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Credit Constraints and the Demand for Higher Education in Latin America

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Latin American countries have seen in the past two decades a record number of students enrolling in tertiary education. However, the gap in tertiary enrollment rates between the highest and lowest income quartiles has also widened—something that could be attributed to tightening credit constraints. Our paper attempts to provide evidence on the extent of short-term credit constraints in college enrollment using data from four countries. In addition, we estimate whether the extent of short-term credit constrains has changed in the last two decades. Results suggest that most of the tertiary enrollment gap by income levels can be explained by long-run factors and not by current income. Thus, the existence of credit constraints cannot fully explain the postsecondary participation gaps by income levels. However, the size of the current income effect has increased for the more recent samples, which could be indicating that credit constraints are tightening as access to postsecondary education has increased.

JEL Classification: I22; I28

Keywords: postsecondary education enrollment; credit constraints; family background; Latin America.

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1. Introduction

In response to rising returns to schooling, increases in secondary completion rates, and a rapid expansion of the private sector, Latin American countries have experienced in the past two decades record numbers of students enrolling in tertiary education. As a result, tertiary education attendance rates in the region among 18-23 year olds have increased considerably (Figure 1). This increase, however, masks important inequalities in access to higher education. Figure 2 clearly shows that tertiary participation rates have increased mostly for high income students, resulting in a widening gap in enrollment rates between the highest and lowest income quartiles. This correlation between family income and tertiary participation rates is commonly interpreted as evidence that low income college-age youngsters are credit constrained and cannot finance their tertiary education (Carneiro & Heckman, 2002; Dearden, McGranaham & Sianesi, 2004; Keane, 2002; Leslie & Brinkman, 1987; McPherson & Schapiro, 1991). In this case, financial aid policies directed towards reducing the short-term liquidity constraints, such as grants, scholarships and student loans, could improve tertiary education participation among low-income youth.

Nonetheless, some economists have argued that the existence of short-term credit constraints is only one possible interpretation of the correlation between college attendance and family income. Another interpretation is that the participation gap is mostly determined by long term factors and not by short-term borrowing constraints. This line of argument suggests that inequality in college participation is "driven by unequal human capital accumulation prior to college-going age" (Keane, 2002, p. 293), and that families that have high income during college-going years are more likely to have had high income throughout the youngster's life cycle. Therefore, recent work has found that long-run determinants of family income are correlated with past academic achievement and with school continuation choices (Cameron & Taber, 2004; Carneiro & Heckman, 2002). In this case, the role of financial aid policies aimed at reducing college attendance inequalities between low income and high income youth could be limited.

Some studies have suggested that the second interpretation is more important. For example, Carneiro and Heckman (2002) using data from the 1979 National Longitudinal Survey of the Youth (NLSY) found that most of the family income gap in tertiary enrollment is due to long-term factors and that at most 8 percent of American youth see their tertiary participation affected because of short-term liquidity constraints. Keane and Wolpin (2001) developed a dynamic structural model of schooling behavior and estimated, using the 1979 NLSY, that tight borrowing limits have a negligible impact on final schooling outcomes and that low income youth adjust to tighter credit constraints by increasing their labor supply while in college. Stinebrickner and Stinebrickner (2007), using a unique longitudinal dataset, found that most of the college dropout decisions by low income students can be attributed to reasons other than credit constraints. Winter (2007) developed a computable dynamic general equilibrium model, and his model calibration resulted in enrollment gaps between high and low income students that are considerably reduced once long-term factors are controlled for. But he also found that up to 18 percent of households are borrowing-constrained even with narrow enrollment gaps. A study using a more recent cohort found a dramatic increase in the effect of family income on college attendance for the 1997 NLSY as compared to the 1979 NLSY (Belley & Lochner, 2008). The authors developed an educational choice model that incorporates borrowing constraints and a consumption value of schooling, and concluded that the data are consistent with the hypothesis that more youth are credit constrained in the early 2000s than were in the early 1980s. Lochner and Monge-Naranjo (2008) also interpreted the rising empirical importance of family income in determining college enrollment as consistent with increasingly binding credit constraints when tuition and returns to schooling have been on the rise. Similarly, a study using data from the United Kingdom found that credit constraints appear to have surfaced for individuals who were making their educational decisions in 1986 as compared to those who made their decisions in 1970, although the share of the more recent male cohort being borrowing constrained remained small—2 to 3 percent (Dearden et al., 2004).

Our paper attempts to provide evidence on the extent of short-term credit constraints in college enrollment using data from four different Latin American countries: Chile, Colombia, Mexico and Peru. In addition, we estimate whether the extent of short-term credit constrains has changed in the last two decades. To estimate the role of credit constraints in college attendance we use a direct approach as suggested in Carneiro and Heckman (2002), which implies estimating the effect of current income on college attendance conditioned on long-run factors such as parental education. Lastly, our study can inform policy makers of the potential role of financial aid policies in reducing the income-college attendance gap in the region. To our knowledge, there is no study that has previously attempted to measure the extent of short-term credit constraints in college enrollment in Latin America. A recent study, arguing that tertiary education enrollment rates in Latin America are low compared to high income countries, estimated the affordability of tertiary education in the region (Murakami & Blom, 2008). It found that Latin American families have to pay up to 60 percent of their per capita income for tertiary education per student, as compared to only 19 percent in high income countries, and concluded that the governments in the region could provide student assistance to make tertiary education more affordable. However, from this study one cannot infer that tertiary education enrollment in the region will necessarily increase in response to the availability of student assistance.

Our findings are consistent with Carneiro and Heckman's (2002) argument that most of the tertiary enrollment gap by income levels can be explained by long-run factors (parental education and household assets) rather than by current monetary income. As a result, the existence of credit constraints cannot fully explain the postsecondary participation gaps shown in Figure 2. Interestingly, the size of the current income effect has increased for the more recent samples—an indication that credit constraints may be slightly tightening as access to postsecondary education has increased. This paper is structured as follows. In the next section, we present a theoretical model that provides different interpretations for the correlation between family wealth and tertiary education enrollment. In section 3 we present the empirical model, while in section 4 we describe the data sources and the main variables used in the analysis. Results are discussed in section 5, and we conclude the paper with a summary of findings and implications for policy in section 6.

2. The Theoretical Model

To understand the role of credit constraints in tertiary education enrollment, we rely on a two-period discrete choice model of college attendance that is adapted from work by Cameron and Heckman (1998), Belley and Lochner (2008), Keane (2002), and Keane and Wolpin (2001). In period 1, the model assumes that agents decide whether to attend college and devote time s at direct cost t. If they do not attend college, they are assumed to be working and receiving a wage rate w_1. Individuals can borrow or save, with the net amount borrowed equaling *b.* In addition, they can receive a transfer payment *y* from their parents. Lastly, agents face a discount factor ρ. In period 2, individuals are assumed to receive a wage rate w_2 if they did not attend college and a rate $w_2+\beta$ if they attended college, and to make a fixed payment on the loan at an interest rate *r*. In both periods, agents receive utility from consumption *c*—with the utility function *u(c)* concave—and they also derive utility from college attendance ξ.

The value function for college attendance is given by:

$$(1) \qquad V_s = \max_{h,b} u(y_1 + (1-s)w_1 + b - t) + \xi + \frac{1}{\rho} u(w_2 + \beta - rb)$$

while the value function for those who do not attend college is:

$$(2) \qquad V_0 = \max_{h,b} u(y_1 + w_1 + b) + \frac{1}{\rho} u(w_2 - rb)$$

Then, an individual attends college if and only if:

$$(3) \qquad \frac{\beta}{r} + \frac{\xi}{\omega_1} > t + sw_1$$

where $\omega_1 = u_1(c_1)$ is the marginal utility of consumption in period 1. If there is no utility from school attendance (ξ=0) and agents do not face binding borrowing constraints, the parental transfer *y* does not affect the college attendance decision. This means that without borrowing constraints parental wealth should not affect college enrollment decisions.

If we add borrowing constraints to the model, and assume that the constraint binds and all students can only borrow φt, then the college attendance decision rule becomes:

$$(4) \qquad \frac{\beta}{r}\left[\frac{r\omega_2}{\rho\omega_1}\right] + \frac{\xi}{\omega_1} > \varphi t\left[\frac{r\omega_2}{\rho\omega_1}\right] + (1-\varphi)t + sw_1$$

If the borrowing constraint is binding, we have that $\omega_1 > r\omega_2/\rho$ and a positive relation between parental wealth and children's college attendance may emerge because higher parental transfers lower the effective cost of college.

The model implies several mechanisms through which a positive correlation between parental wealth (such as income and education) and college attendance can be generated:

i. If agents get utility from college enrollment ($\xi>0$,) the parental transfer y enters the decision rule though the marginal utility of consumption ω_1. Since increments in y reduce ω_1 and increase ξ/ω_1, then the decision rule implies that if wealthier parents make larger transfers their children are more likely to attend college *even in the absence of borrowing constraints*.

ii. If there is a binding borrowing constraint, $\omega_1 > r\omega_1/\rho$, and the constraint tightens (that is, φ declines), the effective direct costs of college increase and parental transfers can mitigate this effect by reducing ω_1 and thus affect college attendance decisions.

iii. If wealthier parents make larger college-attendance contingent transfers y_s, they reduce the direct costs of college attendance.

iv. Parental wealth may be positively related to children's skill endowments and increase their returns to schooling β.

Thus, one of the main implications of the model is that a positive correlation between college attendance and parental wealth (income *and* assets) may arise even when liquidity constraints have a negligible impact on tertiary enrollment rates.

3. Empirical Strategy

Our empirical strategy relies on the direct approach to test for short-term credit constraints proposed by Carneiro and Heckman (2002). They argue that observed differences in postsecondary enrollment between those in the top income quartile and those in the bottom quartile are due either to long-run family effects (i.e., parental education) or to short-run credit constraints (proxied by current income). Then, this approach suggests that credit constraints are any gap that remains in postsecondary enrollment by current income level after conditioning on long-run factors. Controlling for parental education and other family background measures, thus, is meant to purge current family income of its long-term family components that also affect educational attainment.

To implement this direct approach, we estimate the following logistic regression for the probability of postsecondary enrollment:

$$\text{(5)} \qquad \Pr(Y_{ijt}) = \gamma C_{ijt} + \beta X_{ijt} + \delta P_{ijt} + \psi O_{ijt} + \tau D_{ijt} + \varepsilon_{ijt}$$

where *i* indexes the individual, *j* the country and *t* the survey year.

The dependent variable *Y* is a dichotomous indicator for participation in postsecondary education (we expand on its computation below). Our explanatory variable is *C*, which measures current household income. We control for a vector of demographic and household characteristics *X*, for measures of permanent household income *P*, for the opportunity costs of schooling *O*, and for direct costs of schooling *D*.

The parameters of interest are γ, which captures the effect of current income on postsecondary enrollment, and δ, which measures the effect of long-run family factors on postsecondary enrollment. In addition, we are interested in the variation in these parameters by country and by year. Lastly, in order to compute the residual effect of current income, we estimate equation (5) sequentially by introducing each block of right-hand variables once at a time.

4. Data

We rely on cross-sectional national household surveys from four Latin American countries. We use Chile's Encuesta de Caracterización Socioeconómica Nacional (CASEN) for the years 2000 and 2006, Colombia's Encuesta de Condiciones de Vida (ECV) for 1993 and 2003, Mexico's Encuesta Nacional de Ingresos y Gastos del Hogar (ENIGH) for 1998 and 2006, and Peru's Encuesta Nacional de Hogares sobre Medición de Niveles de Vida (ENNIV) for 1997 and 2000. In addition, we have obtained information on the number of postsecondary institutions and on tertiary enrollment from the Ministries of Education of each country, either by perusing their websites or by specially requesting the data.

Our sample is composed of individuals who are 18 to 23 years old at the time of the survey. In addition, we restrict our sample to those youngsters who have not already completed tertiary education,[1] who are not household heads,[2] and who have complete information on all variables used in the analysis. The dependent variable is a dummy that equals 1 if the youngster indicates to be attending an educational institution *and* is enrolled at postsecondary levels (tertiary or university), and 0 if the if the youngster indicates to be attending an educational institution but enrolled at levels other than postsecondary *or* if the youngster does not attend an educational institution. The explanatory variable, current income *C*, is proxied using a set of dummy variables for the quartiles of total monetary household income.[3] Table 1 shows the percentage of 18-23 year olds who are enrolled in postsecondary education overall and by income quartile, and how this percentage has changed between survey years. Postsecondary participation rates varied between 12 (Mexico) and 24 percent (Chile) for the earliest set of surveys, and we can observe that they have increased considerably in each country in the last 5 to 10 years—with the exception of Peru where it remained fairly constant. In addition, participation rates have increased for all income quartiles and, with the exception of Chile, the gap between the highest and lowest quartiles has widened considerably.

Long-run factors that proxy permanent income *P* are: a set of dummies that indicate the educational level of the household head (no education to secondary education, with a head with postsecondary education as the reference category), a continuous variable that measures the number of different types of goods a household owns (such has cars, TVs, cell phones, etc.), a dummy that indicates type of home ownership (equaling 1 for home owners and 0 for other types of residential arrangements such as rentals or squatters), a set of dummies that indicate a household's access to basic utilities (whether the household is connected to the public water system and whether it is connected to the public electric grid), and a set

1 Although most 18-23 year olds have not had time to finish postsecondary education in the countries we study—since most college degrees are between four to five years in length—we introduce the restriction of not having already completed tertiary education to avoid categorizing as "attending college" an individual who is enrolled in graduate school. As expected, we lose very few observations by including this restriction.

2 Since we are working with cross-sectional datasets that do not ask about the educational level of the youngster's parents, we instead must rely on the educational level of the household head. Thus, in order to condition on the education of the household head we need to exclude from our sample those individuals who are already household heads at the time of the survey.

3 Total monetary household income includes income from subsidies, cash transfers, and pensions.

of dummies that indicate whether the building materials are "rustic" (for example, whether the floors are of dirt, the walls are of adobe, and the roof is of straw).

As described in the empirical section, we also control for background characteristics by including a continuous variable to measure age, a dummy for whether the individual is female, a dummy for minority status,[4] a dummy for whether the household head is female, and a continuous variable to measure the number of members in the household. Opportunity costs of schooling *O* are measured by the logarithm of average labor income of the population with complete secondary schooling in the individual's state/region and area of residence (urban/rural), and by the unemployment rate in the youngster's state/region and area of residence. Lastly, direct costs of schooling *D* are proxied by using measures of accessibility to postsecondary educational institutions relative to the potential demand for postsecondary education. For computing accessibility, we use the number of tertiary and university institutions per 1,000 18-23 year-olds with complete secondary education in the state/region where the individual included in our sample lives.

Table 2 presents some of the descriptive characteristics of our sample. There have been overall improvements in some of the socioeconomic indicators (current income and education of the household head) with deteriorations in others (share of female-headed households and accessibility to tertiary non-university institutions). In Table 3 we show, using the latest survey available for each country, the characteristics of tertiary education participants and contrast them with those of non-participants. As expected, there are clear differences in household background between participants and non-participants. Eighteen to 23 year olds who enroll in postsecondary education have higher household income, come from smaller households, and their household head has more schooling and is less likely to be female. Interestingly, there are almost no differences in terms of opportunity and direct costs of schooling between postsecondary participants and non-participants—although this could be an artifact of these variables being measured at the state/department/regional level.

[4] This variable is available in Chile and Peru for both survey years, and in Colombia for the latest survey year. In Chile and Colombia it measures self-identification with indigenous groups, while in Peru it measures whether the mother tongue is an indigenous language such as Quechua.

It is important to highlight some of the main limitations of our data. First, we rely on cross-sectional data so we cannot effectively measure the long-term factors that a youngster has at his/her disposal when growing up, but rather we can approximate it by measuring household characteristics that tend to be stable over time—such as the educational level of the household head. This also implies that we must exclude from our sample 18-23 year olds who are household heads, which might result in an overestimation of our tertiary education participation rate. A second limitation is that we do not possess information on an individual's ability level. Since ability is influenced by the families' long-term permanent income, lacking measures of academic ability could result in an overestimation of the effect of short-term credit constraints on postsecondary education enrollment. Lastly, data on postsecondary institutions are not available for all survey years, limiting the time span for which we can study the changing role of income on postsecondary education enrollment.

5. Results

In order to directly estimate the role of short-term credit constraints on college enrollment in Latin America, we estimate logistic regressions adding different blocks of control variables sequentially. As previously mentioned, any remaining current income effect after these controls are added suggests the presence of short-term credit constraints.

In table 4 we present the results of our estimations for all four countries and for two points in time. Column (a) shows the effect of current income on postsecondary education enrollment, without controls. Results suggest that the probability of participation in postsecondary education is between 19.7 (for Mexico in 1998) and 42.3 (for Colombia in 2003) percentage points lower for youngsters from the lowest income quartile as compared to that of youngsters from the highest income quartile (the omitted category); between 11.8 (for Peru in 1997) and 36.3 (for Colombia in 2003) percentage points lower for individuals from the middle-low income quartile compared to those from the highest quartile; and between 7.6 (for Peru in 1997) and 27 (for Colombia in 2003) percentage points lower for youngsters from the middle-high income quartile compared to those from the highest income quartile.

The size of the current income effect does not change much if we control for individual and household demographic characteristics (column b). However, once we condition for long-run family background effects such as the education of the household head and ownership of household assets (column c), we find that the effect of current income is reduced considerably. When comparing the lower income quartile with the higher income quartile all but three coefficients become statistically insignificant. These coefficient, which approximate the presence of short-term credit constraints, are now between one-hundredth and one-third the size of that of column (b) and ranges from 0.0023 (for Chile in 2006) to -0.1087 (for Chile in 2000). The coefficients for middle-low and middle-high income quartiles, as compared to the higher income quartile, also are reduced considerably in both size and significance for almost all cases when controlling for long-run family factors. Further controlling for opportunity costs (column d) and for direct college costs (column e) does not change the estimated current income effect, which remains small and within the same range of variation. For a graphic approximation, panel A and B of figure 3 show, respectively, the unadjusted and adjusted postsecondary enrollment levels by income quartiles. It is clear from this figure that the enrollment gaps by income quartile are considerably smaller in panel B than in panel A.

Table 5 shows the coefficients for the variables used as proxy for permanent income—long-run factors such as education of the household head and ownership of assets—using the most complete model that controls for all confounding factors. Results indicate that both the educational level of the household head and the number of goods a household owns have very large and significant impacts on the probability of college attendance among 18-23 year olds.

We perform two sensitivity tests of our results. In the first test, we restrict the sample to only those 18-23 year olds who completed secondary schooling. Secondary graduates are the only ones who truly qualify for access to postsecondary education. Therefore, using the entire sample of 18-23 year olds might have underestimated the role of current income and overestimated that of long-run factors. Results indicate that the role of current income on postsecondary education attendance remains largely unaltered (compared to results presented in table 4), although the size of about a third of the income quartile coefficients becomes slightly larger in absolute terms. In contrast, the education level of the household head has a larger effect among this restricted sample of secondary school graduates across all countries and survey years (compared to the effect shown in table 5).[5] This implies that long-run factors are even more important among high school graduates, and that what could be driving part of the gap in college participation rates between low income and high income high school graduates is the quality of the schools attended throughout the youngsters' educational lives.

In the second sensitivity test, we replace current household monetary income (*C*) as explanatory variable with the residual between *C* and the predicted household income $(\hat{C})$. To do this, we first compute

(6) $$\hat{C}_i = f(ED_{hi}, WE_{hi}, G_{hi,}, S_i)$$

which is a function of the household head's educational level, his/her work experience (and experience squared), his/her gender, and the household size. Then we compute

(7) $$\mathrm{Res}_i = C_i - \hat{C}_i$$

and use Res_i as our explanatory variable. Since family income at the time when the youngster makes his/her college enrollment decision is strongly correlated with family income throughout the life cycle, this step should clean up of any indirect effect that long-term factors have on postsecondary enrollment through current income. In this case, the coefficients for residual current income $(\hat{C})$ are considerably smaller than those obtained using current income (*C*) while those for long-run factors are slightly larger.[6] These results suggest that we were somewhat overestimating the role of credit constraints in tertiary education participation since our measure of current income was capturing the influence of long-run factors in producing that income.

Thus, the evidence presented in tables 4 and 5 is consistent with Carneiro and Heckman's (2002) argument that most of the tertiary enrollment gap by income levels can be explained by long-run factors (parental education and household assets) than by current monetary income.7 As a result, the existence of credit

[5] Results are not shown here but are available from the author upon request.

[6] Results for the model using a continuous measurement of current income (instead of the model using dummies for income quartiles, as shown in tables 4 and 5) and the model using residual current income are not shown here but available from the author upon request.

[7] Our results are also consistent with findings from Belley and Lochner (2008) and Dearden et al. (2004).

constraints—as measured here by current monetary income—cannot fully explain the postsecondary participation gaps shown in Figure 2. Interestingly, however, is that with the exception of Chile [8] in all countries the size of the current income effect has increased for the more recent sample. This result, consistent with Belley and Lochner's (2008) findings for the US, could be indicating that credit constraints may be slightly tightening as access to postsecondary education has increased. Additionally, most of the recent expansion in tertiary education enrollment in the region has been in the tuition-charging private sector (Inter-American Development Bank, 2008), which could explain this apparent tightening of short-term credit constraints.

6. Conclusions

Latin America has experienced large increases in postsecondary participation rates, but these have been accompanied by a widening enrollment gap between high income and low income youngsters. A traditional interpretation of the widening gap would indicate that short-term borrowing constraints are tight, and becoming increasingly so. Thus, in this paper we have attempted to estimate for the first time the extent of short-term credit constraints to college enrollment using cross-sectional household data from four Latin American countries: Chile, Colombia, Mexico and Peru. In addition, we have estimated whether the extent of short-term credit constrains has changed in the last two decades.

Our results suggest that most of the tertiary enrollment gap by income levels can be explained by long-run factors and not by current income, since the effect of current monetary income is considerably reduced once we control for proxies for permanent income such as education of the household head and ownership of assets. This finding seems to indicate that the existence of short-term credit constraints cannot fully explain the postsecondary participation gaps by income levels. However, it is important to highlight that we have also found that the size of the current income effect has increased for the most recent samples, which could be

[8] The number of postsecondary students in Chile who received some form of financial assistance (grants or loans) grew 42 percent between 2000 and 2006. In addition, this increased financial assistance has disproportionably benefited students from the lowest two quartiles of the income distribution. Thus, the smaller estimated role of short-term borrowing constraints (as measured by the coefficient on lower income quartile) for Chile in 2006 could be a result of these changes in financial aid.

indicating that credit constraints are tightening as access to postsecondary education has increased.

What are the policy implications of these results? At face value, they imply that a public policy based solely on expanding financial aid policies might not bring about increased postsecondary education participation rates among low income youngsters in these Latin American countries. They also imply that government policy aimed at expanding access among low income youth should also focus on improving their college-preparedness by developing their cognitive and non-cognitive potential.

Nonetheless, our findings do not lend themselves to suggest that financial aid policies are completely ineffective for reducing the college attendance gap by income levels in these four Latin American countries. First, all four countries included in this study had financial aid policies—albeit limited—in place during the years for which our current income effects have been estimated. For example, approximately 36 percent of Chilean, 25 percent of Colombian, and 11 percent of Mexican postsecondary students received a loan or a grant in the years 2000, 1997, and 1998 respectively. Thus, we cannot predict the size of the estimated current income effect in the absence of these programs. Second, we have found that short-term borrowing constraints have tightened slightly in the past years in all countries, with the exception of Chile. Given the difficulty of expanding enrollment in the already overcrowded public universities, improved postsecondary access for low-income students might need to increasingly come from tuition-charging private institutions—making credit constraints more binding. Lastly, in this paper we have only provided for the first time a direct and crude approximation of the extent of short-term credit constraints in Latin America. Thus, further evidence in this regard is needed in order to make more definitive policy recommendations. There are alternative approaches as well as a limited number of panel datasets that can be exploited for continue estimating the influence of credit constraints in postsecondary education participation in Latin America, and we intend to explore them in future research.

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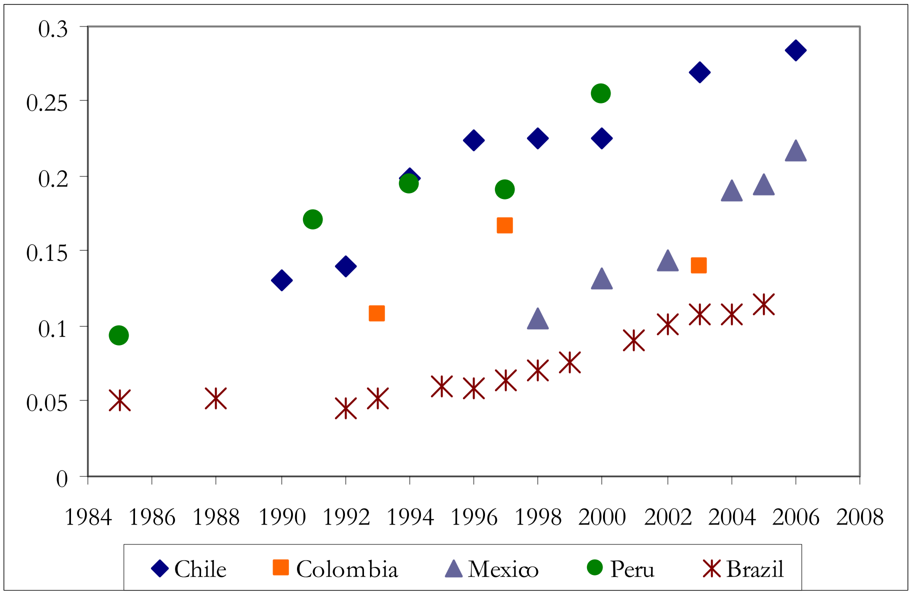
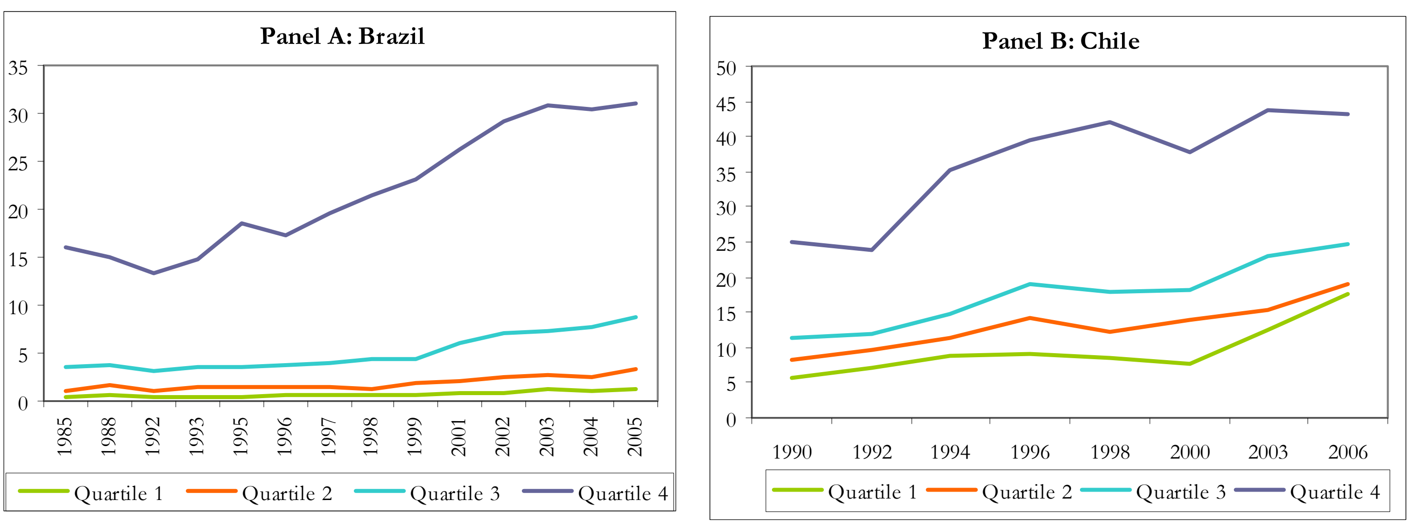
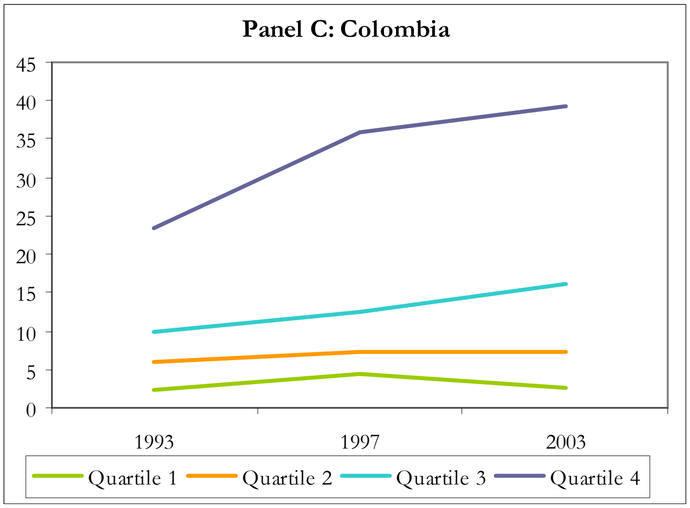
Figure 1: Tertiary participation rates for 18-23 year olds

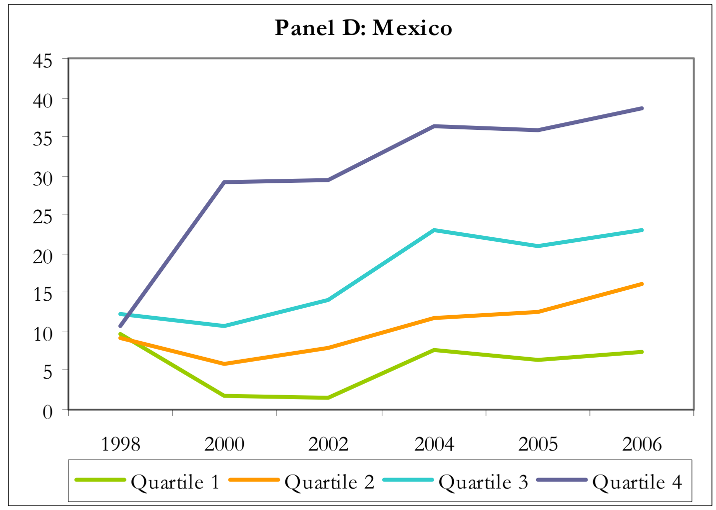


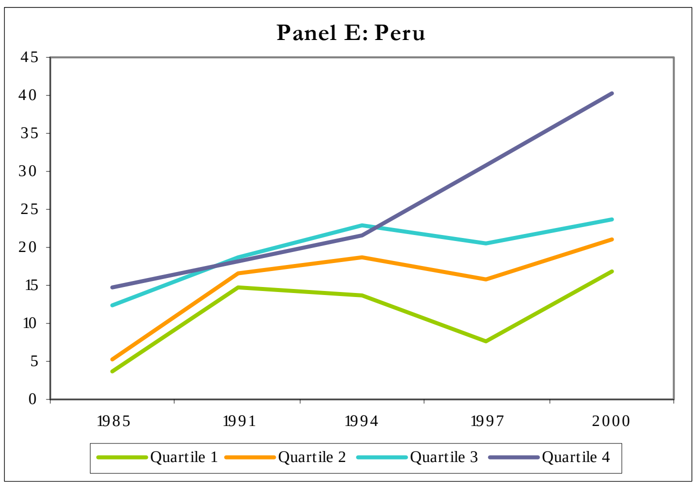
Source: own calculations based on CASEN, ECV, ENIGH, ENNIV and PNAD.

Figure 2: Tertiary participation rates of 18-23 year olds by income quartile









Source: own calculations based on PNAD, CASEN, ECV, ENIGH, and ENNIV.

Table 1: Trends in Postsecondary Education Participation Rates

	Earliest survey[a]	Latest survey[b]
Chile	*23.9%*	*29.4%*
Quartile I	10.1%	19.6%
Quartile II	18.1%	20.9%
Quartile III	23.4%	28.5%
Quartile IV	43.7%	48.5%
QIV - QI	33.6	28.9
Colombia	*13.1%*	*22.6%*
Quartile I	4.1%	6.7%
Quartile II	9.0%	12.7%
Quartile III	12.9%	22.0%
Quartile IV	26.5%	49.0%
QIV - QI	22.4	42.3
Mexico	*11.9%*	*23.7%*
Quartile I	3.7%	11.0%
Quartile II	8.2%	18.0%
Quartile III	12.0%	26.3%
Quartile IV	23.5%	39.6%
QIV - QI	19.7	28.6
Peru	*21.3%*	*20.0%*
Quartile I	11.3%	8.9%
Quartile II	19.3%	16.4%
Quartile III	23.5%	18.9%
Quartile IV	31.2%	35.3%
QIV - QI	19.9	26.4

[a] Chile: 2000, Colombia: 1993, Mexico: 1998, Peru: 1997

[b] Chile and Mexico: 2006, Colombia: 2003, Peru: 2000

Notes: all data weighted for national representativeness

Source: own calculations based on CASEN, ECV, ENIGH, ENNIV

Table 2: Descriptive Statistics, Selected Variables

	Chile		Colombia		Mexico		Peru	
	2000	**2006**	**1993**	**2003**	**1998**	**2006**	**1997**	**2000**
Minority	4.5%	6.9%	n/a	11.4%	n/a	n/a	9.0%	7.6%
Log household income	12.895	13.169	12.871	13.697	8.335	10.379	7.463	7.113
Female-headed household	23.5%	29.3%	27.3%	36.9%	18.7%	26.9%	16.7%	17.8%
Head with no education	3.4%	2.7%	15.2%	11.5%	15.5%	11.0%	7.2%	6.1%
Head with primary education	37.6%	36.7%	57.6%	46.7%	58.4%	43.6%	40.9%	41.6%
Head with secondary education	41.9%	43.3%	17.7%	29.9%	17.8%	27.0%	35.6%	33.3%
Log opportunity cost	12.340	12.449	13.191	12.279	7.912	8.317	6.319	6.236
Unemployment rate	10.5%	7.5%	5.7%	12.4%	2.6%	3.3%	4.7%	5.3%
Accessibility of universities	0.103	0.158	0.075	0.140	0.606	0.476	0.036	0.041
Accessibility of technical colleges	0.309	0.285	0.165	0.061	0.020	0.018	0.608	0.620
Sample size	19,087	21,178	9,224	6,797	4,233	7,252	1,858	1,943

Note: all data weighted for national representativeness. Source: author's calculations based on CASEN, ECV, ENIGH and ENNIV.

Table 3: Tertiary Education Participants vs. Non-participants - Descriptive Statistics, Selected Variables

	Chile 2006		Colombia 2003		Mexico 2006		Peru 2000	
	Participants	**Non-participants**	**Participants**	**Non-participants**	**Participants**	**Non-participants**	**Participants**	**Non-participants**
Female	50.1%	46.8%	54.6%	46.7%	48.1%	50.0%	52.4%	47.4%
Minority	4.7%	7.8%	4.9%	13.2%	n/a	n/a	2.9%	8.8%
Lowest income quartile	16.6%	28.4%	7.4%	30.2%	11.6%	29.2%	11.1%	28.2%
Middle-low income quartile	17.7%	27.9%	14.1%	28.2%	19.0%	26.9%	20.3%	26.0%
Middle-high income quartile	24.3%	25.4%	24.4%	25.2%	27.7%	24.2%	23.3%	25.0%
Female-headed household	25.8%	30.6%	33.6%	37.9%	24.3%	27.7%	16.1%	18.2%
Household members	4.57	5.21	4.85	5.74	4.83	5.93	5.78	6.60
Head with no education	0.8%	3.6%	2.0%	14.3%	2.9%	13.6%	1.4%	7.2%
Head with primary education	16.4%	44.9%	27.2%	52.3%	23.1%	49.9%	23.0%	46.3%
Head with secondary education	47.4%	41.6%	36.3%	28.1%	31.2%	25.7%	32.3%	33.6%
Log opportunity cost	12.464	12.443	12.335	12.263	8.376	8.299	6.298	6.221
Unemployment rate	7.7%	7.5%	13.6%	12.1%	3.4%	3.2%	6.3%	5.1%
Accessibility of universities	0.158	0.157	0.132	0.142	0.466	0.479	0.039	0.042
Accessibility of technical colleges	0.284	0.285	0.061	0.061	0.017	0.018	0.606	0.623

Note: all data weighted for national representativeness. Source: author's calculations based on CASEN, ECV, ENIGH and ENNIV.

Table 4: The Changing Role of Current Income on College Attendance
Marginal Effects

	(a)	(b)	(c)	(d)	(e)
Panel A - Chile					
Year 2000					
Lower	-0.3312***	-0.3408***	-0.1087**	-0.1166***	-0.1121***
	(0.0165)	(0.0174)	(0.0348)	(0.0343)	(0.0333)
Middle-low	-0.2509***	-0.2577***	-0.0147	-0.0218	-0.0213
	(0.0184)	(0.0191)	(0.0296)	(0.0296)	(0.0291)
Middle-high	-0.1986***	-0.2069***	-0.0445*	-0.0466*	-0.0465*
	(0.0182)	(0.0191)	(0.0254)	(0.0254)	(0.0249)
Year 2006					
Lower	-0.2889***	-0.3128***	0.0023	-0.0176	-0.0183
	(0.0159)	(0.0164)	(0.0245)	(0.0245)	(0.0247)
Middle-low	-0.2754***	-0.2913***	-0.0616**	-0.0717***	-0.0734***
	(0.0163)	(0.0169)	(0.0223)	(0.0224)	(0.0226)
Middle-high	-0.1993***	-0.2068***	-0.0538**	-0.0609**	-0.0608**
	(0.0173)	(0.0178)	(0.0203)	(0.0204)	(0.0205)
Panel B - Colombia					
Year 1993					
Lower	-0.2236***	-0.2016***	-0.0259	-0.0316	-0.0326
	(0.0220)	(0.0219)	(0.0209)	(0.0254)	(0.0262)
Middle-low	-0.1752***	-0.1642***	-0.0169	-0.0209	-0.0217
	(0.0244)	(0.0222)	(0.0148)	(0.0181)	(0.0188)
Middle-high	-0.1361***	-0.1275***	-0.0190	-0.0229	-0.0235
	(0.0262)	(0.0228)	(0.0161)	(0.0194)	(0.0199)
Year 2003					
Lower	-0.4234***	-0.4234***	-0.0738**	-0.0858**	-0.0838**
	(0.0207)	(0.0240)	(0.0335)	(0.0384)	(0.0362)
Middle-low	-0.3628***	-0.3656***	-0.0484*	-0.0570*	-0.0577*
	(0.0224)	(0.0244)	(0.0281)	(0.0316)	(0.0294)
Middle-high	-0.2703***	-0.2672***	-0.0261	-0.0328	-0.0339
	(0.0246)	(0.0257)	(0.0243)	(0.0274)	(0.0254)
Panel C - Mexico					
Year 1998					
Lower	-0.1972***	-0.1682***	-0.0139	-0.0139	-0.0127
	(0.0187)	(0.0181)	(0.0194)	(0.0210)	(0.0208)
Middle-low	-0.1522***	-0.1333***	-0.0011	-0.0004	0.0006
	(0.0217)	(0.0190)	(0.0118)	(0.0135)	(0.0138)
Middle-high	-0.1145***	-0.0928***	0.0037	0.0049	0.0068
	(0.0220)	(0.0198)	(0.0109)	(0.0127)	(0.0139)
Year 2006					
Lower	-0.2861***	-0.3107***	-0.0194	-0.0274	-0.0283
	(0.0192)	(0.0196)	(0.0268)	(0.0277)	(0.0281)
Middle-low	-0.2159***	-0.2355***	-0.0185	-0.0216	-0.0219
	(0.0195)	(0.0203)	(0.0202)	(0.0208)	(0.0210)
Middle-high	-0.1334***	-0.1336***	0.0023	0.0010	0.0004
	(0.0216)	(0.0227)	(0.0180)	(0.0180)	(0.0181)

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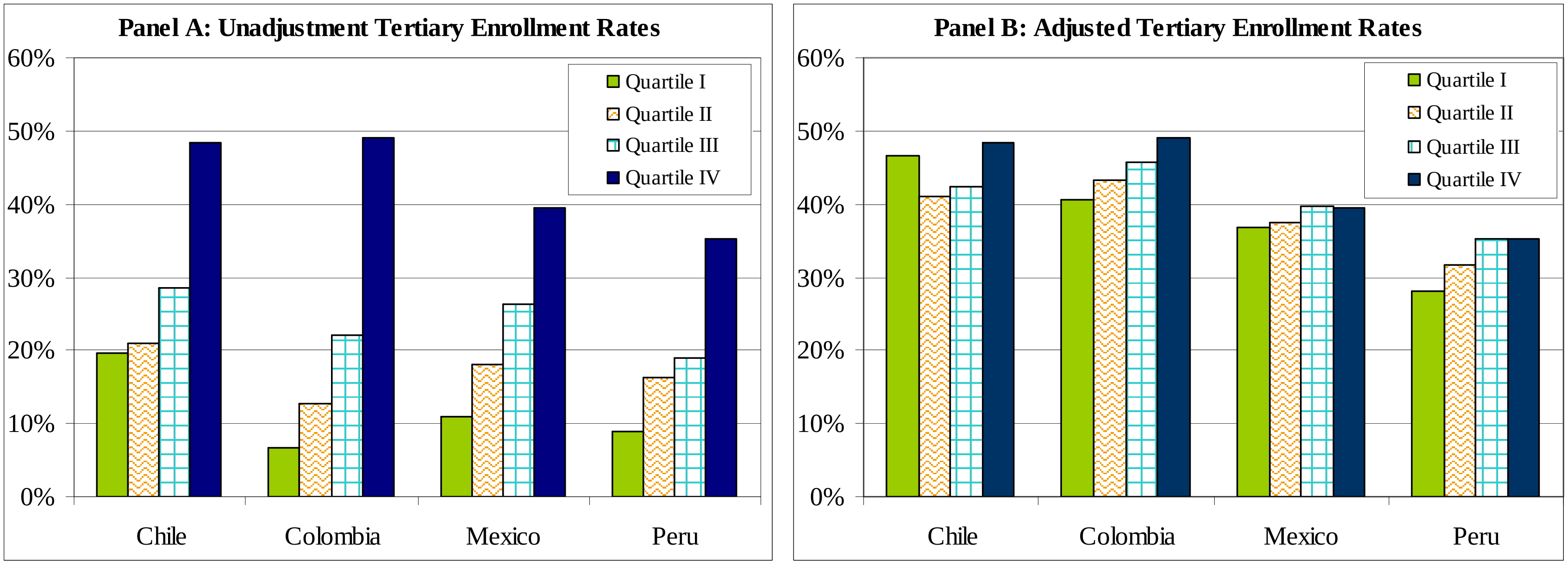
Table 4, continued: The Changing Role of Current Income on College Attendance - Marginal Effects

	(a)	(b)	(c)	(d)	(e)
Panel D - Peru					
Year 1997					
Lower	-0.1990***	-0.2235***	-0.0293	-0.0265	-0.0322
	(0.0272)	(0.0294)	(0.0328)	(0.0321)	(0.0321)
Middle-low	-0.1186***	-0.1485***	0.0073	0.0111	0.0037
	(0.0297)	(0.0311)	(0.0290)	(0.0289)	(0.0281)
Middle-high	-0.0763**	-0.1029***	-0.0039	-0.0045	-0.0065
	(0.0309)	(0.0321)	(0.0260)	(0.0255)	(0.0254)
Year 2000					
Lower	-0.2649***	-0.2591***	-0.0690**	-0.0714***	-0.0721***
	(0.0273)	(0.0284)	(0.0309)	(0.0286)	(0.0289)
Middle-low	-0.1906***	-0.1962***	-0.0375	-0.0353*	-0.0358*
	(0.0294)	(0.0287)	(0.0254)	(0.0216)	(0.0219)
Middle-high	-0.1655***	-0.1712***	-0.0566**	-0.0447**	-0.0447**
	(0.0304)	(0.0299)	(0.0253)	(0.0218)	(0.0219)
Demographics	No	Yes	Yes	Yes	Yes
Permanent income	No	No	Yes	Yes	Yes
Labor market	No	No	No	Yes	Yes
Direct costs	No	No	No	No	Yes

Notes: *** logistic coefficient significant at 0.001 level; ** significant at 0.05 level; * significant at 0.1 level
Standard errors in parentheses. Source: author's calculations based on CASEN, ECV, ENIGH and ENNIV

Figure 3: Unadjusted and Adjusted Tertiary Enrollment Rates by Income Quartiles



Note: The adjusted enrollment rates are based on the results from the specification shown in table 4, column (e). Source: own calculations based on CASEN 2006, ECV 2003, ENIGH 2006, and ENNIV 2000

Table 5: The Changing Role of Permanent Income on College Attendance
Marginal Effects

	Chile		Colombia		Mexico		Peru	
	2000	**2006**	**1993**	**2003**	**1998**	**2006**	**1997**	**2000**
Head with no education	-0.3285***	-0.2426***	-0.0389***	-0.1322**	-0.0604***	-0.1460***	-0.1322***	-0.1005**
	(0.0510)	(0.0447)	(0.0322)	(0.0465)	(0.0659)	(0.0815)	(0.0479)	(0.0285)
Head with primary education	-0.2722***	-0.2409***	-0.0364***	-0.1161**	-0.0482***	-0.1335***	-0.0961***	-0.0781***
	(0.0383)	(0.0390)	(0.0298)	(0.0377)	(0.0523)	(0.0717)	(0.0356)	(0.0223)
Head with secondary education	-0.1627***	-0.1435***	-0.0249**	-0.1120***	-0.0300***	-0.0913***	-0.0766***	-0.0539**
	(0.0265)	(0.0250)	(0.0216)	(0.0351)	(0.0329)	(0.0478)	(0.0291)	(0.0181)
Household goods	0.0866***	0.0769***	0.0127***	0.0661***	0.0112***	0.0317***	0.0176***	0.0179***
	(0.0070)	(0.0047)	(0.0095)	(0.0136)	(0.0117)	(0.0139)	(0.0056)	(0.0046)
Home ownership	-0.0205	0.0345**	n/a	0.0105	0.0159	0.0141	0.0399	0.0091
	(0.0214)	(0.0171)	n/a	(0.0218)	(0.0208)	(0.0187)	(0.0247)	(0.0154)
Demographics	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Current income	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Labor market	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Direct costs	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: *** logistic coefficient significant at 0.001 level; ** significant at 0.05 level; * significant at 0.1 level.
Standard errors in parentheses. Source: author's calculations based on CASEN, ECV, ENIGH and ENNIV.