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The Effect of the Availability of Student Credit on Tuition: Testing the Bennett Hypothesis Using Evidence from a Large-Scale Student Loan Program in Brazil

ABSTRACT Exploring the expansion of FIES—a large student lending program in Brazil—we test whether eligibility for subsidized student lending causes tuition to rise, in accordance with the Bennett hypothesis. FIES rules created arguably exogenous variation in eligibility across different majors and higher education institutions, which we exploit in a difference-in-differences framework. Using unique information on tuition, we document that FIES eligibility caused tuition to rise. We then estimate a structural demand model to explore whether a reduction in the sensitivity of demand to price increases is one of the possible mechanisms behind this credit-driven tuition rise. Our results show that FIES expansion is associated with a reduction in the tuition elasticity of demand.

JEL Codes: D04, D12, I22

Keywords: Student lending, tuition inflation, policy evaluation

An extensive literature documents the relation between investment in higher education and development. This literature shows a strong correlation between greater investment in higher education and increases in the skill level of the workforce, research levels, development of new technologies, and productivity gains.[1] Governments in both developed and developing countries have implemented a number of strategies aimed at increasing enrollment rates in higher education.

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1. See Task Force on Higher Education and Society (2000).

Understanding the costs and benefits of such policies is particularly important for policymakers in developing countries. Only a small portion of the workforce in the developing world has completed any form of tertiary education.[2] In Argentina, the rate is just 35.7 percent; in Colombia, 23.4 percent; in Mexico, 18.0 percent; and in Brazil, 17.2 percent.[3] The workforce in developed countries has a significantly higher level of education. According to the Organization for Economic Cooperation and Development (OECD), 37 percent of the population in OECD countries has some form of tertiary education. To increase the educational level of their workforce and catch up with more developed economies, developing countries need to boost enrollment. So far, however, there is no indication that the developing world is catching up. In 2018, the gross enrollment rate in tertiary education in high-income countries was 75 percent; in low- and middle-income economies, 33 percent; and in Latin American and Caribbean countries, 52 percent.[4]

Unequal access is also a major concern. There is a strong correlation between family income and investment in higher education. In Latin America, for instance, students from the bottom two income quintiles represented only 16.8 percent of students enrolled in higher education in 2013.[5]

Though policymakers usually support the idea that greater investment in higher education might bring benefits for society, policy implementation is often constrained by budget considerations. Policymakers are often expected to design strategies capable of expanding enrollment rates, especially for students from low-income families, while not imposing a large fiscal cost on society. Designing policies that meet these requirements is a challenging task. Policymakers often resort to subsidized student lending programs to expand access without covering the full cost of tertiary education.

In the past decade, subsidized student lending programs were created and expanded in several developing countries, including some Latin American economies. In 2005, Chile introduced a student loan program for low-income

2. In China, only 9.7 percent of the population has some form of tertiary education; in India, 10.6 percent; and in South Africa, 7.2 percent.

3. The share of the population aged twenty-five to sixty-four, by educational attainment, in 2018. See stats.oecd.org/Index.aspx?DataSetCode=EAG_NEAC.

4. The gross enrollment ratio in tertiary education is a measure of total enrollment in education expressed as a percentage of total population of the age group which officially corresponds to tertiary education. This information is provided by the United Nations Educational, Scientific and Cultural Organization (UNESCO) Institute of Statistics. See data.worldbank.org/indicator/se.ter.enrr.

5. Ferreyra and others (2017).

individuals with good academic records, namely, the State-Guaranteed Loan Program. In 2010, 42 percent of tertiary education students with some form of student aid had a loan through this program. The government of Colombia, in turn, offers merit-based subsidized loans. In 2011, 20 percent of students enrolled in higher education in Colombia had a government student loan.

These policies were accompanied by a considerable increase in enrollment in the region. From 1999 to 2018, average gross enrollment in higher education went from 23 percent to 52 percent in Latin America.[6] Other outcomes have been mixed. Some countries are facing high default rates, raising questions about the ex post fiscal impact of these programs and the debt burden they impose on students. Ferreyra and others (2017) provide a comprehensive survey of the difficulties faced by Latin American countries in implementing student lending programs.

Tuition inflation raises student indebtedness and the propensity to default. It may also have an impact on the fiscal cost of these programs, which are either explicitly or implicitly backed by taxpayers. In this context, it is important to build a comprehensive body of evidence on the pricing consequences of student lending programs, especially for developing countries.

We measure the causal impact of eligibility for a government student lending program in Brazil on tuition. Policymakers have long raised the concern that student aid may translate into higher tuition, in line with the so-called Bennett hypothesis. In a famous *New York Times* article, William Bennett—then the U.S. secretary of education—asserted that student aid "enabled colleges and universities blithely to raise their tuitions, confident that Federal loan subsidies would help cushion the increase" (Bennett, 2017). For the United States, a small literature finds evidence of the Bennett hypothesis for cohorts that enrolled in higher education in the past decade. The evidence for developing countries is scarce, with Espinoza (2017) being a noteworthy exception. In this paper, we test the Bennett hypothesis using Brazilian data and find strong supportive evidence. In addition, we explore the mechanisms that lead colleges to alter their pricing behavior. Similar to Espinoza's findings, we find evidence that increasing student credit reduced the tuition elasticity of demand.

To test the Bennett hypothesis, we explore the ramp-up of the *Fundo de Financiamento Estudantil* (FIES), a large student lending program funded by Brazil's federal government. FIES offers loans to students enrolled in private

6. Source: UNESCO Institute for Statistics.

higher education institutions (HEIs) in Brazil. Created in 1999, the program did not become practically relevant until 2010, when it went through a major reform. Since then, the volume of FIES loans has increased consistently, from 33,000 new loans in 2009 to 560,000 in 2013.[7] In 2013, the ratio between the number of new loans and the number of students newly enrolled in private HEIs in Brazil was nearly one to three. During this period, there were also changes in the aggregate trend for tuition in Brazil. From 2009 to 2010, the year of the FIES ramp-up, tuition fees dropped 3.2 percent in real terms, remaining flat in 2011. Between 2011 and 2013 tuition fees increased almost 7 percent in real terms. Coincidentally or not, the stock of FIES-financed students jumped from 3.7 percent in 2010 to 13.0 percent in 2013.

In this paper, we evaluate whether the aggregate pattern, compatible with the Bennett hypothesis, persists when we implement a rigorous identification strategy designed to estimate the causal impact of FIES eligibility on tuition. The FIES ramp-up in 2010 had a heterogeneous impact on different major-HEI pairs. The eligibility rules determined by the legislation restricted access to FIES based on an arguably exogenous criterion, that is, a criterion that is unrelated to the pricing trends implemented by different types of institution. Specifically, right after the ramp-up, a major-HEI pair could enroll students financed through FIES only if it was considered of sufficient quality according to evaluations conducted by the Ministry of Education in the previous years. We argue that HEIs were not anticipating the expansion of FIES and had little control over their short-term performance on these quality evaluations. That is, in the first few years after FIES expansion, major-HEIs had no control over their exposure to the program. We define treatment and control groups according to these rules and implement a difference-in-differences (DD) strategy. We argue that our framework meets the assumptions required for identification.

Using a unique data set with annual information on tuition fees at the major-HEI level, we document two facts. First, eligibility for FIES at the major-HEI level had a strong impact on tuition. Our preferred specification shows that eligibility for FIES is associated with a 4.6 percent increase in tuition fees. This result is robust to the inclusion of a sizable set of controls. In our most saturated specification, FIES is associated with a 3.1 percent

7. Source: Ministry of Education, National Institute for Educational Studies and Research (INEP).

increase in tuition.[8] Second, estimating a simple structural demand model, we show that the expansion of FIES is associated with a reduction in the tuition elasticity of demand. This result indicates that a reduction in price sensitivity may be the mechanism behind the credit-fueled increase in tuition.

The evidence we obtain through our reduced-form approach reveals that HEIs increase tuition in response to being eligible to enroll students funded through student lending programs. Our structural analysis uncovers some of the possible mechanisms behind this increase. From a policy perspective, it is important to identify these mechanisms. If the availability of student lending does not change the tuition elasticity of demand, then the FIES-driven increase in tuition probably reflects the increased marginal costs of supplying tertiary education. The policy implications may be different if FIES reduces the tuition elasticity of demand.[9] This lower price sensitivity can increase rents for the tertiary education industry, with at least part of the government subsidies being transferred to private HEIs in the form of higher profits. Though unlikely in the case of Brazil, a lower price sensitivity can also result in overinvestment in tertiary education, an issue if individuals undertake human capital investment with a negative net present value. In Brazil, this is unlikely in light of the high returns to tertiary education.[10]

The next section reviews the literature to position our contribution. We then describe our unique data set, which contains information on tuition at the major-HEI level, as well as a rich set of major-HEI characteristics. Observing fees at this level of disaggregation is crucial to our identification strategy. After presenting the institutional background of FIES, including the operational and normative changes that occurred in early 2010, we outline our reduced-form estimation strategy and present our main reduced-form results. In essence, we explore a rule that produced arguably exogenous variation in

8. These estimates are likely to be a lower bound insofar as our strategy essentially compares the dynamics of tuition fees in eligible and ineligible major-HEI pairs, and prices tend to be strategic substitutes.

9. We cannot establish the reason why credit availability may reduce price sensitivity. One conjecture is that students expect to renegotiate with the government. Another is that students are overconfident about the returns to higher education, or, quite simply, may not fully understand the financial consequences of borrowing. Lusardi, Mitchell, and Curto (2010), for instance, present evidence of significant financial illiteracy among youth in the United States. The population of Brazil is more financially illiterate than that of the United States, suggesting that behavioral explanations are plausible in our setting (Lusardi and Mitchell, 2011).

10. Though still relatively high, there is evidence that premiums have dropped substantially over the last fifteen years (Ferreira, Firpo, and Messina, 2014).

FIES eligibility at the major-HEI level. We then show that our identification strategy is sound and that our results are robust to a different set of assumptions. Finally, we present a structural-form approach designed to investigate the mechanism behind our reduced-form results, wherein we estimate a differentiated-product demand system and find that FIES eligibility is associated with a reduction in the tuition elasticity of demand.

Related Literature

Our work relates to a large literature on the impacts of government-sponsored student lending programs. Most available papers investigate the impact of student credit on measures of student behavior, such as enrollment and dropout. Our work is directly related to a small but growing literature that investigates the effect of credit availability on prices, that is, tuition and other fees.

From a normative perspective, government-sponsored student lending is justified if students are credit constrained and thus underinvest in human capital. A large literature investigates the empirical relevance of borrowing constraints on schooling choices. Results have been mixed. Cameron and Heckman (1998) and Carneiro and Heckman (2002), using the U.S. National Longitudinal Survey of Youth 1979 (NLSY79), do not find evidence of borrowing constraints. Using more recent data, Kane (2006) and Belley and Lochner (2007) find evidence of borrowing constraints for students choosing to enroll in higher education in the United States. Though most of this literature focuses on the United States, there are exceptions. Solís (2017) evaluates the causal effects of two large college loan programs in Chile and finds strong evidence for the credit constraint argument, with access to loans effectively eliminating the income gap in enrollment and attainment in Chile.[11]

Lochner and Monge-Naranjo (2011) posit that the stronger relationship between family income and school attainment for more recent cohorts results from two factors: a substantial increase in both the costs and returns associated with higher education, combined with no change on the limits of government student loans in real terms. They touch on an important aspect that was until recently overlooked in the academic literature, namely, increased

11. It is beyond our scope to provide a thorough revision of the literature on borrowing constraints and schooling choices. Lochner and Monge-Naranjo (2011) provide an extensive review.

tuition costs. From 1984 to 2014, average posted tuition at private four-year institutions in the United States rose 146 percent in real term. At public two-year colleges, tuition increased 150 percent. At in-state public four-year institutions, mean tuition rose 225 percent.[12] Higher returns to education combined with increasing marginal costs may explain the rise in tuition costs, but credit availability is another culprit. Our work contributes to the growing literature investigating the impact of student lending programs on tuition fees.

A few early contributions to this literature are worth mentioning. Hoxby (1997) notes the Bennett hypothesis as a possible explanation for tuition increases in the United States, but finds no supportive evidence before 1991. Other early papers relating credit availability and tuition inflation are McPherson, Schapiro, and Winston (1989), Rizzo and Ehrenberg (2004), and Long (2004). In general, they find weak support for the Bennett hypothesis.

The interest in credit-fueled tuition hikes has increased recently, and a number of papers do find evidence that credit availability causes tuition inflation. Cellini and Goldin (2014), Lucca, Nadauld, and Shen (2019), and Gordon and Hedlund (2019) all find evidence of a credit-driven tuition increase for the U.S. higher education market. There is less evidence in developing countries. One noteworthy exception is Espinoza (2017), who finds evidence that in Chile, schools raised tuition by 6 percent in response to a student lending program.

We make a number of contributions to the literature. First, we complement the scarce evidence on the relation between student credit and tuition for developing countries. The literature generally focuses on the impact of loan availability on tuition in the United States. There is a growing body of evidence on the impact of student loans on enrollment and dropout rates for developing countries (Ferreyra and others, 2017), but more evidence is needed to fully understand the impact of student loans on tuition. Given the growing number of subsidized student lending programs in developing economies, producing evidence on how HEIs in developing countries react to such programs is indispensable from a policy perspective. Espinoza (2017) evaluates how HEIs alter their pricing strategy in response to a student loan program in Chile. The author builds on a structural strategy that allows him to identify a model-based estimate of the program's impact. We, in turn, build on Espinoza (2017) by identifying an exogenous variation in eligibility for

12. Source: College Board.

student loans, which allows us to obtain an estimate of tuition increase that does not depend on modeling assumptions.

Brazil offers an interesting framework for studying the impact of student lending on tuition prices. Relative to the United States, credit markets are shallow, and there is evidence that borrowers are credit constrained (De Mello and Garcia, 2012). Student credit can be particularly relevant in this setting. Brazil's institutional framework may explain why we find strong support for the Bennett hypothesis, while the empirical literature using U.S. data finds either no support (Hoxby, 1997; Rizzo and Ehrenberg, 2004) or only partial support (Singell and Stone, 2007).[13] Our results are in line with Espinoza (2017).

Finally, we focus exclusively on student credit. Most of the aforementioned papers estimate the effect of student credit and financial aid jointly. This distinction is important because the credit channel has broader implications, and it relates to the literature on credit availability and asset prices in general. Ours is one of the few papers that document a possible mechanism behind the nexus between credit availability and tuition.[14] Specifically, we estimate a structural model of demand and document that credit availability is correlated with a reduction in tuition elasticity, which, in turn, could be one of the mechanisms leading to tuition inflation if colleges are not price takers.

Data

Our empirical strategy relies on data from a number of sources. First, we use the *Censo do Ensino Superior* (CES), a data set provided by the *Instituto Nacional de Estudos e Pesquisas Educacionais Anísio Teixeira* (INEP).[15] The CES is a nationwide survey that contains information on all HEIs in Brazil. Information from the CES is available from 1995 through 2017.

The CES contains information at four levels of aggregation: HEI, major-HEI, student, and instructor. At the HEI level, the CES contains information on institutional characteristics, such as the number of employees by type (instructors, professors, administrative staff, and so on), infrastructure, and

13. More recent papers for the U.S. higher education market do find evidence for the Bennett hypothesis (Cellini and Goldin, 2014; Lucca, Nadauld, and Shen, 2019; Gordon and Hedlund, 2019).

14. Espinoza (2017) is another example.

15. INEP is an independent government agency linked to Brazil's Ministry of Education.

financial statements. Each HEI is defined by ownership and geography. We use the term higher education institution to refer to the school-city unit. We call the entity that owns the school unit the owner institution. More precisely, the operations of an owner institution under the same brand in cities A and B constitute two different HEIs. Typically, an owner institution has several HEIs.

At the major-HEI level, the CES contains information on the number of credits required for graduation, the minimum length of each program, the number of applicants, the number of enrolled students, the number of dropouts, and the number of graduating students. Each major is grouped into one of eight broad fields of study according to a Ministry of Education classification (for example, humanities, engineering, health, and so on) and further subdivided into twenty-two more specific fields of study. In Brazil, students declare a major for the entrance exam (that is, before being admitted). If a specific owner institution operates under the same brand in different cities, each city-level operation is considered a different HEI. Different majors represent different fields of study (for example, business administration, law, medicine, or nursing). A major-HEI pair represents a given field of study offered by a specific HEI.

The CES also contains detailed information on students and instructors.[16] Student data include demographics and information on financial aid by source and type. Crucial for our purposes, we have information on the number of students that have an FIES loan at the major-HEI level. For instructors, we have data on demographics, education, and employment type (part-time versus full-time).

From the Brazilian Ministry of Education, we use data on two different measures of major-HEI quality, both from standardized evaluations. The first is the National Student Performance Examination (ENADE), which is administered to freshman and senior students. The ENADE evaluation groups major-HEIs into three broad areas, and each year one of these areas is subject to a student's performance assessment. Thus, undergraduate students enrolled in specific major-HEIs in Brazil are assessed every three years. After each assessment, major-HEIs receive a grade that reflects the average academic performance of its students. This grade can range from one to five, in increasing order of quality. The Ministry of Education considers grades of three or higher to be acceptable.

16. Starting in 2009.

The second quality measure is the Preliminary Major Score (CPC). As with the ENADE evaluation, major-HEIs are grouped into three broad areas, and each area is evaluated every three years. The CPC evaluation considers three dimensions. The first is the quality of faculty, as measured by three proxies: the proportion of instructors with a Ph.D., the proportion of instructors with a least a master's degree, and the proportion of full-time instructors. The second is the quality of physical and academic resources. The performance of each major-HEI on this dimension is determined by enrolled students' subjective assessment. Finally, academic performance is measured as the average performance of enrolled students on the ENADE exam. It also considers a measure of added value: the difference between the actual performance of senior students on the ENADE exam and the expected performance given their socioeconomic background. For each dimension, the Ministry of Education assigns a grade from one through five. A major-HEI's CPC is given by the average of these three grades.

Obtaining data on tuition is essential for our analysis. In Brazil, tuition is defined at the major-HEI level. Tuition fees vary considerably across HEIs and across majors within HEIs. Information on tuition at the major-HEI level is not publicly available. We overcome this limitation by accessing a unique database from Hoper, a consultancy firm specialized in the education sector. The data cover 82 percent of HEIs in Brazil and contain information on tuition at the major-HEI level from 2009 through 2013. Our tuition data set is consistent with publicly available data for a subset of major-HEIs and years, and our empirical results are robust to using this alternative data source (see the online appendix for details).[17]

We also use information from the *Relação Anual de Informações Sociais* (RAIS), a data set organized by Brazil's Ministry of Labor. The RAIS contains detailed information on all wage earners in the formal sector, from which we construct city-level annual series of instructor and staff salaries.[18]

The final sample is an unbalanced panel containing 17,945 observations of the quadruple: year (from 2009 through 2013), HEI, city, and major-HEI. The panel is unbalanced because tuition information is not available for all the major-HEI pairs included in the sample in the 2009–13 period. There is no

17. Supplementary material for this paper is available online at http://economia.lacea.org/contents.htm.

18. While the Brazilian labor market has a large informal sector, HEIs operate in the formal sector.

TABLE 1. Descriptive Statistics

Variable	*Descriptive statistic*
Tuition (in 2008 reais)[a]	561.15
	(343.68)
Enrolled students (total)[a]	348.87
	(484.29)
Students with FIES loan (total)[a]	27.21
	(68.28)
Percentage of students with FIES loan	7.34
	(11.02)
Major-HEI quality assessment score	2.35
	(0.61)
Senior students (total)[a]	56.38
	(80.81)
Freshman students (total)[a]	113.37
	(160.22)
Applicant students per available slot[a]	1.83
	(3.41)
Faculty quality[b]*	0.64
	(0.16)
Faculty (total)[b]	620.25
	(957.98)
Degrees (total)[b]	60.83
	(108.47)
Administrative staff (total)[b]	594.82
	(968.72)
No. observations	17,945

Notes: This table presents descriptive statistics for the final sample used to obtain the main results of this paper. The final sample covers the period between 2009 and 2013 and consists of 17,945 major-HEIs. For each of the variables included in the table, we present their average value at the major-HEI or HEI level. Standard errors are in parentheses. An asterisk (*) represents the proportion of faculty with at least a master's degree. a. Variables at the major-HEI level. b. Variables at the HEI level.

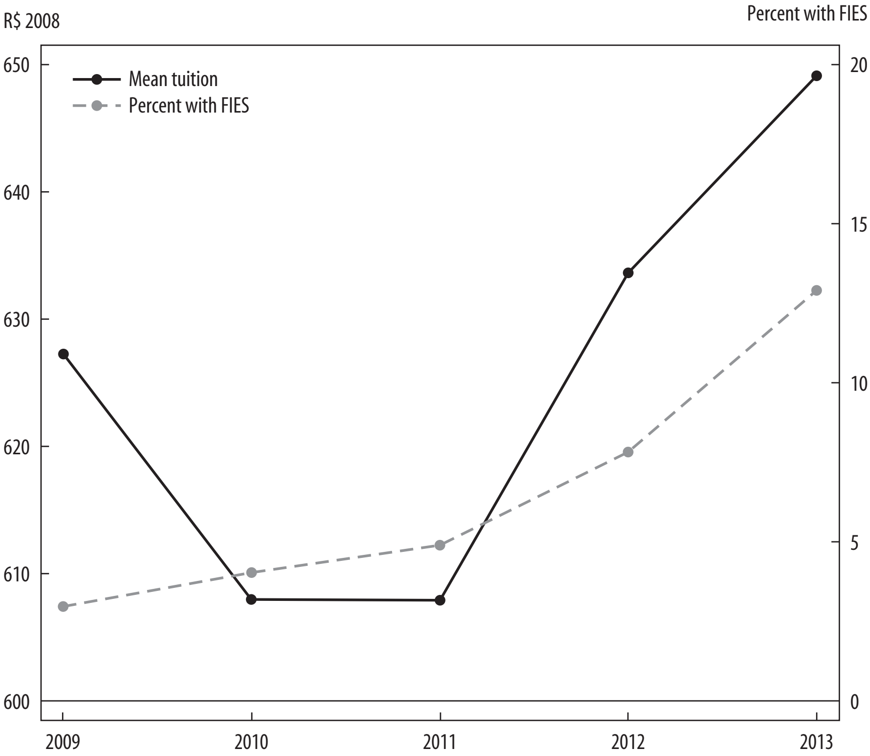
reason to believe that the missing information and the policy we are evaluating are correlated.[19]

Table 1 contains summary statistics for our final sample. The average tuition fee in the 2009–13 period was R$561 per month (in 2008 reais). This is 35 percent higher than the minimum wage, and it amounts to approximately U.S. $1,530 annually.[20] On average, 7.3 percent of students enrolled in a

19. In the next sections, we test the robustness of our results to our sample selection.

20. In the United States, tuition fees are higher. In both countries, tuition represents a similar fraction of per capita income.

FIGURE 1. Aggregate Monthly Tuition and FIES Penetration



Notes: This figure presents the average value of tuition (weighted by number of students) (left axis) and the average share of enrolled students with an FIES loan (right axis) for every major-HEI pair in the final sample for each year from 2009 to 2013. In 2008, R$1.00 was roughly equivalent to U.S. $2.00.

given major had an FIES loan in the 2009–13 period. In 2013, 13 percent of students had an FIES loan (see figure 1).[21] The remainder of table 1 contains information on quality proxies and HEI size. Major-HEIs enroll, on average, 350 students. The average number of total instructors at the HEI level is 620. This number includes instructors hired on either a part-time or full-time basis, as well as instructors on leave. Among the faculty, 64 percent of instructors

21. Within the universe of HEIs, approximately 6 percent of students had FIES financing in the 2009–13 period. In 2013, 15 percent of students had FIES financing.

have at least a master's degree. There is evidence of supply constraints for major-HEIs in our sample, with major-HEIs having, on average, 1.83 applicants per available slot.

FIES and the 2010 Intervention

FIES is the Brazilian federal government's subsidized student lending program. FIES loans cover tuition for students enrolled in private HEIs.[22] Both students and major-HEIs have to satisfy certain criteria to be eligible for FIES loans, which cover between 50 and 100 percent of the tuition for enrolling in a specific major-HEI. The fraction of tuition eligible for financing depends on family income and the ratio between tuition and per capita household income.[23] In Brazil, students choose their major on admission.

Loans are distributed by the two largest federal financial institutions (CAIXA and Banco do Brasil). In exchange for providing educational services, HEIs receive treasury bonds called *Certificados Financeiros do Tesouro Série E* (CFT-E), a special issue for FIES financing. The face value of these bonds corresponds to the tuition financed through FIES. The bonds are tradable for social security obligations. No secondary market exists for these bonds, but the government holds repurchase auctions.

22. Public universities are tuition-free. Entrance exams are very competitive, and only the top-performing students are admitted. Because of a lack of government funding at the primary and secondary levels, these high-performing students generally come from affluent families—65 percent of the students in public universities belong to the 40 percent richest of the population, and, given the income restrictions of the program, they are not directly affected by changes in the FIES (World Bank, 2017). Using the CES data, we can evaluate whether public HEIs were affected by FIES. After the FIES expansion, there was no significant change in the demand trend previously observed for public HEIs: the number of applicants to public HEIs increased 29 percent between 2009 and 2010 (the pre-FIES period) and an average of 30 percent between 2010 and 2013 (the post-FIES period). The number of enrolled students in public universities increased in both the pre- and post-FIES periods (8 percent and 4 percent, respectively). We do not consider tuition-free public HEIs in our reduced-form strategy, and we treat them as an outside option to obtain the main results of our structural analysis.

23. FIES coverage varies by household income level: (1) full tuition for students with gross household income of less than 10 minimum wages and a rate between tuition and per capita household income of more than 0.6; (2) 75 percent of tuition for students with gross household income of less than 15 minimum wages and a rate between tuition and per capita household income of more than 0.4 and less than 0.6; (3) 50 percent of tuition for students with gross household income under 20 minimum wages and a rate between tuition and per capita household income of more than 0.2 and less than 0.4.

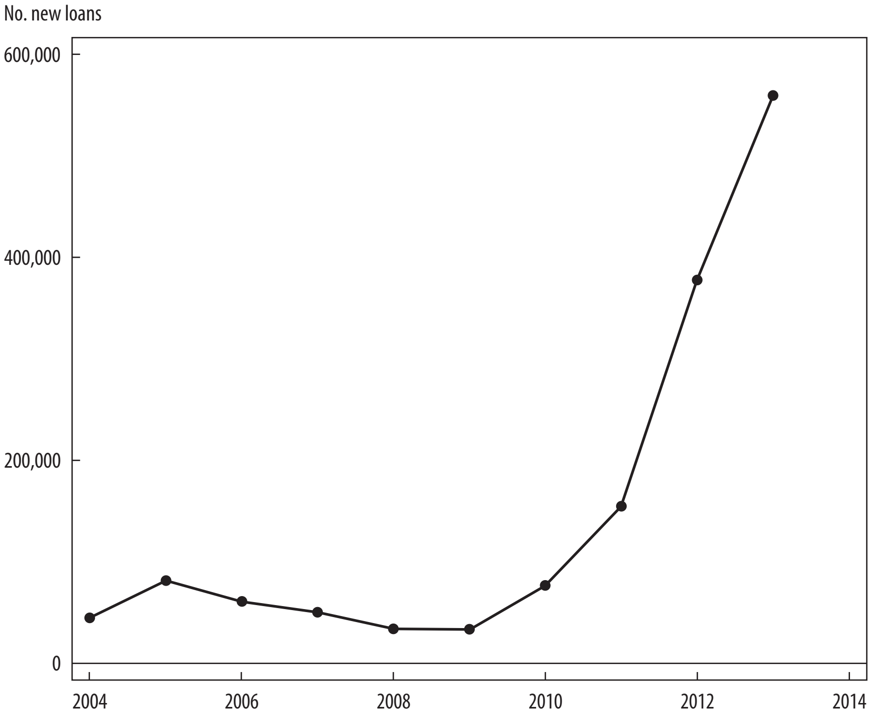
FIES was created in 1999. Until 2010, it was a small program, in terms of both government spending and number of students. In the first semester of 2010, the program went through a significant and unexpected reform aimed at increasing subsidies for enrollment in private higher education.[24] These changes made FIES more accessible to students and more attractive for HEIs. As a result, the number of students with an FIES loan increased considerably, and FIES became one the most relevant sources of funding for higher education in Brazil.

From the students' perspective, overall conditions improved considerably. Government interest rates dropped from 6.50 to 3.50 percent per year. For comparison, the interbank rate set by the central bank was 8.75 percent per year in December of 2009. The reduced FIES rate applied both to new loans and to the stock of previous loans. The government facilitated access to the program. Before the 2010 reform, students could only apply during a specific period of the year (the subscription period). In 2010, the government established a rolling application process. Repayment conditions improved, as well. Interest payments for FIES are due from the moment the contract is signed, but in 2010, the government capped interest disbursements for enrolled students at R$50 (U.S. $12), effectively deferring interest payments until after graduation. The grace period was extended from twelve to eighteen months after graduation. The amortization period was extended from twice to three times the loan period. That means that a student enrolled in a four-year degree would have twelve years instead of eight to repay the loan. The underwriting process was relaxed and streamlined. Before the 2010 reform, borrowers needed a cosigner. In 2010, the government created a subsidized insurance scheme for students of lower socioeconomic status, the *Fundo de Garantia de Operaçoes de Crédito Educativo* (FGEDUC).

FIES also became more attractive for HEIs. The main change was in the frequency of repurchase auctions. Before 2010, there was no rule about the frequency of auctions. After 2010, the government established that repurchase auctions would occur at least quarterly, which reduced the length of accounts receivable to just ninety days. This change had a large impact on working capital costs for HEIs. The FGEDUC also benefited HEIs. HEIs could shift risk to the FGEDUC by paying a premium of up 7 percent of the revenue on the

24. FIES was again completely reformulated in 2015. The federal government introduced a set of new rules intended to reduce the disbursement of public funds and to target FIES subsidies to worse-off students. Since our analysis does not cover this period, we do not detail this reform. More information on the new FIES rules can be found on the NOVO FIES website (http: //fies.mec.gov.br/).

FIGURE 2. FIES: New Loans

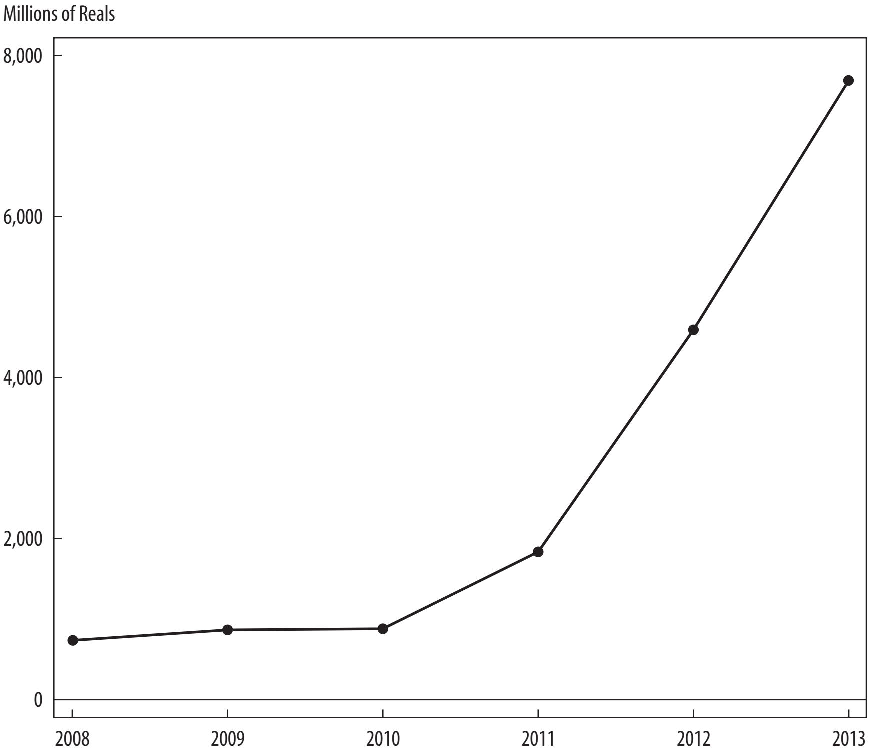


Source: Ministry of Education.

contract. Market analysts considered the 7 percent premium cheap (Itaú BBA, 2013). For HEIs, FIES permitted not only a rapid expansion in enrollment but also a reduction in dropout rates. Incidentally, Itaú BBA (2013) mentions the possibility of an FIES-induced increase in pricing power. Sell-side analysts' models suggested that, relative to no FIES, an entirely FIES-financed class would deliver a double-digit increase in net present value to the HEI.

After the 2010 reform, the government repeatedly expressed its desire to increase enrollment rates through FIES. In the years following the reform, we see a considerable increase in the number of students enrolled with an FIES loan. Figure 2 depicts the number of new FIES loans from 2004 to 2014. Starting from a low level, the number of new loans doubled in 2010 and increased another 90 percent in 2011. From this higher level, new loans jumped 120 percent in 2012 and 40 percent in 2013. Government spending

FIGURE 3. FIES: Government Disbursements for FIES Loans



Source: Government of Brazil, Portal da Transparência (www.portaltransparencia.gov.br).

on FIES loans increased accordingly (figure 3). Over our sample period, an average of 7.34 percent of the students enrolled in a private HEI had FIES financing. In 2013, this percentage was 13 percent.[25]

Reduced Form: Identification and Main Results

For identification, we exploit two types of variation: time-series variation, because of the large and unexpected ramp-up in FIES in the first half of 2010, and cross-section variation, insofar as the new FIES rules created

25. See table 1.

heterogeneity in eligibility at the major-HEI level, which we exploit to define treatment and control groups. With regard to the latter, FIES rules restrict participation in the program to students enrolled in major-HEIs that reach a minimum threshold in standardized quality evaluations conducted by the Ministry of Education. According to FIES legislation, a major-HEI is considered eligible to enroll students with FIES financing if it scores three or more on one of three quality assessments, according to the following order of relevance: (1) Major Score (*Conceito de Curso*, CC); (2) Preliminary Major Score (CPC), if the CC is not available; and (3) National Student Performance Examination (ENADE), if neither the CC nor the CPC is available. The CC is an on-site evaluation and as such is not as broadly assessed as the CPC or ENADE. Since data on CC results are not available, we do not consider CC results in our strategy. We use the CPC as our main proxy for eligibility. For some major-HEIs, the CPC is not available, but the ENADE grade is. In these cases, we consider the ENADE grade as the eligibility proxy. Major-HEIs not yet subject to their first evaluation are considered eligible pending evaluation.[26]

We identify the causal impact of being eligible for an FIES loan by exploiting time-series and cross-section variation through a DD framework. Treatment is defined according to FIES eligibility at the major-HEI level in 2010; that is, we include in the treatment group all major-HEIs with a CPC grade of three or higher in 2010, as well as unevaluated major-HEIs.[27] We focus on 2010 eligibility because we have evidence that HEIs were not familiar with the FIES expansion or with the new eligibility rules when the grades for the 2010 quality evaluations were set. Because HEIs have some control—although limited in the short run—over their quality assessment, we could have selection bias in our sample if we allow treatment status to vary after policy implementation.

In our main identification strategy, the DD, we compare the change over time in tuition at eligible major-HEI pairs (treatment group) with the change in tuition at ineligible major-HEI pairs (control group). Thus our results identify the impact of being eligible to enroll students financed through FIES on tuition. The pretreatment period is 2009–10; the posttreatment period is

26. The government only remits FIES payments if the owner institution has no pending tax-related debt. In practice, industry reports suggest that this fact is not inconsequential. We do not explore it because we do not observe tax-related debt for owner institutions.

27. We consider the ENADE grade when information on the CPC is not available.

2011–13. We estimate the average treatment effect on the treated group by exploring variations of the following basic model:

$$(1) \qquad \log(\text{TUITION})_{jt} = \theta + \alpha D_t + \beta \text{TREAT}_j + \varphi D_t * \text{TREAT}_j + \rho \mathbf{X}_{jt} + \mu_t + \lambda_j + \varepsilon_{jt}.$$

The dependent variable is the natural logarithm of posted tuition charged by major-HEI j in year t (2009 through 2013). D_t is a dummy variable equal to one in posttreatment periods (2011 through 2013) and zero otherwise. $\mathbf{TREAT}_j$ is a dummy variable equal to one if the major-HEI j was eligible for FIES in 2010 and zero otherwise. $\mathbf{X}_{jt}$ represents a vector of time-varying major-HEI and HEI characteristics; μ_t, time (year) fixed effects; λ_j, unit (major-HEI) fixed-effects; and ε_{jt}, the unobserved error term. The parameter φ is the eligibility effect, our object of interest.

We estimate extensions from the basic model, including field of study–year and city-year fixed effects. For every specification, we cluster standard errors at the HEI level, a higher level than the treatment level (major-HEI). Thus our estimated standard errors are, if anything, conservative relative to the common practice when estimating treatment effects with DD. The fraction of treated units varies by specification, ranging from as high as 96 percent (monopoly markets) to 73 percent when we restrict the sample to major-HEI pairs with grades two and three. For the main sample, eligible major-HEIs amount to 90 percent of total major-HEI pairs.[28]

Table 2 shows our estimates of the parameters in equation 1.[29] The first row shows the estimates of the eligibility effect parameter, φ, our main parameter of interest. Consistent with equation 1, we include time and major-HEI

28. We do not implement an RDD as our main identification strategy because the assumptions required for identification in an RDD framework do not necessarily hold in our case. Specifically, an RDD approach requires that no other relevant factor—except treatment status—presents a discontinuity across the eligibility threshold. In our case, eligibility is determined according to a quality evaluation that is continuous, but made public as a discrete number. Thus, going from an evaluation of two to an evaluation of three on a five-point scale will alter students' and HEIs' behavior around the cutoff beyond the eligibility to receive FIES funding. If that is the case, RDD is not a valid identification strategy. Regardless, we implement an RDD estimation as a robustness test (see the online appendix). The results are consistent with our main results.

29. We lose observations in our reduced-form estimations because we exclude singletons from our analysis. Maintaining singleton groups in linear regressions where fixed effects are nested within clusters can overstate statistical significance and lead to incorrect inference (Correia, 2016).

TABLE 2. Reduced-Form Estimation: Difference-in-Differences

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	0.046**	0.043**	0.048***	0.048***	0.047**	0.031***
(0.018)	(0.018)	(0.018)	(0.018)	(0.020)	(0.010)	
Summary statistic						
No. observations	15,219	15,219	15,219	15,219	15,218	15,019
R^2	0.924	0.925	0.925	0.925	0.926	0.953
Covariates, HEI level	No	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	No	No	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	No	No	No	Yes	Yes	Yes
Field of study–year fixed effects	No	No	No	No	Yes	Yes
City-year fixed effects	No	No	No	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: This table shows the results of the difference-in-differences (DD) specification of equation 1. The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The treatment group includes major-HEIs eligible to enroll students with FIES (according to the quality evaluations conducted by the Ministry of Education); the control group, ineligible major-HEIs. The pretreatment period consists of the years that precede FIES expansion (2009 and 2010). The posttreatment period consists of the years after the expansion (2011, 2012, and 2013). The estimated coefficients associated with the elegibility effect variable represent the impact of being eligible for FIES on log(tuition). We include time and major-HEI (unit) fixed effects for all the specifications presented in this table. In column (2), we include a set of time-varying covariates at the HEI level: faculty quality, number of majors offered by the HEI, number of employers hired as administrative staff size, and number of faculty members. In column (3), we include time-varying covariates at the major-HEI level: number of enrolled students, number of applicant students per available slot, and a measure of major-HEI quality. In column (4), we include a measure of market concentration (Herfindahl-Hirschman Index). In column (5), we include field of study–year fixed effects. Finally, in column (6), we include city-year fixed effects. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

fixed effects for all specifications. Column 1 does not contain any controls. In this case, FIES eligibility increases tuition by 4.6 percent in real terms over the 2011–13 period. We include an increasing set of controls in columns 2 through 6. When we include time-varying covariates at the HEI, major-HEI, and major-HEI-market levels (columns 2 to 4), $\hat{\varphi}$ barely moves.[30] In column 5, our preferred specification, we include field of study–year fixed effects. Being eligible for FIES causes tuition to increase 4.6 percent even when we consider only within field of study–year variation. Column 6 contains the most complete specification, which includes city-year fixed effects. The model becomes saturated ($R^2 = 0.953$). Eligibility for FIES causes tuition to increase by 3.1 percent in real terms in this specification.

30. We include a large set of time-varying controls for quality (namely, applicant-to-vacancy ratio, Ministry of Education quality measure, and percentage of faculty with at least a master's degree) and size (number of degrees, number of administrative faculty, total faculty, and market concentration—the HHI index).

Reduced Form: Threats to Identification

Equation 1 is a reduced-form object. The error ε_{jt} contains unobserved time-varying supply and demand shifters that affect prices. Interpreting φ presents a number of challenges. First, the causal interpretation of our results relies on the absence of selection bias. Our results depend on the fact that major-HEIs assigned as treatment did not choose to be part of the treatment group. This assumption would be violated if HEIs anticipated the FIES ramp-up and started to implement quality-enhancing strategies to qualify for the program. Although possible, the anticipation story is implausible. Anecdotal evidence suggests that market participants did not anticipate the FIES reform. Many large education providers waited to see how credible the government's commitment to FIES lending was. For instance, Itaú BBA, a large investment bank in Brazil, only started producing reports on the impact of FIES and the higher education sector in Brazil in 2012.

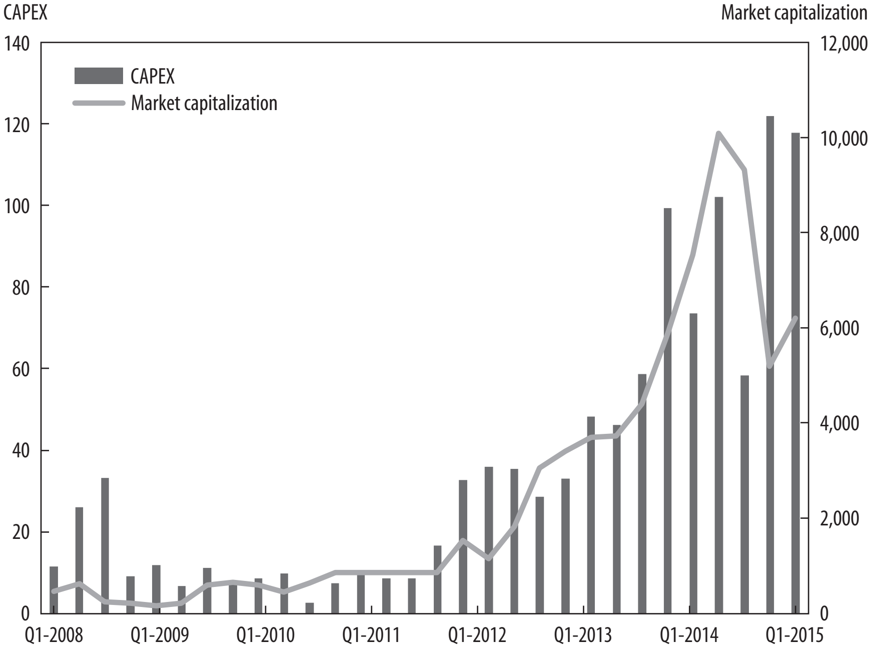
From Bloomberg and Economatica, we obtain financial data on publicly traded education providers. Data on the largest private higher education provider in Brazil, Kroton, also suggest that investors did not anticipate the FIES reforms.[31] Figure 4 depicts Kroton's capital expenditures and market capitalization. Listed since 2007, Kroton is the largest private tertiary education provider in Brazil. After 2010, Kroton's aggressive expansion strategy heavily relied on FIES.[32] If the company was expecting a boom from FIES, one would expect a surge in capital expenditures before 2010. Figure 4 shows no such surge. Investors do not seem to have anticipated any value in FIES. From January 2008 through the end of 2010 (a year after the changes in FIES were implemented), Kroton's market capitalization barely moved. Interestingly, the market capitalization of Kroton starts surging when the number of students covered by FIES increases sharply, that is, in 2012.

We also use our data to investigate whether schools changed their behavior right before the FIES expansion in order to improve their performance on the Ministry of Education's quality assessments. Figure 5 shows the evolution of CPC scores over time, the quality indicator that determines eligibility for FIES. We compute the kernel estimate of the density function of the CPC for every year from 2009 through 2013. If colleges were increasing quality

31. Data from other large listed private universities, such as Anhanguera and Estácio, show similar trends (available on request).

32. Market capitalization is the equity value on the São Paulo Stock Exchange. By the market capitalization criterion, Kroton was one of the three most valuable private sector listed universities in the world in 2013 and 2014.

FIGURE 4. Kroton: Market Capitalization and Capital Expenditures
US$ millions



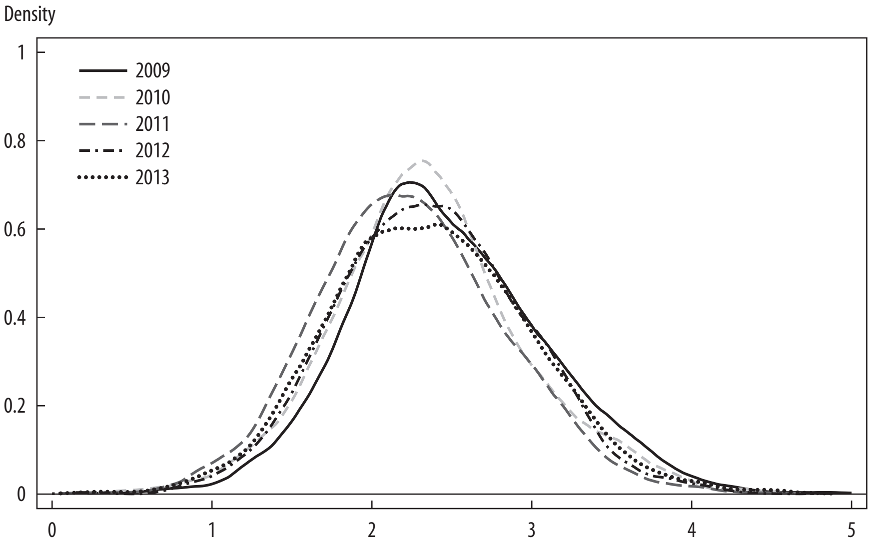
Source: Bloomberg and Economatica.

in anticipation of the FIES ramp-up, we would expect a shift in the densities of CPC toward higher grades over time. We see no discernible changes from year to year.[33]

Another threat to our causal interpretation comes from the quality of our eligibility proxies. As previously mentioned, FIES determines eligibility based on the performance of major-HEIs on three different quality evaluations conducted by the Ministry of Education: the CC, the CPC, and the ENADE evaluations. CC grades are not readily available, so we have to use the CPC and ENADE evaluations to determine treatment and control groups.

33. We also estimate the densities of CPC separating major-HEIs into treatment and control groups. The anticipation story would be most damaging to a causal interpretation of our results if treated units were increasing quality more rapidly than control units. After 2010, if anything, the densities of the control units shift to the right more rapidly than the densities of the treatment units. After 2010, ineligible major-HEIs units increased their CPC scores, possibly responding to competitive pressures from FIES-eligible HEIs. The densities of treated units do not change meaningfully after 2010. Results from this exercise are available on request.

FIGURE 5. Major-HEI Quality Score: Kernel Density Estimation



Notes: This figure shows the trend through time of the quality indicator that determines eligibility for FIES. Specifically, it shows the kernel estimate of the density function of the CPC quality evaluation, if available, or the ENADE score, if the CPC is not available, for every year from 2009 through 2013.

We believe that the CPC and ENADE provide a good approximation of actual eligibility status for two reasons. First, since the CC evaluation is not as broadly assessed, in many cases eligibility is indeed determined by the CPC or ENADE grades. Second, the CC grade is a quality evaluation and thus is likely correlated with the CPC and ENADE grades. We use our data to test the quality of our eligibility proxy. Specifically, we evaluate whether major-HEIs in the treatment group are actually more likely to enroll students with FIES financing than major-HEIs in the control group. To evaluate the quality of the proxies, we use a very conservative measure. We say a unit in the treatment group is a complier if that unit enrolls at least one student with an FIES loan in each period. A unit in the control group is a complier if that unit does not enroll any student with an FIES loan in the period. Our eligibility proxies are solid. Between 2011 and 2013, the rate of compliance for units in the treatment group varies from 60 to 80 percent. Alternatively, the rate of compliance for units in the control group starts at 50 percent, but falls to 40 percent by the end of the period.

From a broader perspective, identification in a DD framework relies on the assumption that the variable of interest—tuition—is following the same trend for treatment and control units before the policy implementation—before FIES expansion—and would have continued following parallel trends if there had not been an intervention, that is, in a counterfactual scenario. We cannot test how reasonable the counterfactual parallel-trend hypothesis is. However, if information were available, we could test the hypothesis of no differential pretrend. Unfortunately, information on tuition at the major-HEI level is only available for two years before the intervention. Without a large enough preintervention period, we are limited in our ability to perform a direct test for the presence of differential preintervention trends.

There are, however, complementary exercises that can be performed to check for evidence that schools were following different trends in their pricing strategy before the intervention. We perform two exercises that, when combined, provide evidence for the parallel-pretrends assumption. In the first exercise, we use information at the HEI level to directly test for the presence of differential trends in tuition for a five-year period before the intervention. We have financial statements at the HEI level from the Higher Education Census. We consider the *receita própria* or "revenue from tuition" variable, which measures revenues from tuition fees.[34] We divide this variable by the number of students enrolled in the HEI and consider the new variable to be an approximate measure of average tuition. Starting in 2006, we test for the presence of differential preexisting trends in our proxy for tuition.

To investigate whether our general conclusion—that eligibility for FIES causes tuition to rise—is sensitive to changing the unit of analysis, we present, first, an estimate of the impact of FIES on tuition considering our tuition proxy measured at the HEI level. Column 1 of table 3 presents the results for this exercise. The estimated impact of FIES is again positive. We have less variation at the HEI level, but we can still reject the null hypothesis at the 5 percent level. The magnitude of the estimated treatment effect in table 3 is not comparable to the estimate in table 2 because the treatment is now continuous. An increase in FIES penetration from 0 to 1 is associated with a 16.3 percent increase in the revenue per student.[35]

34. The variable receita própria includes fees and fund transfers related to scholarships and tuition loans.

35. For brevity, we omit the estimated coefficients of control variables, which are available on request. They have the expected signs. Quality commands price. Tuition fees are lower at HEIs with more students, suggesting economies of scale at the HEI level.

TABLE 3. Reduced-Form Estimation: Placebo and Pretrend Tests

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*
Eligibility effect	0.163**	0.118	0.177	0.255	0.205	−0.079	−0.092
	(0.070)	(0.116)	(0.131)	(0.160)	(0.190)	(0.177)	(0.116)
Summary statistic							
No. observations	10,324	6,636	5,257	4,164	2,685	2,173	6,636
R^2	0.445	0.429	0.454	0.561	0.727	0.682	0.429

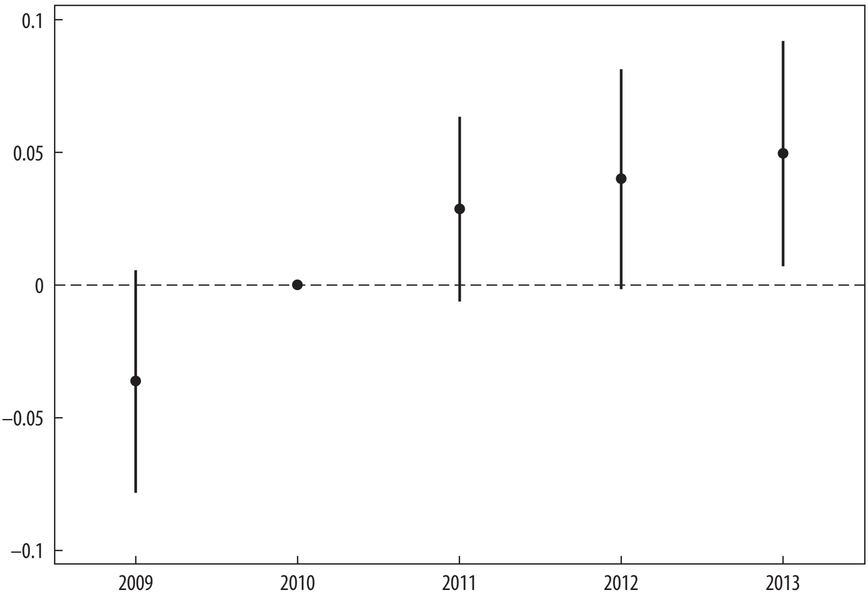
* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the natural logarithm of a variable representing revenue per student at the HEI level (HEI revenue divided by the number of enrolled students). The difference-in-differences (DD) strategy is implemented at the HEI level. Treatment is defined as a continuous variable—specifically, the proportion of majors that are eligible to enroll students funded by FIES. The estimated coefficients associated with the eligibility effect variable represent the impact of being eligible for FIES on log(revenue per student). In the first column, we define the pretreatment period as the years that precede FIES expansion (2006 to 2010) and the posttreatment period as the years after the expansion (2011, 2012, and 2013). In the following columns, we implement a placebo test using different combinations of pre- and posttreatment periods to evaluate whether there is evidence that the revenue per student variable followed differential trends by eligibility status in the years that preceed the expansion of FIES. In column 2, the pretreatment period is defined as the years 2006 and 2007 and the posttreatment period as 2008 to 2010; in column 3, pretreatment is 2006 and 2007 and postreatment is 2008 and 2009; in column 4, pretreatment is 2007 and 2008 and postreatment is 2009; in column 5, pretreatment is 2007 and postreatment is 2008; in column 6, pretreatment is 2006 and postreatment is 2007; in column 7, pretreatment is 2006 to 2009 and postreatment is 2010. Year and higher education institution fixed effects are included for all specifications. We also include the following covariates: number of enrolled students, number of applicant students per maximum cohort size, number of majors offered by the HEI, number of employers hired as administrative staff size, number of faculty members, and faculty quality. Standard errors are in parentheses.

Extending the sample backward allows us to perform an exercise that serves as both a placebo and a test of differential pretreatment trends. Columns 2 through 7 report the estimates from our placebo experiment. The exercise pretends that the FIES ramp-up occurred at a different time, and it restricts the posttreatment sample to periods preceding the actual treatment. This exercise effectively tests for the presence of pre-2010 differential trends in tuition costs according to the degree of eligibility in 2010. If the treatment and control groups were following different trends prior to 2010, we would find the "false" impact of FIES when performing the placebo exercise. We do not find a statistically significant effect for any of the periods we consider.

In a second exercise, we use a DD specification that allows for time-varying treatment effect. Estimating the time-varying impact of eligibility allows us to assess whether the impact of eligibility on tuition follows any specific trend (for example, increases or decreases over time). This type of exercise can also be used to assess how reasonable the assumption of parallel pretreatment trends is. If the results show that the impact of eligibility on tuition is not significant for pretreatment years, this result can be considered further evidence in favor of the parallel-trends assumption.

FIGURE 6. Time-Varying Difference-in-Differences Specification



Notes: This figure shows the result of a difference-in-differences strategy that considers the possibility that the treatment effect—the impact of being eligible for FIES on tuition—varies with time. The dots in the graph represent the point estimates of the eligibility effect for each year; the lines represent a 90 percent confidence interval. For this figure, we follow the specification of table 2, column 5; that is, we include time–field of study fixed effects and covariates. We include the following covariates at the HEI level: faculty quality, number of majors offered by the HEI, administrative staff size, and faculty size. At the major-HEI level, we include the following covariates: number of enrolled students, number of applicant students per available slot, a measure of major-HEI quality, and a measure of market concentration (Herfindahl-Hirschman Index).

Figure 6 presents the results of this exercise. There is no significant difference in tuition between units in the treatment and control groups just before the FIES expansion. This result provides further evidence for the parallel-pretreatment-trend hypothesis. The impact of FIES eligibility on tuition also increases with time. Considering that the number of new FIES loans increased during the period, this result is consistent with the idea that the impact of eligibility depends on the actual number of enrolled students with FIES loans.

Another threat to identification arises from the fact that the control group may be affected by treatment, changing the interpretation of the DD estimates. Consider a duopoly market with one major-HEI eligible for FIES. We would get a positive DD estimate if tuition increases more at eligible major-HEI pairs than at ineligible pairs. The interpretation of the results and

TABLE 4. Reduced-Form Estimation: Monopoly Markets

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	0.061***	0.059**	0.053**	0.053**	0.048*	0.034
	(0.023)	(0.023)	(0.025)	(0.025)	(0.028)	(0.041)
Summary statistic						
No. observations	1,094	1,094	1,094	1,094	1,073	1,056
R^2	0.952	0.952	0.952	0.952	0.955	0.989
Covariates, HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	Yes	Yes	Yes	Yes	Yes	Yes
Field of study–year fixed effects	No	Yes	Yes	No	Yes	Yes
City-year fixed effects	No	No	Yes	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in table 2, but the sample is restricted to monopoly markets (that is, city–field of study pairs with only one major-HEI). The DD thus compares eligible and ineligible monopolies and isolates the eligibility impact from competition effects. See the notes to table 2 for additional details. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

the welfare implications are different if competition drives prices down at ineligible major-HEI pairs. In particular, the Bennett hypothesis refers to a generalized increase in tuition (at least an average increase). We estimate equation 1 focusing only on monopoly markets (a market as a city–field of study pair). The DD estimate now compares eligible and ineligible monopolies, avoiding potentially confounding competition effects. Table 4 contains estimates of equation 1 restricting the sample to monopoly markets, that is, city–field of study pairs with only one HEI (eligible or not). We lose precision in some cases because not only is the sample size smaller, but the number of units in the control group is severely reduced (4 percent of the 1,094 total observations). Still, the estimates from this exercise are similar to our estimates in table 2.

Reduced Form: Robustness Tests

We implement a few modified versions of the specification presented in table 2 to assess the robustness of our results. In these exercises, we vary either our sample of analysis or the criterion used to determine treatment status. The objective is to investigate whether our results are sensitive to any of these choices.

First, our panel is unbalanced, which poses several problems. The proportion of eligible and ineligible major-HEI pairs may change over time, or

TABLE 5. Reduced-Form Estimation: Balanced Panel

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	0.045*	0.035	0.037	0.039	0.033	0.038
	(0.027)	(0.025)	(0.027)	(0.026)	(0.027)	(0.026)
Summary statistic						
No. observations	1,530	1,530	1,530	1,530	1,524	1,519
R^2	0.908	0.911	0.911	0.911	0.917	0.939
Covariates, HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	Yes	Yes	Yes	Yes	Yes	Yes
Field of study–year fixed effects	No	Yes	Yes	No	Yes	Yes
City-year fixed effects	No	No	Yes	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in table 2, but the sample is restricted to the major-HEIs for which information is available for every year from 2009 to 2013. See the notes to table 2 for additional details. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

the proportions of majors may change because of demand shocks specific to certain markets. To test how robust our results are to sample selection, we estimate equation 1 restricting the sample to a balanced panel (see table 5). The number of observations drops significantly (the sample size is less than a tenth of the original), and we are unable to precisely estimate the impact of eligibility, with one exception. Nevertheless, estimated coefficients are mostly consistent with their counterparts in table 2.

Second, we estimate the eligibility impact for a sample that includes only major-HEIs right next to the eligibility cutoff—that is, major-HEIs with a CPC grade of two or three. Restricting the sample to major-college pairs with a CPC grade of two and three makes the treated and control units more similar on observable characteristics. The estimated coefficients, shown in table 6, are again consistent with their counterparts in table 2.

We also consider alternative proxies for eligibility. To obtain our main results, we use the Ministry of Education's quality evaluations in 2010 as a proxy for eligibility. Our argument is that at the time these evaluations were conducted, major-HEIs were not expecting the FIES expansion. Evaluations are conducted every three years and include factors that are not easily manipulated in the short term, such as the academic performance of senior students. The 2010 quality evaluation is a good proxy for eligibility because major-HEIs are limited in their ability to influence this outcome in the short run. Therefore, our results should not be sensitive to using earlier quality evaluations as eligibility proxies.

TABLE 6. Reduced-Form Estimation: Major-HEI with CPC 2 and 3

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	0.040***	0.037***	0.045***	0.045***	0.044***	0.023**
	(0.014)	(0.014)	(0.015)	(0.015)	(0.016)	(0.010)
Summary statistic						
No. observations	6,633	6,633	6,633	6,633	6,628	6,524
R^2	0.907	0.907	0.907	0.907	0.910	0.938
Covariates, HEI level	No	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	No	No	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	No	No	No	Yes	Yes	Yes
Field of study–year fixed effects	No	No	No	No	Yes	Yes
City-year fixed effects	No	No	No	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in table 2, but the sample of treatment and control units is restricted as follows: the treatment group includes only major-HEIs that were barely eligible for FIES (that is, major-HEIs with a quality evaluation of three in 2010), while the control group consists of major-HEIs that almost reached the eligibility threshold (that is, major-HEIs with a quality evaluation of two in 2010). See the notes to table 2 for additional details. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

To evaluate the robustness of this assumption, we estimate our preferred model considering earlier quality evaluations as eligibility proxies. Specifically, we estimate the impact of FIES eligibility using eligible major-HEIs in 2008 and 2009 as treatment groups, with the corresponding inelegible major-HEIs in the same years as the control group. We opted to exclude the unevaluated major-HEIs in this exercise. Since major-HEIs are evaluated every three years, including major-HEI units that had no evaluation more than two years before the expansion of FIES as part of the treatment group would likely increase the share of noncompliers in the treatment group. Table 7 presents the results of this exercise. For the 2009 evaluation, results are positive and significant for all the specifications considered. For the 2008 evaluation, we are only able to precisely estimate the impact of eligibility for our preferred specification—the one that includes all the time-varying covariates and field of study–year fixed effects. For both the 2008 and 2009 evaluations, the magnitude of the results is only slightly smaller than the magnitudes reported in table 2.

To determine whether including unevaluated major-HEIs in the treatment groups is driving our general results, we estimate a DD specification for a sample that excludes these units from our analysis (table 8). When we consider only major-HEIs that had a quality evaluation in 2010, we reduce our sample to approximately 60 percent of its original size. Despite having fewer observations, our results do not change considerably. The specifications in

TABLE 7. Reduced-Form Estimation: 2008 and 2009 CPC

	2008 CPC			*2009 CPC*		
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	0.016	0.041*	0.016	0.039**	0.049**	0.021*
	(0.021)	(0.022)	(0.016)	(0.018)	(0.019)	(0.011)
Summary statistic						
No. observations	9,369	9,368	9,131	10,637	10,636	10,414
R^2	0.910	0.913	0.949	0.913	0.915	0.950
Covariates, HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	Yes	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	Yes	Yes	Yes	Yes	Yes	Yes
Field of study–year fixed effects	No	Yes	Yes	No	Yes	Yes
City-year fixed effects	No	No	Yes	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD-specification is the same as in table 2, but the variable used to determine eligibility status is changed as follows: the treatment group includes major-HEIs that would be eligible to enroll FIES students according to their quality evaluation in 2008 (columns 1–3) and in 2009 (columns 4–6). The pretreatment period consists of the years that precede FIES expansion (2009 and 2010). The posttreatment period consists of the years after the expansion (2011, 2012, and 2013). The estimated coefficients associated with the eligibility effect variable represent the impact of being eligible for FIES on log(tuition). In columns 1 and 4, we include major and major-HEI level covariates: faculty quality, number of majors offered by the HEI, number of employers hired as administrative staff size, number of faculty members, number of enrolled students, number of applicant students per available slot, a measure of major-HEI quality, and a measure of market concentration (Herfindahl-Hirschman Index). In columns 2 and 5, we we include field of study–year fixed effects. In columns 3 and 6, we include city-year fixed effects. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

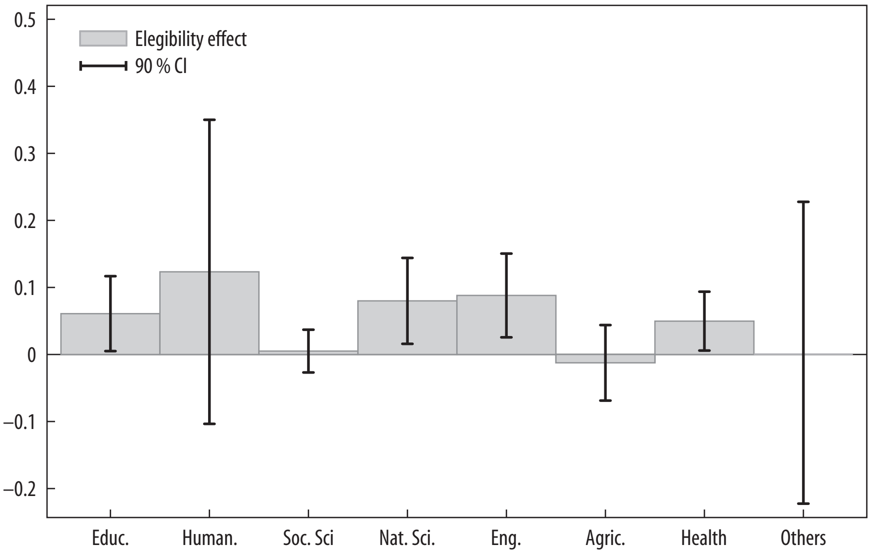
TABLE 8. Reduced-Form Estimation: Only Evaluated Major-HEIs Pairs

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*
Eligibility effect	(0.040***	(0.037***	(0.045***	(0.045***	(0.044***	(0.023**
	(0.014)	(0.014)	(0.015)	(0.015)	(0.016)	(0.010)
Summary statistic						
No. observations	6,633	6,633	6,633	6,633	6,628	6,524
R^2	0.907	0.907	0.907	0.907	0.910	0.938
Covariates, HEI level	No	Yes	Yes	Yes	Yes	Yes
Covariates, major-HEI level	No	No	Yes	Yes	Yes	Yes
Covariates, major-HEI-market level	No	No	No	Yes	Yes	Yes
Field of study–year fixed effects	No	No	No	No	Yes	Yes
City-year fixed effects	No	No	No	No	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in table 2, but the composition of the treatment group is changed to include only the eligible major-HEIs for which we have information on their 2010 quality evaluation (that is, major-HEI with grades three, four, or five on their quality evaluation). The control groups consists of major-HEIs that were not eligible for FIES in 2010 (obtained grades one or two). See the notes to table 2 for additional details. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

FIGURE 7. Reduced-Form Estimates by Broad Field of Study



Notes: This figure shows the results of a difference-in-differences strategy (equation 1) estimated separately by field of study. The height of each bar represents the point estimates of the eligibility effect for each field of study. The lines represents 90 percent confidence intervals. For this figure, we follow the specification of table 2, column (5): we include time–field of study fixed effects and covariates. We include the following covariates at the HEI level: faculty quality, number of majors offered by the HEI, number of employers hired as administrative staff size, and number of faculty members. At the major-HEI level, we include the following covariates: number of enrolled students, number of applicant students per available slot, a measure of major-HEI quality, and the Herfindahl-Hirschman Index.

columns 1 to 5 find a somewhat stronger effect of eligibility on tuition than previously estimated. After including city-year fixed effects, we estimate a slightly smaller impact than in table 2. To test whether the eligibility impact is driven by major-HEIs being perceived as higher quality due to the Ministry of Education evaluation, we exclude evaluated major-HEIs from our treatment group. These results (available upon request) are again consistent with the results in table 2. The estimates show across the board that FIES increases tuition prices.

We also estimate the DD specification splitting the sample by field of study. We consider the following fields: education, humanities, social sciences, natural sciences, engineering, agriculture, health, and other. For brevity, we present only the estimates associated with our preferred specification (column 5 of table 2) (see figure 7 and table 9). FIES eligibility increased tuition for all but one of the eight fields of study. This result suggests that

TABLE 9. Reduced-Form Estimation: Treatment Effect by Field of Study

Explanatory variable	*Education* (1)	*Humanities and arts* (2)	*Social sciences, business, and law* (3)	*Natural sciences, math, and computer sciences* (4)	*Engineering and construction* (5)	*Agricultural sciences* (6)	*Health* (7)	*Other* (8)
Eligibility effect	0.061*	0.123	0.005	0.080**	0.088**	−0.012	0.050*	0.003
	(0.034)	(0.137)	(0.019)	(0.039)	(0.038)	(0.034)	(0.027)	(0.136)
Summary statistic								
No. observations	2,325	231	6,339	1,355	1,241	291	2,973	346
R^2	0.863	0.913	0.896	0.874	0.868	0.917	0.937	0.923

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in column 5 of table 2, but we split our sample by field of study. Major-HEIs eligible to enroll students with FIES—according to the quality evaluations conducted by the Ministry of Education—are included in the treatment group. Ineligible major-HEI are included in the control group. The pretreatment period consists of the years that precede FIES expansion (2009 and 2010). The posttreatment period consists of the years after the expansion (2011, 2012, and 2013). The estimated coefficients associated with the eligibility effect variable represent the impact of being eligible for FIES on log(tuition). We include time, major-HEI (unit), and field of study–year fixed effects for all the specifications presented in this table. We also include major-level and major-HEI-level covariates: faculty quality, number of majors offered by the HEI, number of employers hired as administrative staff size, number of faculty members, number of enrolled students, number of applicant students per available slot, a measure of major-HEI quality, and a measure of market concentration (Herfindahl-Hirschman Index). Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

TABLE 10. Reduced-Form Estimation: Heterogeneous Treatment Effect

Explanatory variable	*Selective major-HEI (1)*	*Enrolled students (2)*	*Applicants per slot (3)*	*HHI (4)*	*Faculty quality (5)*	*For-profit HEI (6)*
Eligibility effect	0.053**	0.047**	0.052**	0.051***	−0.095	0.066**
	(0.022)	(0.023)	(0.023)	(0.019)	(0.063)	(0.028)
Het. treatment effect	−0.028	−0.003	−0.004	−0.019	0.191	−0.051*
	(0.025)	(0.024)	(0.005)	(0.039)	(0.117)	(0.029)
Summary statistic						
No. observations	15,218	15,218	15,218	15,218	15,218	14,922
R^2	0.926	0.926	0.926	0.926	0.927	0.926

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. The DD specification is the same as in column 5 of table 2, but we include the possibility that the treatment effect—the impact of being eligible for FIES—varies across major-HEI characteristics. See the notes to table 10 for additional details. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

the estimates in table 2 are not driven by a composition effect of differential tuition trends across different fields of study. For four of the fields of study considered, the coefficient is precisely estimated.

We explore whether major-HEIs with different characteristics react differently to FIES expansion. While heterogeneous effects do not pose a threat to our identification strategy, understanding whether specific types of major-HEIs are driving our results is crucial for determining the policy implications of the FIES expansion. Policy implications might be different if, for instance, the price increase is mainly a result of the pricing strategy of for-profit HEIs or of major-HEIs with excess demand. We consider the following variables in our heterogeneous treatment analysis: a dummy variable for selective major-HEI, the number of enrolled students, the number of applicants per available slot, the Herfindahl-Hirschman index, faculty quality, and a dummy variable for for-profit HEIs.[36] Table 10 summarizes the results obtained from this heterogeneity analysis. The eligibility effect is not significantly altered by including five of the six major-HEI characteristics. The only characteristic

36. We define selective major-HEIs as programs that have at least two applicants per available slot. With regard to the for-profit dummy variable, the for-profit sector constitutes a large share of the higher education market in Brazil. In 2013, 46 percent of the students enrolled in the private sector were enrolled in a for-profit HEI. The literature documents how for-profit HEIs can provide suboptimal returns for their enrolled students (see Deming, Goldin, and Katz, 2013). If our result is mainly driven by the pricing strategies of for-profit HEI, it would provide further evidence of predatory behavior by the for-profit sector.

that has a significant impact on the treatment effect is the for-profit dummy variable, suggesting—somewhat surprisingly—that the impact is stronger for major-HEIs in the nonprofit sector. We lose our ability to precisely estimate the eligibility effect when we include faculty quality. This is likely related to the fact that faculty quality is one of the factors determining a major-HEI's performance on the Ministry of Education quality evaluation that defines eligibility.

Finally, we test the robustness of our results by implementing two alternative identification strategies. First, we estimate the effect of eligibility around the minimum quality threshold via a regression discontuity design (RDD). Second, we implement a matching DD approach. Matching DD models can be useful if treatment and control groups may differ in ways that could affect their trends over time. Results—which are consistent with the results in table 2—are presented in the online appendix.

Reduced-form results show that eligibility for a large subsidized student lending program—FIES—caused tuition to rise in Brazil. Specifically, eligibility for FIES caused major-HEIs to increase tuition by an average of 4.7 percent under our preferred specification and by 3.1 percent using the fully saturated specification (table 2). This result is mostly robust to changes in sample composition (see tables 4, 5, 6, and 8) and to different definitions of our eligibility proxy (table 7). Finally, we find evidence that the eligibility impact on tuition is consistent across different field of studies (table 9) and is unaltered even when we consider the possibility that treatment effects vary over time (figure 6 and table 9) or with major-HEI characteristics (table 10).

Thus far, we have shown that there is a causal relation between tuition at the major-HEI level and eligibility for a large government-funded student loan program. To estimate this causal association, we evaluated the difference through time in the tuition set by major-HEIs that are eligible and ineligible to enroll students with FIES. This comparison identified the impact of student credit on tuition at their equilibrium level—the final outcome of major-HEIs' pricing strategy. The previous analysis did little to explain how this equilibrium was achieved. There are a number of reasons why major-HEIs eligible to enroll students financed through a government-backed student loan program might increase tuition after a credit shock. In the short run, major-HEIs might be limited in their ability to adjust production factors in order to absorb the increased demand from students, putting upward pressure on costs. As such, it might be the case that the marginal cost of providing higher education increases relatively more for eligible major-HEIs. Another possible mechanism is related to major-HEIs' ability to set prices above the marginal cost.

It may be the case that the credit shock affects the market power of eligible major-HEIs or, correspondingly, their ability to establish a given markup between price and the marginal cost of providing education. Higher markups can reduce consumers' (in this case, students') welfare.[37] Also, with higher markups, at least part of the government subsidy is transferred—in the form of higher profits—to major-HEIs that increased their market power.

From a policy perspective, it is important to understand whether increases in market power are one of the mechanisms explaining the tuition increase. Directly estimating markups is a challenging task (see Basu, 2019). Instead, we focus on estimating one of the factors that determine HEIs' ability to increase markups, namely, the sensitivity of demand to price changes. All else being equal, major-HEIs can set higher prices when demand is less likely to fall substantially after the price change. In the next section, we build a structural demand model for higher education to estimate the impact of FIES on the price elasticity of demand.

Structural Form: Structure and Identification

Our reduced-form results suggest that FIES eligibility induced major-HEIs to increase their posted tuition. In this section, we explore one possible mechanism behind this result. Specifically, we show that FIES is associated with a reduction in the tuition elasticity of demand.

A credit-driven reduction in demand price elasticity may occur for several reasons. Evidence from mortgage and automobile loan markets supports the idea that, if the gains from acquiring a good or service are sufficiently high, credit-constrained individuals are less sensitive to interest rates (Adams, Einav, and Levin, 2009). In Brazil, gains from tertiary education are so large that the net present value of tertiary education is still positive for a wide range of increases in tuition (Ferreira, Firpo, and Messina, 2014). Students may become price insensitive if they anticipate that they will not repay the debt in full because the government cannot credibly collect. In contrast to the United States, the tax authority in Brazil has a limited ability to collect debt related to student loans. Debt collection may be particularly problematic when aggregate shocks render a large fraction of borrowers delinquent (Farhi and Tirole, 2012). There are also possible behavioral explanations. Price

37. The higher markups might affect both students that have access to credit and students that do not (Espinoza, 2017).

insensitivity may arise if borrowers do not understand interest rates and the future consequences of borrowing. Lusardi, Mitchell, and Curto (2010) show that financial illiteracy is widespread among U.S. youth. Brazil has worse financial literacy than the United States, implying that a behavioral explanation is particularly plausible in this case (see Lusardi and Mitchell, 2011).

Changes in demand tuition elasticity can only affect tuition if suppliers have some pricing power. We argue that this is the case in Brazil. In the national market for private higher education, large conglomerates—such as Kroton, the second-largest listed education company in the world (in 2013)—coexist with numerous small institutions. Data from the 2012 Education Census show that the ten largest groups had 20 percent of enrolled students at the national level. A little over half of the institutions in the sample had fewer than 1,000 students. Nevertheless, concentration is high in the largest local markets. In 2012, the ten largest groups had 32 percent of enrolled students in the states of São Paulo and Rio de Janeiro (the largest and third-largest markets, respectively), 49 percent in Mato Grosso do Sul, and 61 percent in Rio Grande do Norte. There is also variability in quality, which suggests vertical differentiation. In 2012, the average ENADE grade, a proxy for quality, was 2.6, with a standard error of 0.75. High concentration at the local level and vertical differentiation suggest the presence of pricing power.

We consider the following framework for estimating demand in a market with differentiated products.[38] Each major-HEI represents a different product. Consumers' (or students') indirect utility is a function of major-HEI characteristics. Let $t = 1, \dots, T$ be T markets, $j = 1, \dots, J$ be J different major-HEI pairs, and $i = 1, \dots, I$ be I students. We define the relevant market as the state-year pair.[39] The indirect utility of student i enrolled in major-HEI j in market t, U_{ijt}, is given by

$$\delta_{jt} \equiv \mathbf{X}_{jt}\beta - \alpha p_{jt} + \xi_{jt} + \epsilon_{ijt}; \tag{2}$$

$$U_{ijt} = \delta_{jt} + \epsilon_{ijt}. \tag{3}$$

38. In Brazil, universities do not usually price discriminate against students. Even if that were the case, FIES rules require that students enrolled with FIES pay the minimum price paid by other enrolled students. The possibility of price discrimination would make our model more complex.

39. It is reasonable to consider a state as the relevant market for higher education in Brazil, since it is uncommon for students to move to a different state to attend private higher education. In 2013, for instance, the median proportion of students enrolled in major-HEIs located in the same state in which they were born was of 82 percent.

Here, δ_{jt} represents the mean utility from attending major-HEI j at market t, p_{jt} is the tuition of major-HEI j at market t, and α is the marginal effect of tuition on indirect utility. Typically, one assumes that students and HEIs observe all the relevant major-HEI characteristics, but the econometrician does not. Thus, $\mathbf{X}_{jt}$ is the vector of characteristics observed by the econometrician, and ξ_{jt} is the vector of variables not observed by the econometrician but observed by the HEI and the student. Finally, ϵ_{jt} is an individual major-HEI-specific error, which is observed only by the individual. The outside option—in this case, choosing not to go to college or choosing to enroll in a public university—completes the specification. We define the demand for the outside option as the difference between the total number of individuals aged fifteen to twenty-four years in a given market and the total number of individuals enrolled in private universities in that market.

We assume students choose only one major-HEI pair, which is a reasonable assumption. We integrate out with respect to individual shock ϵ_{ijt}. For our first set of results, we make the convenient assumption that ϵ_{ijt} is independent and identically distributed (i.i.d.) across major-HEIs (j), years (t), and individuals (i). We also assume that ϵ_{ijt} follows a Type I extreme value distribution (EVI). The i.i.d. and EVI assumptions yield a multinomial logit model of demand. The logit model allows us to derive a closed-form formula for market shares. The market share of major-HEI j in market t, s_{jt}, is given by

$$s_{jt} = \frac{\exp(\delta_{jt})}{1 + \sum_{k=1}^{J} \exp(\delta_{kt})}. \tag{4}$$

Own-price elasticity of demand, that is, the percentage variation in demand in response to a 1 percent increase in tuition, is given by

$$\frac{\partial s_{jt} p_{jt}}{\partial p_{jt} s_{jt}} = -\alpha p_{jt}\left(1 - s_{jt}\right). \tag{5}$$

Let s_{0t} be the market share of the outside option. Taking logs on both sides of equation 4 and subtracting the log of the outside option gives us equation 6, a linear regression model:

$$\ln\left(s_{jt}\right) - \ln\left(s_{0t}\right) = \mathbf{X}_{jt}\beta - \alpha p_{jt} + \xi_{jt}. \tag{6}$$

The market share s_{jt} is an observed quantity. It represents the ratio between the number of students newly enrolled at major-HEI j in market t and the total number of potential students in that same market t.[40]

Inspecting equation 6 shows how difficult it is to identify its parameters. ξ_{jt}—the error term—is (potentially) observed by the HEI and the students, and thus "priced into" tuition p_{kt}. Many unobservable factors can affect students' decisions and tuition, such as convenient location and advertising expenses. Because we do not observe ξ_{jt}, p_{jt} is endogenous. Identification in this case relies on using appropriate instruments for the price variable p_{jt}. We use four cost shifters as instruments: two accounting measures of cost (namely, total expenses with faculty's salary and current expenses at the HEI level) and two indirect measures of cost (the median salary of college instructors and of college administrative staff at the city-year level).[41] Identification comes from the assumption that these cost shifters have no impact on a student's indirect utility, once we control for quality.

The assumptions we use to build the logit demand model are not innocuous. First, they impose a somewhat unrealistic pattern on own-price elasticities. Under these assumptions, own elasticities increase with price. Our assumptions also place a priori restrictions on cross-price elasticities, that is, on demand changes that result from changes in competitors' prices. Specifically, the i.i.d. assumption implies that the cross-price elasticity between any two major-HEI pairs is driven by their markets shares (Berry, 1994). Considering the limitations of the logit model, we also estimate a nested logit model, in which we classify products—major-HEIs—into broader groups. The decision process of consumers (in this case, students) is sequential: students choose first a group and then a product. With this nested model, we allow consumer preferences to be correlated across majors within a defined group, while

40. Measuring quantities in this industry is not trivial. In Brazil, enrollment requires having a high school diploma and, typically, passing entrance exams. Students declare their majors on registering for the entrance exams. There are excess vacancies for some major-HEI pairs, and all eligible applicants are approved. In this case, it is straightforward to measure quantity as the number of enrolled students. However, a little more than half of the major-HEI pairs in the sample have more candidates than available slots, in which case demand is rationed, and quantity is arguably better measured by the number of newly enrolled students. We obtain qualitatively similar results using the number of enrolled students or the number of applicants as measures of quantity demanded. Results are available on request.

41. Including year-city fixed effect precludes the use of these two instruments.

maintaining the assumption that consumer utility is i.i.d. across groups (and has a type I extreme value distribution).

We consider every major-HEI pair as part of one group. Here we define the broad field of study as the relevant group. Following Berry (1994), we calculate the share of each major-HEI within its group and estimate the nested logit model including a new variable: the logarithm of the within-group share. This variable is endogenous, and we need to include appropriate instruments. From the RAIS data set, we obtain the median wage (at the market level) of workers employed in occupations related with each of the eight broad fields and each of the twenty-two specific fields of study as defined by the Ministry of Education. Each of our major-HEI pairs is identified within a broad group and a specific field of study. We need an instrument that shifts demand behavior within each group, which we define as the ratio between the median wage in the specific field of study and the median wage in the broad field of study. Since we control for the median wage in the broad field of study, it is reasonable to assume that the exclusion restriction is satisfied.

We estimate the parameters of equation 6 using two-stage least squares (2SLS) for the logit and nested logit specifications. We calculate standard errors clustering all specifications at the city level. To capture quality and other major-HEI characteristics that influence demand, we include the same large set of controls from our reduced-form specification (see table 1 for covariates' descriptive statistics). We also include the aforementioned set of instruments. Table 11 presents descriptive statistics for the instruments considered in our structural analysis. Finally, we include time and field of study fixed effects.

Table 12 presents the results of the first stage. For both the logit and nested logit specifications, the first stage shows that we have strong instruments. For all cases, the signs of the estimated coefficients are as expected, and we have large values for the partial F test. Table 13 presents the estimated parameters of our demand system. The estimated coefficient on tuition is negative and significant for all specifications.

Our ultimate goal is to explore whether credit availability had an impact on the tuition elasticity of demand. Following, Espinoza (2017), we calculate demand tuition elasticities for two separate periods, before and after the expansion of credit. Alternatively, we could have included FIES availability as a factor influencing students' utility, which would have allowed us to estimate the FIES impact on tuition elasticity directly. We opted not to follow this strategy, however, since it would raise additional endogeneity concerns. If unobserved factors influence both students' utility and the availability of

TABLE 11. Descriptive Statistics: Instruments

Variable	Descriptive statistic
Median wage in field of study ratio	1.045
	(0.427)
Median wage of faculty	1,143
	(475.5)
Median wage of administrative staff	497.1
	(113.5)
log(expenses including faculty)	5.825
	(0.872)
log(expenses including maintenance costs)	5.312
	(1.452)
No. observations	17,945

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: This table presents descriptive statistics for the instruments used for the estimation of the structural demand model. We use four cost-shifters as instruments for tuition: two accounting measures of cost (total expenses with faculty and with maintenance at the HEI level), and two indirect measures of cost (the median salary of college instructors and of college administrative staff at the city-year level). We also use an instrument for the nested logit specification: the ratio between the median wage in the specific field of study and the median wage in the broad field of study. The final sample covers the period between 2009 and 2013 and consists of 17,945 major-HEIs. For each variable, we present their average value at the major-HEI level. Standard errors are in parentheses.

TABLE 12. First-Stage Structural Form Estimation: Logit and Nested Logit Models

	Logit		*Nested Logit*	
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*
log(expenses including faculty)	69.2403***	48.5653***	69.4996***	48.6922***
	(10.854)	(7.793)	(10.793)	(7.762)
log(expenses including maintenance costs)	12.0736***	5.8531**	12.0529***	5.8340**
	(3.928)	(2.856)	(3.931)	(2.854)
Median wage of faculty	0.0465***	0.0099	0.0667***	0.0468***
	(0.016)	(0.013)	(0.016)	(0.013)
Median wage of administrative staff	0.2612***	0.1327***	0.2380***	0.2628***
	(0.048)	(0.040)	(0.048)	(0.040)
Median wage in field of study ratio			19.7512***	11.4228
			(6.983)	(7.964)
Summary statistic				
No. observations	17,945	17,945	17,945	17,945
Partial *F* test	33.38	11.80	31.59	10.35
Year fixed effects	Yes	Yes	Yes	Yes
Field of study fixed effects	Yes	Yes	Yes	Yes
Covariates	No	Yes	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: This table presents the estimates of the first stage of a demand model based on equation 6. The dependent variable is the logarithm of tuition measured in 2008 Brazilian reais. Columns 1 and 2 present the results of a basic logit model. We use four cost-shifters as instruments for tuition: two accounting measures of cost (total expenses with faculty and with maintenance at the HEI level), and two indirect measures of cost (the median salary of college instructors and of college administrative staff at the city-year level). In column 1, we include time and field of study fixed effects. In column 2, we include the same set of covariates used in column 4 of table 2. Columns 3 and 4 present the results of a nested logit model, in which we allow consumer preferences to be correlated across majors within a defined group, while maintaining the assumption that consumer utility is i.i.d. within groups. For these specifications, we include an additional instrument: the ratio between the median wage in the specific field of study and the median wage in the broad field of study. Column 3 includes time and field of study fixed effects, and column 4—our preferred specification—includes the same covariates included in column 2. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

TABLE 13. Structural Form Estimation: Logit and Nested Logit Models

	Logit		*Nested Logit*	
Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*
Tuition (in 2008 reais) (log)	−0.003**	−0.004**	−0.002**	−0.003*
	(0.001)	(0.002)	(0.001)	(0.001)
Within-group market share in field of study (log)			0.269	0.614**
			(0.211)	(0.310)
Summary statistic				
No. observations	17,945	17,945	17,945	17,945
Cragg-Donald Wald *F* statistic	324.2	161.4	217.9	123.3
Year fixed effects	Yes	Yes	Yes	Yes
Field of study fixed effects	Yes	Yes	Yes	Yes
Covariates	No	Yes	No	Yes

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: This table presents the estimates of a demand model based on equation 6. Columns 1 and) present the results of a basic logit model. To estimate the logit specification, we implement a two-stage least-squares regression. We use four cost-shifters as instruments for tuition: two accounting measures of cost (total expenses with faculty and with maintenance at the HEI level) and two indirect measures of cost (the mean salary of college instructors and of college administrative staff at the city-year level). In column 1, we include time and field of study fixed effects. In column 2, we include the same set of covariates used in column 4 of table 2. Columns 3 and 4 present the results of a nested logit model, in which we allow consumer preferences to be correlated across majors within a defined group, while maintaining the assumption that consumer utility is i.i.d. within groups. We estimate the nested logit model including the logarithm of the within-group share in the specification. To properly estimate the impact of this variable, we include an additional instrument, the ratio between the median wage in the specific field of study and the median wage in the broad field of study. Column 3 includes time and field of study fixed effects, and column 4—our preferred specification—includes the same covariates included in column 2. Robust standard errors, computed with observations clustered at the HEI level, are in parentheses.

FIES, we would need to include additional instruments in our specification to properly identify the effect of interest.

To calculate elasticities, we use the estimated parameters from a nested logit specification (column 4 of table 13). Table 14 shows the results of this exercise. The values are mostly consistent with theory. When we compare tuition elasticity before and after the FIES expansion, we find a reduction in the price elasticity of demand in every field we considered. This result indicates that one of the likely mechanisms behind the FIES-induced price increase is a reduction of demand elasticity.

Conclusion

Over the last few decades, policymakers in developing countries have created and expanded subsidized student lending programs as a way to increase higher education enrollment rates. Understanding whether these programs

TABLE 14. Structural Form Estimation: Mean Elasticity of Demand by Field of Study

Field of study	*Pre-FIES (2009)* *(1)*	*Post-FIES (2013)* *(2)*
Education	−0.987	−0.956
Humanities and arts	−1.313	−1.315
Social sciences, business, and law	−1.224	−1.238
Natural sciences, math, and computer science	−1.975	−1.879
Engineering and construction	−1.696	−1.656
Agricultural sciences	−2.185	−1.966
Health	−1.890	−1.842
Others	−1.202	−1.189

* $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

Notes: This table presents the mean elasticity—the percentage change in demand in response to a 1 percent change in price—for each field of study in our sample considering two different moments in time: before the expansion of FIES (2009) and after the expansion of FIES (2013). We use the estimated parameters in column 4 of table 14 to calculate elasticities for each major-HEI at each moment in time. We then aggregate elasticities to obtain a mean elasticity for each year/field of study.

cause tuition to rise is of first-order importance. Higher tuition increases the debt burden on students. As Espinoza (2017) shows, credit-driven tuition hikes can also affect the welfare of students that do not participate of these programs. Finally, higher tuition may also affect the fiscal cost associated with the programs, depending on the program design.

In this paper, we explore the expansion of FIES, a major federal government student lending program in Brazil, to evaluate whether eligibility for student loans at the major-HEI level causes tuition to rise. Using a unique data set with annual information on tuition at the HEI and major level, we conclude that eligibility for FIES caused major-HEIs to increase tuition. Specifically, we show that eligibility for FIES caused major-HEIs to increased tuition by an average of 4.7 percent.[42] This result is robust to variations in sample composition, to the possibility of heterogeneous effects, and to different strategies for the definition of control and treatment groups.

From a policy perspective, it is important to understand if increases in market power are one of the mechanisms explaining the tuition increase. We therfore investigate whether the expansion of student credit had an impact on the tuition elasticity of demand. We estimate a demand system for differentiated products based on the classic demand models from the industrial organization literature. Our results show that the tuition elasticity of demand

42. Considering our preferred specification.

fell after the expansion of FIES. A credit-driven reduction in demand price elasticity may occur for several reasons. For instance, there might be behavioral explanations, or students may be anticipating that they will not repay their debt in full. In this paper, we focused on the effect of the reduction in elasticity on the pricing strategy of higher education institutions. We leave the mechanism through which the reduction in elasticity occurs to future research.

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