

Labor Market Effects of Short-Cycle Higher Education Programs

Challenges and Evidence from Colombia

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Abstract

This paper estimates the labor market effects of enrolling in a short-cycle program in Colombia. Following evidence for the U.S., increasing access to short-cycle degrees might attract some students who would not have enrolled in higher education otherwise (i.e., the expansion or democratization margin), while also inducing other students to divert from bachelor's- and into short-cycle- degrees (i.e., the diversion margin). To identify responses along these margins, this paper uses an Instrumental Variables strategy and exploits local variation in the supply of short-cycle programs for the universe of high school graduates in 2005. Having at least

one higher education institution specialized in short-cycle degrees within a 10 km radius of the student's high school municipality increases enrollment in short-cycle programs by 3 percentage points, or 30 percent of the sample average. Results indicate that this enrollment increase is largely driven by students who would divert from bachelor's to short-cycle degrees due to changes in the local supply of short-cycle program. For these students, SCPs improve participation in the formal labor market among females, although they lead to lower monthly wages among males.

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Labor Market Effects of Short-Cycle Higher Education Programs: Challenges and Evidence from Colombia

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1 Introduction

The effects of higher education expansions and the role of degrees of short duration (such as two-year college degrees or short-cycle degrees) as pathways to provide skills fast for a larger share of students have received increased attention: for the U.S., recent evidence shows that community, or two-year, colleges can benefit students in terms of educational attainment and earnings (Acton, 2020; Bettinger and Soliz, 2016; Denning, 2017; Jepsen et al., 2014; Marcotte, 2019; Minaya and Scott-Clayton, 2020; Mountjoy, 2019). Moreover, these effects are mostly driven by students who would not have enrolled in higher education in the absence of short-term degrees.

In this paper, we estimate labor market effects of enrolling in a short-term program using local variation in the supply of higher education for the universe of high school graduates in 2005 in Colombia. Following a long-standing literature on two-year college enrollment for the U.S. (Acton, 2020; Denning, 2017; Leigh and Gill, 2003; Mountjoy, 2019; Rouse, 1995), we argue that access to short-term programs (henceforth, short-cycle programs¹) might attract some students who would have not enrolled in higher education otherwise (*i.e.*, the *expansion* or *democratization* margin), while also inducing other students to divert from bachelor’s- and into short-cycle- degrees (*i.e.*, the *diversion* margin). Students along these margins likely differ both in their characteristics and in the gains, or losses, from choosing a short-cycle program.

Latin American countries have experienced a fast and large expansion of higher education enrollment (Camacho et al., 2017; Carranza and Ferreyra, 2019; Ferreyra et al., 2017, 2020; González-Velosa et al., 2015). While short-cycle programs are less popular in the region than in the rest of world, Colombia is an exception, as its short-cycle programs currently attract about a third of higher education enrollment. Given recent efforts in Colombia and other Latin American countries to expand higher education access (Ferreyra et al., 2017), uncovering effects along those two margins of choice warrants attention. Nonetheless, the identification of effects at each margin is not straightforward and poses multiple challenges. First, students self-select into their preferred choice, which biases OLS estimates of the effects of short-cycle programs on labor market outcomes. Second, while a standard Instrumental Variables (IVs) approach can help overcome self-selection, it does not identify effects along each margin but rather a combination of the two margins. In other words, it only identifies the effect of short-cycle programs relative to a next-best option that combines not enrolling in higher education and enrolling in a bachelor’s program. Only additional (but often restrictive) assumptions can decom-

¹UNESCO uses “short-cycle tertiary education programs” for International Standard Classification of Education (ISCED 2011). This encompasses community college programs lasting one or two years in the U.S., and the two- or three-year programs in Colombia discussed in this paper.

pose this effect into margin-specific effects (Hull, 2018; Kline and Walters, 2016; Lee and Salanié, 2020 and Mountjoy, 2019).

Our main contribution lies in providing evidence of the expansion and diversion effects for a context such as the Colombian higher education system. While the economic returns to higher education in Colombia have been previously studied, the evidence either focuses on the effects of new *versus* existing programs (Camacho et al., 2017), selective colleges (Barrera-Osorio and Bayona-Rodríguez, 2019), or programs within fields (Ferreyra et al., 2020 and González-Velosa et al., 2015). Unlike previous studies, and for simplification, we abstract from within-field heterogeneity and provide Local Average Treatment Effects (LATEs) for students along the two well-defined and policy-relevant margins of choice: expansion and diversion into short-cycle degrees.

To understand the policy relevance of these margins, consider an expansion in the local supply of short-cycle programs. Thanks to this expansion, students might enroll in higher education even though they might not have done it otherwise. In principle, we would expect positive labor market effects for these students due to the accumulation of human capital. Other students, in contrast, might divert from a bachelor's program onto a short-cycle program. These students might experience worse labor market outcomes as short-cycle than bachelor's graduates because they might be "undermatched" to the short-cycle program and might therefore not reach their full productive potential. Nonetheless, they might also experience better labor market outcomes because they might attain a better match through the short-cycle program. This would be the case, for instance, if they were only able to enter a low-return bachelor's program. In other words, whether students gain or lose from short-cycle programs depends on what their next best would have been, and is an empirical question.

We use an Instrumental Variables (IVs) approach that uses outcomes-by-choice interactions and propensity scores (as in Heckman and Pinto (2018) and Mountjoy (2019)) to estimate LATEs of short-cycle programs *versus* the next-best. A critical advantage of this strategy is that it allows us to isolate the effect of a single instrument and, in this way, improve over standard Two-Stage Least Square (2SLS) methods with multiple endogenous choices and multiple IVs.² Using administrative data for the universe of high school graduates in Colombia in 2005, we track students from their high school

²Heckman and Urzúa (2010) show that if an instrument affects only one choice, IVs estimate the effect of one option *versus* the next-best. Also, Mountjoy (2019) shows that multivariate 2SLS (multiple choices and multiple IVs) estimate effects that are a combination of many margins of choice and thus do not necessarily identify well-defined effects. For instance, if there are three margins of choice (not enrolling in higher education vs. enrolling in short-cycle programs, not enrolling in higher education vs. enrolling in bachelor's programs, enrolling in bachelor's programs vs. enrolling in short-cycle programs) and multiple instruments, then multivariate 2SLS would estimate a combination of effects for those three margins.

exit exam to their enrollment choice and match them to labor market participation data from 2008-2013 for those working in the formal sector. We link the resulting dataset to information on higher education supply at the municipality level to exploit local variation in a binary variable that indicates whether there is at least one higher education institution specialized in short-cycle degrees within a 10 km radius of the student's high school municipality.

According to our estimates, enrollment in short-cycle programs increases by 3 percentage points due to the existence of such institutions. This is a sizable effect, as it represents a 30% increase over the sample enrollment rate in short-cycle programs. Moreover, almost 89% of this enrollment increase stems from students who would divert from bachelor's programs into short-cycle degrees (*i.e.*, the diversion margin). The remaining 11% of students would react along the expansion margin by entering higher education, yet this effect is not statistically significant. We also observe a higher diversion share among men than women (99% and 75%, respectively). Further, we uncover positive effects of choosing a short-cycle program on labor market participation rates for women. In contrast, we observe negative effects on average monthly wages among men.

The fact that diversion effects are greater than expansion effects in Colombia contrasts with the U.S., where the reverse takes place (Acton, 2020; Denning, 2017; Leigh and Gill, 2003; Mountjoy, 2019).³ This may be because students who enter community college in the U.S. typically intend to transfer to a four-year institution (Acton, 2020; Denning, 2017; Kane and Rouse, 1999). In other words, they view the community college expansion as an additional opportunity to enter four-year colleges. In Colombia, in contrast, short-cycle programs are viewed as terminal degrees and rarely offer a pathway towards a bachelor's program.⁴

The labor market effects of diversion are similar in Colombia and the U.S.: they are positive for formal employment but, as in the US, they are negative on wages (Mountjoy, 2019). The positive employment effects accrue mostly to females, while the negative wage effects accrue mostly to males. Although diverting students are, on average, less disadvantaged than short-cycle students in general, diverting females are actually more disadvantaged than diverting males. It is possible, then, that the diverting females improve their labor market participation because they would have otherwise attended a low-quality bachelor's program. By a similar reasoning, the diverting males may lose

³For instance, for the state of Texas, Mountjoy (2019) finds that about two-thirds of students would react to changes in proximity to two-year college by switching along the expansion margin; the remaining 30% would divert from four-and into two-year college. Denning (2017) finds that lower two-year college tuition induces students who would not have enrolled otherwise to attend a two-year college, with no evidence of students substituting two- for four-year college degrees.

⁴In our sample, we observe that less than 20% of students who choose short-cycle degrees end up transferring to a bachelor's program.

in terms of wages because the short-cycle program may be an "undermatch" relative to their ability.

Our results are broadly consistent with those from the United States. We estimate effects for students who would react to local differences in the existence of short-cycle programs. That is, we are comparing the enrollment choice of students residing in a municipality with at least one institution offering short-cycle programs vs. students in municipalities without these programs. Other sources of variation might have larger (or smaller) effects on enrollment in short-cycle programs than our finding of an increase in 2.9 percentage points. For instance, previous studies for the U.S. use variation in distance to college (Mountjoy, 2019) or changes in tuition (Acton, 2020; Denning, 2017) with effects between 3-5 percentage points on two-year college enrollment.⁵ Our results fall in this range even though they correspond to the Colombian context and exploit a different source of variation.

This paper is organized as follows. The following Section describes the higher education system in Colombia and the data. It presents an exploratory analysis of enrollment choices and labor market participation as well as results using a standard Instrumental Variables strategy. Section 3 discusses how to identify shares of students along the expansion and diversion margins, as well as their LATEs. Last, we present conclusions and discussion in Section 4.

2 Institutional Information and Data

The higher education System in Colombia offers different types of programs that can be classified in bachelor's programs which last four to five years, and short-cycle programs which last two to three years (similar to US two-year or community college programs). The latter are divided in technical and technological degrees. In terms of higher education institutions (HEI), there are Universities, University Institutes, Technological Institutes, and Technical Professional Institutes. Universities and University Institutes can offer either bachelor's or short-cycle programs, Technological Institutes and Technical Professional Institutes can only offer short-cycle degrees. In our estimation strategy, we exploit variation in the supply of short-cycle programs by focusing on two sets of variables: (i) whether SC programs are offered at the local level (in the municipality where the student attended high school), (ii) the type of HEI available at the local level (distinguishing between HEI that can only offer short-cycle programs and HEI that can offer either

⁵For Texas, Mountjoy (2019) finds that an increase of 10 miles in the distance from the student's high school to the nearest two-year college decreases enrollment in two-year college by 4.3 percentage points. Acton (2020) finds that a decrease in local community college tuition in \$1000 increases enrollment in these institutions by 3.5 percentage points.

bachelor’s or short-cycle programs).

2.1 Data

We use administrative data on test scores, enrollment choices, higher education trajectories, and wages in the formal labor market for the universe of high school graduates in 2005. First, individual records from the high school exit exam (Saber 11, in Spanish) contain information on test scores for math, language, science, physics, history, chemistry, geography, and philosophy. These records also include data on student’s age and gender, as well as socioeconomic characteristics before enrolling in higher education, such as mother’s education, household income level, and the number of siblings. Higher education enrollment, trajectories, and completion come from the System for Dropout Prevention of Higher Education (SPADIES, in Spanish).

We combine the administrative records at the student level with information on higher education supply from the National System of Information on Higher Education (SNIES, in Spanish), which contains the number of institutions and programs, their year of creation, field and academic level, and total enrollment from 1998-2013. We also use municipal-level data on labor market participation from the 2005 Census of Population, from IPUMS International ([Minnesota Population Center, 2019](#)). In addition, we link our dataset to the Municipal Panel of the Center of Studies on Economic Development (CEDE, in Spanish), which contains information on population (urban, rural, and total), distance to the main city in the department, area in squared km, poverty level, and the fraction of unemployment among households.

Wages and occupation in the formal labor market for those who graduate from a higher education program are available in the Labor Market Observatory for Education (OLE, in Spanish) records from 2008-2013. For high school graduates and those who did not complete a higher education program, we use data for Colombia from the Socioeconomic Database for Latin America and the Caribbean (SEDLAC) to estimate labor market participation and experience between 2008-2013. In particular, we follow the labor market trajectories of students within that period, 2008-2013, using data from OLE and the Integrated Household Survey (GEIH, in Spanish) in SEDLAC. We use the latter to estimate labor market participation of high school graduates and higher education dropouts with a linear probability model for the cohort of individuals 14-24 years of age in 2005. We estimate each equation separately for students who did not enroll in higher education and for higher education dropouts.⁶

The estimating sample consists of high school graduates in 2005 who were between

⁶For more details on our imputation strategy see Appendix A2.

14-24 years of age at the time of the high school exit exam. About 53.8% of high school graduates in the sample are women, 27.2% come from low-income households, and 42% have mothers with less than primary education. More than half of students in the sample (61.2%) never enrolled in a higher education program.⁷ Of the remaining students, 38.8% enrolled in higher education with 9.5% in short-cycle programs while 29.3% selected a bachelor's program.⁸ In terms of higher education completion, about 4.7% of students enroll in a short-cycle program but do not complete it, and only 3.6% of the sample has a short-cycle degree. In contrast, 14.9% of students obtain a bachelor's degree. Half (50.1%) of the students in our sample participate in the formal labor market in 2013.

Students who enrolled in short-cycle programs (henceforth, SC) are older, belong to larger low-income households, and have less educated mothers, on average, than students who enrolled in bachelor's programs (BP). Moreover, students who did not enroll in higher education (NE) stem from the most disadvantaged backgrounds and have the lowest average performance on the high school Exit Exam (Table 2). In turn, BP students have higher test scores, on average, than SC students across all academic subjects of Saber 11. The most considerable differences between SC and BP students in their Saber 11 performance correspond to Chemistry, Biology, and Reading; on average, BP students outperformed SC students in at least 0.3 standard deviations.

About 16.4% of students in SC programs transfer to a bachelor's program, and less than half of those who transfer complete the program. In contrast, only 1.2% of students who enrolled in BP completed an SC (see Table 2). While the fraction of dropouts is similar across SC programs and BP (about 50%), completion rates are higher among BP students by almost 14.5 percentage points. Last, those who do not enroll in higher education (NE students) have the lowest share, at 48%, of students working in the formal labor market in 2013. NE students, who have been out of school for longer than SC and BP students, also have the highest average years of experience. In contrast, average monthly wages are almost twice as high among BP students than among NE students.

Table 3 presents the average of some characteristics of the student's high school municipality by type of higher education enrollment. On average, NE students live in municipalities with lower shares of urban/rural population, higher poverty, and higher distance to the main city than SC and BP students. In turn, SC students live in denser, slightly more urbanized, and better-connected municipalities with lower poverty rates than BP

⁷These are students who do not match to SPADIES, the dataset of higher education enrollment, or OLE, the dataset of labor market earnings. It is worth noting that SPADIES does not contain information for students who enroll in SENA, a public institution that provides post-secondary technical and technological training. In this sense, the share of our sample classified as "not enrolled in higher education" also contains students who attend SENA. Nonetheless, we do not have information available to identify these two groups of students separately.

⁸We restrict the first enrollment to be no later than five years after high school graduation.

students. Hence, on average, the differences in local characteristics between SC and BP students go opposite to what we observed in terms of individual and household characteristics. Thus, while SC students are more disadvantaged in terms of average household characteristics and test scores, they also stem from municipalities with better local conditions, on average, than BP students.

2.2 Exploratory Analysis

To determine the role of short-cycle degrees on student’s labor market trajectories we focus on two outcomes: the probability of working in the formal labor market and average monthly wages in 2013. This year corresponds to the last year of student-level labor market data we have available from OLE.

To explore how SC enrollment affects labor market trajectories, we first estimate the following equation by OLS:

$$Y_i = \beta_0 + \beta_{SC}D_{SCi} + \beta_X X_i + u_i, \quad (1)$$

where Y_i denote labor market outcomes for student i , and D_{SCi} is a binary variable that takes the value of one if student i enrolls in a short-cycle program. X_i includes age, gender, number of siblings, household income level, mother’s education, test scores in Saber 11, and characteristics of the municipalities where the student attended high school (as detailed in Table 3).

One important feature of equation (1) is that the omitted category is a mixture of students who did not enroll in higher education (NE) and students who enroll in BP. Also, u_i is an unobserved component which contains students’ preferences and unobserved ability, among other factors. Since these factors can also explain enrollment choices, we see the results we present here as descriptive rather than causal evidence of the labor market effects of short-cycle degrees. In Table 4 we observe a negative association between enrolling in a SC program and working in the formal labor market relative to a mixture of enrolling in bachelor’s programs or not entering higher education. In contrast, enrolling in a SC degree is associated with an increase in average monthly wages (both for the period of 2008-2013 and during 2013) vs. not enrolling in a SC program.

2.3 Standard Instrumental Variables

Students who choose to enroll in a short-cycle program (SC) differ in observable and unobservable characteristics from students who choose not to enroll in higher education (NE) or enroll in a bachelor’s program (BP). Preferences, motivation, and abilities (which

we can only be measured with noise using test scores) can explain enrollment choices and labor market outcomes. To deal with self-selection into short-cycle programs, we use an instrumental variables (IVs) strategy. In a standard 2SLS set up, equation (1) represents the second-stage, or the outcome equation. The first-stage corresponds to the following:

$$D_{SCi} = \gamma_0^{SC} + \gamma_1^{SC} Z_i + \gamma_X^{SC} X_i + \epsilon_{SCi}, \quad (2)$$

where Z is an IV, X_i is a matrix of student-level and municipal-level characteristics, and ϵ is an unobserved component. We define the instrument in Z as follows:

Local supply of higher education by type of program: we use data from Saber 11 to identify the municipality where the student attended high school. We use this information to determine if, in 2004, there were any SC or BP institutions near the students' high school municipality. Further, we distinguish between higher education institutions (HEIs) that only offer SC programs, those that only offer BP programs, and those that offer both types of programs. We use the variation in local availability by type of HEI to define a binary instrument, Z , that takes the value of one if there is an HEI specialized in SC programs within a radius of 10km from the student's HS municipality. We also define variables for the availability of other institutions availability (those only offering bachelor's programs and those offering either kind).

The IV of existence of SC programs at the local level follows the logic of cost-shifters, as in [Card \(1995\)](#) and [Mountjoy \(2019\)](#). That is, we assume that students with an HEI near their high school municipality might have lower costs of enrolling in higher education. In contrast, students with no HEI locally would face higher costs. These costs can be in terms of access (*e.g.*, having an HEI nearby reduces transportation costs) or information (*e.g.*, if there is an HEI near, the student has fewer barriers to learn about the supply of programs). [Table 5](#) shows the average of HEI availability in 2004 for the total sample. About 78.3% of students went to high school in a municipality with an HEI that offers both SC and BP or only BP. This fraction reduces to 61.1% for HEI that only offers SC programs.⁹ Moreover, we observe that almost 74.2% of SC students went to high school in a municipality where there was at least one SC HEI, and this share reduces to 65.4% for BP and 57% for NE students.

We also use data on enrollment in SENA in 2004 to capture public availability, free of charge, of technical and technological programs. Around 75% of students attended high school in municipalities with positive SENA enrollment, and this share is higher among

⁹Although these local differences in the supply of HEI can proxy for the costs that students face, there can be heterogeneity in the tuition that different types of HEI can charge. There is no publicly available data on tuition by HE programs in 2005, however more recent data ([González-Velosa et al., 2015](#)) shows that tuition in private HEI can be twice as high as that of public HEI. Using imputed data for tuition, we observe that SC programs charge about half of the average tuition among BP.

BP and SC students than among NE students (see Table 5). We include this variable in our regressions since our group of NE students consists of high school graduates who did not continue their education and high school graduates who enroll in SENA. In our 2SLS specification, we exclude from the outcome equation all the variables of supply of higher education, namely enrollment in SENA in 2004 and the existence of HEI not specialized in short-cycle programs. This approach translates into 2SLS with one endogenous variable and multiple IVs.

The effects of enrolling on SC can largely depend on the counterfactual option or what students would have chosen in the absence of SC programs. For instance, students who would not enroll in higher education if SC were not available could derive larger (or smaller) gains from choosing an SC than students who would have chosen a bachelor’s program instead of SC. In our setting, we can define two groups of SC students according to their counterfactuals: (i) students who would not have enrolled in higher education and (ii) students who would have enrolled in BP. Previous literature (Rouse, 1995; Mountjoy, 2019) defines the former as the *democratization* or *expansion margin* (students enter the HE system via short-cycle programs), while the latter is referred to as the *diversion margin* (students divert from BP and into short-cycle programs). Changes in the instrument of local existence of SC could shift students into SC from either of those margins (*e.g.* from NE towards SC, or from BP towards SC). Students who would change their enrollment choices as the instrument changes are defined as compliers. There are two complier groups in our setting, namely compliers along the expansion margin and compliers along the diversion margin.

Compliers along the expansion or diversion margins can experience different effects of enrolling in a short-cycle program. Let $LATE_{SC-BP}$ represent the local effects of choosing SC for complier students along the diversion margin; $LATE_{SC-NE}$ represents the local effects of choosing SC for complier students along the expansion margin. Previous literature (Heckman and Urzúa, 2010; Mountjoy, 2019) has shown that, in contexts where compliers are switching from different initial states, standard (univariate) IV estimates the effect of one option, (*e.g.* SC) *versus* the next-best (*i.e.*, a mixture of NE and BP),

$$\beta^{SC} = \omega LATE_{SC-BP} + (1 - \omega) LATE_{SC-NE}, \quad (3)$$

where ω is the share of compliers along the diversion margin.¹⁰ Our main goal is to estimate ω and the LATE of short-cycle programs vs. the next-best. It is worth noting

¹⁰Mountjoy (2019) presents a formal derivation of these effects (*i.e.*, of one option *versus* the next-best) in the context of two-year college in the U.S. Moreover, he also shows that multivariate IV (*i.e.*, where there are two or more endogenous choices) with multiple instruments does not necessarily identify a well-defined effect but rather a combination of students shifting along many margins.

that, depending on the IV, either the share of compliers along the diversion or along expansion margins might be negligible (*e.g.*, $\omega \rightarrow 1$). In that case, the LATE of short-cycle programs vs. the next-best would largely reflect the LATE for students along the diversion margin.

Table 6 shows the results of estimating equation (1) and (2) with standard 2SLS, for one endogenous variable and multiple excluded instruments. Taken as given, the results suggest that among compliers SC enrollment increases formal experience in 1.2 years. The 2SLS results in Table 6 for formal labor market participation have the opposite direction than the OLS results in Table 4, and only the 2SLS effect on experience is statistically significant. An important caveat to the interpretation of these results is that changes in the variable of the existence of other (not SC) HEIs might induce students to switch along the diversion margin, but towards BP (opposite to the direction of our IV of SC HEIs). As a result, there might be two-way flows of students switching into and away from short-cycle programs. Hence, we cannot interpret the results in Table 6 as the LATEs of SC programs since they might be capturing effects from students who would substitute SC for BP (instead of the opposite). In the next section, we employ a strategy to estimate LATEs of short-cycle programs and show evidence of expansion and diversion effects in the Colombian context.

3 Self-selection into Short-Cycle programs: expansion *versus* diversion

Uncovering LATEs of SC programs for students along the expansion and diversion margins is both an empirical and policy-relevant challenge. First, without additional assumptions, standard IV does not identify group specific sub-LATEs. Furthermore, there should be as many instruments as margins of choice for the identification of sub-LATEs. More important, SC programs might be a better match for students along the expansion or diversion margins. Overall, LATEs of one option *versus* the next-best might not provide enough information about the type of students who might, or not, benefit from a SC program.

In this section we employ the methodological framework in Heckman and Pinto (2018) and Mountjoy (2019) to, first, estimate the share of expansion-compliers and diversion-compliers induced by the variation in local availability of SC HEIs. Then, we estimate LATEs of SC programs vs. the next-best alternative on labor market outcomes. In particular, we find no statistical evidence that changes in the availability of SC HEI shift students along the expansion margin (*i.e.*, $1 - \hat{\omega} = 0$). We show how, in turn, for students

along the expansion margin other local conditions, such as employment prospects, matter for the decision of whether or not to enroll in college.

3.1 Complier Shares

In this section we present our identification strategy to estimate ω in equation (3). Let D_{di} denote the student's enrollment choice. That is, D_{di} are binary variables such that $D_{di} = \mathbb{1}[\text{student } i \text{ chooses enrollment option } d]$ for $d \in \{SC, BP, NE\}$. Choices depend on students' characteristics, in X_i , such as household income, maternal education, gender, age, and household composition, as well as test scores in the high school Exit exam, and on the existence of at least one HEI specialized in SC at the student's high school municipality (Z). In what follows, we implicitly condition on X_i to simplify notation.

Following the logic of cost-shifters, we assume that changes in Z would induce some students to change their enrollment choices towards SC programs. Let $D_{di}(Z)$ denote potential choices, or the enrollment status student i would choose at different realizations of Z . For instance, if $D_{SCi}(1) = 1$ the student would choose a SC program when there was a SC HEI in her high school municipality. We formalize the potential changes in enrollment choices as the instrument changes as follows:

$D_{NEi}(0) \geq D_{NEi}(1)$, which means that having a SC HEI available should make NE weakly less attractive for students.

$D_{BPi}(0) \geq D_{BPi}(1)$, which means that having a SC HEI available should make BP weakly less attractive for students.

$D_{SCi}(0) \leq D_{SCi}(1)$, which means that having a SC HEI available should make SC weakly more attractive for students.

The combinations of these enrollment changes as a result of changes in the instrument satisfy the property of *monotonicity* (Imbens and Angrist, 1994) such that all students are (weakly) induced towards the same option (SC, in this case). Using these assumptions on potential enrollment choices, we can identify the total share of compliers, as well as the shares of margin-specific compliers. Formally,

$$\omega = \frac{E[D_{BP}|Z = 0] - E[D_{BP}|Z = 1]}{E[D_{SC}|Z = 1] - E[D_{SC}|Z = 0]}$$

$$1 - \omega = \frac{E[D_{NE}|Z = 0] - E[D_{NE}|Z = 1]}{E[D_{SC}|Z = 1] - E[D_{SC}|Z = 0]},$$

where the total share of compliers is given by $E[D_{SC}|Z = 1] - E[D_{SC}|Z = 0]$ which denotes the share of students changing their choices towards SC programs as the local

availability of SC HEI changes.

To estimate the complier shares, we use propensity scores for each enrollment choice in $d \in \{SC, BP, NE\}$ with a linear probability model¹¹ as follows:

$$D_i = \gamma_0 + \gamma_1 Z_i + \gamma_X X_i + \epsilon_i$$

where D_i is a categorical variable that takes on three values: not enrollment, SC programs enrollment, and BP enrollment; ϵ_i is an unobserved component. Table 7 and Table 8 show the results from the first-stage, for each enrollment choice. Women have a lower probability of enrolling in SC programs, in 2.6 percentage points vs. male students. Moreover, increases in age, the number of siblings, and maternal education decrease the probability of choosing a SC program. In terms of academic performance in Saber 11, increases in one standard deviation in Math or Physics decreases the probability of choosing SC programs. The opposite effect appears for Reading, History, and Philosophy subjects. In contrast, higher scores in Saber 11 across all subjects increase the probability of enrolling in BP and decrease the probability of NE.

Table 8 shows the effects of municipal characteristics and of the local availability of HEI. In particular, we observe that having a HEI specialized in SC programs in a 10km radius, increases the probability of choosing SC programs in 3.1 percentage points, and decreases the probability of choosing BPs in 2.7 percentage points. Furthermore, we observe that changes in the local availability of SC HEI have no significant effect on the probability of not enrolling in higher education. The magnitude of this effect is also small, at 0.04 percentage points. It is worth noting that other type of HEI (for instance, universities) seem to shift students along the NE-BP margin, as well as along the diversion margin (away from short-cycle programs and into bachelor's programs). Nonetheless, the latter effect is small in magnitude (about 0.9 percentage points).

We use the previous results to estimate the share of $SC - BP$ compliers (or, students switching along the *diversion margin*). Table 9 shows that diversion compliers represent the majority of compliers shifting towards SC, *i.e.*, $\hat{\omega} = 0.881$. We estimate these shares for women and men, and observe similar patterns. Importantly, the prevalence of SC-BP compliers is higher among men (at 3.5%) than among women (at 2%). These results are consistent with the descriptive analysis in the previous section, where we observe that women are less likely to enroll in SC programs. Overall we see no evidence of Z , local availability of SC HEI, inducing students to change their choices along two margins but rather one. We use these findings to argue that LATEs estimated with the variation of Z

¹¹Given that we have a multiple choice model, we also estimate the propensity scores with a multinomial logit. Our results for the effect of the instrument on each enrollment choice are similar with a linear probability model and with a multinomial logit (see Table 17 in the Appendix).

are likely to be mostly driven by diversion compliers that by a combination of subLATES for expansion and diversion compliers.

3.2 LATEs

In this section, we estimate LATEs for compliers induced into short-cycle programs as the local availability of SC HEI changes. Importantly, our first-stage in equation (2) includes Z as well as variables for the availability of other HEIs (*i.e.*, those that offer other programs and are not fully specialized in SC degrees). If some students would also change their enrollment choice as the availability of other HEIs changes, standard 2SLS does not identify the effect of one option *versus* the next-best, instead the effect it uncovers is a weighted average of subLATEs for all compliers. Table 8 shows that as the availability of other HEI changes, some students are induced to switch between BP and NE. Note that these students would not switch towards, or away, of SC programs which is the change in behavior that interests us in this paper. Nonetheless, standard 2SLS would attribute part of the subLATE for the BP-NE margin to the effect of SC programs. To deal with this bias component in standard 2SLS, we employ the assumption of *Partial Monotonicity* in Mountjoy (2019) and Mogstad et al. (2020a,b):

Let Z^- denote other variables that explain choices in equation (2) and are excluded from the outcome equation in (1). Let $D_{di}(Z = z|Z^-)$ denote the potential enrollment choice of student i when $Z = z$, keeping Z^- fixed. In our setting of SC programs, partial monotonicity states that, conditional on Z^- , all students should be induced towards, or away from, SC programs as Z changes: $D_{NEi}(0|Z^-) \geq D_{NEi}(1|Z^-)$, $D_{BPi}(0|Z^-) \geq D_{BPi}(1|Z^-)$, and $D_{SCi}(0|Z^-) \leq D_{SCi}(1|Z^-)$.

With this assumption, we isolate the effect of SC programs *versus* the next-best (which, in our case, denotes a mixture of NE and BP), as a result of the variation in Z :

$$LATE_{SC-/SC} = \frac{E(Y|Z = 1, Z^-) - E(Y|Z = 0, Z^-)}{E(D_{SC}|Z = 1, Z^-) - E(D_{SC}|Z = 0, Z^-)}$$

We estimate this effect with regressions of Y and D_{SC} on Z , Z^- , and X . As we discussed in Section 2.3, $LATE_{SC-/SC}$ is a combination of subLATEs for SC-NE compliers and SC-BP compliers (see equation (3)). However, given that we find no evidence of compliers along the SC-NE margin we attribute most of the effect on $LATE_{SC-/SC}$ to the SC-BP margin.

Table 10 shows the estimates of $LATE_{SC-/SC}$ on the probability of working and average monthly wages in 2013. SC programs increase the likelihood of working in the formal labor market by about 19 percentage points, but this effect is not statistically significant.

In turn, we observe a negative effect on average monthly wages, contrary to our findings with standard 2SLS (in Table 6). In terms of heterogeneity, we observe positive effects of SC enrollment on labor market participation for women and men, but with a higher magnitude and significance among the former. In contrast, we observe negative effects on average monthly wages in 2013 among men and a positive but not significant effect among women. Overall, we observe positive gains of choosing SC programs in terms of labor market participation among compliers, which are likely to be driven by compliers along the diversion margin. Among women compliers, we estimate large and statistically significant gains of switching towards SC programs. In contrast, we observe negative effects on wages for men compliers, who might not experience any gains in terms of increasing their participation in the formal labor market by choosing an SC instead of a BP.

To explain the larger gains for women vs. men, as well as the overall positive effects of SC programs on formal labor market participation, we estimate the average initial (before enrolling in higher education) characteristics of compliers. We focus on their performance in Saber 11, by computing the share of compliers above and below the median test score. Figure 1 shows the share of compliers in each group, for the full sample and by gender. For reference, we added (in green) the share of students who are enrolled in short-cycle programs in our sample and belong to each group. With the exception of Geography, there are no large differences in the distribution of the compliers above and below the median test scores in the full sample vs. the distribution for the sample of short-cycle students. In contrast, men compliers are more likely to outperform women compliers in all subjects, particularly in Math, Biology, and Physics. In addition, Table 11 shows that women compliers are more likely to belong to low income households and have mother's with fewer years of education than men compliers.

3.3 Responses along the *expansion* margin

In the previous section, we find that changes in the existence of short-cycle programs in the student's high school municipality would induce some students to divert from bachelor's programs and into short-cycle programs, but would not induce students along the SC-NE (expansion) margin. These results contrast with findings for the U.S.: for the state of Texas, Mountjoy (2019) shows that about 60% of compliers induced towards two-year college because of variation in distance to this option correspond to expansion margin compliers. Our estimates for Colombia, in turn, show that expansions of the supply of SC HEI are likely to divert some students from BP rather than induce students to enroll in HE with a SC program.

In this section, we explore the local labor market conditions for short-cycle degrees as a

potential driver to induce students along the SC-NE margin. There are other constraints that might restrict NE students from choosing SC programs. For instance, tuition and other enrollment and attendance costs could deter NE students from enrolling in a HE program.¹² Given data limitations on HE costs, we explore the role of local employment prospects which could also affect the student’s decision of not to enroll in higher education programs.

We use Census data for 2005 from IPUMS on municipal employment rates by educational level, for individuals 28-33 years old. IPUMS contains data for 445 municipalities, of which only 59 are also in our sample. As a result, we lose about 45% of our original estimating sample. Nonetheless, we perform the analysis in this subset to explore if students respond to local changes in employment when making enrollment decisions. Our main hypothesis is that changes in employment rates for SC degrees could make enrolling in a SC program more attractive. In particular, we define a binary instrument that takes the value of one if the local employment rate (at the student’s high school municipality) is higher than the department employment rate, for individuals of ages 28-33 with SC degrees. Formally, let Z_E be the binary instrument of local employment rate above the department level.

Our smaller sample largely differs from the total sample in terms of the availability of HEI; the vast majority of students in the subsample has available (in a 10km radius) a HEI that offers other types, not SC, of programs. Hence, we further restrict the subsample to students who had a HEI offering other programs in a 10 km radius, which removes 1% of the observations. Among these students, we estimate propensity scores for each choice (SC, BP, and NE) on student’s characteristics, fixed effects for region and metropolitan areas, as well as the IV of local employment in Z_E and a similar variable of employment rates for HS graduates.

Table 12 shows that students in municipalities with higher SC employment rates have a higher probability of enrolling in SC programs, on 8.8 percentage points, *versus* students in municipalities with lower than at the department level SC employment rates. There is no evidence of effects of differences in the local *versus* department level employment rates on the probability of choosing BP, rather the increase in SC enrollment seems to be driven by a decrease in the probability of not enrolling in higher education. The latter effects is of almost 10 percentage points. Hence, favorable employment conditions for SC degrees could drive students along the expansion margin.

¹²Although we have no direct measures of tuition for our cohort, in 2005, we explore results with two sets of proxies: imputed tuition to 2005 prices using data from 2009, and the public character of some HEIs. The public character of HEI is a potential proxy for tuition, given that public HEI are typically cheaper than private HEI (by almost half, González-Velosa et al. 2015). Our results using these measures do not present different enrollment patterns than those we described in this paper.

While the effect of Z_E on the expansion margin is an important result, we find no empirical evidence to be confident in estimating LATEs using the variation of Z_E . First, the direction of the effect of Z_E on the probability of choosing BP is the opposite to what we would expect: rather than decreasing enrollment in BP, such that students substitute BP for SC, it increases it. Second, the results of diagnostics for weak instruments on the first-stage do not pass strict criteria (although some statistics are above the frequently used criteria of $F > 10$). Hence, while we present results in Table 13 of estimating the effect of SC programs on labor market outcomes by OLS and with Z_E as instrument, we see these results as informative rather than causal effects. Overall, SC might not improve the labor market participation and experience of students who would have not enrolled in higher education but switch to a SC program because of changes in local employment prospects.

4 Discussion

This paper discusses the empirical challenges associated with the estimation of the economic returns to short-cycle degrees, namely self-selection of students into multiple enrollment options. We focus on the context of Colombia; in our setting, students can choose not to enroll in higher education, to enroll in a bachelor's program, or to enroll in a short-cycle program. In line with the literature on returns to higher education in the U.S., (Mountjoy, 2019; Rouse, 1995; Denning, 2017), we argue that the effects of short-cycle programs largely differ depending on the fallback alternative of students. That is, if short-cycle programs were not available some students would have chosen a bachelor's program while other students would have deter from entering higher education. These students, that differ in terms of their fallback option, can derive different effects from choosing a short-cycle degree. We use an Instrumental Variables (IVs) approach, and exploit variation in the supply of short-cycle degrees at the municipal level. In particular, we define a binary IV that takes the value of one if the student when to high school in a municipality where short-cycle degrees were offered in a radius of 10km.

We uncover three main findings: only 3% of students react to the variation in the local supply of short-cycle programs. But we find that among these students almost 89% would divert from bachelor's programs and into short-cycle degrees. The remaining 11% corresponds to students who would be induced to enter the higher education system with a short-cycle degree as a result of changes in the local supply. We also observe similar patterns among men and women, with shares of 99% and 75%. These findings are in stark contrast with previous evidence for the U.S.: for the state of Texas, Mountjoy (2019) finds that about two thirds of students would react to changes in proximity to two-year college

by switching along the expansion margin (*i.e.*, from not enrolling in higher education towards enrolling in a two-year college); the remaining 30% would divert from four-and into two-year college (along the diversion margin). Along these lines, [Denning \(2017\)](#) finds that reduced tuition for two-year college increases higher education enrollment, rather than decreasing enrollment in universities.

Moreover, we estimate that short-cycle degrees can improve the labor market participation rates of students along the diversion margin. These results are larger for women, and we find no significant effects on labor market participation among men. In addition, we observe detrimental diversion effects on average monthly wages for men. That is, men who would have enrolled in bachelor's programs in the absence of short-cycle programs would have obtained higher average wages in the former than in the latter, on average. Along these lines, [Mountjoy \(2019\)](#) finds that students along the expansion margin are most likely to benefit from two-year colleges in terms of average earnings and years of education, than students along the diversion margin. It is worth noting that while we find negative effects of SC enrollment on average monthly wages, these results are only significant among men who react to the variation in the instrument.

Differences in the role of short-cycle degrees *versus* two-year college degrees might explain our findings in contrast with the literature: while two-year college can still serve as a pathway into four-year college enrollment, short-cycle degrees are mostly focused in preparing students for the transition into work, rather than into bachelor's degrees. Moreover, students from countries such as Colombia might face stricter constraints which prevents them to enter the higher education system, compared to students in the U.S. As part of our exploratory analysis, we also offer evidence to show that local employment prospects can be an important driver to induce students to enroll in higher education via short-cycle degrees.

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5 Tables and Figures

Table 1: Summary Statistics

Variable	Mean	SD
Female	0.538	0.499
Age at Saber 11	17.019	1.583
Siblings	2.682	1.639
<i>Household Income Level</i>		
<1 Minimum Wages (MW)	0.272	0.445
1-2 MW	0.432	0.495
2-3 MW	0.159	0.366
>3 MW	0.137	0.344
<i>Mother's Education Level</i>		
Primary	0.423	0.494
Secondary	0.354	0.478
Short-Cycle program	0.109	0.312
At least Bachelor's program	0.113	0.317
<i>Higher Education Enrollment</i>		
Not enrolled (NE)	0.612	0.487
Short-Cycle program (SC)	0.095	0.293
Bachelor's program (BP)	0.293	0.455
<i>Educational Attainment</i>		
High School Graduate	0.612	0.487
Short-Cycle Incomplete	0.047	0.211
Short-Cycle Complete	0.036	0.186
Bachelor's Incomplete	0.156	0.363
Bachelor's Complete	0.149	0.356
<i>Formal Labor Market Outcomes</i>		
Works in 2013	0.5004	0.499
Avg. Monthly Wage (2013) [N=161,154]	764,961.6	626,623.6
N	322,537	

Note: The sample corresponds to the universe of students who took the high school Exit Exam in 2005. The information on educational attainment comes from SPADIES, and the labor market outcomes for higher education graduates is from OLE. For high school graduates and those with higher education incomplete, we impute formal labor market participation and experience using household survey data from SEDLAC. Monthly wages from OLE are in COP which are available for students who were working in 2013.

Table 2: Average characteristics, by type of enrollment

Variable	Short-Cycle Program	Bachelor's Program	Not Enrolled
Female	0.480	0.548	0.542
Age at Saber 11	16.777	16.543	17.284
Siblings	2.329	2.102	3.014
<i>Household Income Level</i>			
<1 MW	0.186	0.148	0.344
1-2 MW	0.506	0.349	0.461
2-3 MW	0.206	0.201	0.131
>3 MW	0.103	0.302	0.063
<i>Mother's Education Level</i>			
Primary	0.366	0.217	0.531
Secondary	0.432	0.323	0.357
Short-Cycle program	0.136	0.179	0.072
At least Bachelor's program	0.066	0.281	0.040
<i>Standardized Test Scores from the High School Exit Exam</i>			
Math	0.004	0.366	-0.148
Reading	0.132	0.561	-0.240
Biology	0.044	0.511	-0.218
Physics	0.005	0.345	-0.144
History	0.070	0.469	-0.202
Chemistry	0.030	0.510	-0.215
Geography	0.057	0.430	-0.182
Philosophy	0.064	0.412	-0.173
<i>Education Attainment</i>			
SC Incomplete	0.494		
SC Complete	0.342	0.012	
BP Incomplete	0.102	0.501	
BP Complete	0.062	0.487	
<i>Formal Labor Market Outcomes</i>			
Works in 2013	0.512	0.536	0.481
Avg. Monthly Wage (2013)	911,114.4	1,127,179	548,288
N	30,514	94,583	197,440

Note: The sample corresponds to the universe of students who took the high school Exit Exam in 2005. The information on educational attainment comes from SPADIES, and the labor market outcomes for higher education graduates is from OLE. For high school graduates and those with higher education incomplete, we impute labor market participation and experience using household survey data from SEDLAC. Monthly wages from OLE are in COP.

Table 3: Average characteristics of HS municipality, by type of enrollment

Variable	Short-Cycle	Bachelors'	Not Enrolled	All
	Program	Program		
Total Population (in millions)	2.356	2.110	1.661	1.863
Area, in squared km	856.81	972.048	891.198	910.72
Ratio of Urban/Rural pop.	146.61	137.32	107.29	120.09
Linear distance to the main city, km	22.627	22.882	34.545	29.997
Poverty Index	0.358	0.364	0.399	0.385
Homicide rate (per 1,000 inhabitants)	0.350	0.350	0.386	0.372
GDP per capita (millions COP)	9.289	8.603	8.095	8.357
N	30,514	94,583	197,440	322,537

Note: The sample corresponds to the universe of students who took the high school Exit Exam in 2005. The information on local characteristics comes from the Municipal Panel from CEDE (*Centro de Estudios de Desarrollo Económico*, in Spanish).

Table 4: OLS Results: Labor Market Outcomes

Variable	Prob(Working)	Avg. Monthly Wage
	2013	2013
$\hat{\beta}_{SC}$	-0.0442*** (0.0047)	81,508.673*** (5,271.618)
N	322,537	161,154
R^2	0.4833	0.2103

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include baseline characteristics at the individual and municipal level, as well as department fixed effects. The last two columns show results for average monthly wages in COP. Each outcome is estimated separately.

Table 5: Availability of Higher Education Institutions (HEIs), by type of programs offered

Variable	Short-Cycle Program	Bachelors' Program	Not Enrolled	All
Z: HEI only offers SCP	0.742	0.654	0.570	0.611
Other type of HEI	0.849	0.851	0.740	0.783
Enrollment in SENA > 0 in 2004	0.798	0.819	0.710	0.750
N	30,514	94,583	197,440	322,537

Note: The sample corresponds to the universe of students who took the high school Exit Exam in 2005. The information on higher education programs and institutions is from SNIES (*Sistema Nacional de Información de la Educación Superior*, in Spanish).

Table 6: 2SLS Results: Labor Market Outcomes

Variable	Prob(Working) 2013	Avg. Monthly Wage 2013
Short-cycle program ($\hat{\beta}_{SC}$)	0.4262*** (0.1425)	170,582.592 (170607.298)
<i>First-stage: Prob(Enrolling in a short-cycle program)</i>		
Z: Short-cycle HEI ($\hat{\gamma}_1^{SC}$)	0.0311*** (0.003)	0.0293*** (0.004)
N	322,537	161,154
Kleibergen-Paap F-stat	43.139	30.6
Cragg-Donald F-stat	113.177	58.822
Effective F-stat	46.374	

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include baseline characteristics at the individual and municipal level, as well as department fixed effects. Each outcome is estimated separately.

Table 7: First-stage: individual characteristics

Variable	Short-Cycle Program	Bachelor's Program	Not Enrolled
Female	-0.0257*** (0.0014)	0.0316*** (0.0020)	-0.0058*** (0.0021)
Age at Saber 11	-0.0075*** (0.0004)	-0.0202*** (0.0006)	0.0277*** (0.0007)
Siblings	-0.0053*** (0.0007)	-0.0097*** (0.0012)	0.0150*** (0.0014)
<i>Household Income</i>			
1-2 MW	0.0110*** (0.0022)	0.0196*** (0.0034)	-0.0305*** (0.0038)
>2 MW	0.0026 (0.0032)	0.0629*** (0.0058)	-0.0655*** (0.0066)
<i>Mother's level of education</i>			
Secondary	0.0168*** (0.0022)	0.0530*** (0.0032)	-0.0698*** (0.0036)
Higher Education	-0.0180*** (0.0036)	0.2380*** (0.0070)	-0.2200*** (0.0068)
<i>Standardized Test Scores from the High School Exit Exam</i>			
Math	-0.0043*** (0.0006)	0.0207*** (0.0008)	-0.0164*** (0.0009)
Reading	0.0058*** (0.0007)	0.0453*** (0.0010)	-0.0511*** (0.0010)
Biology	-0.0017** (0.0007)	0.0291*** (0.0009)	-0.0274*** (0.0010)
Physics	-0.0038*** (0.0006)	0.0151*** (0.0008)	-0.0113*** (0.0009)
History	0.0010 (0.0007)	0.0245*** (0.0009)	-0.0255*** (0.0009)
Chemistry	-0.0031*** (0.0007)	0.0281*** (0.0010)	-0.0250*** (0.0010)
Geography	-0.0014** (0.0006)	0.0221*** (0.0008)	-0.0207*** (0.0009)
Philosophy	0.0016*** (0.0006)	0.0163*** (0.0009)	-0.0180*** (0.0009)
N	322,537	322,537	322,537

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include department fixed effects. Each choice is estimated separately with a Linear Probability Model.

Table 8: First-stage (cont'd): HS municipality characteristics and proximity to HEIs

Variable	Short-Cycle Program	Bachelor's Program	Not Enrolled
Total Population (in millions)	0.0194*** (0.0029)	-0.0083** (0.0040)	-0.0111** (0.0046)
Area, in squared km (10,000)	0.0541*** (0.0120)	0.0350 (0.0241)	-0.0891*** (0.0245)
Urban/Rural population (per 100)	0.0094*** (0.0019)	-0.0032 (0.0030)	-0.0062** (0.0029)
Distance to main city (100 km)	-0.0049* (0.0027)	0.0160*** (0.0043)	-0.0111** (0.0046)
Poverty Incidence	0.0540*** (0.0180)	-0.0046 (0.0269)	-0.0493* (0.0282)
GDP per capita (millions, COP)	0.0003 (0.0004)	0.0007 (0.0004)	-0.0010* (0.0005)
Homicide rate, per 1000 inhabitants	0.0044 (0.0041)	-0.0247*** (0.0067)	0.0203*** (0.0070)
	<i>Instrument Z :</i>		
SC HEI in 10 km radius	0.0311*** (0.0033)	-0.0274*** (0.0048)	-0.0037 (0.0052)
	<i>Availability of other HEI:</i>		
Not specialized in SC, or only BP	-0.0089*** (0.0028)	0.0192*** (0.0043)	-0.0103** (0.0048)
SENA enrollment > 0, 2004	-0.0101*** (0.0025)	0.0137*** (0.0038)	-0.0035 (0.0042)
Constant	0.2337*** (0.0132)	0.5434*** (0.0188)	0.2228*** (0.0208)
N	322,537	322,537	322,537

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include department fixed effects. Each choice is estimated separately with a Linear Probability Model.

Table 9: Share of compliers along the expansion and diversion margins, for the full sample and by gender

	Expansion margin Compliers: NE-SC	Diversion margin Compliers: BP-SC	$\hat{\omega}$
All	0.0036 (0.0052)	0.0273*** (0.0048)	0.881*** (0.164)
Female	0.0067 (0.0066)	0.0208*** (0.0061)	0.754*** (0.227)
Male	0.0001 (0.0064)	0.0347*** (0.0057)	0.996*** (0.185)

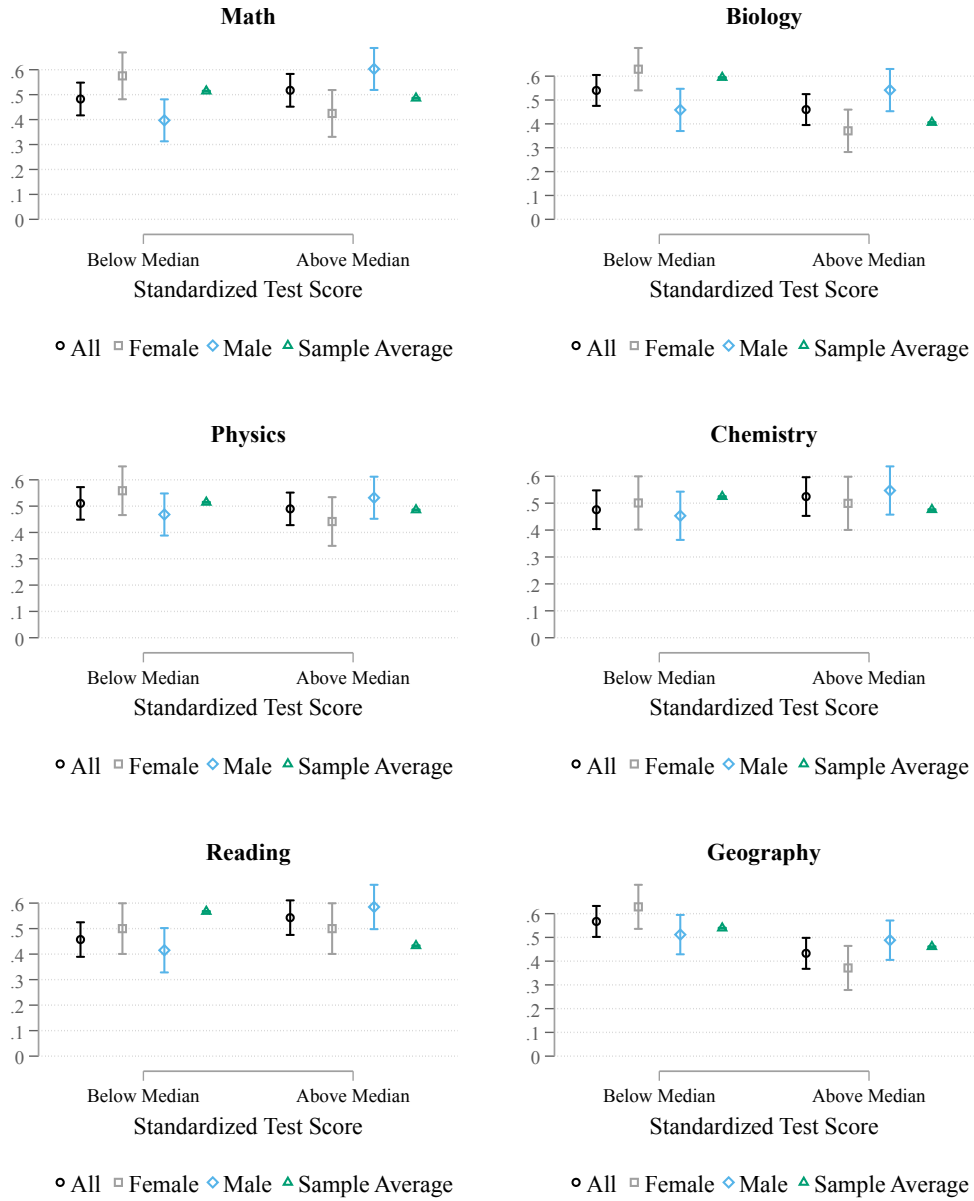
Note: The Table shows the estimated shares of students who react to the variation in the local supply of short-cycle HEI by: switching from not enrolling towards short-cycle programs, and by diverting from bachelor's and into short-cycle programs. $\hat{\omega}$ is the ratio of the effect in the third column ("Diversion margin"), over the sum of column two and three. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level.

Table 10: Local Average Treatment Effects for compliers along the diversion margin, full sample and by gender

$LATE_{SC-/SC}$	Prob(Working) 2013	Avg. Monthly Wage 2013
All	0.1819 (0.1590)	-397,294.221* (236,072.105)
Female	0.4912** (0.2257)	323,098.777 (686,136.872)
Male	0.0298 (0.1600)	-479,047.728** (204,213.606)
Observations	322,537	161,154

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. $/SC$ refers to a combination of other (not short-cycle programs) enrollment options (such as choosing a bachelor's program and choosing not to enroll in higher education). Thus, $LATE_{SC-}$ denotes the local avg. treatment effect of short-cycle programs vs. the next-best alternative. The last column denotes effects for average monthly wages in COP.

Figure 1: Share of compliers in the total sample and by gender, below and above the median of standardized test scores



Note: The figure shows the share of students that belong to each test score group, among compliers in the full sample and by gender. The green triangles denote the share of students enrolled in short-cycle programs in the sample that belong to each standardized test score group.

Table 11: Average household level characteristics of compliers, full sample and by gender

Variable	All	Female	Male	Sample Average
Household Income: <2 MW	0.506 (0.0829)	0.565 (0.0976)	0.452 (0.0967)	0.691
<i>Mother's level of education</i>				
Primary	0.247 (0.0967)	0.250 (0.1201)	0.245 (0.1010)	0.366
Secondary	0.521 (0.1015)	0.553 (0.122)	0.491 (0.1088)	0.432
Higher Education	0.232 (0.068)	0.197 (0.0812)	0.264 (0.0823)	0.201

Note: Standard errors in parentheses, clustered at the high school level. Urban municipality refers to the student's high school municipality where the share of urban population over the total population is above 83% (the sample average). Column (1) - (3) show the proportion of compliers that belong to each category, for the total sample and by gender. The last column corresponds to the sample average of each characteristic among students enrolled short-cycle programs.

Table 12: Subsample first-stage: HEI availability and Local employment conditions

Variable	Short-Cycle Program	Bachelor's Program	Not Enrolled
<i>Local employment rates > department rate for ages 28-33, by education level</i>			
High School	-0.0507 (0.0482)	-0.0826*** (0.0207)	0.1332*** (0.0343)
Short-Cycle programs (Z_E)	0.0884*** (0.0235)	0.0111 (0.0293)	-0.0995*** (0.0270)
N	150,525	150,525	150,525
Compliers NE-SC/total variation	0.887*** (0.275)		

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include metropolitan area (Carranza and Ferreyra, 2019) fixed effects, and controls at the municipal level. Each choice is estimated separately.

Table 13: Subsample 2SLS Results: Labor Market Outcomes

Variable	Prob(Working) 2013		Wage 2013	
	OLS	2SLS	OLS	2SLS
Short-cycle program ($\hat{\beta}_{SC}$)	0.0085 (0.0061)	-0.1267 (0.3182)	64,703.074*** (6157.3397)	1,973,675.198 (1308890.5649)
<i>First-stage: Prob(Enrolling in a short-cycle program)</i>				
Z_E : local employment rate of short-cycle grads. ($\hat{\gamma}_E^{SC}$)	0.0774*** (0.0141)		0.0457 (0.0291)	
N	150,525		84,069	
Kleibergen-Paap F-stat	20.967			
Cragg-Donald F-stat	6.371			
Effective F-stat	16.359			

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include metropolitan area (Carranza and Ferreyra, 2019) fixed effects, and controls at the municipal level. Each choice is estimated separately.

Appendix

A.1 Data restrictions

The master sample consists of the universe of test-takers of Saber 11 in 2005, who were between 14-24 years of age and account for 93.5% of the full set of students taking the test in that year. We restrict the sample to students who have data on all socioeconomic variables and test scores, which removes only 0.3% of students. We matched data on enrollment on higher education from SPADIES to the master sample, and restrict the start date of the first program recorded to be after the semester of 2005 when the student took the test, which removes 1.1% of the observations. We set an enrollment window of five years after graduation, and remove students who started their program on or after 2012 (about 3% of the sample). We also restrict the graduation year to be no less than two years- for those in short-cycle programs- and four years- for those in four-year college (removes 1.5% observations). In addition, we set the graduation age to be between 19-30 which reduces the sample by 0.1%. We remove 0.004% of the sample which corresponds to those who switched, or enrolled, in more than four programs between 2005 and 2015.

We further restrict the sample of those who graduated and have data on wages from OLE to start working at, or after, 20 years of age (removes 0.25% of the sample). For almost 4.8% of students who appear in OLE, enrollment and graduation do not match with SPADIES: two percent show as graduated in OLE but only enrolled in SPADIES, while the remaining 2.8% graduated according to OLE but do not report ever enrolling in

a higher education institution in SPADIES. We removed the latter 2.8% from the sample, given that they never report enrolling in SPADIES, but since OLE contains information on wages we assume that the 2% who enrolled but never graduated in SPADIES actually did so. Lastly, 4% of those who graduated in SPADIES do not appear in OLE which means that for this fraction of students there is no information on wages. The final dataset consists of 369,541 students, of a total of 403,209 initially in the master sample (about 91% of students who took Saber 11 in 2005 are in our estimating sample).

Finally, we restrict our estimating sample to high schools with more than 20 students. We do so since our instruments are constructed based on the municipality of the student's high school, and we estimate with clustered standard errors at this level. This restriction leaves an estimating sample of 322,537 observations.

A.2 Imputation of wages and labor market participation

To impute participation in the labor market and average monthly wages, we use household survey data between 2008 and 2013. We use the set of homogenized household surveys in SEDLAC, which correspond to the Integrated Household Survey (GEIH, in Spanish). We restrict each sample to those individuals who were 14-24 old in 2005. Let Y denote labor market variables such as participation and wages.

$$Y = \beta_0 + \beta_X X + u$$

For labor market participation, we estimate the equation above with a LPM. X contains sex, age, number of household members, household income level, a dummy that takes the value of one if the individual lives in an urban area, and regional fixed effects (results in Table 14). For average wages, we estimate the equation above with a quantile regression and predict coefficients for the 25th, 50th, and 75th percentile. In this case X contains sex, age, age squared, number of household members, household income level, a dummy that takes the value of one if the individual lives in an urban area, and regional fixed effects (results are in Table 15 and 16).

Table 14: Regression Results: LPM of probability of working in 2013

Variables	HS Graduates	HE Incomplete
Male	0.350*** (0.014)	0.135*** (0.017)
Age	0.013*** (0.003)	0.027*** (0.003)
Number of members in main household	-0.019*** (0.005)	-0.028*** (0.005)
1 – 2 MW	0.080*** (0.027)	0.140*** (0.043)
2 – 3 MW	0.154*** (0.031)	0.254*** (0.043)
> 3 MW	0.190*** (0.031)	0.340*** (0.040)
Urban area	0.020 (0.022)	-0.030 (0.035)
Region: Oriental	0.068** (0.028)	0.040* (0.023)
Region: Central	-0.023 (0.025)	0.038* (0.021)
Region: Pacífica	0.000 (0.028)	-0.015 (0.025)
Region: Bogotá	0.067* (0.034)	0.067** (0.027)
Observations	4,722	8,141

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses.

Table 15: Regression Results: hourly wage in main occupation

Variables	High School Graduates		
	q25	q50	q75
Male	461.627*** (100.972)	434.299*** (105.942)	172.838 (143.657)
Age	437.637 (284.343)	131.550 (273.086)	-6.856 (328.158)
Age ²	-8.053 (5.356)	-1.763 (5.101)	0.952 (6.104)
Number of members in main household	-117.627*** (16.687)	-121.458*** (17.445)	-173.951*** (24.616)
1 – 2 MW	1,203.005*** (112.909)	1,318.647*** (128.563)	1,507.471*** (158.184)
2 – 3 MW	1,419.765*** (150.009)	1,598.469*** (138.681)	2,040.413*** (180.284)
> 3 MW	2,023.496*** (142.385)	2,102.490*** (153.329)	2,847.308*** (277.423)
Urban area	-236.529** (102.791)	-315.042** (151.966)	-756.664*** (237.149)
Region: Oriental	315.999** (135.960)	367.526*** (123.872)	197.873 (142.295)
Region: Central	155.621 (128.751)	236.217 (152.566)	354.958** (164.369)
Region: Pacifica	-110.810 (135.958)	-113.379 (116.014)	-47.057 (231.480)
Region: Bogotá	655.060*** (189.608)	510.973*** (150.499)	326.354* (196.395)
Constant	-4,739.014 (3,726.144)	-359.963 (3,609.618)	2,686.551 (4,344.162)
Observations	3,090	3,090	3,090

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level.

Table 16: Regression Results: hourly wage in main occupation

Variables	Higher Education Incomplete		
	q25	q50	q75
Male	240.217*** (89.161)	342.924*** (86.817)	423.750** (184.014)
Age	680.753** (265.982)	210.822 (295.035)	-61.922 (509.594)
Age ²	-11.772** (4.981)	-2.368 (5.542)	3.341 (9.509)
Number of members in main household	-208.221*** (26.766)	-277.203*** (21.133)	-438.176*** (41.265)
1 – 2 MW	1,503.986*** (182.953)	1,853.146*** (159.451)	1,342.269*** (265.630)
2 – 3 MW	2,170.080*** (192.120)	2,771.923*** (173.614)	2,532.960*** (278.575)
> 3 MW	3,077.149*** (151.087)	3,808.032*** (160.508)	4,681.549*** (285.479)
Urban area	-408.050*** (117.800)	-308.839*** (92.751)	-229.820 (269.067)
Region: Oriental	341.604** (141.283)	368.984*** (134.636)	687.261** (293.303)
Region: Central	31.376 (119.496)	-175.898* (90.773)	-285.977 (193.988)
Region: Pacifica	152.830 (139.976)	183.826 (154.098)	103.817 (225.862)
Region: Bogotá	378.608*** (127.699)	243.773** (124.243)	155.136 (227.950)
Constant	-8,028.344** (3,506.050)	-1,587.394 (3,893.089)	3,238.408 (6,742.760)
Observations	5,378	5,378	5,378

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level.

Table 17: First-stage: Multinomial logit model

Variable	Short-Cycle Program	Bachelor's Program	Not Enrolled
	<i>Instrument Z :</i>		
SC HEI in 10 km radius	0.0301*** (0.0032)	-0.0249*** (0.0051)	-0.0052 (0.0053)
	<i>Availability of other HEI:</i>		
Not specialized in SC, or only BP	-0.0096*** (0.0036)	0.0163*** (0.0045)	-0.0067 (0.0049)
SENA enrollment > 0, 2004	-0.0090*** (0.0031)	0.0114*** (0.0040)	-0.0024 (0.0043)
N	322,537	322,537	322,537

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Standard errors in parentheses, clustered at the high school level. All regressions include department fixed effects, and controls at the municipal level.