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Quality of Schooling, Returns to Schooling and the 1981 Vouchers Reform in Chile

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Abstract

This paper exploits unique information on cognitive ability to examine the importance of schooling and non-schooling cognitive skills for heterogeneous individuals using instrumental variables estimation. Using a binary instrument based on the 1981 reform in Chile, the authors find that the main beneficiaries of the reform were those who at the time were pupils in basic schooling (ages 6-13). For this treated group of pupils, only a negligible part of the estimated return to schooling is due to classical ability bias. The labor market reward to an additional year of schooling is a measure of the "true" non-cognitive return to schooling. However, once the treated group is expanded to include secondary school students, the pure return to schooling decreases dramatically, while the return to schooling cognitive and non-schooling cognitive skills increases accordingly, suggesting that a large part of the estimated return in an earnings function is due to classical ability bias. For this treated group (mixture of basic school and secondary school age students), the labor market rewarded cognitive skills (especially those acquired through schooling) significantly.

This paper—a product of the Education Team, Human Development Network—is part of a larger effort in the department to document the returns to schooling. Policy Research Working Papers are also posted on the Web at http://econ.worldbank.org. The author may be contacted at hpatrinos@worldbank.org.

Quality of Schooling, Returns to Schooling and the 1981 Vouchers Reform in Chile

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Introduction

When estimating returns to schooling using a Mincer type earnings function, the disturbance term will capture individual unobservable attributes and effects which, in general, tend to influence the schooling decision, hence resulting in a correlation between schooling and the error term. Cognitive (as well as non-cognitive) ability is such an unobservable. If schooling is endogenous, then estimation by Ordinary Least Squares (OLS) will yield biased estimates of the return to schooling.

The return to investing in education, based on past empirical studies, is known to differ among individuals with different cognitive skills. For example, evidence from the United States (Ingram and Neumann 2005) shows that over the past decades, individuals with college education but without specific skills experienced the lowest benefits from investing in education. Therefore, individuals with low ability may not benefit as much from investing in education, compared to individuals in the upper part of the ability distribution. For the latter, ability is expected to interact positively with education resulting in higher benefits from education investments.

When a true measure of ability is an omitted variable in the earnings equation, one of three different approaches have been used in the empirical literature to capture the "true" return to education. The first approach uses twins, to arrive at a measure of the causal return to education. For example, Ashenfelter and Rouse (1998) and Rouse (1997) using data from the United States, have compared the earnings of twins with different educational levels, and reported an estimate of the return to education that is about 30 percent smaller than the OLS estimate. The second approach uses achievement test scores measuring cognitive ability and employs them as additional controls in the earnings function. The third approach uses sources for exogenous variation in educational attainment, such as institutional changes in the schooling system in the form of changes in compulsory schooling laws, abolition if fees etc, as well as other “natural variations” (for example, school construction projects) affecting the schooling decision, to estimate a causal return to education effect using instrumental variable estimation. This is the approach followed in this paper.

Most of the estimates of the return to education based on “natural experiments” report a higher return to education (rather than a lower one), compared to OLS-based estimates of the return to education (see, for example, Angrist and Krueger 1991; Kane and Rouse 1993; Card 1995; Harmon and Walker 1995; Meghir and Palme 2005; for developing countries, see Duflo 2001; Patrinos and Sakellariou 2005; Sakellariou 2006). The dominant explanation for these seemingly counter-intuitive results is that institutional changes in the school system (such as compulsory schooling laws) affect the schooling decision of a subset of individuals who, otherwise, would not have pursued a higher level of education and not the average individual. Furthermore, individuals affected by such reforms tend to have a higher return to education than the average individual.

To incorporate cognitive ability, we rely on the Chilean data from the International Adult Literacy Survey (IALS). The IALS was created to generate comparable literacy profiles across national, linguistic and cultural boundaries. The IALS data have been used in several studies, with little evidence on Chile, the only country in the survey outside Europe and North America.

Blau and Khan (2001) examined the role of cognitive skills in explaining higher wage inequality in the United States. Leuven *et al*. (2004) used IALS data for 15 countries (including Chile) and explored the hypothesis that wage differentials between skill groups across countries are consistent with a demand and supply framework. They find that cognitive achievement is an important determinant of earnings in all countries examined except Poland and Finland. They also find that about one-third of the variation in relative wages between skill groups across countries is explained by differences in net supply of skill groups. Green and Riddell (2002) used the measure of literacy in the IALS dataset to examine the influence of cognitive and unobserved skills on earnings in Canada. They find that cognitive skills contribute significantly to earnings and that their inclusion in earnings equations reduces the measured impact of schooling. Their findings suggest that cognitive and unobserved skills are both productive but that having more of one skill does not enhance the other's productivity. Devroye and Freeman (2001) used the IALS survey and found that skill inequality among advanced countries explains only about 7 percent of the cross-country differences in earnings inequality. They also find that the bulk of cross country differences in earnings inequality occur within skill groups, not between them.

The 1981 Capitation Grant (Vouchers) Scheme

We use the school reform of 1981 to identify a binary instrument and estimate returns to schooling using instrumental variables, with and without controlling for cognitive skills. Chile, among developing countries, was a pioneer and has gone farther than any other country to redefine the roles of the state and private sector in education, by separating the financing from the provision of education. Consequently, the dropout rate declined sharply (from 8 percent in 1981 to 2.7 percent in 1982 for basic education, and from 8.3 percent to 6.2 percent for secondary education). As a result of the capitation grants paid to schools and the opening of the education system to the private sector, Chile was able to absorb the pressure of secondary school expansion. There was a large shift of students to the private subsidized schools (mainly in major urban areas), whose enrollment increased by 93 percent between 1980 and 1985, at the expense of municipal schools. Despite a decline in public spending for education (from 4.0 percent of GDP in 1981 to 2.6 percent in 1990), student intake increased by 42 percent in higher education and by about 15 percentage points in secondary education (Delannoy 2000). As a result of the sweeping 1981 reform, the average years of schooling peak for the cohorts around and immediately after the reform (see Chart 1).

Chart 1: Average Years of Schooling by Cohort (age in 1981)

Overall, the reform improved the efficiency of the education system (number of students educated per unit of public spending increased significantly over the 1980s, without a significant change in average quality). However, this was achieved at the expense of equity, as subsidized private schools serving a better-informed, better-off population benefited from the opportunities offered by the capitation grants system at the expense of municipal schools which served the less well-off section of the population.

The Chilean voucher system has been studied extensively. Some find that the voucher system had positive impacts on test scores and pre-college examinations (Gallegos 2004, 2002; Contreras and Macias 2002; Sapelli and Vial 2002; Sapelli 2003; Gallegos 2004). Yet others found that there was no impact on test scores, repetition rates or secondary school enrollment rates (Carnoy and McEwan 2000; Hsieh and Urquiola 2003). In addition, Gauri (1998) found that school choice had led to increased social and academic stratification. Bellei (2005) outlines three principal reasons why it is difficult to make comparisons between public and private schools in Chile and how they explain the widely diverging results in individual analyses, all stemming from the lack of random assignment of students to schools: (i) private schools tend to be located in urban areas and serve middle to middle-high income students; (ii) there are wide differences in the level of resources available to schools, even among the same types of schools; and (iii) there is very little information about how families select schools and how private schools select students. Thus, it is difficult to measure the supply of private schools, control for school resources, and the estimates are riddled with selection bias. The studies also differ in the ways that they use control variables such as parental education, school socio-economic status, student characteristics, test-score variation, and so on. Gallegos (2006) explains that the differences in results can be attributed to changes in the voucher and education systems in the mid 1990s. Hoxby (2003) reiterates that existing studies lead to inconclusive evidence of impact due to a lack of random assignment, thus making it difficult to determine whether variation in school choice is endogenous, and lack of pre-treatment data. This paper does not attempt to look at the impact of the reform in this light, but rather use the reform itself as an instrument to estimate the causal effect of education on earnings of the group affected by the 1981 education reform, as well as the effect of the reform on quality of learning.

Data

The International Adult Literacy Survey (IALS) was carried out in 20 countries[1] between 1994 and 1998, a project undertaken by the governments of the countries and three intergovernmental organizations[2]. It is a carefully designed, innovative survey of adult populations, and goes beyond just measuring literacy capabilities to assessing how these capabilities are applied to everyday activities. The IALS was followed by an extensive quality review (see Murray *et al*. 1998) which, after comparing the distribution of the characteristics of the actual and weighted samples, concluded that the actual and weighted samples were comparable to the overall populations of the IALS countries. The questionnaire also included questions about labor market status, earnings, education as well as demographic characteristics.

The data includes three scales as measures of literacy skills: prose literacy, document literacy and quantitative literacy, each in the 0-500 range. Prose literacy tests the understanding and use of texts such as editorials, news stories, fiction and poems. Document literacy tests skills required to locate and use information contained in a variety of formats, such as job applications, payroll forms, maps and tables. Quantitative literacy tests skills required in making calculations

[1] The countries are: Belgium, Canada, Chile, Czech Republic, Denmark, Finland, Germany, Hungary, Ireland, Italy, the Netherlands, New Zealand, Norway, Poland, Portugal, Slovenia, Sweden, Switzerland, the United Kingdom and the United States.
[2] OECD, European Union and UNESCO.

after locating numbers embedded in printed materials; examples of such calculations include determining the interest on a loan, calculating a tip and balancing a checkbook.

For Chile, the survey is representative of 98 percent of the population between the ages of 16 and 65 (it excludes residents of institutions and remote areas) and the total number of respondents was 3,583. A four-stage stratified sample was used and stratification was performed according to region and type (urban/rural). As is the case for the rest of the IALS country data, in the Chilean data the three skills are very highly correlated. Consequently, the average of the three scores will be used in the analysis as an aggregate IALS score measure (see also Blau and Khan 2001; Devroye and Freeman 2001; Leuven *et al*. 2004).

As is the case with the other countries in the IALS dataset, for Chile as well, the relation between skill level and years of schooling is positive. However, the slope of the skill-schooling profile is less steep compared to all but three other countries – Germany, the Netherlands and Sweden – while the slope is steepest in the Czech Republic, the United States, Slovenia and Canada (see Leuven *et al*. 2004).

Methodology

In the basic Mincerian human-capital model (Mincer 1974), schooling is assumed to be independent of ability/skills, and that the return from schooling investments is equal for all individuals. However, in the contemporary literature it is acknowledged that the return to schooling must be different for different skill levels. Intuitively, an estimate of the average return to schooling probably over-estimates the return for low skill individuals and under-estimates the return to high skill individuals. One should, therefore, allow skill level to affect the rate of return to schooling investments.

We incorporate the literacy score as a measure of cognitive skills to proxy for unobserved effects. We expect that the inclusion of a direct measure of cognitive skill will reduce the estimated schooling coefficient (more so for the less skilled), so that the coefficient on education then captures the effect of education alone having controlled for cognitive skills.

The effect of introducing ability/skills differences is two-pronged. First, the more able individuals may be able to 'convert' schooling into human capital more efficiently than the less able, and this raises the return to schooling for the more able. In this case, one can conclude that ability and education are complementing each other in producing human capital. On the other hand, the more able may have higher opportunity costs since they may have been able to earn more in the labor market, if ability to progress in school is positively correlated with the ability to earn, and this reduces the rate of return to schooling (Harmon and Walker 2000). Given a distribution of wages, having controlled for cognitive skills, we assume that this distribution reflects the distribution of inherent (non-cognitive) unobserved ability. As a result, lower ability individuals predominate in the lower quantiles of the distribution and higher ability individuals predominate in the upper quantiles of the distribution.

The model in which both the measure of ability and its interaction with schooling affect earnings and the ability-specific return to schooling is outlined below (Griliches 1977; Nordin 2005):

$$Ln\ w_i = \alpha + \beta S_i + \gamma A_i + \varepsilon_i \qquad (1)$$

where w is the hourly wage rate, S is years of schooling completed, A is ability and ε is an independently distributed error term. Allowing the return to schooling to depend on ability:

$$Ln\ w_i = \alpha + \beta(f(A_i))S_i + \gamma A_i + \varepsilon_i \qquad (2)$$

Using the test score as a proxy for cognitive ability:

$$Ln\ w_i = \alpha + \beta(f(T_i))S_i + \gamma T_i + \varepsilon_i \qquad (3)$$

where T is the IALS test score, which is assumed to perfectly measure cognitive ability. Assuming a linear relationship between test score and cognitive ability:

$$f(T_i) = t_0 + t_1 T_i \qquad (4)$$

we obtain the following earnings equation:

$$Ln\ w_i = \alpha + t_0 S_i + t_1 T_i S_i + \gamma T_i + \varepsilon_i \qquad (5)$$

Finally, including experience, its square and other covariates:

$$Ln\ w_i = \alpha + t_0 S_i + t_1 T_i S_i + \gamma T_i + exp + exp^2 + X_i + \varepsilon_i \qquad (6)$$

where X is a vector of other covariates.

The coefficient t_0 could be interpreted as the return to schooling for an individual of mean ability. The coefficient γ measures the approximate percentage change in the hourly wage arising from a one standard deviation increase in the score for an individual with no schooling.[3]

Equation (6) is estimated using both Ordinary Least Squares (OLS) to obtain estimates of average return to schooling, as well as IV to estimate the causal effect of education and skills applicable to a group of compliers associated with the 1981 reform.

Results

The group affected by the education reform, while consisting of students with a variety of backgrounds, is expected to contain a large proportion of students from higher socio-economic backgrounds and from urban areas. First, this is because in rural areas in Chile, low population density does not permit a choice of schools. Second, there is evidence that private-subsidized schools tend to select students to a larger extent than do municipal schools (Contreras, Bustos and Sepulveda 2007). Therefore, better private-subsidized schools (as well as some elite municipal schools) facing excess demand practiced screening. According to Gauri (1998), 28 percent of students in the subsidized sector of Santiago had taken tests to be admitted in their current schools. Furthermore, top-up fees, lack of transparency in enrollment procedures, and the cost of uniforms constituted *de facto* screening devices. As a result, students who attend private-subsidized schools come from higher-income and better educated families compared to municipal schools (but not private non-subsidized schools).

[3] For an individual with mean years of schooling, the impact is measured by $\gamma + t_1$*(mean years of schooling).

We estimate the returns to schooling for those induced to make schooling decisions (involving quantity or quality) by the change associated with our instrument - the 1981 reform. In this study, the instrument corresponds to a policy of capitation grants which induced a group of individuals to opt for the opportunities that the policy made available. The estimates, therefore, identify a policy relevant return to schooling in Chile.

A Note on Identification

Identification is based on variations across cohorts, as in several other empirical investigations. However, unlike some studies (for example, those by Oreopoulos 2006; Duflo 2001) which used a unique set of data or a suitable panel), in our case we only have one cross-section. Furthermore, ours being a specialized dataset, we lack certain variables which may have allowed for an alternative identification strategy, such as information at the community level. Not being able to identify (unrestricted) age effects separately from cohort effects, as they are collinear in one cross-section, one concern that might arise is that in reality we are using individuals at different points of the life cycle to identify different cohorts and, hence, different ages; age-earnings profiles might vary across age differently by education. This might potentially confound the effect of the instrument with such different profiles. However, we have experience in the regressions, parametrically restricted (that is, quadratic) to still leave room for identification. It is worth emphasizing that we are conditioning on experience profiles and so we identify using variations across cohorts, or around, these experience profiles. It is also worth mentioning that, unlike most other interventions in the education market used in the literature for IV estimation of returns to schooling, which resulted in rather weak instruments, in our case we use a sweeping education reform which results in a binary instrument which is strong enough to allow estimation of IV coefficients with precision.

What is the prior expectation of the magnitude of estimates and in general the nature of the results? Instrumental variables estimates of the return to schooling are expected to differ from OLS estimates. While the standard ability bias would suggest that the OLS estimates are biased upwards, empirical evidence from using a variety of, mainly binary, instruments, several of which are based on education policy reforms, suggest that IV estimates are generally higher that OLS estimates. The dominant explanation for this (as given by Card 2001) is that because of heterogeneity, there is a distribution of returns, and OLS and IV estimates correspond to different weighted averages of this distribution. Therefore, IV estimates can exceed OLS estimates. The fact that the IV estimates are higher than the OLS estimates is interpreted as an indication that the return to the marginal person (the "switchers") is higher than that of the average person. Furthermore, Carneiro, Heckman and Vytlacil (2005) show that the marginal person can have a return that is substantially lower than the return to the average person, and still the return estimate from IV is greater than the corresponding return from OLS.

Assuming that cognitive skills are associated with quality of schooling, and a group of students switched to better quality schools as a result of the reform, we would expect the affected group (especially those who switched to better quality schools at a younger age) to have higher and less dispersed cognitive skills. For example, from the IALS dataset, in some countries (Scandinavian countries as well as the Czech Republic) where we observe high IALS scores and their dispersion is low, the increase in earnings in response to one standard deviation increase in scores is generally low, compared to Chile, which has the lowest scores among the IALS group of countries (see Table A1 in the appendix). Therefore, the returns to cognitive skill scores in the IV regressions are expected to be lower compared to the OLS regressions.

The results from OLS regressions for returns to an additional year of schooling and the effects of controlling for the measure of cognitive skills, using a sample of 22-45 years old group of males working for wages, are presented in Table A2. The dependent variable is the logarithm of hourly wage. Inclusion of the direct measure of cognitive ability reduces the return to schooling by about 34 percent, while a one standard deviation increase in the score increases earnings by 17 percent. On average, the interaction effect of schooling with cognitive skills is moderately positive and nearly significant.

The results from IV regressions using the instrument based on the capitation grants reform are presented in Table 1. The binary instrument initially takes the value of 1 for those who at the time of the reform where 6 years or older (22 years or older in 1997 – the year of the survey) and up to 13 years of age included (29 years old in 1997) and 0 otherwise, that is those in the 8 years of basic education. One-third of the observations belong to the group affected by the reform. Two alternative sample specifications were used: 22-65 and 22-45 years of age. Quantitatively, in both cases the estimates were very similar. However, for the 22-45 group the results were more precise; therefore, we present these results. In the first stage, the correlation of years of schooling with the instrument is strong (Shea's partial R^2 is high), indicating that the instrument is sufficiently relevant to explain the endogenous regressor), while the instrument is uncorrelated with the logarithm of hourly wage.

Table 1: Returns to Schooling and Skills from IV
(male employees 22-45 years; instrument = 1 if age in 1981 between 6 to 13 years)

Variable	(1)	(2)	(3)
Years of schooling	0.122	0.102	0.096
	(5.1)	(2.7)	(2.5)
Experience	0.001	0.009	0.015
	(0.5)	(0.5)	(0.8)
Experience squared	0.0002	0.0002	-0.0000
	(0.3)	(0.4)	(0.1)
Standardized IALS score	-	0.086	0.009
		(1.0)	(0.1)
Stand. IALS score-years of schooling interaction	-	-	0.011
			(1.0)
Constant	4.98	5.19	5.19
	(13.3)	(10.2)	(10.5)
Centered R^2	0.144	0.167	0.174
N	586	586	586
First Stage:			
Shea's partial R^2	0.261	0.139	0.130
F-value	205.5	93.6	86.5
[p-value]	[0.000]	[0.000]	[0.000]
Endogeneity Test:			
Durbin-Wu-Hausman statistic	3.95	2.09	1.63
[p-value]	[0.047]	[0.148]	[0.201]

Note: Theoretical sampling weights included; *t*-values in parentheses.

In the standard Mincerian specification, the IV estimate of the return to schooling is about 37 percent higher that the OLS estimate, and the assumption that the schooling variable is exogenous is rejected at the 5 percent level. Here one needs to consider the composition of the group affected by the reform. This group (which in this case consists of pupils in basic education), contains those who switched to private subsidized schools. While there is no conclusive evidence in the literature that school choice improved the performance of the median student, we assume here, based on the evidence where it has been argued convincingly that the main effect of unrestricted school choice was an exodus of "middle class" students from public sector schooling (Hsieh and Urquiola 2006) and the practice of screening by private subsidized schools (competing for better students). Furthermore, it is possible that parents tended to select schools which provided good peer groups. Based on this evidence, we attribute the higher IV estimate of the return to schooling to those in a treated group with such characteristics, rather than the average individual.

In column (2) of Table 1, we control for cognitive skills. Without controlling for cognitive ability, cognitive ability will be part of the error term. The assumption of independence implies that the instrument must be independent of cognitive ability. Therefore, when cognitive ability is not controlled for, the instrument is compromised. By controlling for cognitive ability we avoid the problem of using possibly invalid instruments and make our estimates more credible (Carneiro, Heckman and Vytlacil 2005). Also, we know that cognitive ability – whether as a result of school quality or other factors – has an impact on labor market outcomes (see, for example, Behrman et al. 2008; Psacharopoulos and Velez 1993, 1992; Hause 1972, 1971).

When we include the measure of cognitive skills in the equation, the estimate of the return to schooling decreases by about 16 percent (which is much less than the reduction one sees in the OLS regressions), while the contribution of one standard deviation in the cognitive score to earnings is one-half the corresponding estimate from the OLS regressions (8.5 compared to 17 percent). We assume, based on the results presented above, that students who switched to private subsidized schools and in 1981 were of basic school age, were part of a rather homogeneous group of students with above average cognitive skills. That is, "better" students formed the bulk of those who switched from public to private subsidized schools. As a result, only a small part of return to schooling estimate is due to classical ability bias. Column (3) indicates that the contribution of cognitive skills is mainly through their interaction with schooling (as was the case in the OLS regressions).

Another instrument (in addition to the policy reform based instrument) is included in the specification presented in Table A3, namely if the pupil's father has at least upper-secondary education. Here the objective is twofold: first, test the over-identifying restrictions and, second, test the hypothesis that if the treated group is an even more privileged one (have taken advantage of the capitation grant system *and* have educated fathers), its cognitive skills will be better and more homogeneous compared to the treated group in Table 1 (which in turn had better and more homogeneous cognitive skills compared to the average individual). Therefore, the contribution of cognitive skills to earnings should be small, while returns to schooling should increase.

Based on the value of Hansen's J-statistic (see Hansen 1982), the null hypothesis (that the instruments are appropriately uncorrelated with the disturbance process) is not rejected (in all cases, that is, columns (1), (2) and (3), p-values are 0.15 or higher). As hypothesized, the contribution of cognitive skills to earnings is nearly zero (column 2). At the same time, the estimate of the return to schooling increases by between 7-19 percent compared to Table 1.

In Table 2 the binary instrument is modified and takes the value of 1 for those who at the time of the reform where 6 years or older (22 years or older in 1997 – the year of the survey) and up to 18 years of age included (34 years old in 1997) and 0 otherwise; that is the group includes all those who could have been affected by the reform, either by switching to private subsidized schools because of their perceived higher quality compared to municipal schools, or by choosing to continue schooling at the secondary level as a result of the increase in the supply of private schools associated with the education reform, which led to a large increase in secondary school participation. Estimates of the return to schooling are now lower by 37 percent in the Mincerian specification and by more than 50 percent when we control for cognitive skills. At the same time, one standard deviation increase in the cognitive score increases earnings by 20 percent, up from 8.5 percent in Table 1. The coefficient estimates are now close to those from OLS; as a result the endogeneity test does not reject the null hypothesis that the difference in coefficients between IV and OLS are not systematic.

Table 2: Returns to Schooling and Skills from IV
(male employees 22-45 years; instrument = 1 if age in 1981 between 6 to 18 years)

Variable	(1)	(2)	(3)
Years of schooling	0.077	0.043	0.042
	(3.2)	(1.3)	(1.3)
Experience	0.003	0.002	0.009
	(0.2)	(0.1)	(0.4)
Experience squared	0.0002	0.0002	-0.0001
	(0.3)	(0.3)	(0.3)
Standardized IALS score	-	0.204	0.051
		(2.9)	(0.5)
Stand. IALS score-years of schooling interaction	-	-	0.015
			(1.4)
Constant	5.66	5.94	5.84
	(14.1)	(12.7)	(12.7)
Centered R^2	0.158	0.183	0.188
N	586	586	586
First Stage:			
Shea's partial R^2	0.300	0.231	0.233
F-value	249.5	174.4	176.1
[p-value]	[0.000]	[0.000]	[0.000]
Endogeneity Test:			
Durbin-Wu-Hausman statistic	0.64	0.45	0.33
[p-value]	[0.42]	[0.50]	[0.56]

Note: Theoretical sampling weights included; *t*-values in parentheses

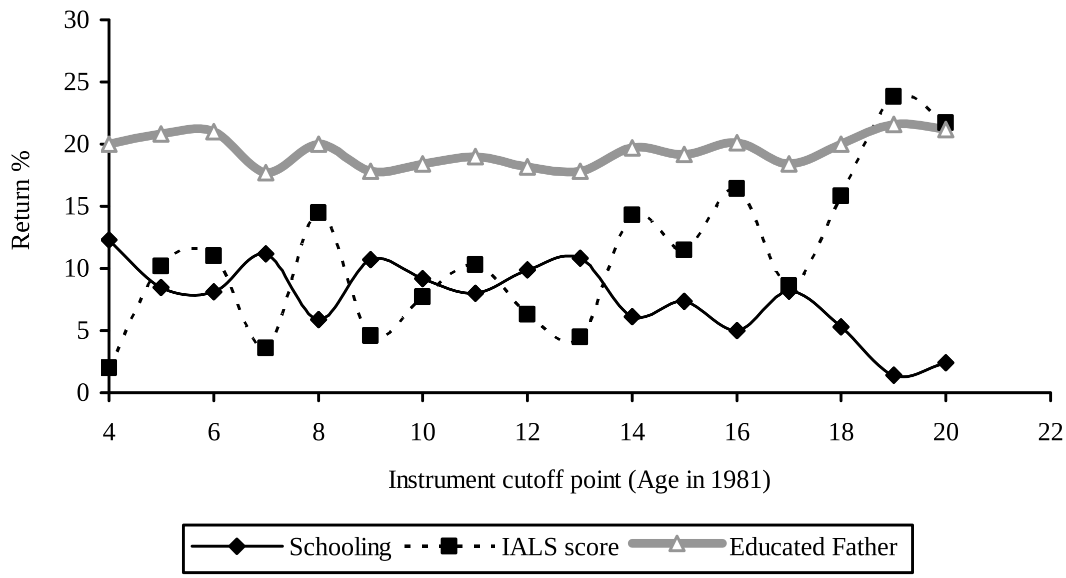
In Table A3a in the appendix we use the expanded age group (ages 6-18 in 1981) with the additional instrument (father at least secondary school). Here again, the effect of including the secondary school age students results in lower estimates for the return to schooling and higher estimates for the return to cognitive score compared to the results in Table 2. However, the return to schooling estimate is still higher and the return to cognitive score is lower than in

Table 2, given the composition of the treated group which consists of students with educated fathers. Once again based on the value of Hansen's J-statistic, the null hypothesis (that the instruments are appropriately uncorrelated with the disturbance process) is not rejected (p-values are between 7 and 19 percent).

Finally, we ask the question: what happens if we progressively expand the treated group to include students who, at the time of the reform were between the ages of 4 (20 years old in 1997) and 20 (36 years old in 1997)? That is, starting with a treated group of very young children in 1981, and progressively expanding the group one year at a time until age 20, tracing the changes coefficient estimates for schooling and cognitive skills. Suppose that such students, before deciding whether to switch schools (based on the comparison between expected benefits and costs) were enrolled in, say, municipal schools. Assuming that municipal schools on average provide lower quality schooling, switching to private-subsidized schools at a younger age (such as the beginning of primary school) should result in a higher and more homogeneous cognitive skill endowment by the end of the schooling period, compared to switching at a later age (such as during high school). We, therefore, would expect that as we progressively increase the instrument upper cutoff age, the treated group becomes less endowed in cognitive skills (because they have received most of their schooling before 1981) and less homogeneous in these skills. As a result, as we add the marginal individual in the treated group, the return to cognitive skills will be increasing and the return to schooling will be decreasing. This is because part of a given estimate of the return to schooling from the Mincerian specification is due to the independent effect of cognitive skills. Given a group with less and more dispersed cognitive skills, a larger part of the Mincerian estimate of the return to schooling will be due the independent effect of cognitive skills. The opposite is true for a group endowed with more and less dispersed cognitive skills).

Chart 2 confirms this hypothesis. Using the Mincerian specification, the return is around 11-12 percent for the cohorts who in 1991 were of primary school age or lower and subsequently (by the end of basic education) declines sharply. With the addition of IALS scores, the return to schooling fluctuates at about 10 percent for the primary school cohorts and subsequently declines sharply; the opposite pattern is observed for the return to the cognitive score. With the last expansions of the instrument cutoff point, the return to schooling approaches zero, while one standard deviation increase in the cognitive score increases earnings by close to 30 percent. Finally, with the addition of the educated father dummy and IALS scores, the independent effect of having an educated father hovers at about a 20 percent increase in earnings (a little less for the primary school cohorts), while the magnitude and pattern of the returns to schooling and cognitive score are similar to the case where we only added the scores. One can also observe that the decline in the return to schooling estimate (increase in the return to cognitive score) coincides with the period between the end of the two cycles of compulsory basic education in Chile (ages 6-14) and the end of secondary education (ages 14-18).

Chart 2: Returns to one additional year of schooling, one standard deviation of IALS score and having an educated father for different instrument cutoff points



Summarizing, the findings from Tables 1 and 2 suggest that for a group of compliers who were of basic school age in 1981, about 84 percent (schooling coefficient in column 2 divided by schooling coefficient in column 1) of what the labor market rewarded was the non-cognitive contribution of schooling. On the other hand, for the mixed group of compliers which includes those who were of secondary school age, only about 56 percent of the labor market reward was for the non-cognitive contribution of schooling.

Note, however, that when a measure of full cognitive skills is used, one cannot account for the origin of cognitive skills. For example the IALS score used in this paper measures an array of cognitive skills acquired by an individual, including skills acquired outside of school (mainly influenced by the parental environment). Therefore our earlier methodology assumes that all cognitive skills are acquired (or signaled) via schooling.

In what follows, we explore an alternative methodology in obtaining additional information about the relative sizes of the non-cognitive vs. cognitive components of the return to schooling. Following Ishikawa and Ryan (2002) and Pasche (2007) (useful insight can also be found in Tyler (2004)), consider the following theoretical model:

$$W_i = a_0 + a_1 S_i + a_2 SCS_i + a_3 NSCS_i + a_4 NSNCS_i + a_5 F_i + a_6 A_i + a_7 X_i + \varepsilon \qquad (7)$$

In this model, the coefficient a_1 is a measure of the pure return to schooling non-cognitive skills. Note that measures of innate ability and non-schooling non-cognitive skills are not usually available in datasets, including the IALS. If a measure of schooling cognitive skills (SCS) was available, the following equation could have been estimated:

$$W_i = \beta_0 + \beta_1 S_i + \beta_2 SCS_i + \beta 3 F_i + \beta 4 X_i + \varepsilon \qquad (8)$$

where *W* stands for earnings, *S* for years of schooling, *SCS* for schooling cognitive skills, *NSCS* for non-schooling cognitive skills, *NSNCS* for non-schooling non-cognitive skills, *F* for family background variables, *A* for innate ability, while *X* is a vector of other controls (in our case years of potential labor market experience and its square). The coefficient β_1 would be a true measure of the non-cognitive return to schooling. The full measure of cognitive skills (the IALS score in our case) is regressed on quantity (*S*) and quality of schooling (*SQ*), that is years of schooling and type of schooling (that is, high school, university, etc.), as well as family background information:

$$IALS_i = \gamma_0 + \gamma_1 S_i + \gamma_2 SQ_i + \gamma_3 F_i + \varepsilon \qquad (9)$$

The measure of schooling cognitive skills is obtained from equation (9) using the schooling coefficients as follows:

$$SCS = \gamma_1 S_i + \gamma_2 SQ_i \qquad (10)$$

then:

$$NSCS_i = IALS_i - SCS_i \qquad (11)$$

The equation to be estimated now becomes:

$$W_i = \delta_0 + \delta_1 S_i + \delta_2 SCS_i + \delta_3 NSCS_i + \delta_4 F_i + \delta_5 Xi + \varepsilon \qquad (12)$$

If one goes as far as assuming that the literacy score captures all of the productivity enhancing skills acquired through schooling (along with skills acquired elsewhere), the estimated coefficient δ_1 would represent the returns to the market signaling value of an additional year of schooling (after controlling for schooling and non-schooling cognitive skills), while δ_2 and δ_3 would represent respectively the returns to those skills acquired at school and elsewhere. However, it would probably be more realistic to assume that the literacy score does not capture *all* the basic skills acquired in school; therefore, *S* contains those productivity enhancing skills acquired in school and are not captured in the literacy score (see also Tyler 2004). The estimated coefficient δ_1 will then represent a mixture of returns to those skills acquired in school and not captured by the literacy score and returns to the signaling value of schooling.

In estimating equation 10, one is likely to encounter multicollinearity problems due to the significant correlation between years of schooling and the derived schooling cognitive skills variable. Pasche (2007), using Swiss data, estimated equation 10 while encountering only moderate multicollinearity problems. In our case, however, the problem is severe. Variance inflator factors (VIF) for the two collinear variables are in excess of 50, rendering the estimated coefficients of years of schooling and *SCS* uninterpretable.

We will now attempt to overcome the multicollinearity problem by modifying the methodology presented in equations 7-12, in particular instead of deriving an estimate of schooling cognitive skills using the significant schooling coefficients, we will derive an estimate of non-schooling cognitive skills instead. For this, we explore a suitable set of variables in the dataset to be used as non-schooling additional controls in equation 9. The set of non-schooling controls identified includes: fathers' and mothers' education, reading books at home frequently, urban residence and age, of which the first four enter with a statistically significant coefficient. Using the derived schooling and non-schooling cognitive skills in the earnings function confirms that there is no evidence of significant multicollinearity (apart from what arises of including experience and its square in the regression – which is unavoidable). The Variance Inflator Factors (VIF) are now between 2 and 3 (see Table A4 in the appendix, column 4) and the mean VIF is well below 10.

From Table 3, the average marginal return to schooling from the standard specification is 9 percent. There is a significant decrease of about 25 percent in the years of schooling coefficient once we control for full cognitive skills. After controlling for both types of cognitive skill (column 4), both schooling and non-schooling cognitive skills are responsible for the overall decrease in the coefficient of years of schooling, with schooling cognitive skills being more important (schooling cognitive skills are rewarded approximately 1.5 times more than non-schooling cognitive skills).

Table 3: Returns to Schooling from OLS (male employees 22-45 years)

Variable	(1)	(2)	(3)	(4)
Years of schooling	0.090	0.068	0.076	0.069
	(8.0)	(4.9)	(5.5)	(4.7)
Experience	0.012	0.011	0.012	0.008
	(0.8)	(0.7)	(0.8)	(0.5)
Experience squared	-0.0002	-0.0002	-0.0003	-0.0002
	(0.6)	(0.4)	(0.6)	(0.4)
Standardized IALS	-	0.128	-	-
		(2.7)		
Standardized SCS	-	-	0.111	0.111
			(2.3)	(2.3)
Standardized NSCS	-	-	-	0.075
				(1.4)
Father >= Post-Secondary	0.236	0.194	0.251	0.127
	(2.8)	(2.3)	(2.9)	(0.9)
Constant	5.34	5.56	5.48	5.57
	(26.6)	(25.8)	(25.1)	(24.1)
Adj. R^2	0.197	0.207	0.228	0.228
N	490	490	467	467

Note: Theoretical sampling weights included; *t*-values in parentheses

In Tables 4 and 5 we use the specifications in columns (1)-(4) of Table 3 and estimate IV regressions using the binary instrument involving a group of compliers who were of basic school age in 1981. The first two columns are the same as in Table 1 and 2 with the addition of father's education in the controls. The third column replaces full cognitive skills with *SCS* while the fourth includes both *SCS* and *NSCS*. The findings in Table 4 confirm the earlier conclusion that for this group of compliers, the coefficient of years of schooling is mainly a measure of the non-cognitive return to schooling. The coefficient remains largely unchanged when we control separately for schooling cognitive skills and non-schooling cognitive skills which enter the equation with coefficients which are statistically near zero.

Table 4: Returns to Schooling and Skills from IV – Male Employees 22-45 years
(Instrument=1 if age in 1981 between 6 and 13 years)

Variable	(1)	(2)	(3)	(4)
Years of schooling	0.117	0.108	0.117	0.117
	(5.5)	(3.2)	(3.5)	(3.0)
Experience	0.019	0.018	0.019	0.019
	(1.2)	(1.1)	(1.2)	(1.1)
Experience squared	-0.0002	-0.0002	-0.0003	-0.0003
	(0.5)	(0.5)	(0.6)	(0.6)
Standardized IALS	-	0.045	-	-
		(0.6)		
Standardized SCS	-	-	0.029	0.029
			(0.4)	(0.4)
Standardized NSCS	-	-	-	0.003
				(0.0)
Father >= Post-Secondary	0.191	0.178	0.205	0.201
	(2.1)	(2.0)	(2.2)	(1.4)
Constant	4.95	5.05	4.95	4.95
	(15.0)	(11.2)	(10.9)	(9.4)
Centered R^2	0.194	0.202	0.221	0.221
N	490	490	467	467
First Stage:				
Shea's partial R^2	0.274	0.165	0.168	0.145
F-value	(182.7)	(95.7)	(93.0)	(78.1)
[p-value]	[0.000]	[0.000]	[0.000]	[0.000]
Endogeneity Test:				
Durbin-Wu-Hausman statistic	2.22	1.65	1.81	1.75
[p-value]	[0.137]	[0.200]	[0.179]	[0.186]

Note: Theoretical sampling weights included; *t*-values in parentheses

The results in Table 4, for the group of compliers of basic or secondary school age in 1981 are again drastically different. Overall, the IV estimates of the coefficient of years of schooling are now lower compared to the corresponding OLS estimates (Table 2). Controlling for the two types of cognitive skills decreases the coefficient of years of schooling by about 27 percent compared to the standard specification in column 1. At the same time, one standard deviation increase in the score increases earnings by about 13 and 9 percent for schooling cognitive and non-schooling cognitive skills.

Table 5: Returns to Schooling and Skills from IV
(male employees 22-45 years; instrument = 1 if age in 1981 between 6 to 18 years)

Variable	(1)	(2)	(3)	(4)
Years of schooling	0.079	0.053	0.064	0.058
	(3.9)	(1.9)	(2.3)	(1.9)
Experience	0.009	0.008	0.010	0.006
	(0.6)	(0.5)	(0.6)	(0.4)
Experience squared	-0.0002	-0.0002	-0.0003	-0.0001
	(0.6)	(0.4)	(0.6)	(0.3)
Standardized IALS	-	0.158	-	-
		(2.3)		
Standardized SCS	-	-	0.132	0.131
			(2.0)	(2.0)
Standardized NSCS	-	-	-	0.092
				(1.3)
Father >= Post-Secondary	0.256	0.200	0.264	0.110
	(2.8)	(2.3)	(2.9)	(0.8)
Constant	5.52	5.75	5.63	5.72
	(17.6)	(15.1)	(14.8)	(13.6)
Centered R^2	0.202	0.214	0.235	0.237
N	490	490	467	467
First Stage:				
Shea's partial R^2	0.305	0.239	0.244	0.231
F-value	(213.2)	(152.0)	148.6	(137.8)
[p-value]	[0.000]	[0.000]	[0.000]	[0.000]
Endogeneity Test:				
Durbin-Wu-Hausman statistic	0.507	0.348	0.219	0.170
[p-value]	[0.478]	[0.555]	[0.640]	[0.680]

Note: Theoretical sampling weights included; *t*-values in parentheses

The evidence could be taken to suggest that the treated group which benefited most from the 1981 “voucher” reform, besides coming mainly from a better socioeconomic background and from urban areas, were enrolled in the early stages of schooling. They benefited from better quality schooling which endowed them with better schooling cognitive skills (compared to the untreated). As a result, given the “abundance” of schooling cognitive skills in this group, the labor market reward for an additional year of schooling reflects the “true” non-cognitive return to schooling.

However, the cohorts who obtained early childhood schooling before 1981 do not seem to have benefited as much. The reform facilitated the absorption of the large secondary school expansion, despite the decline in public spending for education. However, because of the secondary school expansion, lower ability students (from “middle class” families or otherwise) may have been admitted into the system, students who had earlier received their basic education in public schools. This group entered the labor market less endowed in schooling cognitive skills. As a result, the labor marker rewarded these skills significantly; failure to control for such skills in the earnings function results in a return to schooling estimate which is seriously biased upwards. The non-cognitive component of the return to schooling for this group is rather small (about half of that of the younger cohort).

The IV regressions using a treated group which consists of a mixture of basic and secondary age students in 1981, result in estimates of gross returns (column 1) as well as schooling returns with the cognitive skills component netted out (column 4) which are smaller that the corresponding OLS estimates, at about 8 and 6 percent respectively. On the other hand, comparing the results from using a treated group consisting of only basic school age students to those using a mixture of basic and secondary age students in 1981, reveals considerable heterogeneity in the returns to schooling in Chile.

Conclusions

We use the sweeping school reform of 1981 in Chile to identify a binary instrument and estimate returns to schooling from IV, with and without controlling for cognitive ability. Given that ability bias needs to be dealt with, accounting for cognitive skills avoids the problem of using a possibly invalid instrument.

The 1981 capitation grant scheme facilitated the absorption of the pressure of secondary school expansion which resulted from the capitation grants paid to schools and the opening of the education system to the private sector, and resulted in large increases in student intake as well as a sharp decline in the dropout rate, especially in basic education. Overall, the reform improved the efficiency of the education system, possibly at the expense of equity.

The results suggest that those who in 1981 switched to private subsidized schools were mainly urban “middle class” students leaving public schools, possibly in search of better peer groups. We find evidence suggesting that for students with such characteristics, who switched during their basic education, the labor market reward for an additional year of schooling is a measure of the non-cognitive return to schooling. This is not the case, however, for older cohorts (those in secondary education). The large secondary school expansion seems to have attracted a heterogeneous group of students who had earlier received their basic education in public schools. For this group of students the labor marker rewarded these skills significantly, while the non-cognitive component of the return to schooling for this group is rather small.

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Appendix

Table A1: Country Sample Means: Years of Schooling and IALS Score

Country	Mean Years of Schooling	IALS/10	Effect of IALS score on earnings (rank)
Canada	13.1 (3.5)	2.92 (0.60)	7
Chile	9.6 (4.0)	2.15 (0.58)	1
Czech Republic	13.0 (2.8)	2.89 (0.47)	11
Denmark	13.2 (3.3)	2.96 (0.40)	13
Finland	12.8 (3.5)	2.99 (0.43)	12
Germany	11.8 (3.5)	2.92 (0.47)	6
Hungary	12.3 (3.1)	2.63 (0.48)	2
Italy	11.1 (4.1)	2.52 (0.62)	13
Netherlands	13.4 (4.2)	2.94 (0.43)	5
Norway	12.1 (2.8)	3.00 (0.41)	10
Poland	11.7 (3.0)	2.43 (0.64)	15
Slovenia	11.5 (3.0)	2.40 (0.61)	9
Sweden	11.8 (3.7)	3.12 (0.51)	8
Switzerland	12.9 (3.4)	2.83 (0.54)	4
U.S.A.	13.9 (3.2)	2.85 (0.65)	3

Source: Leuven *et al.* 2004

Note: Standard deviations in parentheses

Table A2: Returns to Schooling from OLS – Male Employees 22-45 years

Variable	(1)	(2)	(3)	(4)
Years of schooling	0.089	0.058	0.055	0.068
	(8.9)	(4.8)	(4.5)	(4.9)
Experience	0.000	0.001	0.011	0.011
	(0.0)	(0.1)	(0.7)	(0.7)
Experience squared	0.0002	0.0002	-0.0001	-0.0002
	(0.4)	(0.5)	(0.3)	(0.4)
Standardized IALS score	-	0.174	0.037	0.128
		(4.2)	(0.4)	(2.7)
Stand. IALS score-Years of schooling interaction	-	-	0.014	-
			(1.9)	
Father at least Upper Sec. education	-	-	-	0.194
				(2.3)
Constant	5.48	5.75	5.69	5.56
	(30.0)	(30.0)	(29.2)	(25.8)
Adj. R^2	0.156	0.179	0.182	0.207
N	586	586	586	490

Note: Theoretical sampling weights included; *t*-values in parentheses

Table A3: Returns to Schooling from IV
(male employees 22-45 years; instrument = 1 if age in 1981 between 6 to 13 years)

Variable	IV: Reform + Educated father	IV: Reform + Educated father	IV: Reform + Educated father
Years of schooling	0.130	0.120	0.116
	(5.5)	(3.1)	(2.9)
Experience	0.025	0.024	0.029
	(1.2)	(1.2)	(1.3)
Experience squared	-0.0004	-0.0004	-0.0006
	(0.7)	(0.8)	(0.9)
Standardized IALS score	-	0.028	-0.022
		(0.3)	(0.2)
Stand. IALS score-years of schooling interaction		-	0.006
			(0.5)
Constant	4.80	4.91	4.92
	(12.5)	(9.5)	(9.7)
Centered R^2	0.175	0.185	0.190
N	490	490	490
First Stage:			
Shea's partial R^2	0.308	0.169	0.160
F-value	108.1	49.0	46.2
[p-value]	[0.000]	[0.000]	[0.000]
Over-identification Test:			
Hansen's J-statistic	2.09	1.96	1.71
[p-value]	[0.148]	[0.161]	[0.191]
Endogeneity Test:			
Durbin-Wu-Hausman statistic [p-value]	5.38	2.53	1.93
	[0.020]	[0.111]	[0.164]

Note: Theoretical sampling weights included; *t*-values in parentheses

Table A3a: Returns to Schooling from IV
(male employees 22-45 years; instrument = 1 if age in 1981 between 6 to 18 years)

Variable	IV: Reform + Educated father	IV: Reform + Educated father	IV: Reform + Educated father
Years of schooling	0.099	0.067	0.063
	(3.8)	(1.9)	(1.7)
Experience	0.021	0.018	0.025
	(1.0)	(0.8)	(1.1)
Experience squared	-0.0005	-0.0005	-0.0007
	(1.0)	(0.9)	(1.1)
Standardized IALS score	-	0.142	0.043
		(1.8)	(0.4)
Stand. IALS score-years of schooling interaction	-	-	0.011
			(1.0)
Constant	5.23	5.57	5.55
	(12.3)	(10.7)	(10.7)
Centered R^2	0.190	0.205	0.210
N	490	490	490
First Stage:			
Shea's partial R^2	0.339	0.242	0.241
F-value	124.1	77.3	76.9
[p-value]	[0.000]	[0.000]	[0.000]
Over-identification Test:			
Hansen's J-statistic	3.29	2.15	1.70
[p-value]	[0.069]	[0.142]	[0.193]
Endogeneity Test:			
Durbin-Wu-Hausman statistic [p-value]	9.36	7.14	8.79
	[0.053]	[0.210]	[0.186]

Note: Theoretical sampling weights included; *t*-values in parentheses

Table A4: Variance Inflator Factors (VIF)
(equations in Table 3)

	(1)	(2)	(3)	(4)
Years of schooling	1.7	2.6	2.5	2.9
Experience	14.9	14.9	15.0	15.1
Experience^2	14.2	14.2	14.3	14.5
Standardized Score$_{IALS}$	-	2.1	-	-
Standardized SCS	-	-	2.0	2.4
Standardized NSCS	-	-	-	2.0
Father >= Upper Secondary	1.1	1.2	1.1	2.0
Mean VIF	8.0	7.	6.9	6.5