micro-data makes it possible to account for two factors that have a significant impact on the educational decisions: the individual's year of birth and province of birth. With respect to year of birth, the underlying hypothesis is that the conditions affecting the decision on education have changed over time. In particular, access to education has been significantly improved in the last decades. On the other

[1] Germany, Italy, the Netherlands, Spain and U.K., and the period analysed is the 1980s even though the exact period differs for each country.

hand, locating the individual at this province of birth and not at their province of residence, as is the case with aggregate data, makes it possible to compute the actual values for those variables that affect the decision on education. Specifically, when working with regions within a country and, hence, the assumption of labour immobility does not hold, the individual decision on the level of education will not only depend on the wage premium of their province of birth but also on the maximum wage premium attainable in some region in the country. However, given that migration has costs, labour mobility is not perfect and will depend on how tight the barriers to mobility are. In summary, our data allows us to measure the variables that determine individual decisions on education at the moment they took place in the province of birth.

After controlling for individual characteristics, the results show that, although remoteness from economic centres hampers human capital accumulation, the effect decreases inversely with the individual’s year of birth; specifically, the estimated coefficient diminishes from 1.04 at the beginning of the period to 0.22 at the end of the period. Moreover, taking advantage of the peripheral location of Spain within Europe, this paper provides evidence that remoteness with respect to the European economic mass has a greater impact on educational attainment than remoteness with respect to the Spanish economic mass.

The paper is organised in six sections. The following section develops the empirical modelling framework. The third section describes the data used. The fourth section establishes some facts on the spatial distribution of human capital in Spain relevant to the objective of the paper. The estimated model and the results are presented in the fifth section. Finally, the main conclusions of the paper are outlined in the last section.

2. Modelling framework

We have modelled individuals' education decision as follows. The individual born in province "h" facing the education decision chooses higher education if the wage premium is higher than a given threshold. The individual makes the decision by comparing the wage premium in all potential destination provinces (j=1,....,J) and discounts the costs of moving from "h" to "j". The individual's decision will be made according to the following equation:

$$Max\left\{wp_j\left(1-\lambda_{hj}\right)\right\} > f_h \quad \text{where the maximum is evaluated for all "}j\text{"} \qquad (1)$$

In this equation "wp_j" is the wage premium in province "j", "f_h" is the cost of education in *h* plus the individual's discount rate. As the individual might have to migrate once they have completed education in province *h* to earn this wage premium, "λ_{hj}", with a value of between zero and one, is introduced to capture the importance of barriers to mobility. If the individual is born in "h" and remains in "h", we assume that $\lambda_{hh} = 0$.

On the other hand, if we assume that the individual is making their decision for a given spatial distribution of activity, we can assume a relationship between wage premium and market potential:

$$wp = \theta \cdot MP \qquad (2)$$

By substituting (2) into (1) we get:

$$Max\left\{\theta \cdot MP_j\left(1-\lambda_{hj}\right)\right\} > f_h \qquad (3)$$

If we assume no barriers to mobility (3) becomes

$$Max\left\{\theta \cdot MP_j\right\} > f_h \qquad (4)$$

With no barriers to mobility the place of birth does not influence the decision to become educated. Regardless of the province of birth, the incentives to reach higher education are the same for all individuals. Conversely, if $0 < \lambda_{hj} \leq 1$, the incentives to become educated will depend on the place of birth as equation (3) shows.

In terms of a discrete choice model, the equation can be formulated as follows:

$$Y_{ih} = Const. + \beta \cdot Max\left\{\theta \cdot MP_j\left(1-\lambda_{hj}\right)\right\} + \gamma' \cdot X_{ih} + u_{ih} \qquad (5)$$

where "Y_{ih}" takes value one if the individual "i" who was born in the economic space "h" reaches higher education, and zero otherwise. On the other hand, the remaining explanatory variables of the equation are captured by "X".

If there are no barriers to mobility ($\lambda_{hj} = 0$), the equation takes the following form:

$$Y_{ih} = Const. + \beta \cdot Max\left\{\theta \cdot MP_j\right\} + \gamma' \cdot X_{ih} + u_{ih} \qquad (6)$$

The maximum of the expression within brackets is the same for all individuals, so its effects are captured by the constant term of the equation and the explanatory variable "$Max\left\{\theta \cdot MP_j\right\}$" disappears.

In order to estimate (5), the problem is that the maximum value of the expression within brackets is not observable because the barriers to mobility (economic barriers, cultural barriers, language barriers, etc.) are not observable. To estimate the equation, we introduced the null hypothesis of full barriers to mobility. This implies $\lambda_{hj} = 1$ and $\lambda_{hh} = 0$ and in this case:

$$Max\left\{\theta \cdot MP_j\left(1-\lambda_{hj}\right)\right\} = \theta \cdot MP_h \qquad (7)$$

So the decision to become educated is modelled under the hypothesis of no labour mobility:

$$\text{Prob}\left(Y_{ih} = 1\right) = \text{Prob}\left[Const. + \gamma \cdot MP_h + \phi' \cdot X_{ih} + u_{ih} > h^0\right] \qquad (8)$$

where “ h^0 ” is a threshold.

To estimate the model, the discrete choice variable “ Y_{ih} ” is substituted by years of schooling S_{ih}. The years of schooling can be interpreted as the dependent variable of an ordered probit model with a high number of alternatives.[2] These discrete alternatives, years of schooling, are treated as a continuous variable in order to carry out the econometric approximation, as is usual in the literature.

As a consequence, the equation to be estimated takes the following form:

$$S_{ih} = Const. + \varphi \cdot MP_h + \tau' X_{ih} + \varepsilon_{ij} \qquad (9)$$

In this context, the value and significance of φ is proxying the importance of the barriers to mobility for the labour force. Full mobility must have its counterpart in that the market potential of the place of birth is considered on the same footing as the other provinces’ market potential. To take into account changes in the mobility costs, it is also possible to carry out separate estimations by birth cohorts. The temporal evolution of φ proxies the evolution of the barriers to labour mobility. If the barriers to labour mobility decrease over time as a consequence of the integration process, φ must show a downward trend as we estimate (9) for younger cohorts.

3. Data

The data used in this paper come from the Spanish Labour Force Survey (LFS). The LFS is a large sample survey among households that measures the labour status of individuals aged 15 and over and provides information on other characteristics of the population. The survey is conducted by the National Institute of Statistics on a quarterly basis following the methodology established by EUROSTAT for all member states. The

[2] The analytical derivation has been carried out for a dicothomic probit. It is possible its generalization to an ordered probit but in this case the simplicity is lost without an improvement in conceptual terms.

sample size is about 180,000 people, rotating one-sixth of the sample each quarter. In this paper, we make use of the LFS micro-data on the level of education of the individual together with other personal characteristics, such as gender, age and province of birth. One of the advantages of this dataset is the possibility of locating individuals at a relatively detailed geographical level, NUTS 3, which corresponds to the 47 Spanish mainland provinces, the boundaries of which are defined on administrative grounds. Given that the objective of this paper is related to the role of road transport accessibility, individuals located in the Canary Islands, Balearic Islands, and the North-African cities of Ceuta and Melilla are excluded from the sample.

Using micro-data with information on the date and place of birth makes it possible to control for the point in time in which the individuals take their decisions on education as well as the province where the individual made that decision. Given that the factors determining the level of education evolve over time, it might well be that the skill premium and the cost of education in the late 1950s, when those born in the late 1930s took their decisions on education, were significantly different from their current levels. To account for this fact, the explanatory variables affecting educational level are computed according to the individual's year of birth. The anonymisation criteria applied to the LFS micro-data introduces some limitations to our analysis. Specifically, the individual’s year of birth is not included in the micro-data and the age variable is derived in 5-years age bands. This fact does not allow standard cohort analysis and forces us to compare cohorts at five-year intervals. The micro-data of the LFS are available only since 1999, so we started with the 1999 dataset and selected the following available surveys in five-year intervals. Therefore, we used data from the 1999, 2004 and 2009 LFS, always for the second quarter, and constructed five-year cohorts defined according to the 5-year age bands. We selected individuals aged 25 to 64 years old. For each cohort, a reference year was selected and all the explanatory variables referred to this year. Given that annual changes in these variables in a five-year interval proved to be small, subsuming the five-year cohort into a single year is not likely to imply significant loss of information. Table 1 shows the cohorts for each year in the study and Table 2 the number of observations in each cohort.

Table 1. Cohorts according to the year of birth

	1999 LFS		**2004 LFS**		**2009 LFS**	
Age interval	Year of birth	Reference year	Year of birth	Reference year	Year of birth	Reference year
25-29	1970-74	1991	1975-79	1996	1980-84	2001
30-34	1965-69	1986	1970-74	1991	1975-79	1996
35-39	1960-64	1981	1965-69	1986	1970-74	1991
40-44	1955-59	1976	1960-64	1981	1965-69	1986
45-49	1950-54	1971	1955-59	1976	1960-64	1981
50-54	1945-49	1966	1950-54	1971	1955-59	1976
55-59	1940-44	1961	1945-49	1966	1950-54	1971
60-64	1935-39	1956	1940-44	1961	1945-49	1966

Table 2. Number of observations per cohort

	1999 LFS		**2004 LFS**		**2009 LFS**	
Year of birth	Observations	%	Observations	%	Observations	%
1935-39	9,378	10.9	-	-	-	-
1940-44	9,308	10.8	8,277	10.6	-	-
1945-49	11,164	12.9	9,815	12.5	9,006	12.0
1959-54	11,050	12.8	9,763	12.5	9,111	12.2
1955-59	12,037	13.9	10,674	13.6	10,229	13.7
1960-64	12,014	13.9	11,188	14.3	10,764	14.4
1965-69	11,144	12.9	10,412	13.3	10,498	14.0
1970-74	10,232	11.9	9,351	12.0	9,633	12.9
1975-79	-	-	8,800	11.2	8,643	11.6
1980-84	-	-	-	-	6,963	9.3
All	86,327	100	78,280	100	74,847	100

Table 3 describes the mean and standard deviation for the main variables in the model computed for each cohort. The mean years of education for each cohort can be interpreted as the increase in the young population over time. Instead, education access and market potential variables are the actual mean values for the whole country.

Table 3. Mean and standard deviation (in parenthesis) for the variables in the equation (average across all individuals)

Cohort	Years of education	Gender	Education access	Market access total	Market access Spain	Market access Europe
1956	4.43	0.47	0.35	270.58	105.75	164.83
	(3.17)	(0.50)	(0.20)	(29.98)	(14.91)	(24.59)
1961	5.38	0.49	0.31	264.14	110.82	153.32
	(3.66)	(0.50)	(0.17)	(38.09)	(17.44)	(30.50)
1966	6.20	0.49	0.48	277.70	116.65	161.04
	(3.93)	(0.50)	(0.19)	(42.85)	(20.60)	(32.22)
1971	7.17	0.49	0.74	291.40	123.22	168.19
	(4.05)	(0.50)	(0.22)	(49.08)	(25.36)	(34.06)
1976	8.16	0.49	0.52	303.32	130.95	172.37
	(4.07)	(0.50)	(0.16)	(55.29)	(30.74)	(35.39)
1981	9.64	0.49	0.89	315.17	139.04	176.13
	(3.94)	(0.50)	(0.26)	(59.32)	(34.08)	(37.01)
1986	10.17	0.49	0.98	325.98	145.69	180.29
	(3.84)	(0.50)	(0.28)	(63.40)	(37.47)	(38.90)
1991	10.68	0.51	1.16	336.65	150.69	185.95
	(3.80)	(0.50)	(0.28)	(66.18)	(40.42)	(40.56)
1996	11.30	0.51	1.72	342.27	152.50	189.77
	(3.75)	(0.50)	(0.38)	(66.54)	(40.80)	(41.40)
2001	11.34	0.51	2.94	344.94	155.29	189.66
	(3.69)	(0.50)	(0.61)	(67.87)	(42.39)	(41.24)

The dependent variable in the estimated equation is the number of years of education of each individual in the sample. The LFS provides information on the highest level of education attained by the individual up to 19 categories. This categorical variable has been transformed into a continuous one - years of education - by applying the corresponding number of years in the educational system to each category. The Spanish educational system has undergone several reforms over the last few decades. Related to this paper, the most important one is the reform introduced in 1970 that affected the duration of the school cycles. Following de la Fuente and Doménech (2012), we have attributed to each cohort of the population the duration of each cycle under the system in which they were educated.[3] Considering all the individuals between 25 and 64 years old, the average number of years rises from 4.4 for the oldest generation to 11.3 for the youngest. At the same time, the dispersion of this variable has clearly diminished.

In our case, the explanatory variable of interest is market access. Given the peripheral location of Spain in Europe, we define an accessibility measure that accounts for remoteness from the Spanish and European centres of economic activity. Thus, the measure of market access computes ease of access to both Spanish and European markets for each province. Additionally, we consider the possibility that remoteness from European economic activity has a different effect than remoteness from Spanish economic activity, so we split the total market potential into two measures. The first computes ease of access to Spanish markets while the second computes ease of access to European markets.

The human capital equation specified in this study requires the measure of access to both the Spanish and the European markets to be computed over a long time period, from 1956 to 2001. In order to construct a homogeneous variable, we have had to rely on a simpler measure. Ideally, we should use a theory-consistent measure of market access such as the one constructed by Redding and Schott (2003) based on a trade equation. The first simplification in this paper derives from the lack of data on trade flows at a regional level for Spain and the EU over the entire sample period to estimate

[3] Similar results are obtained estimating a probit model, where the dependent variable takes the value 1 if the individual has a university degree.

a trade equation. Instead of using the theory-based measure, we therefore rely on the concept of market potential as defined by Harris (1954). Secondly, because of the lack of data on regional GDP for the entire sample, economic mass in a given region is proxied by its population. As a result, access to Spanish markets for province *i* is computed as the sum of the population over all 47 mainland provinces (NUTS 3 regions) weighted by the inverse of the Euclidean distance between the capital cities of the origin and destination provinces. The measure includes the own area, and the internal distance of each province is approximated through the radius of a circle with an area equal to that of the province The equivalent measure for access to European markets is built up in a homogeneous way for the NUTS 2 regions of the nine nearest European countries accessible by road from Spain. The total number of regions is 93.[4] The market access formula is:

$$MAP_h = \sum_{j=1}^{47} \frac{Pop_j}{d_{hj}} + \sum_{k=1}^{93} \frac{Pop_k}{d_{hk}} \qquad (5)$$

where *h* is the province of origin, *j* the Spanish destination province and *k* the European destination region.

Although market potential can be criticized for being an ad hoc measure, several academic papers support its use. Using the theory-based measure and the market

[4] The countries considered are: Austria, Belgium, Germany, Denmark, France, Italy, Luxemburg, the Netherlands and Portugal. These countries were selected according to three criteria: the distance from Spain, the level of trade between the two countries and the availability of information. Hence, the most remote countries with a low level of commercial trade with Spain were excluded. In 2001 these countries accounted for 55.1% of total Spanish exports. We did some sensitivity tests with respect the regions included, and the results showed no significant variation. For instance, we computed the market potential for Germany, France, Italy and Portugal, which account for 51% of total Spanish exports. It should also be mentioned that the market potential measure for Europe is based on NUTS 2 regions because of the lack of information for all the selected countries and years at NUTS 3 level.

potential variable as an alternative, Head and Mayer (2004) obtain better results with the latter. Breinlich (2006) confirms that substituting market access by Harris market potential computed according to distance or travel time as the explanatory variable in a wage equation yields very similar results.

According to the underlying model, years of schooling also depends on the costs of education. In this study, educational costs are proxied by a measure of supply, computed as the number of non-compulsory secondary schools per 1000 inhabitants at secondary school age. It should be noted that the definition of non-compulsory secondary schools has changed with the reforms of the educational system in such a way that makes comparability over time difficult. Previous to the 1970 reform, secondary education lasted for seven years, from ten to sixteen years of age. With the 1970 reform, however, lower secondary education became compulsory and upper secondary education started at 14 years of age and finished at 17. The last reform in 1990 extended compulsory education until the age of sixteen, so secondary post-compulsory education goes from ages 16 to 17. Nonetheless, the data show that the dispersion across provinces of the number of schools per capita of school-age population has significantly decreased over time.[5]

[5] Alternatively, we proxied the cost of education by the distance to the nearest university to account for the great territorial expansion of the Spanish university system. However, the results were poorer in terms of statistical significance.

4. A descriptive analysis of the provincial differentials in human capital

The increase of the stock of human capital of the Spanish economy has been remarkable. According to de la Fuente and Domenech (2012), the average years of schooling of the adult population -over the age of 25- in 1960 was 4.7 years. Fifty years later, the average years of schooling had increased to 9.4. Similarly, the share of the adult population holding a university degree increased from 2.9 in 1960 to 19.5 in 2010. Nonetheless, according to the same authors, the average years of schooling of the OECD countries in 1960 was 8.1 years and 11.8 in 2010. Even though the gap has decreased both in relative and absolute terms, it is still high.

In this paper, we look at the interprovincial differentials in human capital over time. In order to provide insight into how these differences have evolved, Table 4 shows the main statistics for the number of years of education computed for individuals born in different cohorts.[6] According to the micro-data from the LFS, the absolute difference in schooling between the province with the highest average number of years of education and the province recording the lowest figure is around three years and has persisted over time. However, given the increase in the level of education in all provinces, the relative difference has decreased from 90% for the first cohort to 30% for the most recent one. The coefficient of variation of the provincial distribution of years of schooling has diminished from 0.17 to 0.07 between the first and the last cohort. It seems, however, that during recent years this process has slowed down.

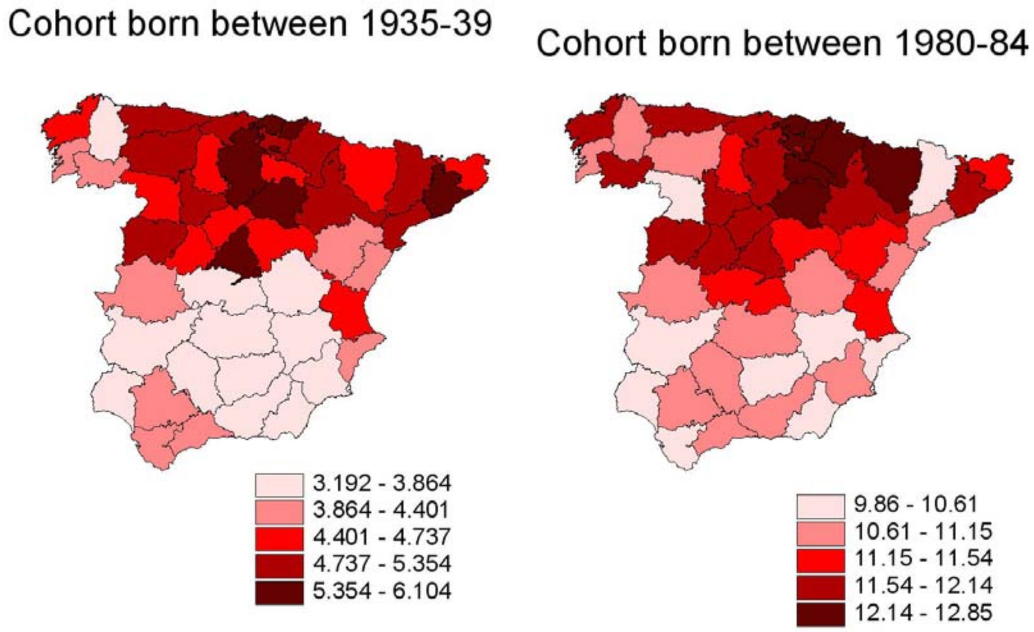
[6] Ideally, we would like to present the educational attainment for the youth cohorts of population observed over time. Unfortunately, these data are not available. Instead, we provide the average level of education of the adult population born in different years as an approximation. The data comes from the LFS of 1999, 2004 and 2009, as described in section 3 of this paper.

Table 4. Average years of schooling per province by cohort

	1956	1961	1966	1971	1976	1981	1986	1991	1996	2001
Mean	4.56	5.53	6.32	7.28	8.28	9.73	10.18	10.61	11.21	11.36
Maximum	6.10	7.60	8.53	9.15	10.16	11.37	11.69	12.05	12.86	12.85
Minimum	3.19	4.20	5.08	5.88	6.68	8.18	8.55	8.84	9.66	9.86
Std. Dev.	0.78	0.95	0.97	0.93	0.92	0.92	0.89	0.83	0.78	0.76
Coef. Var.	0.170	0.172	0.154	0.128	0.111	0.095	0.088	0.078	0.070	0.067
Observations	47	47	47	47	47	47	47	47	47	47

Map 1 depicts the spatial distribution of years of schooling for those born in the first and last cohort, respectively. The first map shows that for those born in the 1930s the spatial inequalities in human capital endowments follow a north-south divide. The more developed and populated northern provinces are those showing higher levels of schooling than the provinces located in the southern part of Spain, where the less developed areas can be found. Almost fifty years later, although the north-south divide is still there, the pattern of human capital distribution is more homogeneous. The data clearly show that a process of spatial convergence in human capital endowment has taken place in recent decades. This has been the result of an extension of the public education system to cover the whole country

Map 1. Spatial distribution of years of schooling



5. Estimated equation and results

According to the modelling framework detailed in section 2, for each cohort in the sample we estimate the following equation:

$$\ln(E_{ih})_c = (e_h)_c + \beta_{1c} \cdot \ln(MAP_h)_c + \beta_{2c} \cdot (SCHOOL_h)_c + \beta_{3c} \cdot (GENDER_i)_c + (u_{ih})_c \quad (5)$$

where:

c is the cohort ranging from 1956 to 2001

E_{ih} is the number of years of education attained by individual i born in province h

MAP_h is the market potential for province h

$SCHOOL_h$ is the number of secondary schools per capita of school-age population in province h

$GENDER_i$ is the gender of individual i

e_h are random variables for the province of birth

u_{ih} is the traditional random disturbance term

Hence, the level of education attained by individuals depends on their cohort of birth and gender, the number of secondary schools per school-age population as a proxy for educational costs, and access to markets, both defined for the individual's province of birth. Observing several individuals in the same province makes it possible to estimate the education equation using panel methodology and thus control for unobserved heterogeneity. Given the characteristics of our data, for each cohort and province, the market potential and the school index are constant, so in a fixed effect model the coefficients for the individual-invariant variables cannot be identified. Instead, we estimate a random effect model. However, in this study, estimating a random effect model may be interpreted as a feasible within estimation. Let us recall that the estimation of the random effect model is equivalent to applying OLS to a transformed equation in which both the dependent variable and the explanatory variables X_{ih} are subject to the following transformation:

$$Y_{ih}^{*} = Y_{ih} - \lambda \overline{Y}_{h}$$

$$X_{ih}^{*} = X_{ih} - \lambda \overline{X}_{h}$$

$$\lambda = 1 - \frac{\sigma_u}{(N\sigma_e^2 + \sigma_u^2)^{1/2}}$$

where N is the number provinces (47 in our case) and σ_u and σ_e are respectively the standard deviation of the error terms and the random effects. If, as in our case, σ_u is very small in relation to σ_e, λ will be near 1 and the transformation almost corresponds to taking differences in relation to the mean, that is to say, the within estimator.[7] Therefore, using panel data methodology allows us to take into account the advantages of this kind of information. As Hsiao (2007) indicates, panel data enables us to get a more accurate inference of model parameters through the increase in the degrees of freedom, and it also enables us to take individual effects into account.

Additionally, we correct the possible correlation of errors between individuals in the same province by computing cluster standard errors.

The results of the estimated regression equations are presented in Table 5 for each of the 10 cohorts. The estimation by cohorts reveals that the explanatory variables have a different effect at different points in time. Starting with the control variables, it is interesting to note that being male has a positive though decreasing effect on education until the beginning of the 1980s; from that point on, the sign of the coefficient reverses and being female increases the number of years of education for the youngest generations. Secondly, as expected, the ratio of secondary schools per capita of school-age population has a positive and significant effect on educational level for all the cohorts. That is, decreasing the cost of access to education raises its level. As explained in section 3, the ratio of access to secondary education is not homogeneous over time. So in order to compare the coefficients we have computed the elasticities which show that the effect of access to school increased until the third cohort, corresponding to those born in the second half of the 1940s, and from that point on the effect dropped slightly

[7] For example, for the first equation corresponding to the 1956 cohort, the estimated λ coefficient, according Table 5, is 0.0522. Thus, the variables are transformed as $Y^*_{ih} = Y_{ih} - 0.948 \cdot \overline{Y}_h$ and $X^*_{ih} = X_{ih} - 0.948 \cdot \overline{X}_h$, rather similar to the within transformation.

Table 5. Estimation results for the regression equations by cohorts (I), (t-statistics in parenthesis)

Dependent variable=ln(years of study)

	1956	1961	1966	1971	1976	1981	1986	1991	1996	2001
Gender	0.1392	0.1766	0.1528	0.1060	0.0615	0.0048	-0.0314	-0.0757	-0.0982	-0.1169
	(12.21)	(15.36)	(16.08)	(13.43)	(7.43)	(0.63)	(-5.10)	(-12.94)	(-22.84)	(-13.79)
Education access	0.2555	0.3319	0.3757	0.2217	0.2933	0.1487	0.1062	0.1007	0.0303	0.0407
	(2.94)	(2.94)	(4.35)	(3.45)	(4.08)	(4.13)	(3.35)	(3.91)	(1.93)	(4.06)
Ln(Market Access)	1.0408	0.8204	0.6700	0.5865	0.5013	0.4094	0.3673	0.2964	0.2654	0.2162
	(5.84)	(6.49)	(6.49)	(5.99)	(6.65)	(6.60)	(5.87)	(5.35)	(4.40)	(4.45)
Constant term	-4.6575	-3.2480	-2.3709	-1.7277	-1.0785	-0.3112	0.0278	0.4907	0.8073	1.0465
	(-4.76)	(-4.8)	(-4.3)	(-3.33)	(-2.64)	(-0.92)	(0.08)	(1.55)	(2.36)	(3.78)
Observations	9378	17585	29985	29924	32940	33966	32054	29216	17443	6963
rho	0.0375	0.0320	0.0208	0.0195	0.0151	0.0152	0.0227	0.0230	0.0290	0.0191
sigma_u (within)	0.1033	0.1006	0.0832	0.0799	0.0667	0.0557	0.0636	0.0608	0.0654	0.0511
sigma_e (between)	0.5228	0.5531	0.5705	0.5668	0.5382	0.4493	0.4172	0.3965	0.3779	0.3661
R2-between	0.6689	0.6996	0.7646	0.7197	0.7502	0.7170	0.6173	0.5546	0.3422	0.4061

Note: rho is the share of the estimated variance of the overall error accounted for by the within components. Standard errors are computed using the option of clusters by provinces until the last cohort. The low value for the reference year of 1996 should be taken with caution given that its level of statistical significant is low, p-value=0.053.

Table 6. Elasticity of years of schooling with respect to access to secondary schools

1956	1961	1966	1971	1976	1981	1986	1991	1996	2001
0.09	0.10	0.18	0.16	0.15	0.13	0.10	0.12	0.05	0.12

Going to the variable of interest of this study, Table 5 shows that accessibility to markets has a positive and significant effect on the level of education, although it decreases steeply over time. For individuals born in the late 1920s, the elasticity of the number of years of education with respect to access to markets is around 1; from that point on, the coefficient continuously diminishes until reaching 0.22 for the youngest generation. According to these results, market access would have played a significant role as a determinant of the regional differences in human capital accumulation in Spain. However, the magnitude of this role has diminished over time, reaching rather low values for the most recent years of the sample. These results are in agreement with the theoretical reasoning outlined in section 2. Let us recall that the estimated coefficient for the access to market variable captures the importance of barriers to mobility for the labour force. The decreasing temporal trend estimated for this variable might reveal a reduction in barriers to labour mobility as a consequence of the process of economic integration and globalisation of the Spanish economy. It can be safely assumed that linguistic, cultural and information barriers, among others, have dropped for the younger cohorts and, consequently, this has tended to equalise the incentive for education for the populations of the different Spanish provinces.

Additionally, we re-estimated the equations distinguishing between access to Spanish and European markets. As can be observed in Table 7, the Spanish market potential,

Table 7. Estimation results for the regression equations by cohort (II), (t-statistics in parenthesis)

Dependent variable=ln(years of study)

	1956	1961	1966	1971	1976	1981	1986	1991	1996	2001
Gender	0.1392	0.1765	0.1528	0.1061	0.0615	0.0049	-0.0314	-0.0758	-0.0982	-0.1170
	(12.23)	(15.37)	(16.09)	(13.43)	(7.43)	(0.63)	(-5.10)	(-12.94)	(-22.84)	(-13.81)
Education access	0.2705	0.3928	0.3915	0.2164	0.2490	0.1254	0.0867	0.0881	0.0210	0.0353
	(3.51)	(2.99)	(4.36)	(3.20)	(3.18)	(3.47)	(2.89)	(3.68)	(1.57)	(3.50)
Ln(Market access Spain)	0.1559	0.0569	0.0728	0.0461	0.1316	0.0800	0.0604	0.0331	0.0138	0.0241
	(1.51)	(0.47)	(0.86)	(0.55)	(1.86)	(1.22)	(0.97)	(0.53)	(0.24)	(0.71)
Ln (Market access Europe)	0.7738	0.5855	0.5084	0.4967	0.3902	0.3417	0.3123	0.2598	0.2414	0.1895
	(6.28)	(8.07)	(8.69)	(6.91)	(5.68)	(5.90)	(5.58)	(5.17)	(4.74)	(3.63)
Constant term	-3.5112	-1.9077	-1.5397	-1.1650	-0.8415	-0.0987	0.2464	0.7024	1.0320	1.2073
	(-4.2)	(-2.79)	(-3.14)	(-2.69)	(-2.43)	(-0.33)	(0.79)	(2.36)	(3.24)	(5.21)
Observations	9378	17585	29985	29924	32940	33966	32054	29216	17443	6963
rho	0.0330	0.0280	0.0180	0.0155	0.0144	0.0142	0.0207	0.0205	0.0267	0.0179
sigma_u (within)	0.0966	0.0939	0.0772	0.0710	0.0650	0.0540	0.0607	0.0574	0.0626	0.0494
sigma_e (between)	0.5228	0.5531	0.5705	0.5668	0.5382	0.4493	0.4172	0.3965	0.3779	0.3661
R2-between	0.7121	0.7434	0.8014	0.7788	0.7690	0.7537	0.6611	0.6084	0.4068	0.4422

Note: rho is the share of the estimated variance of the overall error accounted for by the within components. Standard errors are computed using the option of clusters by provinces though positive, is never statistically significant, whereas the European market potential is always significant but with a clear downward trend with the successive cohorts. These results are again in line with what could be expected in terms of barriers to labour mobility. Given that the barriers to labour mobility within a country are relatively low, at least for highly educated population, we would expect a scant impact of remoteness from Spanish economic activity centres, whereas higher mobility barriers between countries would help to explain differences in the individual incentive to education. The estimated coefficients for access to market behave as expected.

A potential shortcoming of our econometric specification is that we do not control for some individual determinants of investment in education, namely parents' level of education. Unfortunately, the LFS does not provide information on parental education. Nonetheless, we have done some robustness tests using the Spanish sample of EU-SILC (European Union –Statistics of Income and Living Conditions) as an alternative. This survey allocates individuals at NUTS 2 level which in Spain corresponds to Comunidades Autonomas, a level of territorial disaggregation that is too high to carry out spatial analysis. However, this sample provides information on parental education. Therefore, we have re-estimated the education equation with EU-SILC data including the parents' level of education as regressors. Given the number of observations and the low level of variability in the data, there are only 15 NUTS-2 regions, we have not estimated a different equation for each cohort. Instead, we have estimated a single equation for all the sample and we have included the cohort as a control variable. As Table A.1. in the Annex shows, market access has a positive effect on the level of education even when parental education is controlled for, although the estimated coefficient is somewhat lower.

6. Conclusions

The objective of this paper is to investigate whether remoteness from markets has any impact on individual incentives for education and how this impact has changed over time. Making use of micro-data, we estimate a set of equations that model individuals' educational decisions at different points in time as a function of individual characteristics, ease of access to education and access to markets. The explanatory variables are measured according to the place and time the decision was actually made. The first result to be noted is that increasing ease of access to education always has a positive effect on the number of years of schooling. With respect to the variable of interest in this paper, our study provides evidence that remoteness from the markets helps to explain the spatial differences observed in individuals' education decisions in Spain. Nonetheless, the magnitude of the impact has significantly decreased over time. This result is in accordance with our initial hypothesis that as barriers to mobility decrease and migration becomes easier, the importance of accessibility to markets is reduced. The underlying hypothesis is that as barriers to mobility decrease, labour mobility of the educated population increases and individuals decisions on education are no longer based on the wage premium paid in their own province but on the maximum wage premium attainable in the other economic spaces, corrected by the cost of mobility.

This hypothesis is also supported by the fact that when market potential is split between access to Spanish and European markets, the results clearly show that the distance to European economic mass is what really hinders the accumulation of human capital. That is, labour mobility across provinces within a country has always been possible, while mobility across regions of different countries means higher costs. Furthermore, the continuous drop in the coefficient for the successive cohorts that affects the European market potential is in accordance with an increase in the labour mobility between countries for the educated population. In any case, the fact that the European market potential for the last cohort is still significant confirms the lack of perfect labour mobility.

Given that human capital endowment is one of the main factors influencing the economic development of a region, the results of this paper provide effective guidelines for reducing the negative effect of remoteness on the level of education of individuals by reducing the costs of access to the education system and encouraging labour mobility across regions. Improving transport facilities is also a way of reducing the negative effect of remoteness on the skill premium and on incentives to achieving higher levels of education.

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Table A.1 Estimation results for the regression equations using EU-SILC
(t-statistics in parenthesis)

Dependent variable=ln(years of study)		
Gender	0.0245	0.0191
	(3.38)	(2.50)
Cohort		
1966	0.0426	0.0395
	(2.00)	(1.53)
1971	0.0756	0.0728
	(3.49)	(2.57)
1976	0.2133	0.2219
	(12.89)	(11.05)
1981	0.1949	0.1999
	(9.93)	(7.28)
1986	0.2297	0.2482
	(10.05)	(7.86)
1991	0.2414	0.2853
	(7.63)	(6.72)
1996	0.1415	0.2163
	(3.53)	(3.48)
Ln (father's schooling)	0.3660	-
	(18.89)	-
Ln (mother's schooling)	0.1995	-
	(10.50)	-
Education access	0.1729	0.1924
	(5.18)	(3.97)
Ln (market access)	0.1748	0.2506
	(2.13)	(2.53)
Constant term	-0.2239	0.1686
	(-0.50)	(0.30)
Observations	15908	15908

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