The Effect of Education on Financial Market Participation: Evidence from Chile

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Abstract

The low level of participation in financial markets by households is one of the empirical regularities that challenge some of the standard models in financial economics. Education is one of the variables mentioned in the literature as a potential factor explaining this limited participation. In this paper we estimate the causal effect of education on financial market participation using Chilean household surveys and exploiting the 1981 educational reform in Chile as a source of variation for educational levels. We estimate that graduating from high school increases the probability of financial market participation by 3 percentage points. The size of this effect is economically important in the context of the Chilean financial market where participation in 2009 only reached 9.9%. Also, this effect represents almost half of the difference in participation between individuals who completed high school and those who dropped out. Exploring some of the potential mechanisms, we find that education has a larger impact on the probability of holding simple financial assets, suggesting that formal education brings households closer to financial markets either by increasing the households' need for these assets or by turning the households into more attractive clients for the financial markets. Further, we show that education has a positive effect on individuals' willingness to take risk and their financial literacy.

1 Introduction

Empirical evidence shows a low participation of households in the financial market. In the case of the United States, according to the Survey of Consumer Finances from 2001, only 48.6% of households held stocks, either directly or indirectly (Bucks et al., 2006).[1] In Chile the financial participation rates are even lower; according to the results of the Household Financial Survey (Encuesta Financiera de Hogares - EFH) from 2007, only 14.4% held some type of financial asset.[2] This evidence represents a difficult puzzle to explain for conventional portfolio theories, since according to them, an expected utility maximizer will always hold some portion of each of the assets available in the economy. That is, no matter how risk averse than the agents, everyone should keep some percentage of their wealth in financial assets.[3]

In the literature, it has been proposed that the low level of education held by individuals can be part of the explanation for this puzzle. In this context, the present research, motivated by the work of Cole et al. (2012), seeks to estimate the causal effect of education on participation in the financial market by Chilean households, using the exogenous variation generated by the educational reform of 1981 as a means of identification, in addition to investigating the possible mechanisms through which education affects participation. The study of financial market participation is important for several reasons. On the household side, participation in this market provides them with asset accumulation and consumption smoothing, which in turn has a positive impact on their welfare. With respect to the financial system as a whole, the degree of financial participation affects asset prices, the equity premium and market volatility. In addition, participation may affect regulatory decisions since the holding of financial assets affects household tax planning.[4]

Estimating the effect of education on financial participation is a challenge, because a simple OLS estimate would present endogeneity problems due to the presence of unobservable variables, such as the ability or family background of the individuals. To overcome this difficulty, an estimation strategy is developed for instrumental variables, exploiting the heterogeneous impact of the 1981 educational reform on schooling. This reform transformed the Chilean educational system, transferring the administration of public educational institutions to municipalities and encouraging the provision of publicly funded educational services by the private sector through a voucher system. As a result of this reform, private provision of educational services, and especially subsidized establishments, expanded significantly. Furthermore, the effect of this reform was reflected in an increase in the graduation rate and increased average education of Chileans. Additionally, there were strong differences across the different regions of the country in the reform's impact on composition of enrollment. The idea of the estimation strategy is that the variation of the reform's impact on regions and cohorts enables identification of the causal effect of education on financial participation. Given the characteristics of the sample used and the selected econometric specification, a biprobit model

[1] Indirect ownership includes participation in mutual funds, pension accounts, etc.

[2] This broad definition of financial asset holdings includes: savings accounts, voluntary pension savings, time deposits, stocks, mutual funds, holdings in companies, among others.

[3] For an analytic proof of this statement, see Haliassos and Bertaut (1995) section II.A.

[4] For example, in Chile voluntary pension savings is subject to a special tax regulation.

is used for the IV estimation as recommended Chiburis et al. (2011) by Chiburis et al. (2012).[5]

Studying the impact of the 1981 education reform on the schooling of individuals, we find that this is mainly reflected in a significant increase in the probability of finishing secondary education (meaning 12 years of schooling). Taking into account this result and using a biprobit estimation strategy, we find that finishing secondary education increases the probability of participating in the financial market by 3%, which is a significant change if we consider that financial participation in Chile is only 9.9% in the sample. Examining mechanisms, we find that secondary education has a greater effect on the participation in low-complexity assets, suggesting that an important transmission channel would be the approach to the financial market generated by formal education. Furthermore, the results show that education increases an individual's willingness to take risks and increases the likelihood of being employed in the formal sector of the economy, suggesting that both mechanisms increase financial participation. Finally, there is an examination of the effect of education on financial literacy, a term that refers to an individual's ability to process financial information, and a positive, large-scale effect is found. This paper joins the growing literature that has made great theoretical and empirical efforts to achieve understanding of the puzzle of low household participation in financial markets. The theoretical literature has focused on rationalizing this behavior, characterized by low participation in the stock market (Mankiw and Zeldes, 1991) and low rates of savings in financial instruments, through the introduction of fixed costs and alternatives to the conventional portfolio theory. In the empirical literature, studies have aimed at the identification of various factors that help explain this puzzle, but while some of these studies have identified causal effects using instrumental variables, most could only establish correlations. Early research includes that of Haliassos and Bertaut (1995), who explore possible explanations to the puzzle of participation, such as liquidity constraints, heterogeneous beliefs, risk aversion and alternatives to expected utility theory. The results of this study seem to support deviations from the expected utility maximization as potential explanations for the puzzle. In this line of research, Vissing-Jorgensen (2002) suggests that the existence of fixed costs for entering the stock market is the main reason for the low participation in this market. Using data for the U.S. from the Panel Study of Income Dynamics, the author concludes that a cost per annum of $50 USD is sufficient to account for half of the non-participants' decision, suggesting that this type of cost is a simple explanation for the decision of many households to stay out of this market. In a similar vein, Guiso and Jappelli (2005) raise the possibility that households are not aware of the existence of certain types of financial assets, which would help explain the low rates of participation in this market. Specifically, the authors find that in the late '90s, 35% of Italian households did not know of the existence of financial assets such as stocks, mutual funds and investment accounts. [6]

Within the empirical literature, one of the first studies that seeks to establish causal effects is

[5] In our sample the unconditional probability of holding financial assets is below 10% (9.9%). In section 5 we present the 2SLS results for the baseline specification, but they should be interpreted with caution because they are very unstable and sensitive to the exact set of excluded instrument used.

[6] Naudon et al. (2004) formally introduce ignorance as a determinant of low participation in financial asset markets. The hypothesis of this study is that many people are not familiar with financial instruments, so that the optimal response to their own ignorance is to avoid these markets.

that of Cole et al. (2012), who, using exogenous variation in education caused by changes in the minimum compulsory schooling laws in the United States, show that education has a causal effect on financial participation. Specifically, the authors show that one year of schooling increases the likelihood that an individual reports income from financial assets by 7 to 8%, with the other factors remaining constant. This study has the merit of being the first to seek to identify the causal effect of education on financial participation, but its econometric specification fails to correctly identify this effect, because even though they use education, they include other variables that keep the endogeneity problem present.[7]

An issue related to this investigation and which has been extensively studied in recent years is the effect of financial literacy on the various forms of financial participation. Financial literacy is the ability of individuals to understand basic financial concepts, so we thought it might be an important transmission channel of education towards financial participation. Within the literature, van Rooij et al. (2011) found that people with low levels of financial literacy are less likely to invest in stocks, controlling for other factors. Cole and Shastry (2010), using variations in state reforms on financial education requirements in schools to identify the effect of financial literacy on asset accumulation, find that schools' financial education requirements do not affect an individual's propensity to save. Other related studies have found that individuals with lower financial literacy tend not to plan for their retirements, and they borrow at higher rates and keep a lower percentage of their wealth in financial assets (Lusardi and Mitchell (2007) and Lusardi and Tufano (2009), among others).

In the case of Chile, research has focused on identifying the effect of financial literacy on the financial behavior of individuals, which underlines this study's contribution. Landerretche and Martínez (2013), using an IV estimation strategy, obtained results suggesting that a higher level of financial literacy increases the probability of having financial savings, but outside of the pension system. Behrman et al. (2010), also using instrumental variables, find that higher financial literacy is associated with higher wealth accumulation among Chilean households.

With respect to the identification strategy, this study, in addition to the study done by Cole et al. (2012), is related to the literature that exploits changes in state compulsory education laws in the United States to estimate the externalities of education (Acemoglu and Angrist, 2001) and the effect of education on crime (Lochner and Moretti, 2004). Similar strategies were developed for Latin America by Patrinos and Sakellariou (2005), who estimate the return to education in Venezuela using the change in the compulsory education law of 1980 as an instrument, and by Patrinos (2008), who, along the same lines, uses a binary instrument based on the school education reform of 1981 in Chile. Additionally, Rau (2013) estimates the returns to education for Chile through an IV strategy, exploiting exogenous variation between regions and cohorts in the education infrastructure and in compulsory education laws passed between 1929 and 1931.[8]

The rest of the paper is organized as follows. In the next section we review the on the potential

[7] Previous empirical studies have shown that financial participation is correlated with income (Campbell, 2006), education (Bertaut and Starr-McCluer, 2002), social connections (Hong et al., 2004), trust (Guiso et al., 2008), and experience with the stock market (Malmendier and Nagel, 2011).

[8] There is extensive literature that uses exogenous policy changes as instruments for education. The most important studies in this area include Card (2001), Duflo (2001) and Oreopoulos (2006).

relation between financial market participation and education, and also the context of the education reform of 1981 in Chile. In section 3 we discuss the data used in this paper, the empirical methodology and the identification strategy. In section 4 we present the main results of the paper. We then discuss the potential channels and their empirical relevance in our sample in section 5. In section 6 we present the conclusions of our work.

2 The Context: Financial Participation and Educational Reform in Chile

Previous research suggests several mechanisms through which education may affect participation in the financial market. First, education can generate an increase in the income of individuals who save a larger fraction of their wealth in financial instruments. Furthermore, education can expand employment opportunities to sectors of the economy that are closer to the financial system. For example, the opportunity to apply for formal employment introduces individuals to the Pension Fund Administrator system (Administradoras de Fondos de Pensión - AFP), which forces them to participate, indirectly, in the financial system, and to face problems of a financial nature. Similarly, a university degree can lead a person to a job in a large company, facilitating their participation in the financial market.[9] Second, education may increase financial literacy, giving individuals a greater understanding of basic financial concepts, such as interest rates, dividends, returns, etc., which play an important role in the decision to hold financial assets. Third, education may affect people's preferences, through an incease in patience (see Becker and Mulligan, 1997) or a change in the willingness to take risks. Harrison et al. (2002) found that discount rates are negatively correlated with education and results from Halek and Eisenhauer (2001) suggest a negative correlation between risk aversion and education. These changes in beliefs would impact people's financial participation, since increased patience leads to a higher level of savings and a lower level of risk aversion leads households to invest a larger fraction of their wealth in financial assets. Fourth, individuals with higher levels of education may have access to a wider supply of financial products, which facilitates their participation in this market.

These channels suggest that an increase in schooling should increase individuals' financial market participation. In this study, the main objective is to estimate a reduced form equation between financial participation and education, conditional on other characteristics of the individuals

$$y_i = \alpha + \beta educ_i + \gamma X_i + \epsilon_i \tag{1}$$

where y_i is a dummy variable indicating if the individual does participate in the financial market, $educ_i$ is the individuals' level of education and X_i represents a set of additional controls. The parameter β in equation (1) captures the net effect of education on financial participation. Taking into account the aforementioned channels, a positive effect of education on financial market participation

[9] Hong et al. (2004) studies peer effects in this context.

is expected ($\beta > 0$). In addition, it aims to empirically identify some of the previously proposed channels.

2.1 Financial Participation in Chile

In Chile the pension system functions as a compulsory individual capitalization system where individuals save monthly to fund their pensions upon retirement.[10] These funds are managed by the AFPs, who invest in the financial system, so you could say that all payroll workers in Chile indirectly participate in the financial system. But this compulsory participation in the financial system is not sufficient to explain the low voluntary participation observed in Chile, for several reasons. First, conceptually it is problematic to liken participation in pension funds to the direct holding of financial assets, since the nature of the holding assets in pension funds is very different from direct ownership.[11] This is because the funds in the AFP are contingent assets for a particular stage of life that are not payable at any other point in time, and therefore they have radically different payment and liquidity characteristics from direct ownership. Second, AFP pension savings offer a choice of only five investment funds. This restriction may be an obstacle to contributors' portfolio diversification if, for example, the return on AFP investments is highly correlated with the human capital of the individuals. In these cases, the way that affiliates can optimize their portfolio is through direct participation in the financial market.

According to the 2009 CASEN survey (National Socio-Economic Characterization Survey), only 9.9% of people between 24 and 70 years old held some type of financial asset. This broad definition includes housing saving plans, savings in housing fund administrators (Administradoras de Fondos para la Vivienda - AFV), voluntary retirement savings, savings in Cuenta 2 of AFPs (this is one of the first voluntary savings products in Chile's pension system), bank savings account, fixed-term deposits, investments in mutual funds, shares and bonds. Table 1 presents the financial participation patterns obtained from CASEN 2009.[12] We can note that financial participation increases with an individual's level of education, with the university stage generating the greatest difference in terms of participation. In addition, a higher percentage of individuals hold simple assets versus complex assets.[13] Regarding complex assets, we see that the participation of individuals with low education (those with basic education) is practically zero, while for individuals with a college education, participation reaches 4.4%. In the case of simple assets, the degree of participation by sectors with a low level of education is 6.68% and reaches 12.53% among college-educated people.

[10] Affiliation with the AFP system is mandatory for all payroll workers. The funds accumulated by AFP affiliates can be withdrawn only after retirement.

[11] Here, by direct possession we mean all the aforementioned forms of holdings that make up 14.4% in the EFH.

[12] Descriptive statistics in appendix A provide a comparison between the financial participation patterns observed in the CASEN, the EFH and EPS.

[13] We consider as simple assets housing savings plans, savings in housing fund administrators and bank savings accounts as simple assets. Complex assets include voluntary savings plans, savings in Cuenta 2 of AFPs, fixed-term deposits, mutual fund investments, shares and bonds

Table 1: Financial Market Participation Statistics (in %)

Educational Attainment	Participation Simple	Participation Complex	Participation Financial
Less than high school	6.68	0.54	7.22
High School	9.49	1.22	10.71
Tertiary Education	12.53	4.40	16.94
Full sample	**8.60**	**1.35**	**9.95**

Source: Encuesta CASEN 2009

2.2 Chile's Educational System and its Reform in 1981

Until 1980, the administration of the education system was completely centralized in the Ministry of Education. This institution was responsible for establishing the plans and programs for the entire education system, in addition to directly administering fiscal establishments representing about 80% of establishments in the country. This work included the appointment of teachers and school administrators, the allocation and payment of expenses and compensation, etc.

As from 1980, the administration of public educational institutions was transferred to municipalities and incentives were given for the provision of publicly funded educational services by the private sector. This led to three types of schools: municipal, subsidized private and paid private.[14] As part of the reform, a per-student subsidy system was implemented via a voucher scheme, where the government subsidized the schools chosen by guardians, directly based on enrollment numbers. Specifically, the Chilean government gives each establishment a certain amount of resources for each child effectively attending classes. The idea behind this scheme was that the voucher system and private provision of free education would promote competition among institutions to attract and retain students, creating a education market that, through competition, encouraged efficiency and quality in educational services.

As a result of this reform, the private provision of educational services, and especially of subsidized establishments, expanded significantly. In 1981 these schools accounted for 15.1% of enrollments, and in 1995 accounted for 32.8% of total enrollment. By 2005 this type of establishment had 42% of enrollment, while municipal schools fell from 73% in 1981 to 49% in 2005.[15] In terms of the schools' geographical distribution, there were strong differences in the impact of the reform in different regions. In Table 2 we see that in 1996 the subsidized private system was particularly important in the metropolitan region, accounting for 45% of school enrollment, while in other regions more than 70% of enrollment was held by municipal schools.

[14]Municipal schools are funded by per-student subsidy and managed by municipalities. Subsidized private schools are funded by per-student subsidy and managed by private parties. Paid private schools do not receive subsidies paid are paid by parents and managed by private parties.

[15]See Mizala and Romaguera (1998) and Gallego and Hernando (2009).

Table 2: Enrollment by Region in 1996 (in %)

Region	Municipal	Private Subsidized	Private Paid
I	63.7	24.6	11.2
II	70.2	19.2	10.6
III	90.7	9.3	0
IV	69.3	25.2	5.3
V	54.2	33.3	11.2
VI	69.8	20.2	7.7
VII	74.1	20.0	4.6
VIII	65.4	18.7	7.6
IX	56.8	38.7	3.3
X	72.6	21.9	5.3
XI	70.1	28.9	0.0
XII	70.2	16.8	13.0
R.M.	40.6	44.8	12.7

Percentages do not add up to 100% because establishments belonging to corporations are not included

Source: Mizala and Romaguera (1998)

3 Data and Empirical Strategy

3.1 Data

The main database used for this study is the CASEN survey from 2009 (CASEN 2009 hereafter). This survey, conducted by the Ministry of Social Development, is representative at both the regional and national level. It has been conducted biannually or triannually from 1985 to 2011. The main purpose of the CASEN survey is to describe the socioeconomic conditions in Chile and assess the impact of social policies. Interviews are conducted at the household and individual level. The information collected for each household member includes a description of income, employment, housing, educational characteristics, health services, participation in social programs and socioeconomic characterization.

The sample used in the estimates in 2009 includes those between 24 and 70 years old, which is the most active part of the population in financial terms. This sample also includes people who was "affected" by the 1981 education reform and others who were not. In particular, individuals older than 46 years in 2009 were not (or were only minimally) affected by the reform because they were too old in the early 1980s. The younger cohorts were progressively more exposed to the reform, with people 34 years old or younger in 2009 being 100% affected by the reform, because they started receiving primary and secondary education after 1981.

We also created the financial participation variable with information from the CASEN 2009. We constructed a dummy variable that indicates whether the household has any type of financial asset.

Table 3: Financial Participation by Educational Level (in %)

Financial Participation	Years of Schooling: Less than 5	5 to 8	9 to 12	13 to 16	More than 16
Sí	5.73	8.04	10.72	14.52	21.35
No	94.27	91.96	89.28	85.48	78.65

Source: authors' calculations using CASEN 2009

Table 2 summarizes the relationship between financial participation and level of schooling. The financial assets included in our definition of financial participation are: housing saving plans, AFV holdings, voluntary retirement savings (APV), voluntary savings in Cuenta 2 of AFP, bank savings accounts, fixed-term deposits, investments in mutual funds, shares and bonds.

In addition to the CASEN survey, we use the 2006 wave of the *Social Protection Survey* (*Encuesta de Protección Social* - EPS). This survey, developed by the Centro de Microdatos of the Universidad de Chile, is comparable to the *Health and Retirement Study* in the United States, provides information on several outcomes related to our question. Of particular interest for us is the information on financial literacy. Given that the EPS sample size is significantly smaller we use the CASEN 2009 for our main specifications.

Finally, for descriptive purposes, we also use the 2007, 2008 and 2009 versions of the *Encuesta Financiera de Hogares* conducted by the Banco Central de Chile. This survey, similar to the Survey of Consumer Finances conducted by the Federal Reserve of the United States, aims to generate detailed information on the household financial balance and its main sources of income and expenses.[16]

3.2 Empirical Model

This research seeks to identify the effect of education on household participation in the financial system. The data presented in the previous sections and the existing literature strongly suggest that individuals with higher levels of education are more likely to participate in the financial system. However, there are two issues that need further discussion. First, the presence of unobservable factors, such as ability or family background, which affect participation and are in turn correlated with education, implies that OLS estimators fail to capture the causal effect of education on financial participation and instead are contaminated with other effects. Second, in our preferred specifications we have a binary dependent variable and a binary endogenous outcome, namely a dummy variable indicating whether the individual has twelve years of education or more. Because of the potential endogeneity we just discussed we need to use instrumental variables to identify the causal effect of education on financial market participation.

To deal with the identification problem and identify causal effects we use an instrumental variable estimation strategy, exploiting the variation across regions and cohorts in the intensity of exposure to

[16] A complete summary of the variables created and used in the research is found in Appendix A.

the 1981 educational reform. Since the penetration was different in different regions of the country, a person's exposure to the reform is a function of two variables: age and the region in which he or she lived at the time of the reform. Therefore our estimates combine both sources of variation (across regions and across cohorts) as an identification mechanism.

The second problem implies that we need to choose an estimation method that deals with this specific situation: binary outcome and binary endogenous regressor. Most of the literature has tried one of two alternatives for this situation: use 2SLS treating all equations as linear models, or exploit the bivariate probit model (biprobit) with an excluded instrument for the first stage. While 2SLS is simpler, easier to interpret and does not require distributional assumptions, biprobit is more efficient but requires additional assumptions (Angrist and Pischke, 2008).[17] As expected, there is no clear consensus in the literature, but given that financial market participation is below 10% on average for our sample we follow Chiburis et al. (2012) and estimate the model using biprobit.[18]

Finally, it is worth noting that an important difference between the two estimation methods has to do with the effect that their coefficients capture. 2SLS results are consistent estimates of the local average treatment effect (LATE), while the biprobit model provides an estimate of average treatment effect (ATE). This difference between LATE and ATE can explain much of the difference observed between the 2SLS and biprobit estimations.

3.2.1 Econometric Specification

Although biprobit contemplates a joint estimation, we can think of our empirical econometric specification in a similar way as a traditional IV estimation.[19] In this case we have that our outcome equation corresponds to

$$y_i = \alpha + \beta educ_i + \gamma X_i + \epsilon_i \tag{2}$$

where y_i is a dummy variable that indicates whether the household has any type of financial asset, $educ_i$ is a measure of an individual's level of education and X_i is a set of controls that includes a fourth-degree polynomial for age, and dummies for sex and regions. In our study the variable $educ_i$ will be a dummy variable indicating whether the individual has 12 years or more of education.

In a similar fashion the equivalent specification of our preferred first stage in a linear model would be an equation like

$$media_i = \delta + \psi exp_i + \lambda inter_i + \phi X_i + \epsilon_i \tag{3}$$

[17] Chiburis et al. (2012) simulations show that in cases where there are additional controls, the biprobit model performs better than 2SLS for all sample sizes. Additionally, their results also suggest that in cases where the probability of treatment is close to 0 or 1, the bivariate probit results are considerably better in terms of statistical significance. See also Altonji et al. (2005) for a discussion on this issue in the context of the effect of catholic schools.

[18] Our 2SLS estimates are extremely imprecise and change significantly when changing the instruments. Given the above mentioned evidence we further interpret as evidence that we should use biprobit instead; the results are available upon request from the authors.

[19] See Wooldridge (2010)

where $media_i$ corresponds to a dummy variable that takes the value 1 if the individual has 12 or more years of schooling (we will say that the individual has completed high school or educación media as it is called in Chile), exp_i is a variable that measures the degree of the individual's exposure to the reform and $inter_i$ represents a set of interactions between the degree of exposure to the reform and the region. X_i is the set of controls in the second stage that includes a fourth-degree polynomial of age, and dummies for sex and region. We will use different definitions

Unlike the specification used by Cole et al. (2012), this study's base specification does not control for an income polynomial, since introducing income as an additional control would create an endogeneity problem in the specification again. This is because income is also correlated with unobservable variables, such as the ability of individuals or their family background.[20]

3.3 IV Framework and First Stage

The identification strategy used in this research is based on the impact of the voucher education reform on schooling. This reform led to a progressive effect on the younger cohorts. Specifically, we have

- People born before 1963 were not affected by the reform since by 1981 they had completed secondary education;
- People born between 1963 and 1975 were partially affected by the reform, since it was introduced when they were already in primary or secondary education. For this group, exposure to the reform is an increasing function of their year of birth;
- People born after 1975 were fully exposed to the reform, since they started attending school when the reform was already in place.[21]

Since the introduction of a reform of this nature does not occur automatically, one expects to find heterogeneity in exposure to the reform even in post-1975 cohorts. Besides heterogeneity across cohorts, the reform had different levels of impact on different parts of the country. In this way, we exploit two sources of variation, across cohorts and regions, as an identification mechanism.

Our baseline specification considers a biprobit model with an equation for the variable $media_i$, that corresponds to a dummy variable that takes the value 1 if the individual has 12 or more years of schooling, like (3) using as excluded instruments exp_i and $inter_i$ and including as controls all the variables used in the equation for the outcome. The excluded instruments correspond to a variable that measures the degree of the individual's exposure to the reform and to a set of interactions between the degree of exposure to the reform and the region, respectively.[22]

[20] To control for income, we need an additional instrument, one that could not be generated in this study. Later in the paper we do present a specification that includes income as a control for comparability with international results.

[21] Here we use the legal ages for starting primary school (educaci'on b'asica in Chile) and for graduation from high school (educaci'on media) to set the cohorts.

[22] The country is divided into three zones: North (composed of regions I, II, III, IV, XV), Center (regions V, VI, VII, XIII) and South (regions VIII, IX, X, XI, XII, XIV).

The proposed specification is able to identify the effect of the reform on exposed individuals, allowing heterogeneous effects across cohorts and regions in its impact. A key identifying assumption is that the excluded instruments only capture the differential effect of education reform on schooling and do not represent a pre-existing trend. This is achieved thanks to the fourth-degree polynomial of age (included in the set of controls in the second stage) which controls for the trend in education, so that the excluded instruments only capture the differential effect caused by the reform on the schooling of individuals. In addition, interactions between regions and cohorts reduce the probability of confusing the effect of the reform with that of the 1982-1983 recession, unless regional effects have the same patterns and signs. In this way, we can be sure to capture the exogenous effect of the reform, respecting the exclusion condition.

3.3.1 First Stage

We present the results from three alternative sets of instruments for equation (3)

1. In the first specification, the exp_i variable corresponds to a dummy indicating if the individual attended school after 1981 (the individual was supposed to start first grade in 1982). Interactions are not included.

2. Same as the previous one but including interactions of exp_i and $region_i$.

3. In the third specification, the exp_i corresponds to a set of 6 variables indicating the individual's exposure to the reform. We construct six birth cohorts, thus allowing a progressive effect of the reform on the younger cohorts. Interactions are not included.

4. Same as the previous one but including interactions of exp_i and $region_i$.

We present the first stage estimates for the four specifications in table 4. It is important to mention that all the specifications control for a flexible trend in schooling using an age polynomial. We see that in the four estimates, the instruments are significant and display the expected signs. Specifically, in columns 1 and 2, the Post Reform variable estimator is significant and positive, indicating that the reform generated an increase in the probability of completing secondary education. In column 2, we notice that the impact of the reform was smaller in the south of the country, while the impact in the north and central areas was similar. In columns 3 and 4, given that the base was defined as those individuals with a greater degree of exposure to the reform, the Degree of Exposure variables are negative and significant. This indicates that the people with less exposure to the reform have a progressively lower probability of completing secondary education.[23]

[23] The individuals with a higher degree of exposure to the reform correspond to those younger than 29 years old in 2009.

Table 4: First Stage Estimates

Dependent variable: $media_i$, dummy for having 12 or more years of education.

Variables	1	2	3	4
Degree of exposure 1			-0.0814***	-0.0887***
			(0.0280)	(0.0299)
Degree of exposure 2			-0.0901***	-0.0560**
			(0.0239)	(0.0274)
Degree of exposure 3			-0.0815***	-0.0463**
			(0.0186)	(0.0231)
Degree of exposure 4			-0.0545***	-0.0182
			(0.0137)	(0.0203)
Degree of exposure 5			-0.0321***	-0.00372
			(0.00989)	(0.0191)
Exposure 1 * Center				-0.0350***
				(0.0125)
Exposure 1 * South				-0.0631***
				(0.0127)
Exposure 2 * Center				-0.0240*
				(0.0133)
Exposure 2 * South				-0.0802***
				(0.0135)
Exposure 3 * Center				-0.0288*
				(0.0152)
Exposure 3 * South				-0.0772***
				(0.0157)
Exposure 4 * Center				-0.0283*
				(0.0170)
Exposure 4 * South				-0.0575***
				(0.0178)
Exposure 5 * Center				0.00136
				(0.0122)
Exposure 5 * South				-0.0216*
				(0.0129)
Post	0.0154**	0.0245**		
	(0.00743)	(0.0107)		
Post * Center		-0.00419		
		(0.00916)		
Post * South		-0.0182*		
		(0.00957)		
N	135,360	135,360	135,360	135,360
Cragg-Donald	0.038	169.15	31.11	93.97

We only show the estimated coefficients for the excluded instruments. For the degrees of exposition specifications the base group corresponds to those 28 years old or younger in 2009.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

3.3.2 Exploring the Effect of the Education Reform on Educational Attainment

Before moving to the estimation of the model it is useful to explore the identification of the relation between the reform and education. An important first step is to identify how educational attainment was affected by this reform. Following Duflo (2001) we estimate a series of regressions like

$$esc_{ik} = \alpha + \gamma_k post_i + \delta X_i + \epsilon_i \tag{4}$$

where esc_{ik} corresponds to a dummy that takes the value 1 if the individual i completed k or fewer years of schooling, for all values of k between 0 and 19. The variable $post_i$ is a *dummy* is a dummy indicating whether the individual is fully exposed to the reform (if they started first grade after 1981) and X_i represents the set of controls used the "second" stage. The specification of equation (4) is less detailed than equation (3) as we just want to show the effect of the reform on educational attainment.

In figure 1 we plot the estimated γ_k, representing the estimated impact of the reform on every level of education. The shape of the figure indicates on which educational level the reform had an effect. The coefficient values for $k < 12$ are negative. We see that for 12 years of education the effect of the reform is positive, i.e., the reform increased the probability of graduation from high school school. This indicates that the reform implicitly moved schooling from the educaci'on b'asica (8 years of schooling) to graduation from the high school level (educaci'on media). The coefficients over 13 years of schooling are not significant, indicating that the reform had no effect on higher education, thus providing additional evidence to support an identification strategy that focus on high school graduation as the educational outcome to be used.[24]

3.3.3 Placebo

We explore the validity of our identification strategy using two placebo tests. First, we estimate equation (4) redefining the variable $post_i$ to correspond wrong years for the reform, starting in the year 1982 and up to 1987. We observe that the effect on high school graduation dissipates as we move away from the true reform year. This indicates that, after controlling for the trend in schooling, the effect of the reform is observed only around the true date.

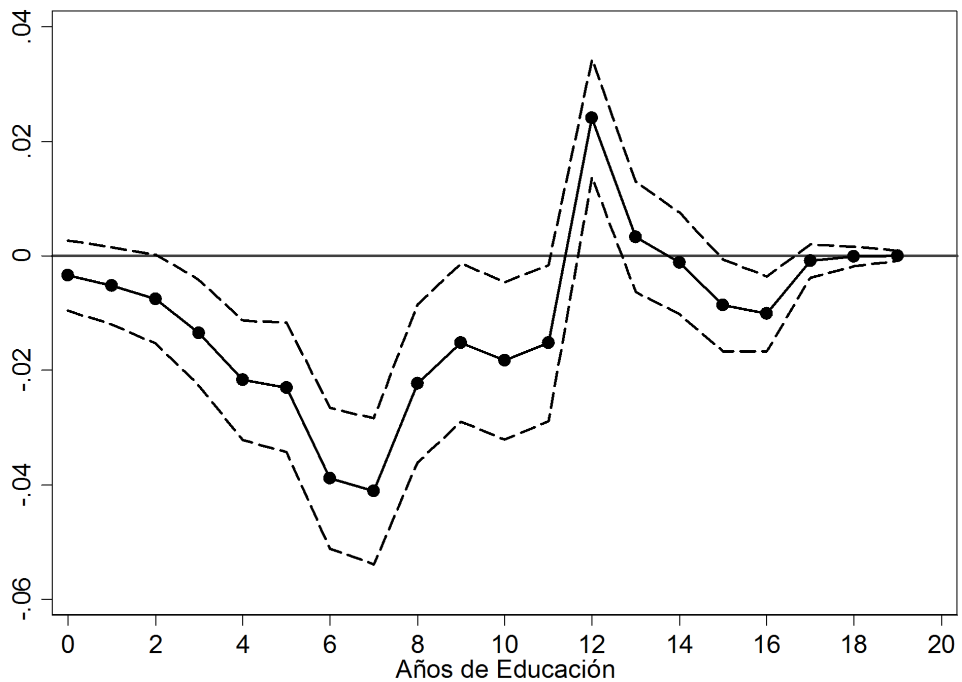
Second, we reestimate equation (4) adding a placebo dummy for a reform corresponding to a wrong year,

$$media_i = \alpha + \gamma_k post_i + \delta_k postfalse_{ik} + \gamma X_i + \epsilon_i \tag{5}$$

where all variables are defined as before except for the $postfalse_{ik}$ variables are defined as dummies that take the value 1 if the individual attended school after year 19k, with k =73, 74, 75 , 76, 77. We estimate the equation once for each of the placebo dummies. If the reform produced an effect beyond the tendencies in education, the $postfalse_{ik}$ dummy coefficients should not be significant once controlled for the real reform dummy.

[24] A similar exercise confirms that the reform had a weak and very imprecisely estimated effect on years of education when using the year of reform as identification variable.

Figure 1: Estimated coefficientes γ_k of the Linear Probability Model (4) with Confidence Intervals at 95%



The results, presented in table 5, confirm that the *post* dummy effect is positive and significant in all the specifications as previously mentioned; while the $postfalse$ dummy effect is not significant in any of them. Column 1 shows the results of the estimate of the equation (4.3) without including the postfalsa variable. We see that the post 81 coefficient remains stable in all the specifications. These results support the proposed identification strategy, as they indicate that the effect of the reform effectively occurs in the expected generation.

4 Results

In section 3, it was shown that the identification strategy used is robust and that the education reform of 1981 had a significant effect on the probability of completing secondary education. In this section, the results of the second stage of the estimate are presented, showing the estimated effect of secondary education on financial participation.

Table 5: Identification Strategy: Simplified First Stage and Placebo Specifications
Dependent variable: $media_i$, dummy for having 12 or more years of education.

Variables	1	2	3	4	5	6
Post	0.0153**	0.0160*	0.0155*	0.0161**	0.0167**	0.0160**
	(0.00820)	(0.00820)	(0.00743)	(0.00787)	(0.00757)	(0.00744)
Post 73		0.00160				
		(0.00754)				
Post 74			0.000525			
			(0.00768)			
Post 75				0.00220		
				(0.00766)		
Post 76					0.00634	
					(0.00753)	
Post 77						0.00848
						(0.00748)
N	135,360	135,360	135,360	135,360	135,360	135,360
R^2	0.160	0.160	0.160	0.160	0.160	0.160

All regressions include the additional controls in equation (4). The p-values for the *Post* in estimates 2 and 3 are 0.051 and 0.054 respectively. Robust standard errors in brackets.
*** p<0.01, ** p<0.05, * p<0.1

4.1 Baseline Specification

The results of the estimates from the Probit model are presented in column 1 of Table 6, where we see that the secondary education dummy variable is significant and positive. Columns 2 to 5 correspond to the results of the Biprobit model according to the different specifications of the first stage. Here we notice that the coefficients are significant and positive in all the estimates and the results are robust to the different specifications from the first stage. In the case of the Biprobit model, the effect of secondary education on financial participation is lower than in the Probit estimate which does not consider the problem of endogeneity. This suggests the presence of a bias in the variable omitted in the Probit estimate, meaning that the secondary education coefficient in this estimate was capturing the effect of an unobservable variable as well as the effect of the education, which would explain its upward bias.

The biprobit estimates suggest that the completion of secondary education raises the probability of participating in the financial market by 3%. Given that the participation in this market is 9.9% in the CASEN 2009, a 3% increase in participation probability is not just statistically significant but also economically important. In this sense, we can say that education is a relevant factor to help explain part of the question of the observed participation in Chile. We can do a simple back of the envelope exercise to measure the contribution of education. Financial participation among those with complete secondary education is 13.41%, while the participation of those with incomplete secondary education is 7.61%. Therefore the estimated causal effect of 3% can explain approximately

Table 6: Probit and Biprobit Estimates of the Effect of Education on Financial Market Participation
Dependent variable: Financial market participation

Variables	1 Probit	2 Biprobit	3 Biprobit	4 Biprobit	5 Biprobit
$media_i$	0.0528***	0.0279*	0.0314**	0.0299**	0.0366**
	(0.00178)	(0.0149)	(0.0152)	(0.0151)	(0.0151)
N	135,360	135,360	135,360	135,360	135,360
Excluded instruments					
Dummy Post Reform	–	Yes	Yes	No	No
Post reform * Zones	–	No	Yes	No	No
Degree of exposure	–	No	No	Yes	Yes
Exposure * Zones	–	No	No	No	Yes

The dummy coefficients by region, gender and age polynomial are omitted. Columns 2,3,4 and 5 correspond to the different specifications of the first stages. The interactions included are the north, center and south zone dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.
*** p<0.01, ** p<0.05, * p<0.1

half of the differences observed in the data. It is worth mentioning again that biprobit captures an ATE effect, so any use of specific initial values should be interpreted with caution.

4.2 Adding Income as an Additional Control

In our original specification we do not include income as an additional control because we are afraid about potential endogeneity; thereby reintroducing bias to the estimates. With the simple objective of obtaining results comparable to those of Cole et al. (2012), we run the same regressions described previously, adding a cubic polynomial of income as a control.[25] Columns 1 and 2 of Table 7 show the results of the estimate from the Probit model. Here we see that after controlling for income, the coefficient of the secondary education dummy falls from 0.0528 to 0.027. This suggests that our baseline estimates reflect that the income level is one of the mechanisms through which education increases financial market participation. Columns 3 and 4 show that in the case of the biprobit model, there is practically no difference between the estimates thus indicating that our results are not fully explained because the instruments are indeed affecting through income and not through education alone.[26]

[25] The variable used is "Work Income" from the CASEN 2009, which corresponds to the income received by those working in their main occupation in the form of salaries or wages, earnings from independent work or self-provision from assets produced in the home.

[26] These results are still not directly comparable for a number of reasons: 1) There are still some differences in the specification, 2) the method of estimation is different, and, 3) the structure of financial markets differ between countries.

Table 7: The Effect of Education on Financial Market Participation: Adding Income as a Control
Dependent variable: Financial market participation

Variables	1 Probit W/o Income	2 Probit With Income	3 Biprobit W/o Income	4 Biprobit With Income
$media_i$	0.0528***	0.0270***	0.0366**	0.0387**
	(0.00178)	(0.00183)	(0.0151)	(0.0176)
N	135,360	135,360	135,360	135,360

The dummy coefficients by region and gender and age and income polynomials are omitted. For the two biprobit estimates the set of excluded instruments includes the degree of exposure to the reform and their interactions with north, center and south. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.
*** p<0.01, ** p<0.05, * p<0.1

The estimates shown in Table 7 are considerably lower than the results obtained for the United States by Cole et al. (2012). In their work, the authors find that an increase in standard deviation in schooling increases the probability of participation in the financial market by 18%.[27] In contrast, the results in our paper show that completing high school increases the probability of participating in the financial market in 3%. One relevant factor which could help to explain part of this difference is that the participation in the financial market in the United States is approximately 40%; which is considerably higher than in Chile where it is just below 10% in our sample.

5 Transmission Channels: An Initial Exploration

One drawback of our baseline specification is that likely captures several channels through which education can affect financial market participation. In this section we attempt to explore some of these potential channels. For this purpose, alternative regressions are estimated, refining the dependent variables using the CASEN 2009 and 2011, and also creating an index of financial literacy with the Social Protection Survey (EPS).

5.1 Separación por Tipos de Activos

With the purpose of better identifying the effect of the reform on financial participation, assets are separated into two categories: simple assets and complex assets. We consider housing savings plans, savings in housing fund administrators and bank savings accounts as simple assets. Complex assets include voluntary savings plans, savings in Cuenta 2 of AFPs, fixed-term deposits, mutual fund investments, shares and bonds.

[27] In the work of Cole et al. (2012) the median value of schooling is 12.9 years and the standard deviation is 2.7 years.

The idea is that the effect of education on financial participation should be greater in the less sophisticated assets, given that the education reform principally affected the probability of completing secondary education. This is particularly the case for the level of education we are looking at, because complex assets require either high income levels or a significant level of relevant knowledge. In a way, the participation in simple assets corresponds to an initial approach towards the financial system, which is exactly where we believe that the reform should have an effect, if any.[28] Therefore we expect that the effect of the secondary education dummy will be greater for the simple than for the complex assets.

In table 8 we present the results from estimating the biprobit model separately by type of assets. For reference purpose we present in Column 1 the estimates using general financial participation as the dependent variable (this corresponds to the same estimates in column 5 of Table 1). Columns 2 and 3 show the results for simple and complex assets, where we see that they are positive and significant. Just as we expected, the high-school completion coefficient is larger for simple assets than for complex assets.

Table 8: The Effect of Education on Financial Market Participation: Separating by Type of Financial Asset

Dependent Variable: Financial Participation

Variables	1 Any Asset	2 Simpler Assets	3 Complex Assets
$media_i$	0.0366**	0.0433*	0.0139***
	(0.0151)	(0.0238)	(0.00292)
N	135,360	135,360	135,360

All equations estimated using bi probit model. The dummy coefficients by region, gender and age polynomial are omitted. The biprobit estimates include the following excluded instruments: Degree of exposure to the reform and their interactions with north, center and south dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.

*** p<0.01, ** p<0.05, * p<0.1

Additionally, using the 2011 CASEN we can measure whether a household has or not insurance policies (beyond the mandatory health insurance or third party protection for cars).[29] In this case we repeat the same type of analysis as we do for the 2009 CASEN, with exactly the same instruments and endogenous dependent variable, but instead of financial market participation we include a dummy variable that indicate whether any member of the household has health insurance

[28]Following this line of reasoning we can argue that this separation also serves as a falsification exercise, because finding a larger effect in complex assets should cast doubts on the channels we propose.

[29]Unfortunately this last wave of the CASEN did not ask the same questions about financial assets and thus we were not able to construct a (syntehtic) panel for our study. In this case we also repeat our analysis of the years of education to make sure our first stage also works and confirm that this is indeed the case. The results are available upon request from the authors.

Table 9: Probit and Biprobit Estimates of the Effect of Education on Having Insurance Policies

Variables	Probit 1	Biprobit 2	Biprobit 3	Biprobit 4	Biprobit 5
Panel A. Dependent Variable: Health Insurance					
$media_i$	0.108***	0.0989***	0.0640***	0.0998**	0.120***
	(0.00196)	(0.0198)	(0.00405)	(0.0207)	(0.0241)
N	109,904	109,904	109,904	109,904	109,904
Panel B. Dependent Variable: Life Insurance					
$media_i$	0.141***	0.144***	0.115**	0.144**	0.150**
	(0.00298)	(0.0287)	(0.00805)	(0.0294)	(0.0339)
N	58,026	58,026	58,026	58,026	58,026
Excluded instruments					
Dummy Post Reform	–	Yes	Yes	No	No
Post reform * Zones	–	No	Yes	No	No
Degree of exposure	–	No	No	Yes	Yes
Exposure * Zones	–	No	No	No	Yes

The dummy coefficients by region, gender and age polynomial are omitted. Columns 2,3,4 and 5 correspond to the different specifications of the first stages. The interactions included are the north, center and south zone dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.

*** p<0.01, ** p<0.05, * p<0.1

and another dummy for life insurance. As we can see in table 9, there is a positive and statistically significant effect of education on the probability of having (voluntary) insurance, a result that is in line with our initial hypothesis. Furthermore, in this case the effect is larger than for financial market participation, even for the case of simple assets, a fact that is somewhat surprising.[30]

5.2 Risk Aversion and Employment Characteristics

In this section we explore two other plausible channels according to the existing literature. First, we test the effect of education on the willingness of individuals to take risks. Conventional portfolio theories indicate that, no matter how averse to risk the agents are, all should maintain some percentage of their wealth in financial assets. In spite of this, we believe that there are two reasons for which the modification of the willingness to take risks can represent a relevant transmission channel through which education affects financial participation. First, lower levels of risk aversion will cause people

[30]We are not aware of any theoretical reason to expect this. Just basic introspection leads us to think that the financial assets included in our 2009 CASEN estimations are probably products individuals would acquire before extra insurance.

Table 10: Effect of Education on Moving to Another Region, Pension Contributions and Employment Status

Dependent variables: Moved to another region, Pension Fund Contributions and Employed

Variables	1 Has changed region	2 Pension Fund Cont.	3 Employed
$media_i$	0.00770**	0.158***	0.347***
	(0.00342)	(0.0348)	(0.0466)
N	135,360	135,360	135,360

All equations estimated using bi probit model. The dummy coefficients by region, gender and age polynomial are omitted. The biprobit estimates include the following excluded instruments: Degree of exposure to the reform and their interactions with north, center and south dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.

*** p<0.01, ** p<0.05, * p<0.1

to invest a higher proportion of their wealth in financial assets, thereby raising the increased level of financial participation, likely making the individuals more aware or more eager to understand the financial market. Second, the parameter of risk aversion can play a crucial role in the decision to participate in the financial market given the fixed costs involved in entering it (Vissing-Jorgensen, 2002). This happens because where these types of costs are present, individuals that are sufficiently averse to risk may prefer to remain outside the financial market in case their subsequent earnings do not exceed the cost of entering the market. In these cases, a modification in the willingness to take risks, produced by an increase in education, may affect the individuals' decision regarding financial participation.

The connection in our case comes from previous studies have shown a relationship between education and individuals' preferences. Becker and Mulligan (1997); Harrison et al. (2002) and particularly Halek and Eisenhauer (2001) suggest a negative correlation between risk aversion and education. In our case, as we do not have a good measurement of individuals' risk aversion, we implement an indirect test, taking the probability of people moving from one area of the country to another as an indicator of risk aversion. The idea behind this strategy is that the decision to migrate is risky, and therefore those willing to move have a higher tolerance of risk.[31]

Column 1 of Table 10 shows that the effect of education on willingness to move from one region to another is significant and positive. Specifically, the completion of secondary education increases the probability of moving from one region to another by 0.7%; which suggests that education can reduce an individual's risk aversion. It is helpful to raise a caution regarding this result, as the increase in probability of moving from one region to another could reflect the fact that the completion of secondary education increases individuals' probability of going to universities; which are generally

[31] Heitmueller (2005) argues that risk aversion is an important determining factor in the decision to migrate within the United States.

located in the center zone of the country.

Second, we examine the effect of education on the job characteristics and employment status. We mentioned previously that education could increase possibilities of employment in areas of the economy that are closer to the financial system. For example, formal employment incorporates the worker to the national pensions system (AFP), causing them to compulsorily participate in the financial system and thus have to face financial issues. To test this transmission mechanism, we estimate the effect of education on the probability of having to contribute to a pension fund and on the probability of being employed. In both cases we expect the estimated coefficients to be positive. Columns 2 and 3 of Table 10 show the results of these estimates. We can see that both coefficients are positive and significant, which supports our idea regarding job characteristics and employment status as a transmission channel.

5.3 Financial Literacy as Transmission Channel

Table 11: Efecto de la Educación sobre el Financial Literacy
Dependent variable: Financial Literacy Index

Variables	1 OLS	2 2SLS	3 2SLS	4 2SLS	5 2SLS
$media_i$	0.165***	0.394**	0.370**	0.363**	0.158
	(0.00465)	(0.186)	(0.182)	(0.178)	(0.118)
Cragg-Donald	–	25.26	11.90	7.20	4.25
N	16,443	16,443	16,443	16,443	16,443
Excluded Instruments					
Dummy Post Reform	–	Yes	Yes	No	No
Post reform * Zones	–	No	Yes	No	No
Degree of exposure	–	No	No	Yes	Yes
Exposure * Zones	–	No	No	No	Yes

The dummy coefficients by region, gender and age polynomial are omitted. Columns 2,3,4 and 5 correspond to the different specifications of the first stages. The interactions included are the north, center and south zone dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.

*** p<0.01, ** p<0.05, * p<0.1

In this section we try to explore another channel that has been extensively documented in the literature. Several authors have found a relation between different indexes of financial literacy and a variety of financial outcomes, including financial market participation (van Rooij et al., 2011, see for example). In the context of the relation between formal education and financial market participation, formal education can have a significant impact on financial literacy. In order to decompose the effect of education on financial market participation we want to take a step back with respect to the literature and look at the effect of education on financial literacy. With a causal estimate of this effect we can then calculate how much of the effect of education on financial market participation can be attributed to the effect through financial literacy. For this purpose we construct a financial literacy index using six questions included in the 2006 Encuesta de Protección Social (EPS). All these questions were designed to measure an individual's ability to perform simple calculations and process financial information. The index is calculated as the fraction of the questions correctly answered by the individual.[32]

We can combine our results in Table 11 with a causal estimate of the effect of education on financial literacy to obtain the effect of education on financial market participation that comes through financial literacy. Taking the results from Landerretche and Martínez (2013), who find that a 1% increase in measured financial literacy leads to a 1.5%-3% increase in the probability that a person holds financial assets, and applying our estimated impact on Table 11 we obtain that education increases financial market participation in 1.1% through its effect on financial literacy, a number that represents between 30% and 40% of the total effect estimated in Table 6.

6 Conclusions

The limited level of participation in financial markets by individuals and households in the real world contradicts most of our theoretical understanding about the benefits of financial markets, and our understanding of the determinants of portfolio choice. At the same time, this limited participation can have important effects on the behavior of financial markets.

Among the several explanation that have been proposed to explain this result, education is often mentioned among the important candidates. In this paper we contribute to this literature by estimating the causal effect of education on financial market participation among Chilean households. Given that a simple regression of financial market participation on educational attainment is likely to suffer from omitted variable bias and reverse causality, we design an instrumental variable strategy that exploits a large scale educational reform that took place in Chile 1981 and that significantly expanded the supply of schools in the country. In particular, we exploit the variation across cohorts and geographic zones to obtain plausible instruments for the probability that an individual graduates from high school (educación media in Chile).

We find that finishing high school increases by 3% the probability of participating in the financial market (this is an increase of 3 percentage points). This effect is statistically significant but also economically meaningful. Given that only 9.9% of our sample actually participates in the financial market, our estimated effect implies an increase of about 1/3 of the initial level. Moreover, this result also implies that the causal impact of education can explain approximately half of the difference between the average financial market participation between those have and have not completed high school. Overall, our baseline results are consistent with the hypothesis that lower educational

[32]In the appendix B we present an alternative index that constructed using a methodology more similar to that used in recent papers, see van Rooij et al. (2011). The results using that index instead of the simpler one presented here are qualitatively similar and are presented in the same appendix.

attainment reduces financial participation, and are robust to considering different sets of excluded instruments in the biprobit estimations.

In order to explore the mechanism(s) through which education affects financial market participation we perform some extra estimations. First, we conjecture that high school education should increase participation in simpler assets vis-a-vis more complex assets, and effectively find this pattern in the data. We also observe that high school education is associated with higher probability of one household member having health or life insurance (beyond what is required by law). Second, we also find that education reduces risk aversion (proxied by the decision to move to a different region in the country), increases the probability of contributing to pension funds (likely because of higher probability of having a formal job, forcing the individual to contribute and thus exposing him to the financial market), and increasing the probability of having a job. Finally, we also observe that education causes an increase in simple financial literacy indexes, and that this particular channel explains between 1/3 and 2/5 of the total impact of education on financial market participation. Our results are compatible with models where individuals do not participate in financial markets because they do not understand the products and lack the knowledge required to take advantage of it. Next in the agenda is the study of the consequences this increased participation has, including welfare improvements or even whether conditional on participating people makes fewer mistakes or not.

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A Descriptive Statistics

Table A.1

Descriptive statistics for individual characteristics in 2009 CASEN

Characteristic	Years of schooling	Age	Financial Participation (%)	N
Educational Attainment				
Educación Básica	5.27	50.21	7.22	63,046
	(2.60)	(12.10)	(25.89)	
Educación Media	11.32	41.37	10.71	53,492
	(1.05)	(11.61)	(30.93)	
College	15.61	39.08	16.94	18,822
	(1.56)	(11.88)	(37.51)	
Geographical Zone				
North	9.39	44.68	9.46	16,224
	(4.21)	(12.82)	(29.27)	
Center	9.52	45.08	9.69	67,328
	(4.24)	(12.79)	(29.58)	
South	8.46	45.44	10.46	51,758
	(4.35)	(12.81)	(30.60)	
Whole sample	9.10	45.17	9.95	135,360
	(4.31)	(12.80)	(29.94)	

Standard deviations in parenthesis.

North includes the following regions: I, II, III, IV and XV.

Center includes: V, VI, VII, XIII.

South includes: VIII, IX, X, XII, XIII, XIV.

Source: CASEN 2009

Tabla A.2

Descriptive statistics for household characteristics in 2009 CASEN

Characteristic	Years of schooling	Age	Financial Participation (%)	N
Educational attainment				
Educación Básica	5.26	52.96	12.47	30,361
	(2.56)	(11.16)	(33.04)	
Educación Media	11.25	45.25	18.05	21,247
	(1.08)	(11.17)	(44.91)	
College	15.67	45.00	28.03	6,978
	(1.61)	(11.36)	(44.91)	
Geographical zone				
North	8.98	48.57	15.16	7,195
	(4.19)	(11.95)	(35.86)	
Center	9.06	49.32	16.05	28,533
	(4.27)	(11.76)	(36.71)	
South	8.09	49.29	17.01	22,858
	(4.28)	(11.91)	(37.65)	
Whole sample	8.67	49.22	16.35	58,586
	(4.29)	(11.84)	(36.98)	

Standard deviations in parenthesis.

North includes the following regions: I, II, III, IV and XV.

Center includes: V, VI, VII, XIII.

South includes: VIII, IX, X, XII, XIII, XIV.

Source: CASEN 2009

Tabla A.3

Descriptive statistics for EFH 2007

Characteristics	Years of schooling	Age	Securities and equity	Fixed Income	Financial Participation	Observaciones
Educational attainment						
Educación Básica	5.91	53.36	0.48	5.21	5.70	520
	(2.14)	(11.06)	(6.97)	(22.24)	(23.20)	
Educación Media	11.25	46.53	2.36	10.03	11.53	1,160
	(1.08)	(11.26)	(15.20)	(30.06)	(31.95)	
College	15.53	44.37	11.09	19.32	26.62	1,636
	(1.76)	(11.24)	(31.41)	(39.49)	(44.21)	
Geographical area						
North	10.96	48.78	3.47	9.77	12.49	373
	(3.57)	(11.66)	(18.34)	(29.73)	(33.10)	
Center	11.32	47.65	4.52	11.21	14.06	2,437
	(3.94)	(11.90)	(20.78)	(31.56)	(34.76)	
South	10.30	47.56	4.87	13.33	16.76	506
	(4.35)	(11.31)	(21.54)	(34.03)	(37.38)	
Whole sample	11.05	47.78	4.45	11.47	14.43	3,316
	(4.00)	(11.75)	(20.64)	(31.88)	(35.14)	

Standard deviations in parenthesis.

North includes the following regions: I, II, III, IV and XV. Center includes: V, VI, VII, XIII.

South includes: VIII, IX, X, XII, XIII, XIV.

Calculations made using the weights provided with the survey.

Demographic variables refer to the head of household.

Source: EFH 2007

Tabla A.4

Descriptive statistics EPS 2006

Individual characteristics	Age	Financial Participation (%)	N
Educational attainment			
Educación Básica	55.28	16.52	6,053
	(15.50)	(37.13)	
Educación Media	44.16	26.68	7,300
	(13.91)	(44.23)	
College	40.45	37.70	3,090
	(13.84)	(48.47)	
Geographical area			
North	46.45	21.64	1,848
	(15.55)	(41.19)	
Center	47.17	25.11	10,045
	(15.68)	(43.37)	
South	48.86	26.15	4,550
	(15.77)	(43.95)	
Whole sample	47.55	25.01	16,443
	(15.71)	(43.31)	

Standard deviations in parenthesis.

North includes the following regions: I, II, III, IV and XV.

Center includes: V, VI, VII, XIII.

South includes: VIII, IX, X, XII, XIII, XIV.

Age and financial participation refer to the head of household.

Source: EPS 2006

Tabla A.5

Descriptive statistics for variables created for the paper

Variable	Average	Std. Dev.	Min	Max	N
CASEN 2009					
$media_i$	0.4039	0.4906	0	1	135,360
Overall Financial Participation	0.0995	0.2994	0	1	135,360
In Simple assets	0.0860	0.2804	0	1	135,360
In Complex assets	0.0135	0.1154	0	1	135,360
Change of region (dummy for yes)	0.0217	0.1458	0	1	135,360
Contributions to pension funds (dummy for yes)	0.3690	0.4825	0	1	135,360
Employed (dummy for yes)	0.5845	0.4928	0	1	135,360
EPS 2006					
$media_i$	0.2966	0.4567	0	1	16,443
Financial Participation	0.2501	0.4331	0	1	16,443
Base Financial Literacy index	0.3378	0.2655	0	1	16,443
Alternative Financial Literacy index	3.99e-09	1.2600	-1.5867	2.9658	16,443

CASEN 2009 measured at the individual level.

EPS 2006 are for the head of the household.

Financial participation includes the head of the household and the spouse

Source: authors' own elaboration using CASEN 2009 and EPS 2006

B Financial Literacy Index

The financial literacy index is based on the following 6 questions from the 2006 EPS

1. Si la posibilidad de contraer una enfermedad es de un 10 por ciento, ¿cuántas personas de 1000 contraerían la enfermedad?
2. Si 5 personas tienen los números premiados de la lotería y el premio es de dos millones de pesos, ¿cuánto recibiría cada una?
3. Suponga que Ud. tiene $100 en una cuenta de ahorro, y la tasa de interés que gana por estos ahorros es de un 2% por año. Si mantiene el dinero por 5 años en la cuenta, ¿cuánto tendrá al término de estos 5 años?
   (a) Más de $102
   (b) Exactamente $102
   (c) Menos de $102
4. Digamos que Ud. tiene $200 en una cuenta de ahorro. La cuenta acumula 10 por ciento en intereses por año. ¿Cuánto tendrá en la cuenta al cabo de dos años?
5. Suponga que Ud. posee $100 en una cuenta de ahorro, la que entrega un interés de un 1% anual. Ud. sabe también que la tasa de inflación es de un 2% anual. Después de un año, si retira la plata de una cuenta de ahorro. Ud. podrá comprar:
   (a) Más de $100
   (b) Exactamente $100
   (c) Menos de $100
6. La siguiente frase, ¿es verdadera o falsa? "Comprar una acción de una empresa es menos riesgoso que comprar con el mismo dinero varias acciones de distintas empresas"
   (a) Verdadero
   (b) Falso

In table 12 we present the results of these questions.

For each question we constructed a dummy variable that indicated whether the person had correctly answered that question. The baseline index, used in the regressions in the text, is computed as the simple average of these dummies. The second index performs factor analysis on the six dummies using the iterated principal factor analysis method. The factor loadings are then used to compute the factor scores with the Bartlett methods. In table 13 we present the results from the factor analysis.

Table 14 presents a comparison of basic statistics for both indexes.

Table 12: Fraction of People who Answered Correctly each of the Questions about Financial Literacy

Questions EPS	Correct Answers
Question 1	45 %
Question 2	41 %
Question 3	46 %
Question 4	2 %
Question 5	25 %
Question 6	44 %

Source: authors' own elaboration using 2006 EPS.

Table 13: Factor Loadings for the Alternative Financial Literacy Index

Questions EPS	Factor Loadings
Question 1	0.4783
Question 2	0.4110
Question 3	0.4277
Question 4	0.0953
Question 5	0.2803
Question 6	0.2740

Source: authors' own elaboration using 2006 EPS.

Table 14: Summary Statistics Financial Literacy Indexes

Index	Average	Std. dev	Minimum	Maximum	N
Base Index	0.3378	0.2655	0	1	16,443
Alternative index	3.99e-09	1.2600	-1.5867	2.9658	16,443

Source: authors' own elaboration using 2006 EPS.

Table 15: Efecto de la Educación sobre el Financial Literacy
Dependent variable: Alternative Financial Literacy Index

	1	2	3	4	5
Variables	OLS	2SLS	2SLS	2SLS	2SLS
$media_i$	0.797***	1.867**	1.755**	1.801**	0.807
	(0.0219)	(0.877)	(0.861)	(0.847)	(0.556)
Cragg-Donald	–	31.03	12.79	7.45	4.75
N	16,443	16,443	16,443	16,443	16,443
Excluded Instruments					
Dummy Post Reform	–	Yes	Yes	No	No
Post reform * Zones	–	No	Yes	No	No
Degree of exposure	–	No	No	Yes	Yes
Exposure * Zones	–	No	No	No	Yes

The dummy coefficients by region, gender and age polynomial are omitted. Columns 2,3,4 and 5 correspond to the different specifications of the first stage. The interactions included are the north, center and south zone dummies. The values in this table correspond to the average individual marginal effects. Robust standard errors in brackets.

*** p<0.01, ** p<0.05, * p<0.1

C Additional Figures

Figure 2: Estimated coefficients for variable Post 82 with 95% confidence intervals

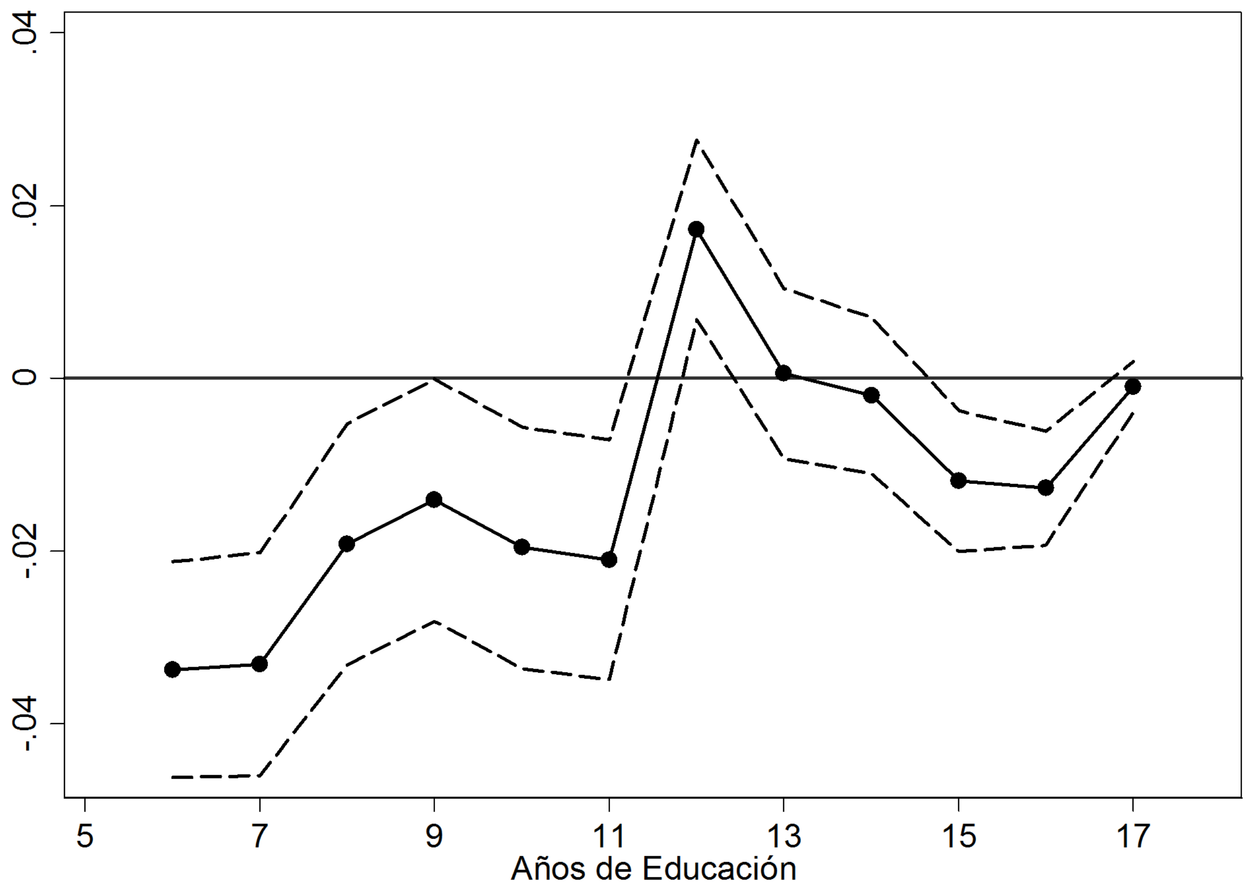

Figure 3: Estimated coefficients for variable Post 83 with 95% confidence intervals

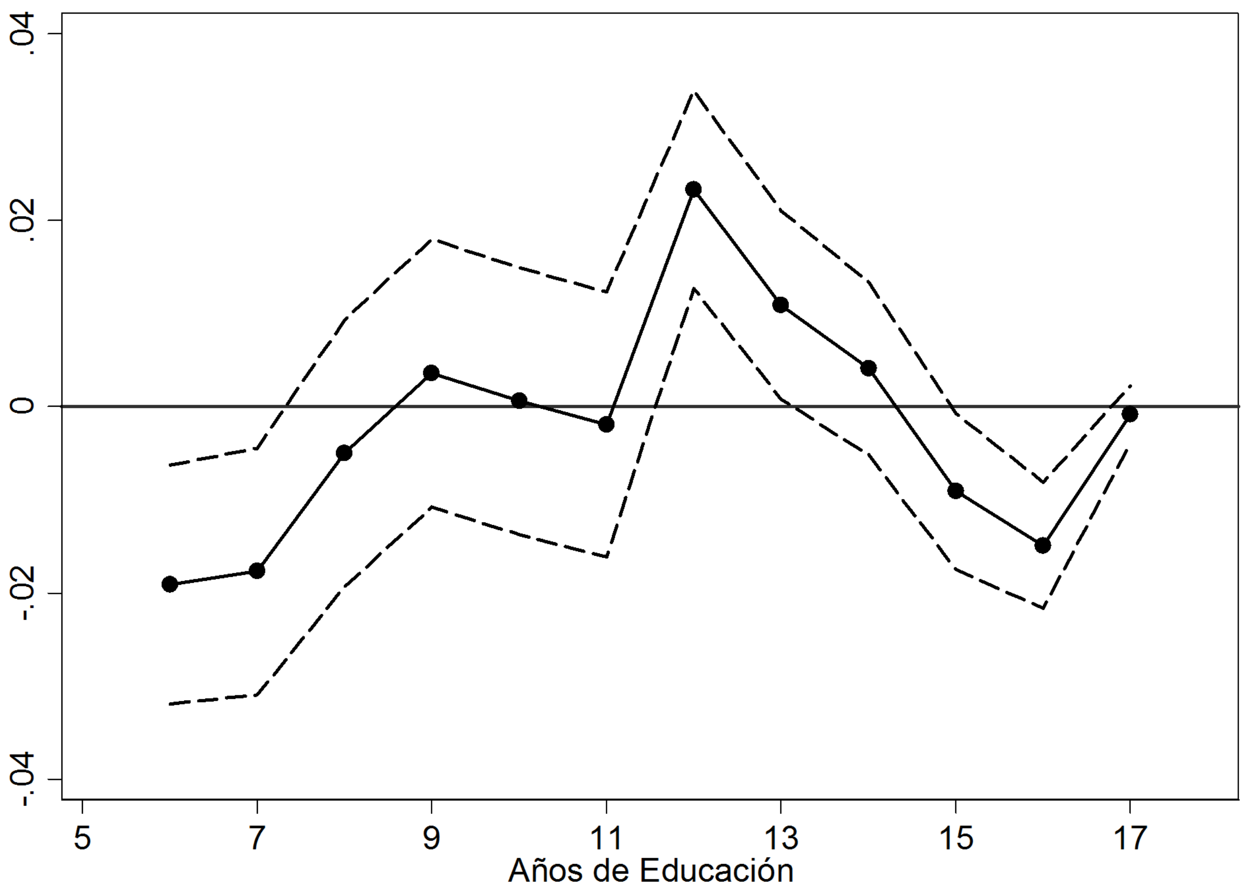

Figure 4: Estimated coefficients for variable Post 84 with 95% confidence intervals

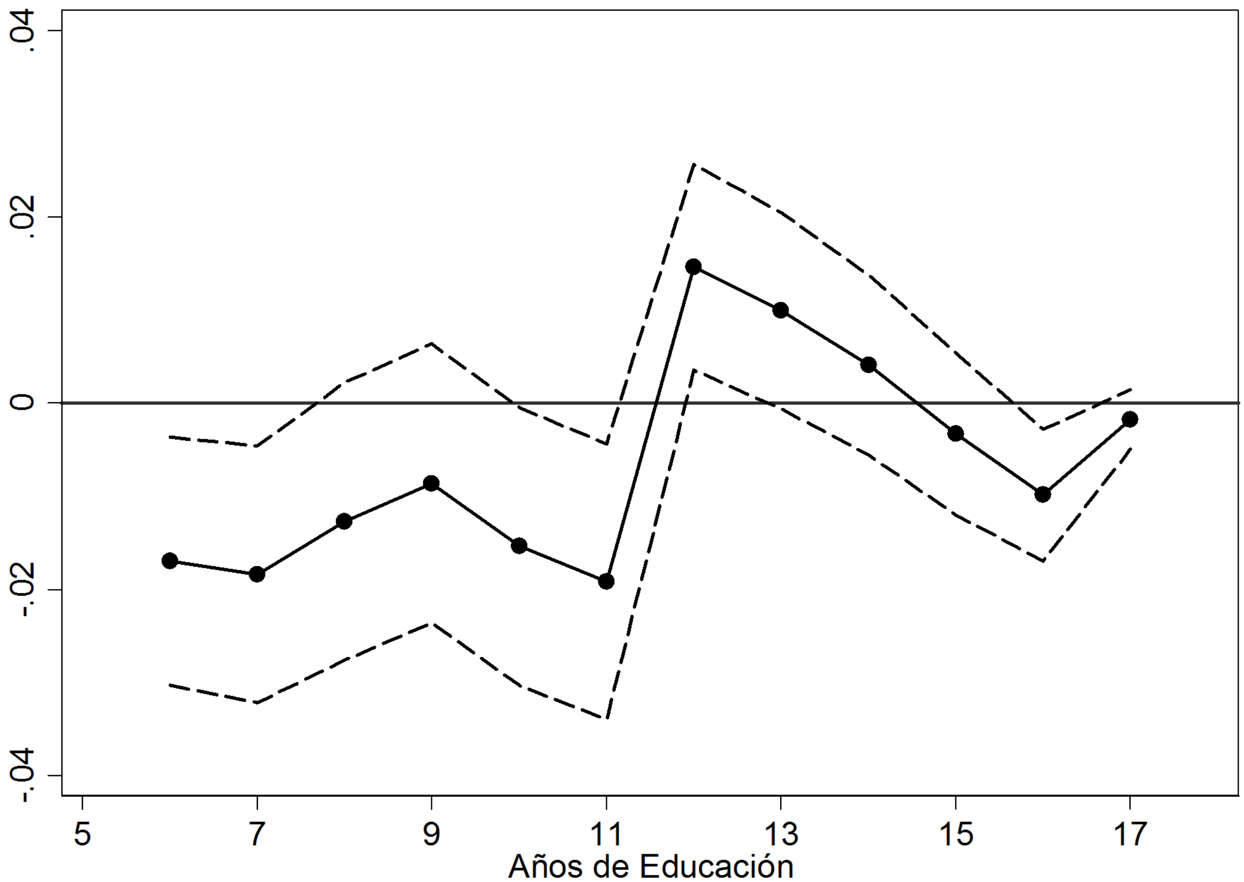

Figure 5: Estimated coefficients for variable Post 85 with 95% confidence intervals

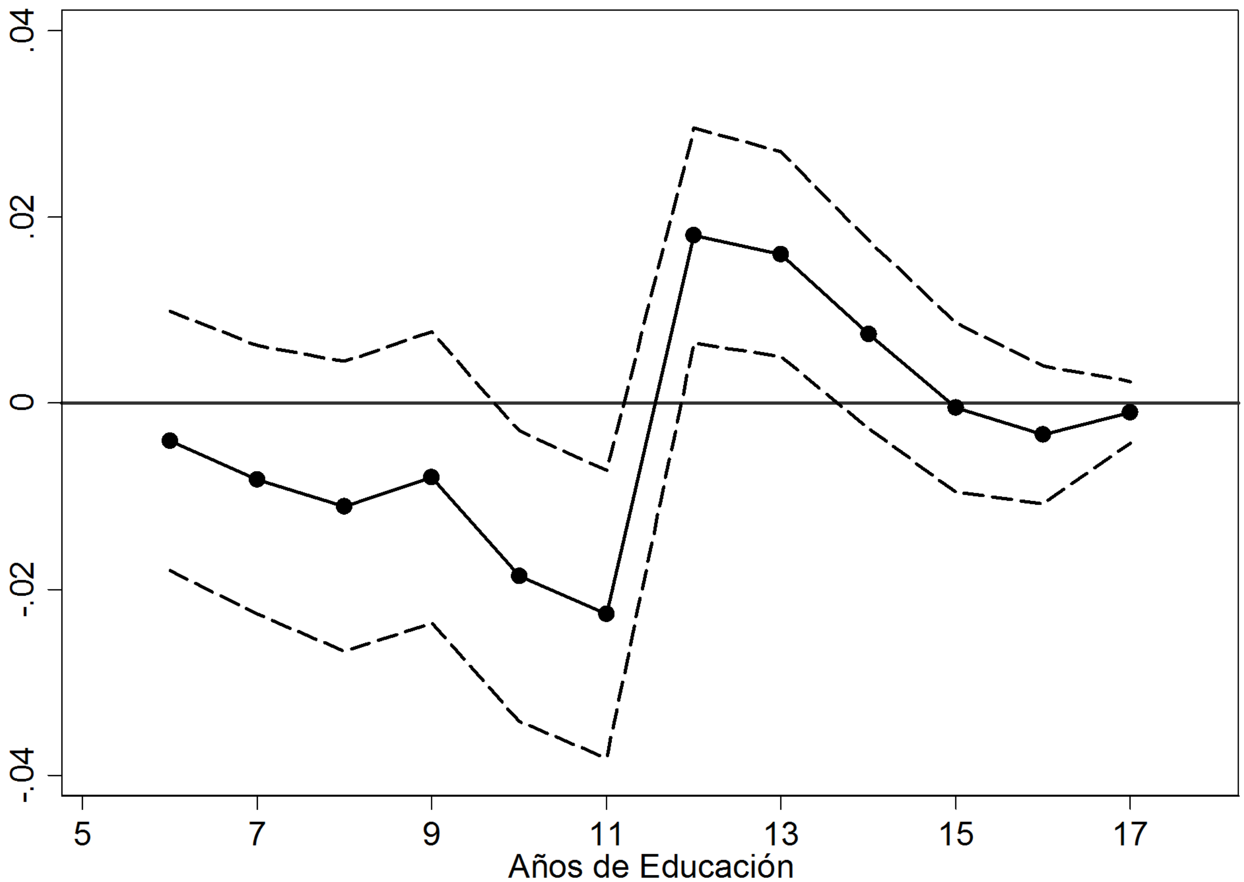

Figure 6: Estimated coefficients for variable Post 86 with 95% confidence intervals

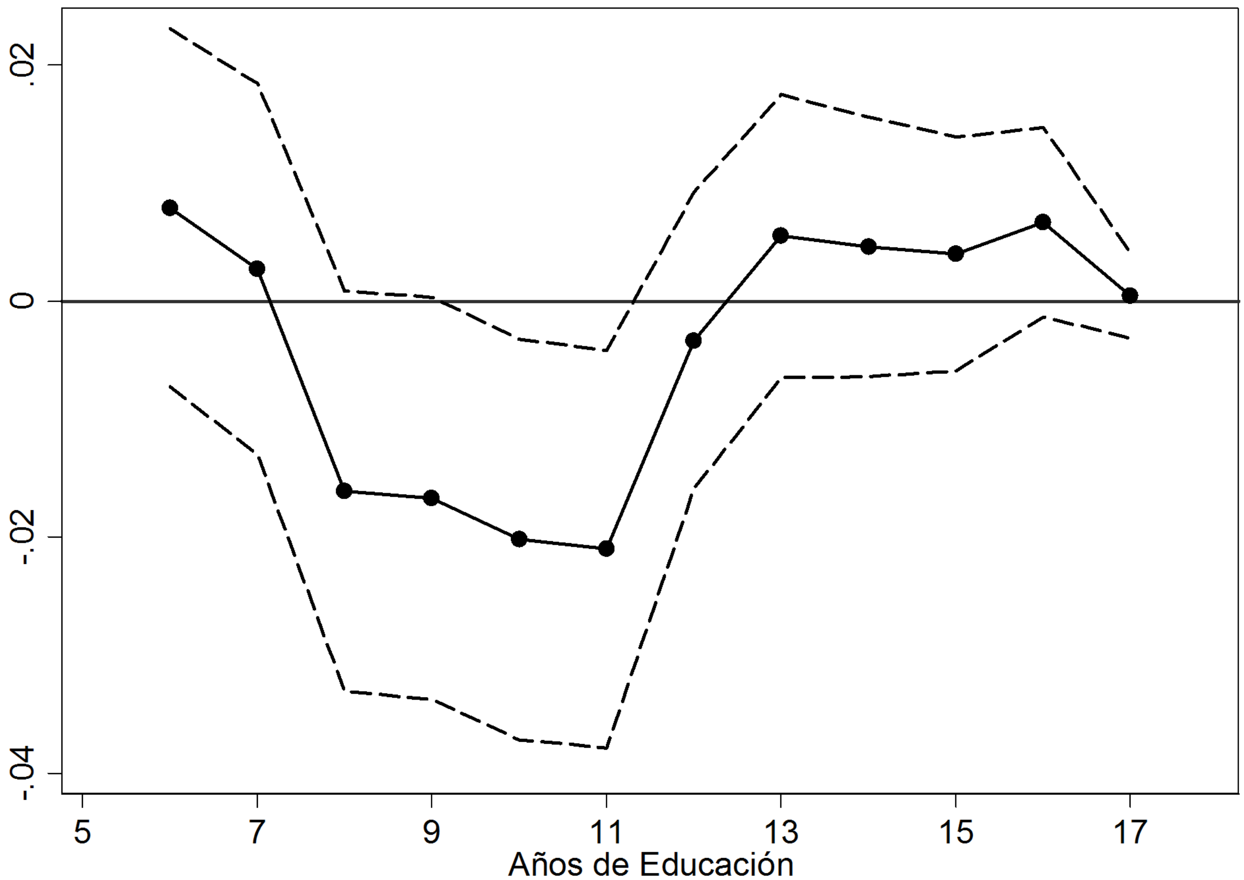

Figure 7: Estimated coefficients for variable Post 87 with 95% confidence intervals