UNIVERSITY OF LEICESTER

School of Business
Economics Division

Access to education and teenage childbearing

Martin Foureaux Koppensteiner, University of Leicester
Jesse Matheson, University of Leicester

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Martin Foureaux Koppensteiner Jesse Matheson†

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Abstract

We look at the effect of expanding secondary school access on teenage childbearing in Brazil. For this purpose we combine information from the Brazilian school census with vital statistics data. Variation in the introduction of schools across municipalities over time is used to estimate the effect of education access on teenage births. Our results show a 4.56% reduction in municipal teenage childbearing following a school introduction. These results suggest that Brazil's secondary school expansion between 1997 and 2010 can account for 25% of a substantial decline in teenage childbearing observed over the same period.

JEL Codes: I20, I26, J13

Keywords: Secondary education, teenage childbearing, Brazil

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† Both authors are based in the Department of Economics at the University of Leicester: mk332@le.ac.uk and jm464@le.ac.uk.

1. Introduction

A recent report by the World Bank (2012) on teenage pregnancy points out a correlation between teenage childbearing and a number of socioeconomic variables including poverty, inequality, public health expenditure and female labour force participation. The report further shows that, despite substantial reductions in teenage pregnancy rates in virtually all countries, we continue to see rates differ vastly between high- and low-income countries, with Brazil—as a middle-income country—being placed somewhere in the middle of the distribution.

Improved access to education, not explored in detail in the World Bank report, potentially provides an important channel through which teenage pregnancy and the above socioeconomic variables are correlated. Casual observation suggests a strong negative relationship between school availability and teenage childbearing in Brazil. In Figure 1 we plot this relationship over time. Starting from 12,684 secondary schools in 1996, the number increased by more than 57%, to 19,964, by 2009. Over the same period, the number of births by teenage girls decreased by 19%. Evidence based on the cross-section presents a very similar picture. In Figure 2A we present the state-level relationship between school density and rate of teen childbearing. There is a clear negative relationship suggesting that a one-unit increase in schools per 1,000 teens is associated with a decrease of about 21 births per 1,000 teens. The negative relationship persists when plotting the data at the level of the municipalities (Figure 2B).

In this paper, we study the effect of improving access to secondary school on teenage childbearing. We provide the first causal evidence of the effect of educational access by exploiting variation from the expansion of secondary schooling across 4,850 Brazilian municipalities over a 14-year period. Our main results suggest that, on average, an additional secondary school will decrease teenage childbearing by approximately 4.56% relative to the mean number of births for a given cohort. We also find that improved access to schools leads to a persistent decrease in childbirths up to age 23 (the oldest age for which we are able to perform our analyses). Back-of-the-envelope calculations based on our estimates suggest that approximately 25% of the total decline in teenage childbearing observed in Figure 1 can be attributed to the expansion of secondary schools.

Brazil is particularly well suited for studying our research question. The school expansion that we examine constitutes one of the largest expansions of secondary schools on record. We use information from 14 waves of the annual Brazilian school census, containing detailed information on the universe of Brazilian schools, to create a new dataset reflecting the availability of secondary schools in every Brazilian municipality between 1996 and 2009. We combine this information with vital statistics data from Brazil capturing the universe of live births over the same period, creating a rich and unique dataset. To our knowledge, this is the first paper to document and utilise data on the rapid growth of secondary schools across Brazil over the last two decades.

In this study, school 'access' is improved through a reduction in geographic distance, thereby reducing a cost associated with school attendance.[1] There is very limited evidence on the effect that the cost of attending school has on teenage fertility. Duflo, Dupas and Kremer (2010) provide a rare exception, using experimental evidence from Kenya. They find a large, negative impact on the fertility of young women from the provision of free primary school uniforms. A much more established literature looks at the relationship between time spent in education and teenage childbearing. Berthelon and Kruger (2011) find that a 20 percentage-point increase in the municipal share of full-day schools (as opposed to half-day schools) decreases the probability of adolescent motherhood by 3.3% in the context of secondary schools in Chile. Several papers use variation from changes to mandatory schooling laws. Black, Devereux and Salvanes (2008) look at variation in teen births resulting from changes to minimum dropout age laws in the US and Norway. Relative to a minimum school-leaving age of 15, they estimate a reduction in teenage childbearing of 4.7% for school-leaving age of 16 and 8.8% school-leaving age of 17. Monstad, Propper and Salvanes (2008) find that the change in the mandatory dropout age in Norway led to postponement of births from teenage years to the mid-thirties. For the UK, Geruso and Royer (2014) estimate that an extra year of schooling—induced by a change in the minimum leaving age—led to a 30% reduction in births at ages 16 and 17.

In this study, we make three key contributions to the existing literature. First, we provide estimates for an explicitly different population than do studies using changes in mandatory schooling laws. While mandatory schooling laws work by constraining individual choice sets (requiring school attendance regardless of other obligations), school expansions work by reducing

[1] This follows a rational similar to that in Currie and Moretti (2003) who look at college attendance in the US.

the cost of school attendance. The resulting group of 'compliers' in our study is strictly different than the compliers in mandatory schooling studies. In our case, compliers are those for whom the perceived return to schooling is relatively high, whereas compulsory schooling laws elicit changes from the bottom of the education distribution (Brunello, Fort and Weber, 2009; Fort, Schneeweis and Winter-Ebmer, 2015). It is not obvious that estimates based on mandatory schooling laws are informative for the margin of the population explored here. Further, Fort, Schneeweis and Winter-Ebmer (2015) find that the observed effect of mandatory laws on fertility is context-dependent. Despite the stark difference in the approach chosen in this paper, our estimates are similar, in sign and magnitude, to the findings of studies using compulsory schooling laws in high-income countries (Black, Devereux and Salvanes, 2008; Monstad, Propper and Salvanes, 2008), providing additional evidence on the role education plays for teenage fertility decisions.

The second contribution arises from our focus on a middle-income, as opposed to high-income, country. This enables us to study a recent and rapid expansion of secondary schools over a relatively short time period, which is not possible in the context of high income countries, where such expansions have taken place decades earlier and often in a more gradual fashion. We are therefore able to provide the first quantification of the role of expanded education access in the substantial reduction of teenage fertility rates observed across many middle-income countries over the last two decades. The Brazilian secondary schooling expansion provides a blueprint for understanding the effect the expansion of the educational system has on fertility. The estimates provided here are therefore of relevance for a large group of low- and middle-income countries currently experiencing or on the brink of similar expansions of their educational systems.

As a third contribution, our novel estimation strategy allows us to provide new evidence on the potential structural channels underlying the education-fertility relationship. We study in more detail the role of human capital versus an incarceration channel—the two channels which in previous studies have been challenging to distinguish between—by separately estimating the introduction of secondary schools in urban settings versus rural settings. We find that in urban settings the effect of improved school access on teenage childbearing persists into early adulthood, whereas this effect dissipates in early adulthood in rural settings. These differences for urban versus rural schools are consistent with differential labour market returns, which would be larger in urban than in rural settings, leading to substitution from childbearing to labour market participation (at least in early adulthood).

The remainder of this paper is organised as follows: In Section 2 we provide a background discussion on the provision of secondary education in Brazil. In Section 3 we discuss the data to be used in the main analysis. In Section 4 we introduce the empirical strategy used for our main result. The main result, as well as a number of robustness checks is presented in Section 5. In Section 6 we discuss and present evidence for the possible mechanisms underlying our main results. Section 7 concludes the article.

2. Background information

Many low- and middle-income countries have, over the last two decades, undergone an expansion of their secondary education system, driven by improvements in primary school completion and an increased demand for a more highly skilled workforce (World Bank, 2005). In Brazil, secondary schooling was an overlooked part of the education system until the beginning of the 1990s (Guimarães de Castro and Tiezzi, 2004). Secondary education was highly geared to the elites preparing for entrance to higher education and considered of little relevance for the education of the broader population. Following the end of the military dictatorship, the new constitution of 1988 made access to secondary education a key aim on the political agenda, mandating it to be available (although not mandatory) for all those completing primary education. Significant changes were made in 1996 when the government of Fernando Henrique Cardoso passed the General Education Law (Lei de Diretrizes e Bases da Educação Nacional(LDB) 1996). The LDB outlined the *progressive universalisation of access to free secondary education* through gradually increasing access to state-funded public secondary schools (Marchelli, 2010). Following the rapid expansion of primary education, which was virtually universal by the early 1990s, secondary education started to expand (Moore, DeStephano, Terway and Balwanz, 2008; Di Gropello, 2006; De Felizio, 2009). The number of students in secondary schools increased from under 5.7 million to 8.3 million (INEP, 2003; INEP, 2011). This increase is mirrored by a steep increase in education expenditure; between 2000 and 2009, Brazil reported the largest increase in education spending as a percentage of total public expenditure for 33 countries for which data is available (OECD 2012).

Secondary education in Brazil is largely the responsibility of the 26 states and is preceded by primary schooling—compulsory for children aged six to 14 years of age—lasting nine years. Of the 15,219 public secondary schools in 1997 for the municipalities in our study, 97% were under state control. State secretariats of education are responsible for the regulation and general management of secondary schools, including the recruiting of teachers and curriculum content (JBIC, 2005). There is no minimum age for initial enrolment to secondary school, but it is targeted at age 15, and students must have completed primary school first.[2] There is no maximum age limit, and because of frequent late enrolment and grade retention in primary schools, age-grade mismatch at secondary school is frequent (Foureaux Koppensteiner, 2014). In 2010, about 30% of students in the first grade of secondary school were above the target age (IBGE, 2012).

The expansion of the secondary system reflects the additional demand brought on by increased primary education enrolment numbers and completion rates (Guimarães de Castro and Tiezzi, 2004). In unison with the secondary school expansion, efforts were undertaken, through new curricular guidelines, to broaden the secondary curriculum towards more professional education, moving away from exclusively preparing for access to the higher education system (MEC, 2002; De Moura Castro, 2009). Thus, secondary schools have been changing their appeal for a larger segment of the Brazilian population.

In Figure 1 we illustrate the increase in secondary schools across Brazil for the period 1997 to 2009. Overall, there was a 57% increase in the number of secondary schools, from 12,684 in 1997 to 19,964 in 2009. This was driven primarily by a 68% increase in the number of publicly funded schools, from 9,068 to 15,219. There was also a notable 31% increase in the provision of private secondary education, from 3,616 to 4,745.[3]

The school expansion had a non-trivial impact on school access across Brazilian municipalities. In particular, there was a remarkable increase in the availability of schools to the poorer northern states of Brazil.[4] In Figure 3 we depict the change over time in the proportion of

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[2] Recently primary education has been extended to 9 years and starts at the age of 6. For the most of our analysis before the mid-2000, primary education started at the age of 7 and lasted for 8 years.
[3] These numbers exclude the 33 municipalities with populations greater than 500,000. These municipalities also exhibited a significant expansion in secondary schools of 58% for publicly funded schools and 24% for privately funded schools.
[4] In appendix figure A1a we map the observed school expansion between 1997 and 2009.

municipalities that had low school density—defined as less than one year-one classroom per 100 youths aged between 13 and 18. The figure shows a steady decline over time in the proportion of low-density municipalities, and this decline is observed for municipalities of all population sizes. The number of municipalities with no secondary school decreased from 401, representing 8% of all municipalities, in 1996 to 11 in 2009.

3. Data

The primary data used in this study comes from two sources: the Brazilian school census (Censo Escolar) and Brazilian vital statistics data from the Ministry of Health. In addition, we use auxiliary data, including population estimates for Brazilian municipalities, from the Brazilian Census Bureau and municipal expenditure data from a variety of sources. These data are discussed below, followed by a discussion of the minimal comparable areas that are used to link geographic areas over time. Population information and descriptive statistics for the municipalities in our sample are reported in Table 1.

3.1 Schooling data

We use 14 waves of the Brazilian school census, collected annually by the Anisio Teixeira Institute of Research on Education (INEP) at the Ministry of Education. The school census includes detailed information on the universe of public and private schools in Brazil, such as enrolment by grade, age and sex, information on the number of classes, the physical characteristics of the schools, as well as information on teachers.[5] We use this information to create a dataset on the number of secondary schools, the number of classrooms and the number of students between 1996 and 2009, collapsing the data by municipality and year. Information on unique municipality codes allows us to locate every school in Brazil to the corresponding municipality. The school census also provides information on primary school enrolment, the availability of nursery classrooms, pre-school classrooms and the location of school in rural or urban areas, which we use as control variables as well as to investigate the channels at work. In Table 1, we show that there are on average three secondary schools per municipality; the vast majority of secondary schools are publicly provided and school density is much higher in urban areas than in rural municipalities.

[5] These data can be downloaded at the website of INEP (www.inep.gov.br).

3.2 Childbearing data

Data on birth outcomes come from the microdata of Brazilian vital statistics, which cover approximately 45 million births occurring between 1996 and 2009. Vital statistics data are based on birth certificates issued by health institutions or midwifes attending homebirths and are collected through the states' health secretariats. The vital statistics microdata are publicly available through the System of Information on Life Births (SINASC) of the IT department of the Brazilian public health system (DATASUS). These data provide information on the age and municipality of residence for the mother, as well as gestational length of the pregnancies, and the mother-reported race of the child.

For each year we collapse these data to create a count of births by municipality and mother's age at conception. We calculate age at date of conception using information on gestational length recorded in the birth certificates. The annual data are then merged to provide a municipal panel of births by mother's age at conception. Brazilian vital statistics data show excellent coverage of all occurring births; information from the 2010 population census shows that more than 99% of all births occurring between 2000 and 2010 were registered and entered into the vital statistics data we use. The advantage of using vital statistics data to learn about fertility in the population comes from the universal coverage of the data for the entirety of Brazilian municipalities over the period of interest. Information about the residence of the mother during pregnancy is particularly important, as information on the place of birth may be misleading if there is a discrepancy between place of residence of the mother and the place of occurrence of birth, which is more likely for relatively small municipalities that do not have clinics with birth facilities.

Figure 5 displays the distribution of births per 1,000 women in Brazil by age of conception (in years) from 1996 to 2009. The figure reveals a peak in fertility in the early 20s, with a substantial fraction of these births occurring before women reach age 20. This confirms that not only are teen childbearing rates substantially higher in Brazil compared to countries such as the US, the UK or Germany, but overall fertility happens at a much earlier age. Around 22% of the 45 million births occurring over this period were births by teenage mothers. With a teenage childbearing rate of 84 per 1,000 women in the 15–19 year age group, Brazil ranges clearly above the OECD average (for example 45, 29 and 10 per 1,000 women in the US, UK and Germany respectively), but below many low-income countries for which data is available (for example, the highest recorded rate of teen pregnancies is 207 per 1,000 women in Niger).[6]

3.3 Auxiliary data

Population estimates

The Brazilian Census Bureau (IBGE) provides official population estimates for each municipality based on the 1990 and 2000 census and the 1996 and 2006 population counts. These data provide population estimates by sex and age group that we use in all the regressions to account for cohort sizes.

Municipality controls

We also use a rich set of municipality-level time-varying data on the characteristics of the municipalities from a variety of sources. These include municipality GDP, the fraction of municipality level expenditure on education, health, welfare, justice and security provided by IBGE. In addition, we include information on the number of health institutions, the number of nurses employed in these institutions, the number of *Bolsa Famíli*a recipients and the total amount of *Bolsa Família* payments in the municipality. These data are available annually for the 1996–2009 time period.[7] We provide details on the source of these data in the appendix.

3.4 Minimal comparable areas

Brazilian municipalities are our primary unit of observation. There are currently 5,570 municipalities in Brazil, and they constitute the country's smallest administrative divisions, similar to US counties. We link information on the availability of schools with the vital statistics data using unique municipality identifiers. Over the period we are interested in a number of municipality boundaries were redefined. Specifically, 533 new municipalities are introduced in 1997, 54 new municipalities are introduced by 2001 and an additional 58 again by 2010. To account for this, we create *minimal comparable areas*. IBGE (2011) provides information on the origins of municipalities since 1996 which we use to build stable geographic area definitions. If, for example, two municipalities were created by splitting one municipality, the two new

[6] Age specific fertility rates, as pregnancies per 1,000 women aged 15–19 for the period 1995–2010 (UN World Population Prospects 2015).
[7] The Bolsa Família programme was introduced in 2004.

municipalities are recoded to the same minimal comparable areas. This results in 4,885 of such units. For simplicity, we will continue to refer to our units of observation as municipalities throughout this article.

Municipality size varies tremendously in Brazil. For the analysis, we restrict the sample to municipalities with populations of less than 500,000 residents (by the 2000 census count), excluding the largest 35 Brazilian cities. This is to ensure that the included municipalities are small enough to credibly link the population of interest to the introduction of secondary schools. We end up with 4,850 geographic units for the analysis, representing 99% of all Brazilian municipalities and 71% of the overall population. Summary statistics with respect to the population distribution for included municipalities are reported in Table 1.

4. Estimation strategy

In this section, we provide the details of our estimation strategy and discuss the key threats to identification. This is followed by an analysis of the response in student enrolment to the school expansion.

4.1 Empirical Strategy

As pointed out in the introduction, a key empirical challenge in estimating the causal effect of school availability on teenage motherhood is addressing the potential confounding effect of unobservable municipality characteristics that are correlated with the number of available secondary schools and with the fertility decisions of teenagers. For example, if the perceived return to education is low, this may impact both the number of secondary schools—in response to low demand—and the timing of fertility. This will create a spurious correlation between incidence of teenage childbearing and low availability of secondary schools. To overcome this issue of erroneous inference, we use variation over time in the number of secondary schools available in each municipality to identify the effect of school access on teenage childbearing. This identification strategy is similar to a difference-in-differences framework.

The following equation describes the relationship between childbearing and schools we estimate:

$$B_{ic}^{a} = \alpha^{a} S_{ic} + \boldsymbol{X_{ic}^{a\prime}} \boldsymbol{\Gamma^{a}} + \delta_{c}^{a} + \eta_{i}^{a} + \epsilon_{ic}^{a} \quad (1)$$

where B_{ic}^{a} is the cumulative number of births in municipality i for cohort[8] c conceived between age 15 and age a. S_{ic} is the number of secondary schools in municipality i in the year that cohort c was 15 years of age. To illustrate, for the cohort that is 19 in 2007, S_{i19} is the number of secondary schools in municipality i in 2003. X_{ic}^{a} is a vector of observable municipal control variables, some of which may vary with cohort age. δ_{c}^{a} captures cohort fixed effects (i.e. time effects) and η_{i}^{a} captures cohort/time-invariant municipality specific unobservables. ϵ_{ic}^{a} reflects idiosyncratic cohort varying unobservables.

The coefficient of interest, α^{a}, is identified by the within municipality change in S_{ic} across cohorts and reflects the average effect of an additional secondary school on cohort childbearing. This identification strategy assumes that unobservables, ϵ_{ic}^{a}, are independent across cohorts. This assumption would be violated if, as an example, cross-cohort peer effects are important in determining teen childbearing.

Our identification strategy relies on a number of lagged variables. Specifically, we include up to the 4$^{\text{th}}$ lag in the number of secondary schools and the 5$^{\text{th}}$ lag in primary school enrolment (the rational for this variable is discussed in detail below.) As the data cover the years 1997–2009, our primary estimates are based on the change in secondary school availability, S_{ic}, between 1998 and 2005, and reflect outcomes, B_{ic}^{a}, for cohorts up to age 19 observed between 2001 and 2009. This provides nine years of outcomes across 4,850 municipalities for a total of 43,650 observations.

There are three specific challenges about which we need to be concerned in the estimation of Equation (1). The first is that birth outcomes and school supply may both be driven by the characteristics of teen cohorts. Strong academic performance in primary school for a particular cohort may lead to the introduction of secondary schools. Strong academic performance may also be the result of cohort-specific unobservable characteristics, which are plausibly correlated with low teen birth. This may lead us to incorrectly infer a negative causal relationship between schools and teen births. To address this, in all regressions we include cohort-specific enrolment numbers for the final-year primary classes. Weaker cohorts, reflected by students having dropped out of

[8] It should be noted that our data do not allow us to link individuals over time. Therefore, *cohort* is defined according to date-of-birth (age) and municipality. As we will discuss shortly, the extent to which we are able to identify the same underlying population year-on-year will be challenged by migration. With this in mind, we use the term *cohort* to mean any two populations, observed at different years, for which all members have the same date-of-birth and live in the same municipality.

primary school or being held back in earlier primary years, will have lower enrolment numbers in their final year. We further provide a robustness check by regressing birth outcomes on the lead values of schools (see Section 5.2, Table 3). The results suggest that cohort birth-characteristics do not predict the introduction of a secondary school.

A second challenge to identification is the potential existence of unobserved municipal-level changes to policy or infrastructure which are correlated with school introductions. If, for example, unknown to the researcher, family planning clinics are introduced alongside secondary schools, we will incorrectly attribute the effect of such clinics to the schools. We take a number of steps to mitigate this. First, we want to rule out that major federal and state health programmes are coordinated with the expansion of schools. In particular, we are interested in programmes that include family planning components, such as the Family Health Programme (*Programa Saúde da Família*).[9] The majority of the rollout of this federal Ministry of Health programme happened around the millennium; we find no evidence of coordination with the expansion of secondary school expansion led by state education secretariats (Rocha and Soares 2010). Second, we attempt to account for possible unobserved municipal changes by controlling for municipal expenditures across a number of services,[10] including health expenditures. Third, we conduct a robustness check in which we estimate Equation (1) for older cohorts (see Section 5.2, Table 3). We do not find a systematic change in births for older age groups with school introductions. Fourth, we conduct an event study to look at the possibility of systematic pre-trends (see Section 5.3, Figure 5). The event study strongly suggests that birth outcomes did not systematically vary prior to the introduction of secondary schools.

A third challenge to identification might arise because of selective migration. This might be a concern if we think that the introduction of a secondary school may induce an inflow of academically engaged (and less sexually active) teens and/or an outflow of non-academic (and more sexually active) teens, or visa-versa. Because Equation (1) reflects the total number of births, estimates will not be impacted by flows of non-sexually active teens, only by those who are

[9] Another federal health programme, the School Health programme (*Programa Saúde na Escola*), was introduced only in 2007 at the end of our period of interest. The Stork Network (*Rede Cegonha*), which has a focus on maternal and child health, and includes a family planning component, was launched in 2011 after our period of interest.
[10] Specifically, we control for (per-capita) spending on welfare, education, health, transportation and housing, as well as per-capita and per-recipient *Bolsa Família* transfers, per-capita income and per-capita total municipal spending. See Appendix C for data sources.

sexually active. In Appendix B we provide an analysis of this possibility based on information from the 2010 Brazilian population census. We find that a) females who migrate between ages 10 and 18 are significantly more likely, relative to other ages of migration and the general population, to be teenage parents; b) migrants tend to leave municipalities with low school growth in favour of municipalities with high school growth; c) the difference in growth between current and previous municipality is higher for teenage parents than for non-teenage parents. All of these facts work against finding a negative relationship. This suggests that migration may lead us to underestimate the magnitude of a negative impact of schools on teenage childbearing.

4.2 School expansion and enrolment

Before considering the effect of the expansion of secondary schools on teen childbearing, it is useful to understand whether school expansions had a meaningful impact on enrolment in secondary school. It is possible that the introduction of a new school leads to the reallocation of existing students among schools, but has little effect on aggregate enrolment numbers. This would call into question the validity of the hypothesised relationship between childbearing and school attendance. To examine this we look at enrolment using data from the school census.

Using the school census, we consider a change in aggregate enrolment following the introduction of a secondary school into a municipality, estimating the following regression equation:

$$e_{it}^k = \vartheta_1^k S_{it} + \rho_i^k + \theta_t^k + \tau_{it}^k, \qquad (2)$$

where e_{it}^k is the number of students of type $k \in \{male, female, age15, \dots\ age18, adult\}$ enrolled in secondary school in municipality *i* in year *t*. S_{it} is the number of year-one secondary classrooms in municipality *i* and year *t*. θ_t^k captures type-specific time trends in enrolment, ρ_i^k capture sex-specific municipality fixed effects, and τ_{it}^k captures all other unobserved influences on student enrolment. The estimated correlation between the change in classrooms and enrolment, $\hat{\vartheta}_1^k$, provides information about the average, contemporaneous change in total municipal enrolment following the addition of a year-one classroom.

Estimates are reported in Table 2 for total enrolment (Column 1), female enrolment (Column 2), male enrolment (Column 3) and enrolment by age in columns 4 to 9. We find a strong, positive, correlation between schools and enrolment numbers for both sexes; the introduction of a secondary school leads to an average enrolment increase of 27.3 female and 16.9 male students.[11] While school enrolment for adults is common in Brazil, the increase in enrolment appears to be largely caused by an increase in teenage enrolment, particularly for ages 15, 16 and 17. For age 18, the increase in enrolment is still positive, but less than one fifth of the magnitude of age 17.We find that the expansion did not have a significant effect on adult secondary school enrolment.

These results suggest that the introduction of schools had a non-trivial impact on enrolment of ages 15–17, and this effect was particularly pronounced for females.

5. Results

In this section we provide the main results of the paper. This is followed by robustness checks and an event study to test for pre-trends.

5.1 Main results

In Table 3 we report estimates of α^a, from Equation (3), for each maximum age of conception $a = \{15, 16, \dots, 19\}$. In the final column, we also include the results of a regression looking at all births conceived before age 15. Regressions include controls for primary school enrolment (year 8 corresponding to cohort age 14), contemporaneous nursery and preschool availability, the number of males and females in each cohort/municipality combination, and time and municipality fixed effects. Panel A reports the preferred specification, using the number of schools available at age 15. Panels B and C report alternative measures of the secondary school expansion, using the number of year-one secondary classrooms and the number of secondary teachers.

The results in Table 3 suggest that the secondary school expansion had a significant negative effect on teenage childbearing. The estimate in Column 1 suggest that the addition of a secondary school decreased the total number of births conceived by age 19 by 4.44 births; a 4.56% decrease relative to the mean number of births.[12] Similar results are found when the number of year-one classrooms or the number of secondary teachers is used.

[11] Data before 2007 does not permit us to examine enrolment by both sex and age.
[12] A given cohort in a municipality has, on average, 97.57 births between the ages of 15 and 19.

For all specifications, the estimated effect of a school on the number of births conceived before age 15 (reported in the final column) is statistically indistinguishable from 0, economically small and positive. This suggests that the school expansion did not have a meaningful impact on the childbearing behaviour of young women before the age they would have been expected to enter into secondary school (although it should be noted that the number of births resulting from conceptions before age 15 is a small proportion of the total number of teenage births).

These estimates suggest that the school expansion had a sizeable impact on teenage fertility in Brazil. To think about the economic impact, we can consider how much of the 19% decrease in births conceived between ages 15 and 18 over time can be explained by the school expansion. The magnitudes presented in Table 3 suggest that the school expansion can explain 25% of the total decrease in fertility observed in Figure 1.[13]

We also estimate Equation (1) by the parent-reported race of the child. The corresponding estimates for α^a can be found in Table A2 in Appendix A. We find some heterogeneity by race. Estimates for childbirths identified as Asian, black and white are significant and negative, similar to those reported in Table 3. However, the results for one of the largest groups, mixed race (*Pardo*), and the smallest group, indigenous, are less clear. These estimates are positive and largely insignificant. We interpret these results cautiously, as information on race is self-declared (by the mother/parents). Previous literature has noted that the reporting of Pardo, as opposed to black or white, varies geographically and inter-temporally, and is correlated with parental education (Marteleto, 2012). This suggests that our racial categories may be a function of parental education.

5.2 Robustness checks

Here we perform a number of robustness checks directed at the concerns that we outlined in Section 4.1. The results of these checks are presented in Table 4.

The first threat to identification concerns the possibility that unobservable cohort characteristics led to the municipal expansion of secondary schools, and this is not sufficiently controlled for by the inclusion of primary school enrolment in estimating Equation (1). To test this

[13] Between 1997 and 2009, the total number of annual births decreased from 394,192 to 318,666, or a reduction of 75,562 births. The number of schools increased by 7,280, with each leading to an average reduction in births conceived by between age 15 and 18 of 2.625. The proportion of the total decrease in births that we can attribute to schools is therefore 19,110/75,560 = 0.253.

we regress the number of classrooms on lagged birth outcomes (as well as the controls included in Equation 1). Columns A1 and A2 include all births for ages 14 to 19 in first and second lags, while columns A3 to A6 consider lagged births for different age groups. Only the estimated coefficient in A5 is statistically significant, and all coefficients are small in magnitude. Taken literally, the largest coefficient (in column A6) implies that one additional birth for ages 13–14 (almost a 25% increase over the mean) will lead to an increase of 0.01 schools. We interpret these results as evidence that school introductions were not driven by patterns in teenage childbearing.

The second threat to identification concerns the existence of unobservable influences on childbearing that are correlated with school introductions. To test this we estimate Equation (1) for older age groups. It should be emphasised that there are no explicit age restrictions to entering secondary education in Brazil. Therefore, there is no discontinuous age cut-off at which we can say the ‘treatment’ will apply. However, we expect that secondary school access is much less likely to impact on older individuals than younger individuals. Similar estimates for births at ages 30 and older will be suggestive of an unobserved municipal influence on birth outcomes. We estimate Equation (1) for birth outcomes for ages 35–40 (Panel B) and 5-year age groups from 20 to 44 (Panel C). The estimates for all groups over 25 years of age are small in magnitude and, with the exception of 35–39, statistically insignificant. We interpret this as evidence in support of the assumption that school introductions can be treated as exogenous.

As an additional robustness check, we also add municipality-specific time trends. The results are reported in Table A1 in Appendix A. Estimates for school introductions are qualitatively similar (with the exception of older birth-ages when using classrooms as the measure of school availability) to the results in Table 3, and several remain statistically significant. However, it should be pointed out that estimates are reduced in magnitude by approximately one quarter.

5.3 Event study

As a final check on the validity of our estimation strategy, we explore the dynamics of teenage childbearing relative to the introduction of secondary schools. If the decrease in childbearing is—conditional on observable factors—the result of school introductions, then we will not expect to see a systematic trend in childbearing prior to the introduction of a school. We conduct an event study to check for the existence of pre-trends in childbearing.

The complexity in conducting an event study in our framework is that some municipalities experience multiple 'events'—by having schools introduced at different points in time—and an 'event' may vary in magnitude, as some municipalities receive multiple schools at once. To account for this we follow the methodology outlined in Sandler and Sandler (2014). For each age group, a, we run the following regression:

$$b_{it}^{a} = \sum_{d=-D}^{D} \lambda_{d}^{a} 1[t - e_i = d] * \Delta S_{it} + \boldsymbol{X_{it}^{a\prime} \Omega^{a}} + \nu_t^a + \xi_i^a + \mu_{it}^a. \qquad (3)$$

The outcome, b_{it}^{a}, captures the number of births for age group a in municipality i in year t. $1[t - e_i = d]$, a binary indicator equal to 1 when the event (a school introduction), e_i, is d periods away, is interacted with ΔS_{it}, capturing the change in schools. λ_d^a reflects the average births for age group a d periods away from a unit change in school availability. The vector X_{it}^{a} includes the same municipality-time varying controls included in Equation (1). The regression also includes time and municipality fixed effects, captured by ν_t^a and ξ_i^a. We present the results for each λ_d^a, normalising $\lambda_{-1}^{a} = 0$, in Figure 5.

In Figure 5 we depict the dynamics of childbearing for each of ages 15 to 17. Bars show 95% confidence intervals and the vertical line is set at the reference period (-1). Two important observations can be made from Figure 5. First, for all age groups, there are no pre-trends in childbearing. This provides strong support for the assumptions that there are no unobserved variables that confound our results. Second, for all age groups, a clear downward trend in childbearing is observed following the change in the availability of secondary schools. We interpret these figures as evidence that the estimates we report in Table 3 reflect a causal relationship.

In Appendix A, Figure A2, we include corresponding figures for older age groups, whom we expect to be less impacted in terms of school attendance by the expansion of secondary schools. We report estimates for age of conception at 25 years and 35 years.[14] These figures do not suggest a corresponding pre- or post- trend for childbearing at older ages.

[14] Similar results are found for other older ages.

6. Mechanisms through which school introductions impact teenage births

The introduction of secondary schools across Brazilian municipalities appears to have had a significant negative impact on teenage childbearing. In this section we discuss the possible structural mechanisms that underlie this relationship, and provide new evidence on their relative importance.

The first possible mechanism noted in the literature is the *incarceration effect*. Incarceration has previously been used to describe both an effect arising from mandatory increases to time spent in school (Black, Devereux and Salvanes, 2008) and changes to the length of a school day in non-mandatory schooling (Berthelon and Kruger, 2011). It would seem that there exists an important difference between these types of incarceration that is worth highlighting. Incarceration, as described by Fort, Schneeweis and Winter-Ebmer (2016), broadly reflects an incompatibility between school and childbearing. This can manifest itself in two ways. The incarceration effect can arise purely mechanically by reducing the number of hours young women (and men) have to engage in activities that may lead to pregnancy. This effect operates independent of individual preferences. The incarceration effect may also arise because childbearing makes attending school more costly, and some women choose to forgo children because they allocate more time to school. This effect operates explicitly through preferences. While both of these channels may be in operation in the case of mandatory schooling laws, only the second will operate in the case of the school expansion examined here.

The incarceration effect can be contrasted with a human capital effect. Black, Devereux and Salvanes (2008) refer to two types of human capital effects, the *current human capital effect* and the *future human capital effect*. The current effect reflects the direct impact of human capital acquired in school on childbearing preferences. For example, schools may provide information about the cost of raising a child, deterring students from parenthood (at least in the short run). The future effect reflects the indirect impact of human capital acquired in school on childbearing preferences through increased earnings (or potential earnings) in the labour market. The prospect of higher future labour force earnings makes childbearing relatively more costly, leading some young women to substitute away from parenthood to ensure they are free to participate in the labour force.[15] Both human capital channels have a behavioural implication not present with the incarceration effect. For both effects, we expect a spill-over of the negative effect on childbearing into early adulthood (after the completion of schooling).

A final mechanism through which school may impact fertility works through the marriage market (Fort, Schneeweis and Winter-Ebmer, 2016). If education increases one's value in the marriage market, then greater access to education may lead to earlier marriages, which is expected to result in earlier childbearing. Likewise, secondary schools may promote assortative matching, leading to earlier matches and earlier childbearing. The behavioural implications of this channel are unique in that we expect to observe an increase in early childbearing for teenagers and early adults.

These different effects have overlapping and conflicting implications for observable behaviour, and disentangling these different effects has been a challenge in the literature. Using Norwegian and US mandatory schooling laws, Black, Devereux and Salvanes (2008) find little evidence of an incarceration effect and attribute their results instead to human capital effects. Looking at completed fertility, Fort, Schneeweis and Winter-Ebmer (2016) find a negative causal impact of education for England but a positive causal impact of education in continental Europe. They attribute these differences to differences in the strength of the 'mechanical' incarceration effect and differences in the importance of the marriage-market versus the labour market.

As with previous studies, completely disentangling the different channels is not possible in our setting. However, we can provide some new evidence regarding the relative strength of these effects in the Brazilian context. Our empirical strategy rules out the 'mechanical' component of the incarceration effect which we explained above; observed reductions in childbearing during the teen years are attributable to either women choosing to forgo children in favour of attending school, to the future cost children impose on labour market participation (future human capital effect), or to the change in preferences for children (current human capital effect). On the other hand, changes to childbearing in early adulthood can be attributed to either the human capital effects (reduction)

[15] It is noted here that an increase in human capital would lead to both a human capital and substitution effect. As discussed in Black, Devereux and Salvanes (2008), the general view in the literature on women's labour market participation is that the substitution effect dominates. We believe that this is particularly likely to be true in the short time span we examine here (young women will have limited ability to capitalise on their increased human capital to lead to a significant income effect.)

or marriage markets (increase). While our data do not allow us to look at completed levels of fertility (as the variation in schools we are exploiting was relatively recent), we can look at fertility in early adulthood. Here we examine the outcomes for all births conceived up to age 23.

To examine the mechanisms, we want to estimate how the availability of schools at 15 affects childbearing at different ages. The estimated values for α^a in Equation (1) reflect the effect of schools at age 15 on *total cumulative* births for a cohort. To see the effect, a unit increase in the number of schools has on births at different ages we difference Equation (1) for any two ages:

$$B_{ic}^{19} - B_{ic}^{18} = \sum_{a=15}^{19} b_{ic}^{a} - \sum_{a=15}^{18} b_{ic}^{a} = b_{ic}^{19},$$

which can be written as

$$b_{ic}^{19} = (\alpha^{19} - \alpha^{18})S_{ic} + \boldsymbol{X_{ic}^{19'}}(\boldsymbol{\Gamma^{19}} - \boldsymbol{\Gamma^{18}}) + \left(\boldsymbol{X_{ic}^{19}} - \boldsymbol{X_{ic}^{18}}\right)'\boldsymbol{\Gamma^{18}} + (\delta_c^{19} - \delta_c^{18}) +$$

$$(\eta_i^{19} - \eta_i^{18}) + (\epsilon_{ic}^{19} - \epsilon_{ic}^{18}) \qquad (4)$$

The coefficient on S_{ic} in the above equation reflects the incremental effect of secondary schools on births at each age. To focus on the dynamics of the estimated effect, we present estimated coefficients, $(\alpha^a - \alpha^{a-1})$, for $a = \{15, \ldots, 23\}$, graphically in Figures 6.

Figure 6 provides a visual description of the effect of an additional school as the cohort ages. We see a u-shaped pattern in childbearing by age following the addition of a secondary school. This begins with a modest reduction at 15 years of age, and steadily increase until age 19. After age 19, we see a slow decline in the magnitude of the effect. However, the effect remains negative through to age 23. This suggests that there is, in terms of aggregate births, a human capital effect in the medium term.

By contrasting estimates by urban and rural schools we can shed some light on the relative importance of the human capital effect in terms of how it works through labour market attachment, as opposed to directly changing preferences for children. To do this, we note that the earned income return to secondary school is high in urban areas relative to rural farming areas. Therefore, we would expect that the future human capital effect, which works through this channel, will be stronger for urban than for rural schools. However, the current human capital effect, which operates through changing preferences for childbearing, should be the same in rural and urban areas. We re-estimate Equation (4), including separate schooling variables, S_{ic}^{urban} for urban schools and S_{ic}^{rural} for rural schools.[16] The corresponding coefficients are plotted in Figure 7.

If the future human capital effect is not important, then we expect urban and rural schools will have similar dynamics in terms of birth outcomes. The increasing magnitude of the effect is very similar for both urban and rural schools until age 19. However, after age 19, we see a decline in the magnitude of the effect for rural schools, with a positive (although statistically insignificant) point estimate for the oldest age group. This suggests that any negative human capital effect fades, and is possibly reversed, shortly after the teenage years and after completing secondary schooling age. Urban schools show very modest decreases after age 19, remaining stable at an average effect between -1.0 and -1.5. The downward trend after 19 in Figure 7 appears to be driven by rural schools, suggesting that the future human capital effect plays an important role in the relationship between education and childbearing in urban centres.

7. Conclusion

In this paper we estimate the effect of improved access to secondary education on the childbearing of young women. Combing information from 14 waves of the Brazilian school census and Brazilian vital statistics, we create a novel and rich dataset. Brazil's expansion of secondary schools over time and across municipalities provides quasi-experimental variation from one of the largest expansions of secondary schools on record.

We find that an additional secondary school leads to an average decrease of 4.44 births, or 4.56% relative to the mean, by age 19. Further, we find a persistent negative impact on municipal childbearing for women up to age 23 (the oldest age for we are able to perform our analyses), and this effect is particularly strong for schools placed in urban locations. This result is consistent with an increase in human capital increasing the opportunity cost associated with having children and some women forgoing childbearing as young adults.

In addition to building on the literature examining the causal relationship between schooling and fertility, the current study also informs an active debate regarding the direction of causation

[16] We use the definition of urban-rural by IBGE, the Brazilian census bureau, at the level of the municipalities.

between economic outcomes and teenage childbearing. It is not clear if the negative correlation between long-term economic outcomes is the result of teenage child bearing itself, or if teenage child bearing is a response to low later-life outcomes. The evidence on the effect of teenage childbearing on later life outcomes is mixed. Klepinger, Lundberg and Plotnick (1999) find that by reducing time spent in formal education and work experience, early motherhood reduces later-year earnings. Recent evidence by Hotz, McElroy and Sanders (2005), Fletcher and Wolfe (2009) and Ashcraft, Fernandez-Val and Lang (2013) finds negative but modest effects of teenage motherhood on education and labour market outcomes. Lang and Weinstein (2015) find the negative effect of teen motherhood on education to be particularly large for teens from disadvantaged backgrounds. In contrast, Kearney and Levine (2012) argue that causation runs in the opposite direction. They argue that the substantial variation in US teen pregnancy rates reflects geographic differences in economic opportunity—both perceived and real. In the context of the current study, by decreasing the geographic distance necessary to attend secondary school, the Brazilian expansion improved opportunities for school attendance. To the extent that secondary education increases human capital and labour market earnings, improved access to a secondary school represents a real improvement in economic opportunity. In this sense, our estimated negative relationship between schools and teenage childbearing is consistent with Kearney and Levine's argument.

References

Ashcraft, A., I. Fernandez-Val and K. Lang, 2013. 'The Consequences of Teenage Childbearing: Consistent estimates when abortion makes miscarriage non-random', *Economic Journal*, 123: 875–905.

Berthelon, M.E. and D.I. Kruger. 2011. 'Risky Behavior among Youth: Incapacitation effects of school on adolescent motherhood and crime in Chile', *Journal of Public Economics*, 95, 41–53.

Black, S. E., P. J. Devereux and K. G. Salvanes. 2008. 'Staying in the Classroom and out of the Maternity Ward? The effect of compulsory schooling laws on teenage births', *Economic Journal,* 118: 1025–1054.

Brunello, G., M. Fort and G. Weber, 2009. 'Changes in Compulsory Schooling, Education and the Distribution of Wages in Europe', *Economic Journal*, 119 (536): 516–539.

Currie, J. and E. Moretti. 2003. 'Mother's Education and the Intergenerational Transmission of Human Capital: Evidence from college openings', *Quarterly Journal of Economics*, 118: 1495–1532.

De Felizio, F. 2009. 'O Ensino Fundamental – Desafios desde a Alfabetização até a Transição para o Ensino Médio', in: *Educação Básica no Brasil: Construindo o País do Futuro*, ed: F. Veloso et. al. Elsevier, Rio de Janeiro.

De Moura Castro, C. 2009. 'Desventuras do Ensino Médio e seus Desencontros com o Profissionalizante', in: *Educação Básica no Brasil: Construindo o País do Futuro*, ed: F. Veloso et. al. Elsevier, Rio de Janeiro.

Di Gropello, E. 2006. *Meeting the Challenges of Secondary Education in Latin America and East Asia: Improving Efficiency and Resource Mobilization*. Washington D.C. The International Bank for Reconstruction and Development.

Duflo, E., P. Dupas and M. Kremer. 2010. 'Education and Fertility: Experimental Evidence from Kenya', *MIT Working Paper.*

Fletcher, J.M. and B.L. Wolfe, 2009. 'Education and Labour Market Consequences of Teenage Childbearing: Evidence Using the Timing of Pregnancy Outcomes and Municipality Fixed Effects', *Journal of Human Resources*, 44: 303–325.

Foureaux Koppensteiner, M. 2014. 'Automatic Grade Promotion and Student Performance: Evidence from Brazil', *Journal of Development Economics*, 107: 277–290.

Fort, M., N. Schneeweis and R. Winter-Ebmer, 2016. 'Is Education Always Reducing Fertility? Evidence from compulsory schooling reforms', *Economic Journal*, 126: 1823–1855.

Guimarães de Castro, M.H. and S. Tiezzi. 2004. 'The Reform of Secondary Education', in: *The Challenges of Education in Brazil*. Ed.: C. Brook and S. Schwartzmann. Cambridge University Press, Cambridge.

Geruso, M. and H. Royer, 2014. 'The Impact of Education on Family Formation: Quasi-experimental evidence from the UK', Mimeo.

Hotz, J.V., S.W. McElroy and S.G. Sanders. 2005. 'Teenage Childbearing and its Lifecycle Consequences: Exploiting a natural experiment', *The Journal of Human Resources*, 40: 683–715.

IBGE. 2011. *Evolução da Divisão Territorial do Brasil 1872–2010*. Available at http://www.ibge.gov.br/home/geociencias/geografia/default_evolucao.shtm.

IBGE. 2012. *Estatícas do Sécolo XX.* Database Appendix. Rio de Janeiro.

INEP. 2003. *Sinopse Estatística da Educação Básica 1991–1995*. Ministry of Education, Brasília.

INEP. 2011. *Sinopse Estatística da Educação Básica 2010*. Ministry of Education, Brasília.

JBIC (Japan Bank for International Cooperation). 2005. 'Sector Study for Education in Brazil', *JBIC Sector Study Serie*s 2004–No.2.

Kearney, M. and P.B. Levine. 2012. 'Why is the Teen Birth Rate in the United States so High and Why Does it Matter?', *Journal of Economic Perspectives*, 26: 141–163.

Klepinger, D., S. Lundberg and R. Plotnick. 1999. 'How Does Adolescent Fertility Affect the Human Capital of Young Women?', *The Journal of Human Resources*, 34: 421–448.

Lang, K. and R. Weinstein. 2015. 'The Consequences of Teenage Childbearing before Roe v. Wade', *American Economic Journal: Applied Economics*, 7(4): 169–197.

LDB (Lei de Diretrizes e Bases da Educação Nacional). 1996. Senate of the Federal Republic of Brazil. http://www6.senado.gov.br/legislacao/ListaPublicacoes.action?id=102480.

Marchelli, P.S. 2010. 'Expansão e Qualidade de Educação Básica no Brasil', *Cadernos de Pesquisa* 40: 561–585.

Marteleto, L. 2012. 'Educational Inequality by Race in Brazil, 1982–2007: Structural Changes and Shifts in Racial Classification', *Demography* 49: 337-358.

MEC (Ministry of Education). 2002. Parâmetros Curriculares Nacionais – Ensino Médio. Brasília.

Monstad, K., C. Propper and K.G. Salvanes. 2008. 'Education and Fertility: Evidence from a natural experiment', *Scandinavian Journal of Economics,* 110: 827–852.

Moore, A. S., J. DeStephano, A. Terway and D. Balwanz. 2008. 'The Expansion of Secondary Education and the Need for Teachers: How big is the gap?', *Equip2 Working Paper*, USAid.

OECD. 2012. *Brazil Country Note. Education at a Glance: OECD Indicators 2012*. OECD, Paris.

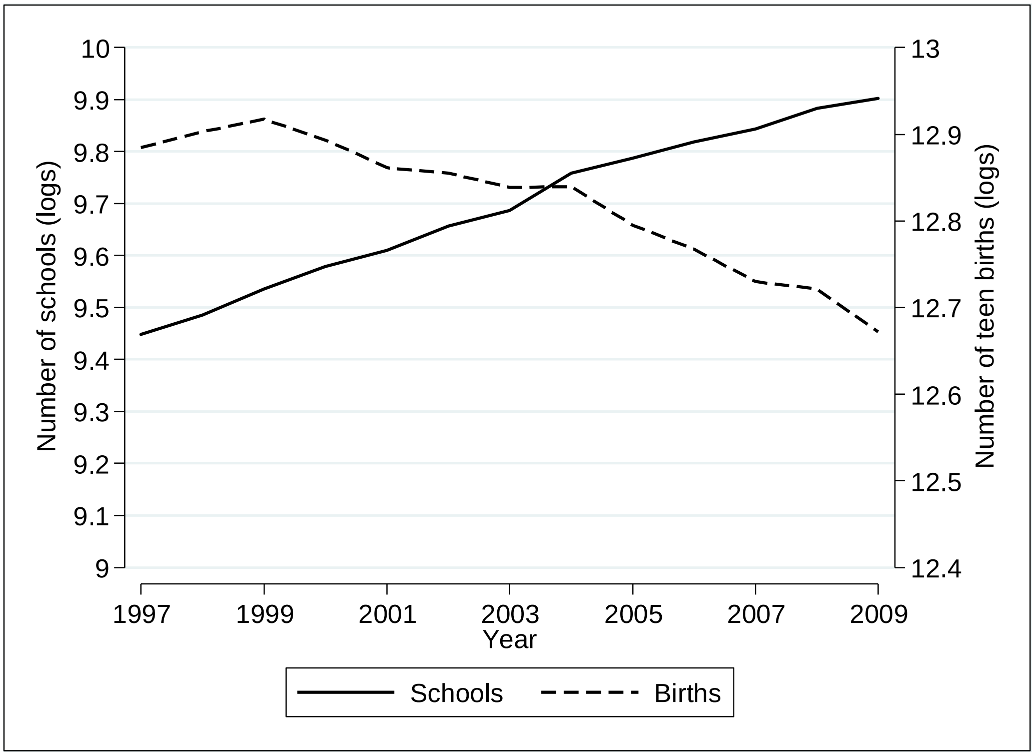
Rocha, R. and R. Soares. 2010. 'Evaluating the Impact of Municipality-Based Health Interventions: Evidence from Brazil's Family Health Program', *Health Economics*, 19: 126–158.

Sandler, D.H., and R. Sandler. 2014. 'Multiple Event Studies in Public Finance and Labor Economics: A Simulation Study with Applications', *Journal of Economic and Social Measurement*, 39: 31–57.

World Bank. 2005. *Expanding Opportunities and Building Competencies for Young People A New Agenda for Secondary Education.* Washington, D.C.

World Bank. 2012. *Teenage Pregnancy and Opportunities in Latin America and the Caribbean. On Teenage Fertility Decisions, Poverty and Economic Achievement.* Washington, D.C.

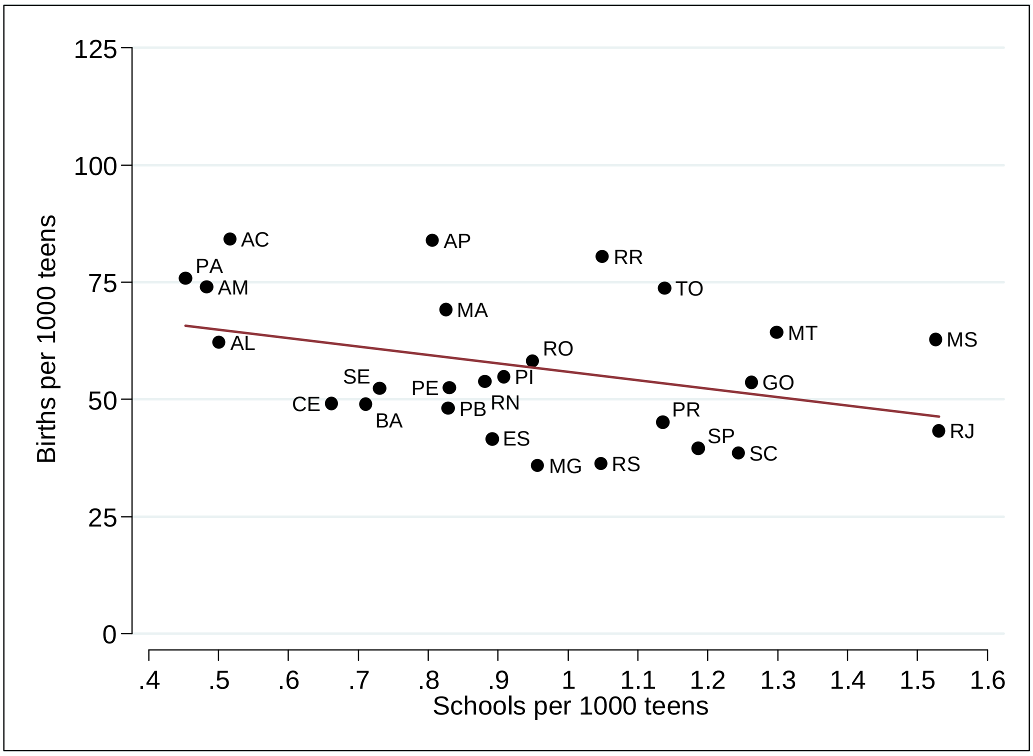
Figure 1: Numbers of secondary schools and teenage births in Brazil over time



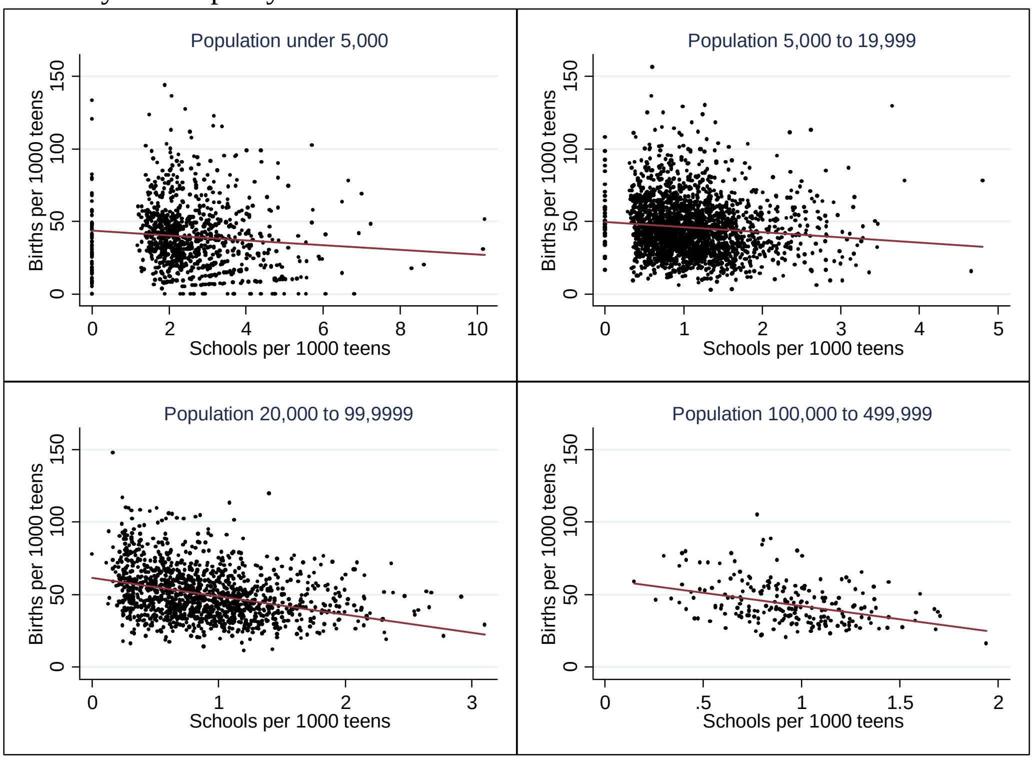
Notes: The figure depicts the annual log number of secondary schools and log births for the period 1997 to 2009. Teen births include all births conceived between the ages 13 to 18. Data for all Brazilian municipalities with populations less than 500,000 included.
Source: School data come from the 1997–2009 waves of the Brazilian School Census; births by age of conception from Brazilian Vital Statistics.

Figure 2: School density and teen birth rates

A. By state



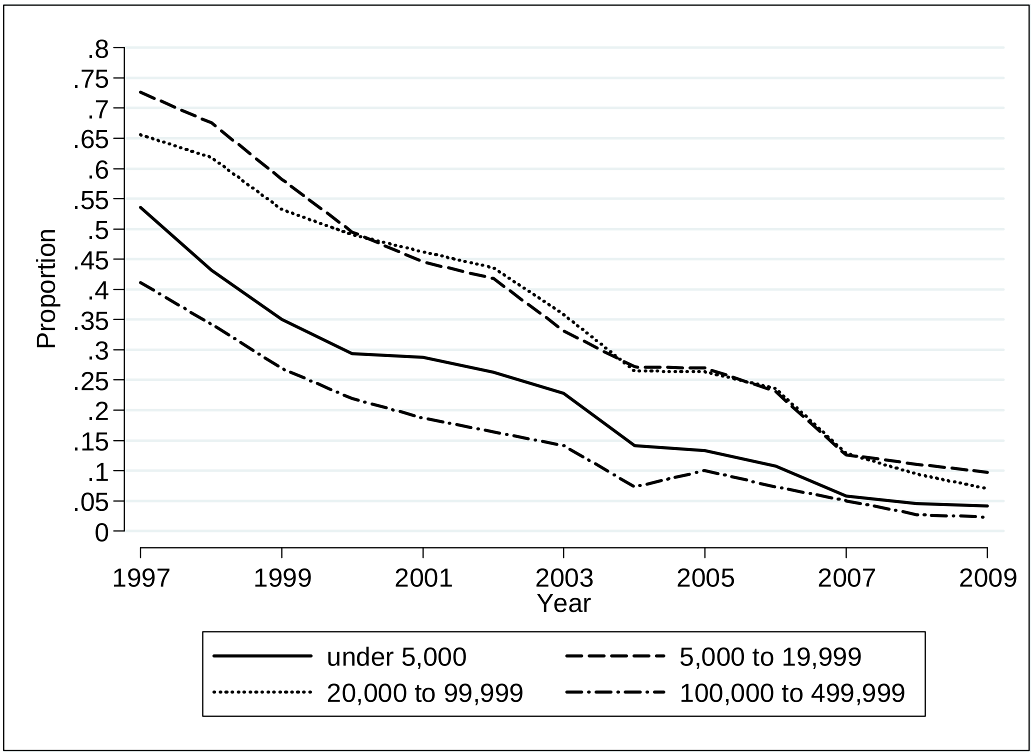
B. By municipality



Notes: Data for 2002 cross-section. Figures capture municipalities with population of less than 500,000. Linear fitted lines weighted by population size.

Source: School data come from the 2002 wave of the Brazilian School Census; official population estimates from the Brazilian Census Bureau; births by age from Brazilian Vital Statistics.

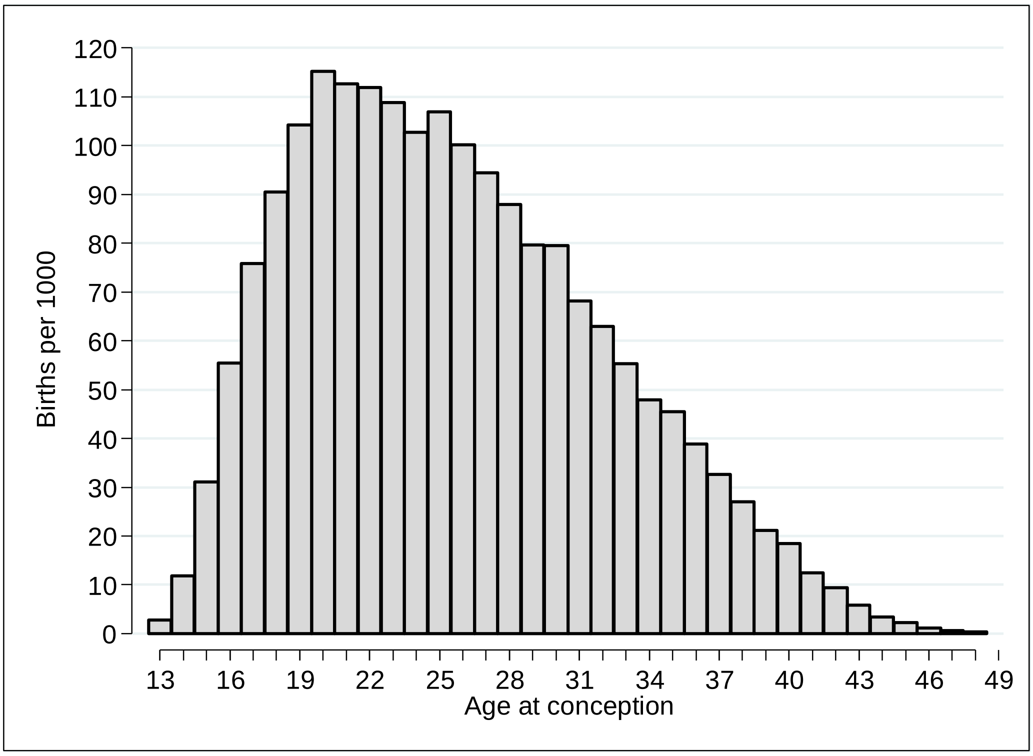
Figure 3: Proportion of municipalities with low secondary classroom density



Notes: Low classroom density is defined as less than one year-1 classroom per 100 youth aged 13 to 18.

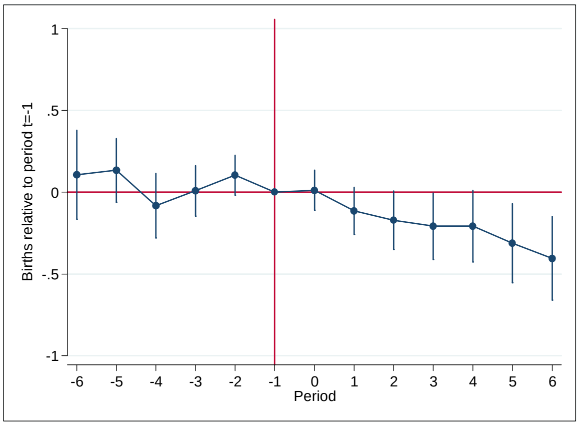
Source: School data come from the 1997–2009 waves of the Brazilian School Census; official population estimates from the Brazilian Census Bureau.

Figure 4: Births per 1000 population by age

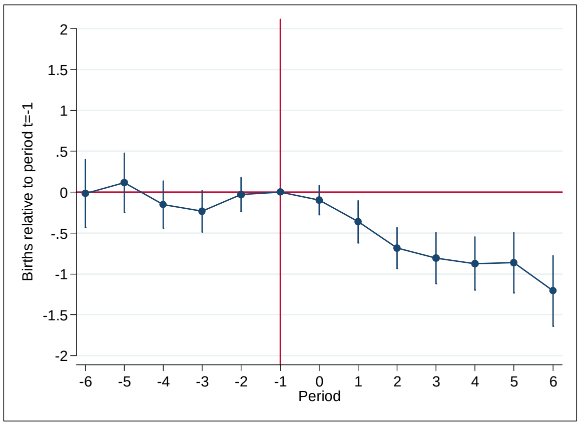


Source: Brazilian Vital Statistics.

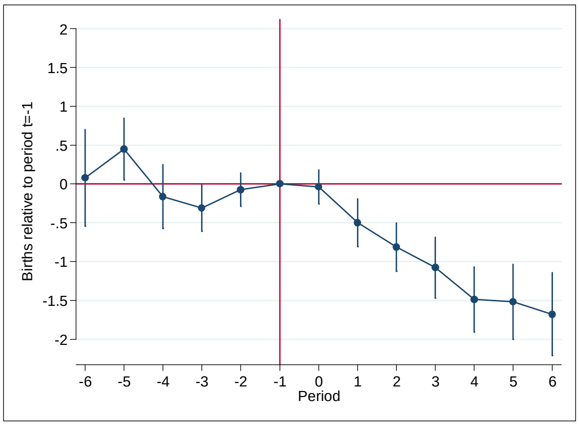
Figure 5: Event study of births before and after school introduction, by age at conception



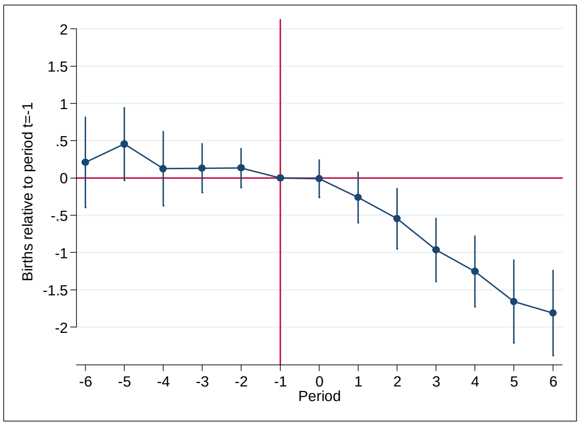
a) Births for conception age 15



b) Births for conception age 16



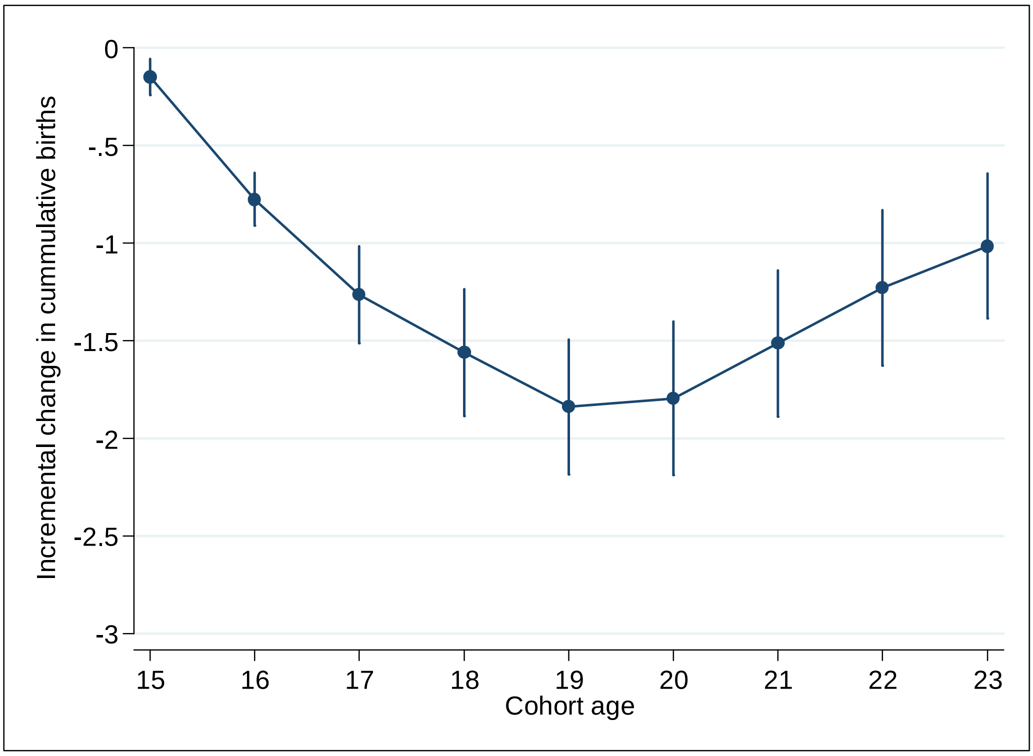
c) Births for conception age 17



d) Births for conception age 18

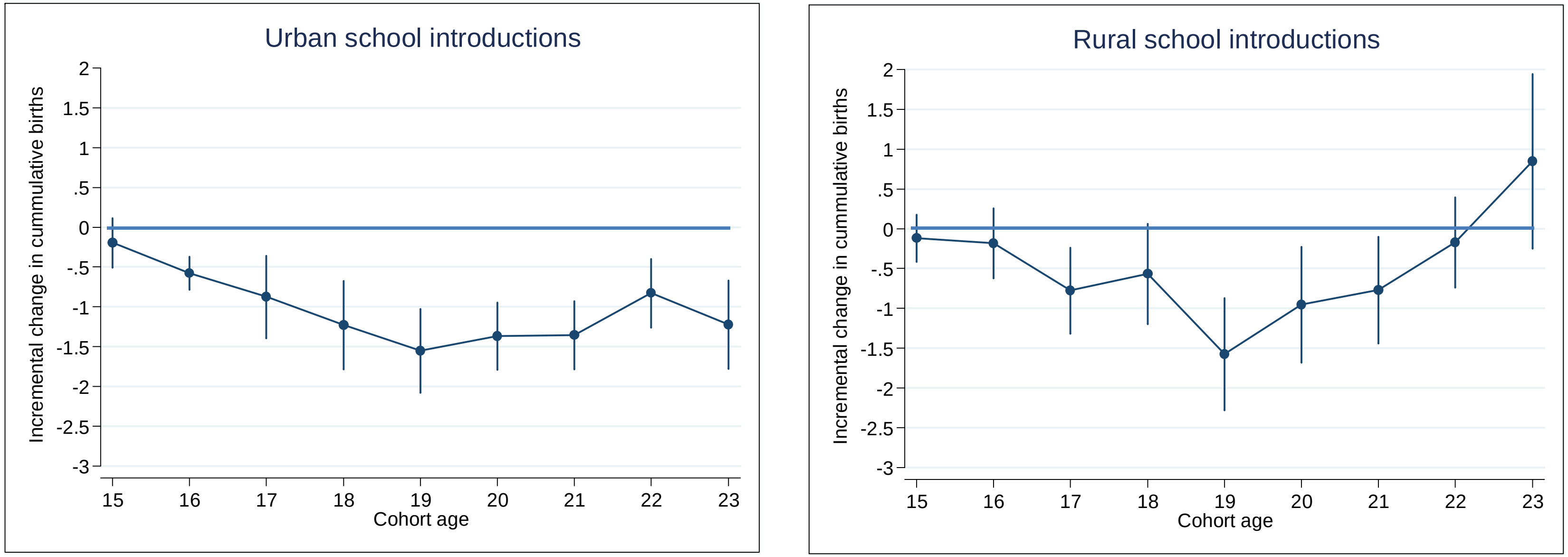
Notes: Bars indicate 95% confidence intervals. Points reflect coefficient estimates from a regression capturing the years from school introduction (relative to births at period -1), detailed in main text. All estimates condition on age-specific male and female populations, primary enrolment, nursery and preschool classrooms, municipality expenditures, municipality and cohort fixed effects.

Figure 6: Secondary schools and incremental cohort births, medium run



Notes: Bars indicate robust 95% confidence intervals. Points reflect the estimated effect of schools introduced at age 15 on births conceived at each age. Estimates condition age-specific male and female populations, primary enrolment, nursery and preschool classrooms, municipality expenditures, municipality and cohort fixed effects.

Figure 7: Secondary schools and incremental cohort births, urban versus rural school introductions



Notes: Bars indicate robust 95% confidence intervals. Points reflect the estimated effect of schools introduced at age 15 on births conceived at each age. Regressions include both urban school and rural school variables and condition on age-specific male and female populations, primary enrolment, nursery and preschool classrooms, municipality expenditures, municipality and cohort fixed effects.

Table 1: Descriptive statistics

	Mean	*Min*	*Max*
Population			
Total	25,958	690	551,857
Teenage	3,290	70	72,727
Cohort	512	10	12,357
	Mean	*SDB*	*SDW*
Births (aggregate cohort)	101.06	(183.89)	[17.66]
Schools	3.31	(5.60)	[1.49]
Rural	0.23	(0.66)	[0.58]
Urban	3.08	(5.36)	[1.27]
Public	2.47	(3.63)	[1.24]
Private	0.84	(2.28)	[0.55]
Schooling controls			
Primary enrolment, final year	416.18	(806.65)	[229.12]
Nursery rooms	9.35	(22.49)	[14.06]
Preschool rooms	37.48	(66.44)	[18.16]
Municipality expenditures			
Bolsa Familia p.c.†	5.92	(3.59)	[2.21]
Bolsa Familia p.r.†	71.45	(10.39)	[12.06]
Welfare p.c.	50.89	(50.41)	[87.83]
Education p.c.	244.36	(119.58)	[196.43]
Health p.c.	180.71	(93.11)	[129.74]
Transportation p.c.	48.86	(64.00)	[57.99]
Housing p.c.	79.39	(69.01)	[93.20]
Total p.c.	559.29	(341.02)	[525.23]

Notes: Between standard deviation (SDB) reported in parenthesis, within standard deviation (SDW) reported in brackets. Aggregate cohort births reflect total cumulative birth for cohorts at age 19. Municipality expenditure data is reported in nominal Brazilian Reais (R$), per-capita (p.c.) and per-recipient (p.r.).
Source: School data comes from Brazilian School Census 1999–2009; official population estimates from the Brazilian Census Bureau; municipal expenditures come from the Ministry of Finance. See Appendix C for details.
†Reported averages and standard deviation based on years 2004 and later. Bolsa Familia was not implemented prior to 2004.

Table 2: Secondary schools and enrolment

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Total	*Females*	*Males*	*Under 15*	*Age 15*	*Age 16*	*Age 17*	*Age 18*	*Adults*
Schools at 15	44.251	27.301	16.949	1.686	23.459	29.080	31.561	5.661	-13.471
	(15.491)***	(7.921)***	(7.638)**	(0.596)***	(1.655)***	(2.911)***	(2.910)***	(2.209)***	(7.422)
R2†	0.179	0.164	0.183	0.036	0.264	0.240	0.263	0.171	0.245
Observations	43,650	43,650	43,650	43,650	43,650	43,650	43,650	43,650	43,650
Municipalities	4850	4850	4850	4850	4850	4850	4850	4850	4850

Notes: Robust standard errors reported in parenthesis. ***, ** and * denote statistical significance at 1%, 5% and 10%. Regressions include male and female population sizes corresponding to each age group, year fixed effects and municipality fixed effects.

†R-squared for within variation reported.

Table 3: Secondary schools and aggregate cohort births

	Age of conception					
A	15–19	15–18	15–17	15–16	15	<15
Schools at 15	-4.444	-2.625	-1.627	-0.860	-0.190	0.045
	(0.603)***	(0.408)***	(0.248)***	(0.128)***	(0.049)***	(0.043)
R2†	0.315	0.286	0.219	0.119	0.024	0.017
B						
Classrooms at 15	-0.662	-0.349	-0.212	-0.089	-0.011	0.024
	(0.199)***	(0.141)**	(0.085)**	(0.044)**	(0.015)	(0.016)
R2†	0.301	0.274	0.207	0.107	0.021	0.016
C						
Teachers at 15	-0.329	-0.192	-0.099	-0.051	-0.010	0.000
	(0.026)***	(0.017)***	(0.012)***	(0.008)***	(0.003)***	(0.002)
R2†	0.356	0.315	0.229	0.123	0.024	0.017
Mean births‡	97.569	69.652	44.402	23.719	8.682	4.523
Observations	43,650	43,650	43,650	43,650	43,650	38,800
Municipalities	4,850	4,850	4,850	4,850	4,850	4,850

Notes: Robust standard errors reported in parenthesis. ***, **, and * denote statistical significance at 1%, 5% and 10%. Estimates condition on cohort-specific primary school enrolment, nursery and preschool classrooms, cohort-specific male and female populations, municipality expenditures, municipality and cohort fixed effects.

†R-squared for within variation reported.

‡Total number of births averaged by cohort and municipality for each age group.

Table 4: Robustness: Child birth as a predictor of schools & older age groups

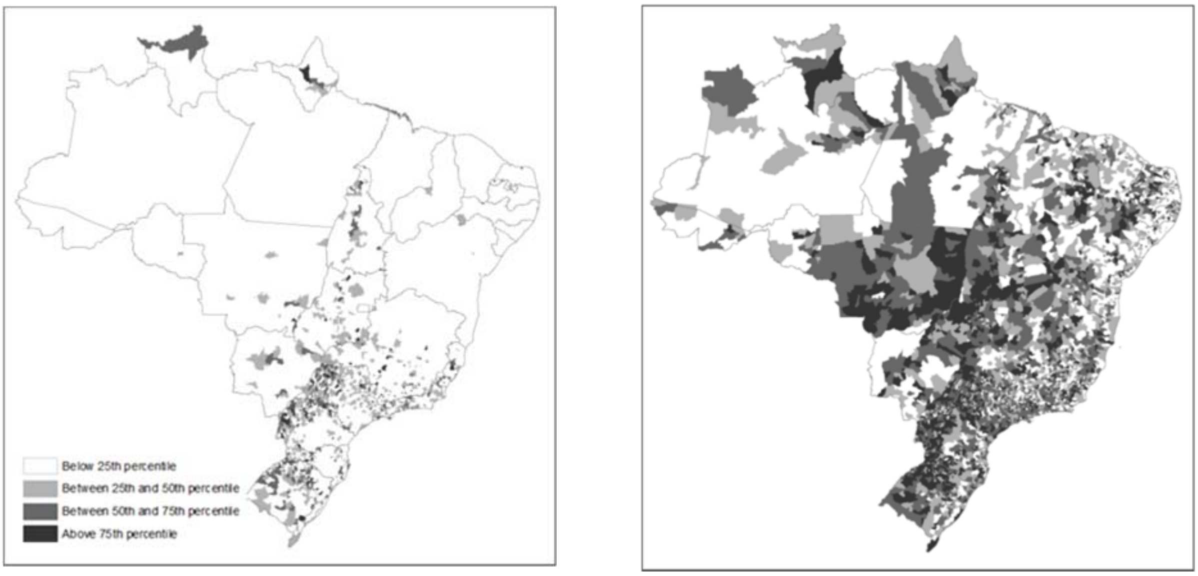
A						
Outcome: Schools	(A1)	(A2)	(A3)	(A4)	(A5)	(A6)
	Ages 14–19 (lagged)	Ages 14–19 (2 x lagged)	Ages 14–18 (lagged)	Ages 14–17 (lagged)	Ages 14–16 (lagged)	Ages 13–14 (lagged)
Cohort births	-0.003	0.000	-0.003	-0.004	-0.009	0.010
	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)***	(0.007)
R2†	0.219	0.185	0.218	0.218	0.220	0.217
Observations	38,800	33,950	38,800	38,800	38,800	38,800
Municipalities	4,850	4,850	4,850	4,850	4,850	4,850
B						
Outcome: Total births	(B1)	(B2)	(B3)	(B4)	(B5)	(B6)
Age groups	35	36	37	38	39	40
Schools (lagged 4 years)	0.019	-0.008	0.032	-0.054	-0.037	0.039
	(0.067)	(0.056)	(0.047)	(0.037)	(0.032)	(0.031)
R2†	0.113	0.068	0.060	0.047	0.037	0.030
Observations	38,800	38,800	38,800	38,800	38,800	38,800
Municipalities	4,850	4,850	4,850	4,850	4,850	4,850
C						
Outcome: Total births	(C1)	(C2)	(C3)	(C4)	(C5)	(C6)
Age groups	15–19	20–24	25–29	30–34	35–39	40–44
Schools (lagged 4 years)	-4.444	-4.955	0.636	0.305	0.328	0.025
	(0.603)***	(0.840)***	(0.626)	(0.286)	(0.135)**	(0.051)
Mean births‡	97.6	140.7	103.9	61.1	30.0	8.3
R2†	0.113	0.068	0.060	0.047	0.037	0.030
Observations	38,800	38,800	38,800	38,800	38,800	38,800
Municipalities	4,850	4,850	4,850	4,850	4,850	4,850

Notes: Robust standard errors reported in parenthesis. ***, **, and * denote statistical significance at 1%, 5% and 10%. Estimates condition on cohort-specific primary school enrolment, nursery and preschool classrooms, cohort-specific male and female populations, municipality expenditures, municipality and cohort fixed effects.

Appendix A: Supplementary analysis

Figure A1a: Classroom density by percentile

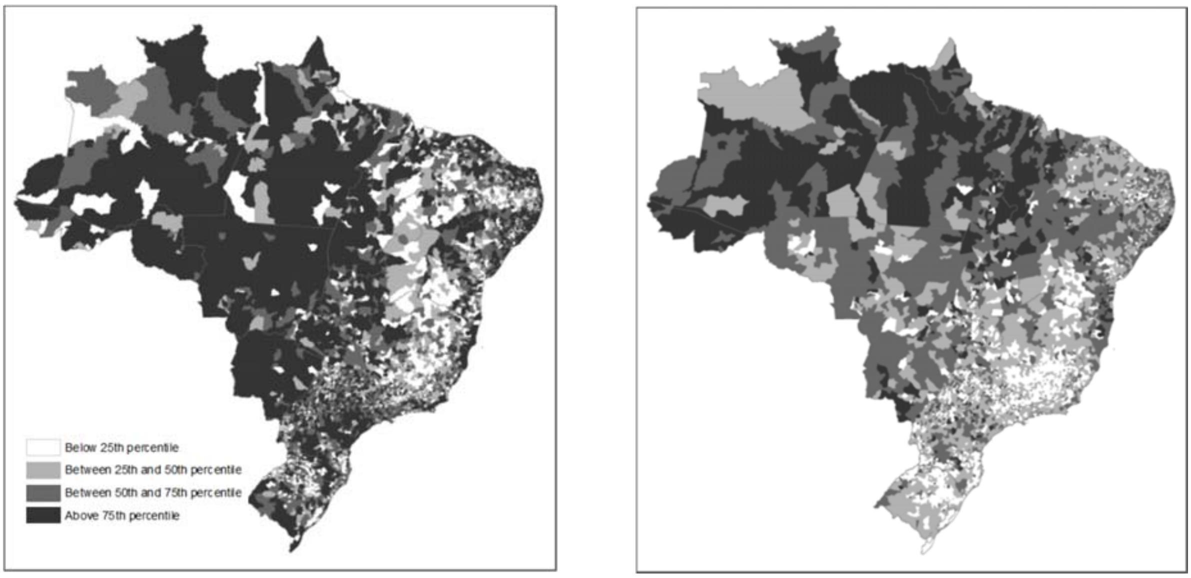
1997 2009



Source: Brazilian School Census 1996 and 2009. Percentiles held constant at 1996 cut-offs.

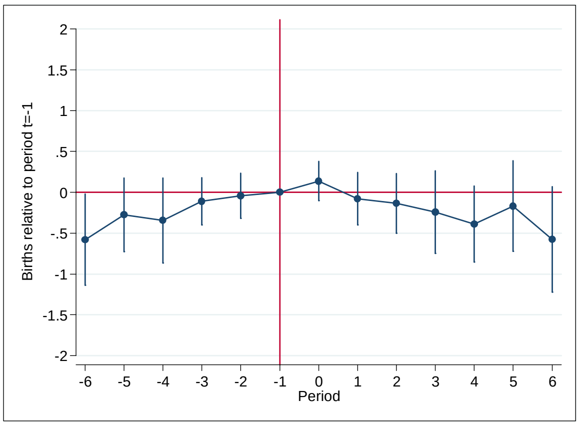
Figure A1b: Births rates by percentile

1997 2009

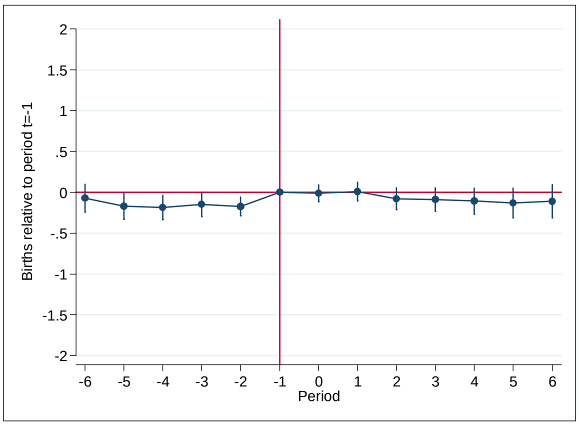


Source: Brazilian Vital Statistics. Percentiles held constant at 1996 cut-offs.

Figure A2: Event study of births before and after school introduction by age



b) Births at age 25



b) Births at age 35

Notes: Bars indicate 95% confidence intervals. Points reflect coefficient estimates from a regression capturing the years from school introduction (relative to births at period -1), detailed in main text. All estimates condition on age-specific male and female populations, primary enrolment (one-year lag), nursery and preschool classrooms, year fixed effects and municipality fixed effects.

Table A1: Secondary schools and aggregate cohort births, municipal time trends included

	Age of conception					
	15–19	15–18	15–17	15–16	15	<15
Schools at 15	-1.050	-0.407	-0.280	-0.299	-0.139	-0.102
	(0.396)***	(0.270)	(0.185)	(0.132)**	(0.060)**	(0.064)
R2†	0.995	0.994	0.992	0.988	0.972	0.950
Classrooms at 15	0.100	0.134	0.077	-0.011	-0.029	-0.006
	(0.105)	(0.074)**	(0.043)**	(0.027)	(0.019)	(0.015)
R2†	0.995	0.994	0.992	0.988	0.972	0.950
Teachers at 15	-0.111	-0.058	-0.019	-0.010	-0.001	-0.003
	(0.019)***	(0.013)***	(0.009)***	(0.006)	(0.004)	(0.003)
R2†	0.995	0.994	0.992	0.988	0.972	0.950
Observations	43,650	43,650	43,650	43,650	43,650	38,800
Municipalities	4,850	4,850	4,850	4,850	4,850	4,850

Notes: Robust standard errors reported in parenthesis. ***, **, and * denote statistical significance at 1%, 5% and 10%. Regressions include cohort primary enrolment, nursery and preschool classrooms, cohort-specific male and female populations, municipality and cohort fixed effects.

†R-squared for within variation reported.

Table A2: Secondary schools and aggregate cohort births, by reported race of child

	Age of conception					
	15–19	15–18	15–17	15–16	15	<15
	Race: White					
Schools at 15	-2.861	-1.748	-0.981	-0.491	-0.098	-0.049
	(0.540)***	(0.375)***	(0.233)***	(0.123)***	(0.046)**	(0.025)**
R2†	0.232	0.207	0.177	0.125	0.046	0.018
Mean births‡	40.227	28.090	17.464	9.066	3.217	1.527
Coefficient/mean	-7.11%	-6.22%	-5.62%	-5.42%	-3.06%	-3.19%
	Race: Black					
Schools at 15	-0.234	-0.213	-0.156	-0.081	-0.028	-0.011
	(0.082)***	(0.072)***	(0.054)***	(0.030)***	(0.010)***	(0.004)**
R2†	0.111	0.098	0.082	0.056	0.022	0.009
Mean births‡	1.860	1.270	0.773	0.403	0.140	0.080
Coefficient/mean	-12.59%	-16.77%	-20.22%	-20.10%	-20.14%	-14.13%
	Race: Asian					
Schools at 15	-0.197	-0.169	-0.108	-0.049	-0.017	-0.013
	(0.088)**	(0.065)***	(0.040)***	(0.019)***	(0.007)**	(0.004)***
R2†	0.148	0.111	0.069	0.037	0.022	0.014
Mean births‡	0.376	0.242	0.139	0.068	0.022	0.012
Coefficient/mean	-52.42%	-69.98%	-77.74%	-73.00%	-78.00%	-103.54%
	Race: Pardo					
Schools at 15	1.798	1.473	0.630	0.236	0.065	0.150
	(0.811)**	(0.630)**	(0.418)	(0.230)	(0.086)	(0.052)***
R2†	0.102	0.071	0.042	0.037	0.039	0.060
Mean births‡	49.368	36.137	23.620	12.937	4.858	2.611
Coefficient/mean	3.64%	4.08%	2.67%	1.83%	1.35%	5.74%
	Race: Indigenous					
Schools at 15	-0.021	0.028	0.056	0.060	0.052	0.051
	(0.045)	(0.053)	(0.053)	(0.047)	(0.030)*	(0.029)*
R2†	0.010	0.008	0.012	0.015	0.023	0.030
Mean births‡	0.750	0.572	0.398	0.235	0.103	0.091
Coefficient/mean	-2.86%	4.88%	14.12%	25.53%	50.63%	56.07%
Observations	43,650	43,650	43,650	43,650	43,650	38,800

Notes: Robust standard errors reported in parenthesis. ***, **, and * denote statistical significance at 1%, 5% and 10%. Regressions include cohort primary enrolment, nursery and preschool classrooms, cohort-specific male and female populations, municipality expenditures, municipality and cohort. †R-squared for within variation reported.

Appendix B: Migration, teenage childbearing and school growth

In Section 4.1 we discuss the potential risk to our identification strategy posed by municipal migration flows within Brazil. In this appendix we explore this threat using data from the 2010 Brazilian population census. The 2010 census provides detailed information, for an 11% population sample, on family structure and migration status.[17] We identify a census individual as a teenage parent if: a) a child is in the same household; b) the individual is identified as the child's parent and not a step-parent; c) the age difference between the individual and the child is not greater than 18 years. Under this strategy, individuals who are teen parents will only be identified as such if they live with their child at the time of the census.

The census provides three categories for an individual's migratory status: 1) Born in municipality and lived there the entire life; 2) Born in municipality but lived elsewhere; 3) Not born in municipality. An individual is identified as a migrant if 1) does not apply. For migrants we have information on the current municipality, previous municipality (only for migrants who moved after 2000) and the age at which they moved to their current municipality. To ensure confidentiality, the publicly available microdata do not reveal municipalities with small populations. Therefore, for migration flows, the sample analysed a sub-set of the larger municipalities used in the main paper.

Our analysis here shows the following: a) Migration taking place between 11 and 18 years of age is associated with higher rates of teenage parenthood; b) On average, the origin municipalities have a lower school growth rate than destination municipalities; c) there is a small, but positive, correlation between the relative school growth rate (of destination municipality to municipality or origin) and teenage childbearing.

In summary, this analysis suggests that any potential threat to our identification will lead us to underestimate the average effect of the school expansion on childbearing.

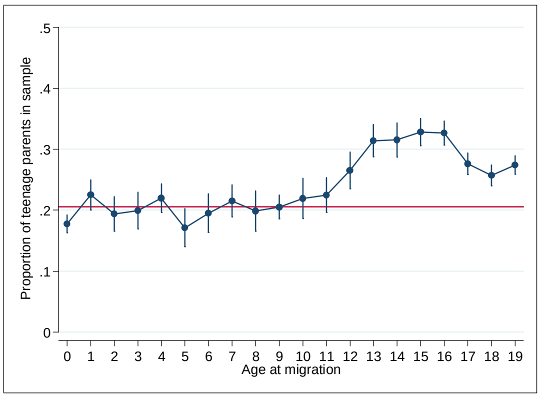
Migration and teenage parenthood

Here we look at the association between migration and teenage parenthood. In Figure B1 we plot, for different ages at census, the teenage birth rate for migrants against age at migration. We show a selection of census ages—19, 24 and 30 for females, 19 and 24 for males—but the pattern is very consistent regardless of the chosen census age. For females these figures suggest that migration between 11 and 18 years of age is associated with significantly higher rates of teenage childbearing than is migration at younger or older ages. Teenage birth rates for migrates at younger or older ages do not significantly differ from the non-migrating population.

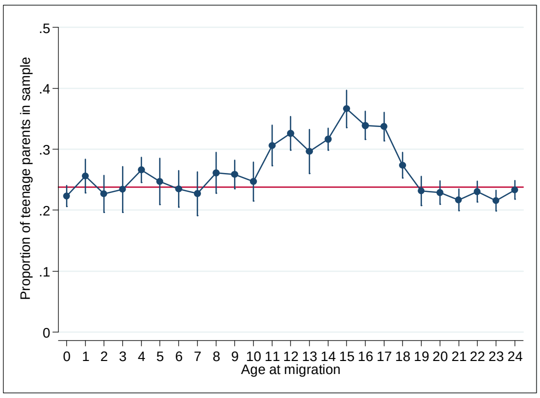
[17] The final sample is based on stratification based on municipality size. Details on the sampling can be found at http://biblioteca.ibge.gov.br/visualizacao/livros/liv81634.pdf.

Figure B1: Teenage parenthood by age of migration

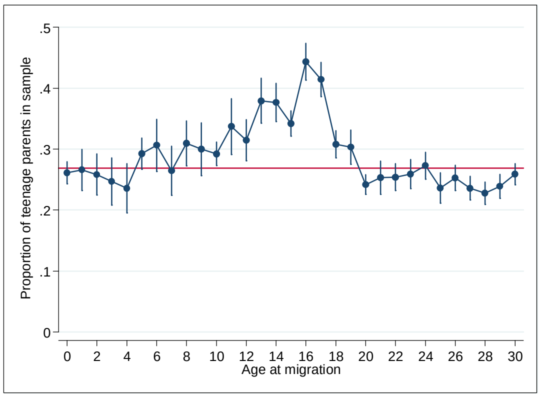
Female, 19 years old at census



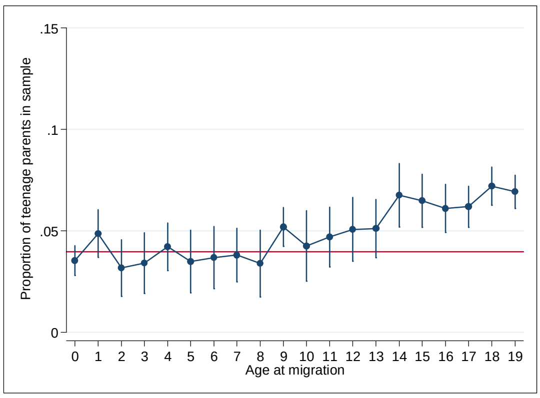
Female, 24 years old at census



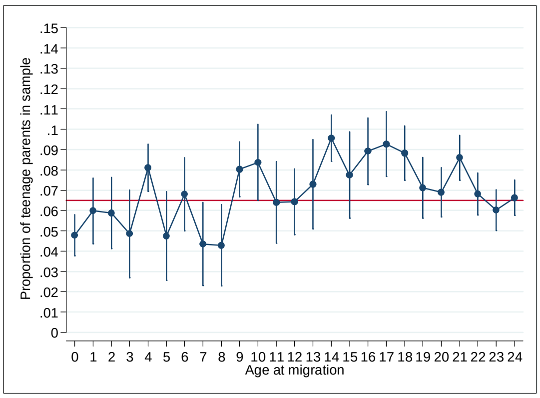
Female, 30 years old at census



Male, 19 years old at census



Male, 24 years old at census



Notes: Markers indicate mean of teenage pregnancies, bars indicate 95% confidence interval. Red horizontal line is the estimated teenage birth rate for non-migrating population of same age.

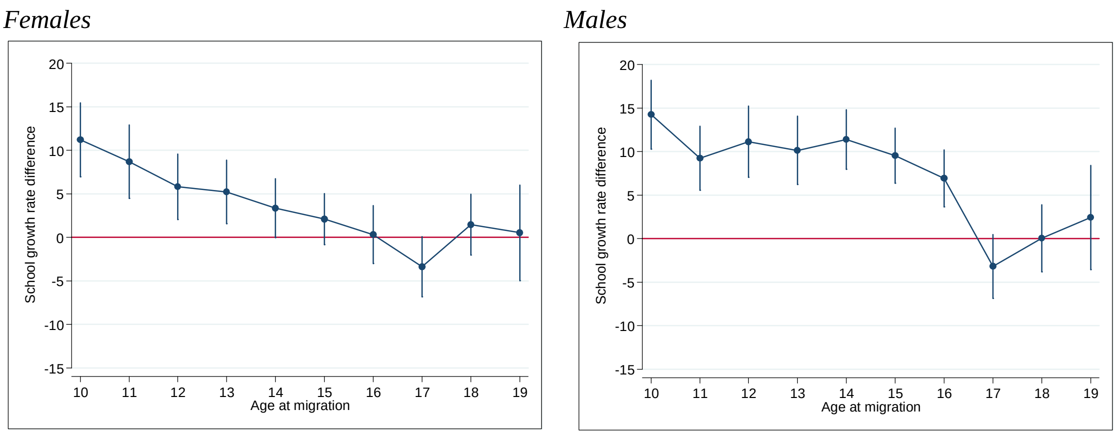
The pattern is similar for males, although smaller magnitude and noisier. This may reflect that children of teenage parents are more likely to live with the mother, leading to a tendency to under-identify teenage fathers.

Migration flows and municipal school growth

To consider the relationship between migration and school growth we define net school growth as the difference in the school growth rate between the origin municipality and the destination municipality. The school growth rate is defined as the percentage growth in schools between 2000 and 2009. The net school growth therefore reflects the percentage point difference between the school growth rate in the origin municipality and the destination municipality.

The average net school growth rate is 4.03 percentage points, suggesting that destination municipalities have a higher average growth rate than origin municipalities. We regress the net school growth on the age at which an individual migrated, for migrants aged 15 to 19 at the 2010 census, and plot the coefficient estimates for each age at migration in Figure B1. The pattern in this figure suggests that migrants at younger ages have, on average, a higher net school growth rate than migrants at older ages.

Figure B2: Migration and net school growth



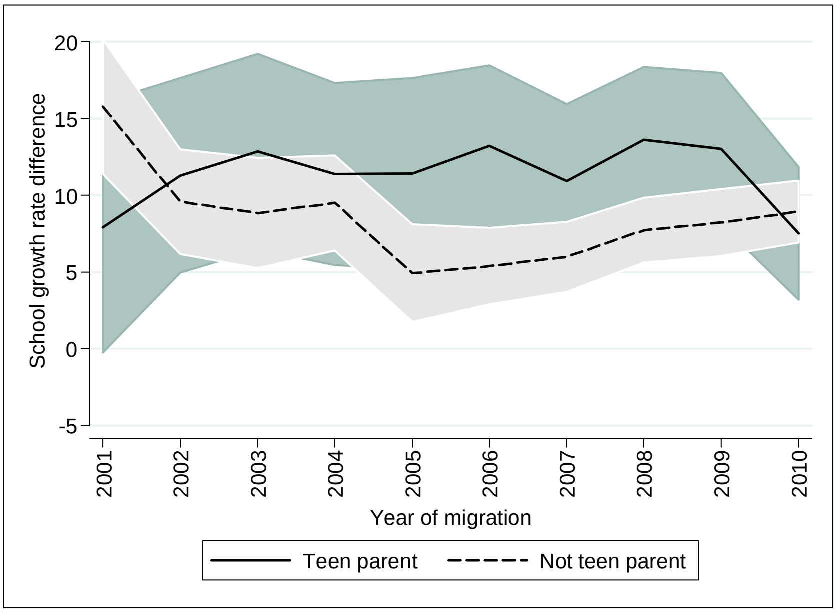
Notes: Markers indicate the percentage point difference in the school growth rate between the municipality left and the current municipality. Bars indicate 95% confidence interval.

The pattern in Figure B2 suggests that, on average, young women in Brazil, during the period considered, migrate to municipalities with higher rates of school growth than their municipality of origin.

Migration flows, municipal school growth and teenage parenthood

Finally, it may be the case that, for reasons unobserved to the econometrician, migrant households with teenagers that have a high propensity to become a teenage parent exhibit differential migration patterns then migrant households with teenagers that have a low propensity to become a teenage parent. We examine this possibility by looking at, for migrants aged 19-24 in the 2010 census, net school growth for teenage parents versus those who do not have a child present born in their teen years. This is summarized in Figure B3 and Table B1.

Figure B3: Net school growth for teenage parents and non-teenage parents



Notes: Shaded areas cover 95% confidence interval for teen parents (green) and non-teen parents (grey).

Two things appear from this analysis. First, there does appear to be a difference in net school growth between the two groups. Second, teen parents appear to, at least for latter years, have higher net school growth. This again is suggestive evidence that migration may lead us to under-estimate the true effect of interest.

Table B1: Net school growth for teenage parents and non-teenage parents

Year of migration	Non-teen parent	Teen parent	Difference	SE
2001	15.763	7.899	7.864	(4.633)*
2002	9.582	11.288	-1.706	(3.998)
2003	8.843	12.846	-4.002	(3.527)
2004	9.505	11.382	-1.876	(3.367)
2005	4.918	11.416	-6.498	(3.457)*
2006	5.386	13.206	-7.820	(3.004)***
2007	5.997	10.924	-4.927	(2.960)*
2008	7.728	13.605	-5.877	(2.720)**
2009	8.221	13.032	-4.810	(2.738)*
2010	8.944	7.516	1.429	(2.298)
Observations	46,513	13,198		

Robust standard errors corresponding to difference reported in parenthesis. ***, **, and * denote statistical significance at 1%, 5% and 10%.

Appendix C: Sources for municipal expenditure data

Variable	Notes	Source	Link
GDP	Municipality gross national product at current prices (R$)	SIDRA/IBGE	http://goo.gl/OpQffe
Municipality spending	Total local government (municipality) expenditure at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Welfare spending	Local government (municipality) expenditure on assistance and welfare at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Education spending	Local government (municipality) expenditure on education and culture at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Health spending	Local government (municipality) expenditure on health and sanitation at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Judicial spending	Local government (municipality) judicial expenditure at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Security spending	Local government (municipality) expenditure on national security and public defense at current prices (R$)	IPEADATA	http://goo.gl/lSI3nz
Health institutions	Number of public health institutions per 1,000 population, including general and specialized hospitals, policlinics, health centers (*posto de saúde*), basic health centers (*unidade básica de saúde*)	CNES/ DATASUS	http://goo.gl/tdwW
Nurses	Number of qualified hospital nurses according to the Brazilian Classification of Professions (CBO-2002) per 1,000 population	CNES/ DATASUS	http://goo.gl/tdwW
Bolsa Família recipients	Number of *Bolsa Família* recipients per 1,000 population	DATASUS	http://goo.gl/tdwW
Bolsa amount	Average *Bolsa Família* amount per recipient at current prices (R$)	DATASUS	http://goo.gl/tdwW