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Education Quality, Income Inequality, and Female Labor Force Participation in Brazil

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This paper examines the impact of education quality on income inequality among men and on female labor force participation. I introduce a new dataset on local education expenditures for a 64-year period. By matching education spending to the time and place where each person went to school, the data allow for a much more granular measurement of human capital differences than measures like level of schooling or years of school attainment. They also permit measurement of human capital differences and evolution over a much longer time period than the data that are typically available. I show that differences in the quality of education received during childhood become significant determinants of income differences among fully employed adult men. In a finding that is new to the literature, I report that school quality differentials are significant determinants of how adult women allocate their time between domestic labor and formal wage work.

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JEL codes: I24, I25, O15, J24, J16, D31, D63, H75, N16

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1 Education Quality, Income Inequality, and Female Labor Force Participation

1.1 A Summary Overview

This paper examines the impact of education quality on income inequality among men and on female labor force participation. I introduce a new dataset on local education expenditures for a 64-year period. The data allow for a much more granular measurement of human capital differences than measures like level of schooling or years of school attainment. They also permit measurement of human capital differences and evolution over a much longer time period than the data that are typically available. I show that differences in the quality of education available during childhood become significant determinants of income differences among fully employed adult men. In a finding that is new to the literature, I report that school quality differentials are significant determinants of how adult women allocate their time between domestic labor and formal wage work.

Brazil is a fitting environment for exploring the topics of education, income inequality, and female labor

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force participation. It is a traditional society with strongly gendered roles and one of the most unequal income distributions in the world. It is one of the largest countries, a federation with strong local governments that have their own sources of revenue and are constitutionally charged with financing schools. This results in ample variation in school funding across time and place. The Brazilian economy went through momentous change over this period, including a record-setting period of growth and structural transformation that was dubbed "miraculous" in its time, rapid evolution from rural to urban, and a swift expansion in school coverage.

The tremendous changes Brazil has undergone since the 1940's are bound to have had an effect on men's relative wages, but also on female labor force participation decisions. Even a simple model with two goods, where the husband's earnings are given, predicts that the share of time women devote to home production depends on the relation between the husband's earnings and her own earnings potential. This interaction should be evident in the data, but if we measure education as 1, 2, 3, ... years of schooling, we will pick up very little variation in human capital– especially during periods when school attainment was low.

Throughout this paper I use "human capital" to describe the set of individual characteristics that affect productivity and are reflected in wage differentials among otherwise identical individuals. Human capital can be enhanced by increasing formal schooling duration, by attending better quality schools, and by learning through labor market experience. The dataset I introduce enables me to tease out a much more nuanced measurement of human capital by controlling for school quality. I show that the measure does in fact contain significant human capital information by estimating classical Mincer equations for prime-aged men and then demonstrating that the quality data significantly boost our ability to explain income differences among these men.

I use coefficient estimates from quality-enhanced Mincer regressions for prime-aged men to define "Hk", a human capital index for women. Very high labor force participation rates among prime-aged men make it unlikely that the relation between their human capital and their market wages is significantly distorted by selectivity and censorship problems. Hk is a prediction of what a woman's wage rate would be, if she were a fully employed, prime-aged man. I employ Hk to examine how human capital affects women's time allocation decisions and how these decisions have changed over time.

I assume that labor markets are "gendered". By this I mean two things: that women bear most of the responsibility for non-marketed domestic production and that women's wages are no larger than men's. In this context, theoretical guidance on how secular growth in women's human capital should affect labor force participation is complicated for two reasons. First, if home-produced goods are normal, then for a given spousal income, increasing a woman's human capital will have income and substitution effects that are opposite. These opposite effects can net-out to supplying more or to supplying less formal market labor. Secondly, spousal human capital levels are likely to be correlated. So, since the theory predicts that her labor force participation decision depends not on her human capital directly, but on the relation between her earnings potential and the earnings of her husband, the effect of secular growth in men's and women's human capital is also not straightforward.

My interest is in explaining how evolving levels of male and female human capital affect income inequality and women's labor force participation. To this end I develop a correlated spousal earnings (CSE) model that allows for spousal incomes to be co-determined. Women draw husbands from distributions that depend on their own human capital. Higher human capital women draw husbands from first order stochastically dominant distributions. At the time of marriage, a woman knows the lognormal distribution of her husband's income, but his actual income is only revealed to her after marriage, and this makes her labor supply stochastic. The CSE model allows for a straightforward and unambiguous estimation of the net effect of human capital on women's time allocation.

CSE estimates reveal that over the 64-year period studied, having more human capital has consistently led to women spending less time on domestic tasks. Classic Heckman 2-step model estimates show sign flips for the woman's human capital in the second stage and for husband's income in the first stage. These sign changes might be caused by a more complex relationship between these two variables than is typically allowed for and not by real patterns in the data.

A major take-away from what this new data elucidates is that unequal access to quality education among children translates a major source of inequality among adult men and women.

1.2 A Detailed Overview

In the classic depiction of economic development, technological change favors industry and services with larger labor productivity increases. Agricultural productivity growth allows a smaller fraction of the labor force to feed the population. Engel curves guide an increasing share of expenditure to urban-based industrial and service sector goods. Technological change and evolving demand conspire to bring about an exodus of the widely scattered agricultural population, funneling it to high density urban areas, pulled by better wages and better and more diverse public goods (Lewis [1954] and Ranis and Fei [1961]). There is a "shift in the locus of employment in contemporary societies from family to firm" (Schultz [1990], 457).

The literature has long recognized this process as inherently inequality-inducing (Kuznets [1955] and [1966], Anand and Kanbur [1993], Bourguignon and Morrisson, [2002]). Development transforms the production landscape from a loosely-knit network of rural, largely self-sufficient nodes of home production to an integrated, urban system where labor specializes and most output is marketed for trade. As countries modernize, economic specialization makes education an increasingly important production input and determinant of income (Autor and Dorn [2013]). The distribution of education becomes an important determinant of international income distribution across countries (Psacharopoulos [1993], Barrow and Lee [1993], and O'Neill [1995]) as well as within them (Schultz [1961] Mincer [1974, Becker [1975, and Lam and Levinson [1991).

Development affects men and women differently, but the literature is less clear on how women fit into the development picture. One long-standing explanation is that responsibility for raising children —and by production complementarity also most household production— has preponderantly fallen on women's shoulders for reasons ultimately having to do with nursing. This idea has been around for a long time (for instance, see Engels, 1884), with more recent work emphasizing the role of agricultural technologies like the plow and crop type (Burton and White [1984, Alesina et al. [2013). Development increases labor specialization and extends the monetization of economic activity, but primarily to goods and labor that are destined for use outside of the home. Since traditional gendered assignment of productive activity places work outside the home in men's camp and domestic production in the women's, these changes affect men's activities more. The classic development story is therefore largely a story about men and male labor, and the study of development and inequality has focused on inequality among men's earnings.

Women's work activities are more difficult to observe, especially in the less developed countries. We know that the rate of formal female labor force participation (FLFP) is U-shaped with respect to development. We know little more than that they tend to withdraw from the labor force, move behind the beaded curtains of home production, and then reappear in formal labor markets gradually. Shultz (1990), Goldin (1995), and Mammen and Paxson (2000) suggest that the relation between economic development and female labor force participation (FLFP) may be U-shaped. Women tend to "disappear from the measured labour force" (Bandeira et al. 2022, p. 2228) and then re-enter at middle stages of development. Using a 43-country set of time-use surveys, Bridgman et al. (2018) find that household production hours account for nearly half of total hours worked around the world and that this makes women's formal labor more sensitive to development than men's. Women's total hours worked tend to be higher than men's at low levels of development. Hours in domestic production for women fall and hours in market activities rise. Systematic patterns in this "marketization" of women's labor are evident within given countries at a given point in time too, for the degree to which women's labor is marketized rises with educational attainment.

We also know that education plays an important role in this process, since advances in education raise the opportunity cost of working at home.

But whatever its origins, even modern advocates of gender equity argue that the "last chapter" in social evolution towards gender equity need not eliminate all differences. Goldin (2014) suggests that we should rather adapt by encouraging more flexible work schedules and reducing the incentives for working long hours, essentially ensconcing some gendered labor force patterns in social policy.

Humphries et al. (2024) study the gender wage gap in Sweden, at the top of the world's per capita income distribution, examining exquisitely nuanced roles of cognitive, grit, and interpersonal skills in gender sorting across college majors, finding that college major choice explains more than half of the wage gender gap in Sweden. The education level, labor market sorting, and gender pay and work gaps that are at work in this paper are of a very different sort. The vast majority of the population I study has not gone to college –indeed, for most of the period I study the majority had not finished high school (Table 2).

This paper follows in the Labor Economics and Economics of Education tradition of employing per capita school spending as an indicator of differential school quality (for instance, Altonji and Dunn (1996), or Nicoletti and Birgitta [2018). I introduce a new dataset with 64 years of local primary and middle school education spending in Brazil. This adds to the literature on economic development, education, and income inequality by providing a very granular measure of human capital differences across men and women over a very long period of time. More importantly, this paper expands our understanding of development by focusing on women's participation in formal labor markets over a critical transitional period in one of the world's largest, most unequal, and "gendered" countries. In particular, it examines how education quality differentials affect women's participation in formal labor markets, highlighting the role of education quality in helping women through the transition from fully specializing in home production to participating in wage labor markets.

The relation between development and FLFP during the development process is the first of two main subjects in this paper. The second subject considered here is the relation between education inequality and income inequality, generally. I will argue that education quality differentials partially explain changing FLFP patterns.

This paper also makes a large contribution to understanding the particular Brazilian context. Brazil's development story is prototypical. As its economy transformed, the country metamorphosized from 55% rural in 1960 to 85% urban in 2015 (PNAD surveys). In the process, it became the most unequal country in the world. Its Gini coefficient of 0.63 has stood as "*almost a historical and worldwide record*" (Lopez-Calva et al. [2012), for a long time time (Ray and Genicot [2023). Given that the Latin America was resource-rich at the starting gate, development in the region has been unspectacular (Edwards [2009) and accompanied by unusually high income inequality (Bourguignon and Morrisson [2002, Pinkovskiy and Sala-i-Martin [2009). Recent explanations for this Latin American growth puzzle have focused on the role of educational quality. Hanushek and Woessman (2012) brought attention to deficiencies in the quality of education in the Latin American region as a whole, while Brotherhood, Cavalcanti-Ferreira, and Santos (2019) have shown that there are enormous differences in pre-college education quality within Brazil and Dureya et al. (2023) lay a large portion of the blame at the feet of unequal access to high-quality public universities.

Brazil also provides a quintessential example of initially low, but rapidly rising FLFP. Differences in years of school attainment have long been recognized as a major cause of inequality in Brazil (Yap [1976], Almeida dos Reis and Paes de Barros [1991], Lam and Levinson [1991], Dureya et al. [2023]). The data I introduce in this paper show that school *quality* variations are also expressed in the earnings differentials among people schooled between 1941 and 2004.

Economic theory predicts that FLFP in traditional societies depends on the relative earnings potential of women *vis a vis* their husbands. School quality affects the opportunity cost of women's time and therefore influences their decision on whether and for how many hours to participate in the formal labor market. I show that education quality variation decreased substantially over this period of Brazil's drastic economic transformation. As education quality grew and the inequality of education across Brazil fell, more women worked, they worked for more hours, and the inequality of earnings fell. During Brazil's transition from a mostly rural, agriculture-based economy to its modern highly urbanized industrial and service-sector foundation, education quality has played a pivotal role in the integration of women into formal wage-labor markets. Nevertheless, school quality inequality remains high, and education quality differentials continue to play a major role in explaining general income inequality and gender differences in labor force participation.

I match a 3-year moving average of *municipio* education spending to the time and place where survey respondents lived when they were in school, as reported in panels of Brazil's large PNAD surveys, spanning a half-century. I then test the explanatory power of this school quality proxy by estimating standard Mincer earnings equations for prime-aged men. Limiting estimates to them minimizes selection bias since this demographic group is nearly fully employed (Heckman [1990, Mulligan and Rubinstein [2008]). I find that school quality is a fundamental determinant of earnings inequality among prime-aged men. Between 1976 and 2015, differences in school quality account for about 11% of the explained variation in men's log wages.

Having established the validity of the school quality measure, I create a human capital index for women based on these quality-augmented Mincer equations for men. The women's human capital index is a prediction of what women would earn if they were prime-aged, fully-employed men. I use this index to examine how human capital differentials affect women's allocation of time between domestic production and the formal labor market.

This paper also contributes to our econometric toolbox. The econometric problems of censorship and selection bias are particularly keen when we study the formal labor market employment of women in traditional societies. Low FLFP leads to many zero hours-worked observations and induces more serious selection bias than is typically found in samples of men (Heckman [1990], Mulligan and Rubinstein [2008]). Some consequences of this bias can be significantly diminished with methods developed by Tobin (1956) and Heckman ([1974] and [1976]). However, neither of these methods results in a straight forward representation of the net effect of human capital on women's time allocation. I posit that this is because neither method allows for correlation between a woman's own potential earnings and the human capital of her husband.

I overcome this problem with a female labor force participation model that accounts for right-censoring and explicitly incorporates correlation between latent spousal earnings in a correlated spousal earnings model (CSE). When a woman marries, she only knows the earnings potential of her husband as a lognormal density function, and she also knows that the median value of his potential earnings is a function of her own human capital endowment. Her husband's actual income is fully manifested only after after she marries him, making the labor force decisions of otherwise identical women stochastic. Unlike the Heckman and Tobit models, the CSE model produces a reduced form, stochastic female labor supply function that can be used to estimate in a simple way the full effect of differences in women's human capital endowments on the allocation of their time between domestic and formal wage labor.

Maximum likelihood estimates of the CSE time allocation model show that higher quality schools produced women who dedicated less time to domestic production throughout Brazil's development over the last half-century. This finding complements Heckman (1976) replications, which I also augment with the school quality information. The Heckman model —which treats marriage, husband income, and fertility as pre-determined— suggests that from 1976 to 1995 women who had more human capital and a husband with higher earnings were less likely to work in formal labor markets, but —conditional on working— having more human capital led them to work for more hours. These effects are reversed after 2005, and it appears that sign of Heckman selection into formal labor also reversed, in a pattern similar to what Goldin (2014) finds for the United States. Higher human capital women with high earning husbands have become *more* likely to work (Mammen and Paxson, 2000 report a similar reversal). I posit that the reversal reflects functional form restriction in the Heckman model that is not supported by changes in real empirical patterns.

Maximum likelihood estimates show that the net effect of having more human capital is to strongly reduce the time women spend on domestic tasks. A new finding is that higher school quality unambiguously and significantly bolsters the FLFP effects of more years of schooling. For robustness, I replicate the results of Heckman (1976) with the Brazilian data panels and show that quality also enhances those results. By repeating the procedure roughly every 10 years for panels that span 1976 to 2015, I shed light on why FLFP behavior has changed over the last half-century and on the important and inequitable consequences of differential access to quality schooling.

In sum, this paper examines how human capital attainment has affected women's participation in formal labor markets, extending the literature on the relation between schooling, income inequality, and education quality. When I combine school quality with years of schooling attained, I obtain a measure of inter-person differences in human capital that is much finer than years of schooling alone, the measure traditionally used in this literature. I find that the quality of education that is available to children becomes a significant determinant of earnings inequality among adult men and of the labor force participation of adult women. The paper contributes to three areas of the literature. It adds to our understanding FLFP by jointly examining topics of selection, human capital, and women's wages over time as in Mulligan and Rubinstein (2008), Shultz (1990), and Goldin ([1995] and [2014]). It contributes new evidence on the importance of taking school quality differentials in into account when explaining income inequality among men, and women's time allocation decisions and thus extends the work of Welch (1966), Behrman and Birdsall (1983), Card and Krueger (1992), Altonji and Dunn (1996), Bedi and Edwards (2002), Hanushek and Woessman (2012), and Brotherhood, Cavalcanti-Ferreira, and Santos (2019). Thirdly, this paper extends our understanding of income inequality in Brazil, specifically, by providing a more granular measure of inequities in access to human capital across place and over more time than has ever been studied before, extending evidence provided by Yap (1976), Almeida dos Reis and Paes de Barros (1991), Lam and Levinson (1992), Fernández and Messina (2018), Correia de Souza and Menezes Filho (2020), and Brotherhood, Cavalcanti-Ferreira, and Santos, (2019). Most importantly, the paper breaks new ground by explicitly considering inequality among women.

The rest of the paper is laid out in classical form. Section 2 describes two data sets, the PNAD survey series and a new data set on Brazilian local education expenditures. Section 2 includes important summary statistics from both data sets and also lays out the method for matching local school spending to individual PNAD respondents.

Section 3 introduces a framework for thinking about women's relation to the labor force changes during development. It includes a simple FLFP model, a characterization of human capital, and a depiction of the gendered nature of labor markets and of the labor force participation decision.

Section 4 describes the econometric problems that arise in analyzing FLFP when spousal human capital is correlated. It then lays out 3 estimation methods for dealing with these econometric problems. First, the new CSE econometric model of female domestic labor supply that allows for correlation between a woman's earning potential and her husbands and also accounts for censorship. The maximum likelihood specification of this model is then contrasted with with Heckman's two-step selection and a more traditional Tobit specification.

Model estimates appear in Section 5, followed by summary of findings and the major conclusions in Section 6.

2 DATA

The Brazilian experience is a very good setting for watching how education shapes the role of women in the labor force. The country's development over the past half century is a quintessential example of Kuznets' modern economic growth paradigm (Kuznets [1966). Brazilian experience is one of tremendous structural change. The share of agricultural value-added in total value added fell by 75%, between 1960 and 2006, from 20.6% of total to 5.1%. In 1975 only 22% of its labor force resided in one of the country's 10 largest cities. But by 2015, this had grown to 66%, trippling in just over 30 years. Brazil is ideally situated for finding the FLFP patterns posited by the literature (Schultz [1990, Goldin [1995, and Mammen and Paxson [2000).

Additionally, Brazil has been the world's poster child of economic inequality for a long time (Baer [2014) and its income inequality has been convincingly blamed on unequal educational opportunities (Almeida dos Reis and Paes de Barros [1991], Lam and Levinson [1992], and Brotherhood, Cavalcanti-Ferreira, and Santos [2019]). Finally, as one of the largest countries in the world, it is diverse in climate and population, and as a federation with a tradition of strong state and local government, it provides fiscal diversity.

The data I use combine individual-level survey data with public education expenditure data. The survey data come from Brazil's long-standing PNAD household survey series. PNAD panels for 1976, 1985, 1995, 2005, and 2015, provide nationally representative surveys with detailed questions that are roughly consistent over a very long period of time.

The school quality dataset consists of detailed public finance information that I have collected over the past 30 years. It is new to the literature. It contains yearly local level expenditures on education at the municipality level for 64 years. I combine the PNAD surveys with a 3-year moving average of local municipality education spending at the PNAD respondent level. I link PNAD respondents to education spending in the location where and at the time when they were 10 years old. I examine the education quality effects on people schooled between 1941 and 2004. This 64-year time span is long enough to see generational turnovers in the labor force and witness the impact of Brazil's tremendous strides in increasing educational attainment. PNAD's large samples make it possible to parse the data finely.

2.1 PNAD Surveys

Tables (1) and (2) provide summary statistics on some of the variables of interest from the five PNAD surveys. Each table has two panels. Panels A contain data for the labor force –defined as people between the ages of 16 and 65– while Panels B contain the same data for young adults, ages 22 to 26. Contrasting Panels A and B for each survey year provides a good sense of change and the rate of change; it is insight-evoking to look at data in the B panels as though they were the derivatives of statistics in the A panels.

Table (1) summarizes a broader set of labor force variables for these age groups. Panel A shows that between 1976 and 2015 the share of working aged men who reported wage work fell from 86.65% to 77.43%, but that the proportion of men who were in the labor force –either working or were looking for work– remained close to 90%. The remainder were either retired or in school.

In contrast to the male pattern, the share of women working or looking for work increased rapidly, from about one, to two thirds of the 16-65 population. The proportion of women working rose from 33% to 54%. During this period when Brazil moved from a six to a five-day work week, the average number of hours worked by employed men fell by more than 27%, from over 43 to under 32 per week. Women's LFP made substantial gains, and the proportion of women who report zero hours of paid work fell by nearly one-third, from 65.2% to 45.5%. Nevertheless, remunerated employment remains a secondary activity for Brazilian women (Berniell et al. [2023]). Mean hours for working women between the ages of 16 and 65 rose from fewer than 15 to 19 per week.

Statistics for the young cohort, in Panel B of Table (1) provide a glimpse of what the future portends. About 90% of 22-26 year old men were active in the labor force, though the proportion that was "temporarily out of work" has grown over time, at the expense of the proportion actually working. Also, weekly hours worked among employed young men fell from the equivalent of five and a half, eight-hour work days to fewer than four work days. The share of young men who report zero hours of paid labor rose from 6.5% to nearly a quarter of the labor force over the 40 years covered by the five surveys.

During the same period, FLFP among the younger group of Panel B rose from 40.6% to 74.3% and the share of young women who report zero hours fell from 60% to 44%. The proportion of 22-26 year old women who reported attending school as their main activity began to be significantly higher for women than men by 1985 and the female/male gap rose to 10% by 2015.

Over the period examined, Brazil was ending its rapid transition from a mostly rural country to a mostly urban one. In 1976 about two-thirds of the labor force was urban-based and by 2015 85% was urban. This transition also involved moving the population from small urban centers into the ten largest ones. These larger cities accounted for 23% of the labor force in 1976, but for 66% in 2015. The declining proportion of people living in places where they were not born suggests that the mass reshuffling of Brazil's population is finally quieting down. Migrants made up nearly half of the labor force in 1976, but their share declined to 43% in 2015 (Panel A). Among the young, this proportion declined from 46.6% to 31.8% (Panel B).

Panel A of Table (2) shows that the labor force matured over the last half century, brought about by decreasing fertility rate and leading to a higher proportion that is pensioned. The mean age of the labor force rose by about 5 years between 1976 and 2015. Most significant for this paper is evidence that educational attainment doubled, from about 4 years of schooling to more than nine. The proportion of the labor force that is functionally illiterate (fewer than 4 years of schooling) declined from nearly 60% in 1976 to less than 15% in 2015. At same time the proportion with at least a high school degree rose from 13.5% to nearly half of the labor force.

[INSERT TABLES (1) AND (2)]

The 1976 educational statistics in Panels A and B of Table (2) are very similar. This means that Brazil's rapid gains in school attainment over the past half century had not yet begun. Mean years of schooling for the age 22-26 group (Panel B) were about the same as the average attainment of the 16 to 65 labor force in 1976 (Panel A). By 1985, however, the panel B group had attained one more year of schooling than the full labor force in Panel A; rapid change in overall educational attainment was taking place. An important fact that stands out is that early gains in mean educational attainment came mainly from more years of schooling at the top of the distribution.

During this Brazilian education revolution, the attainment of working aged women caught up and surpassed men's. It was about 5.5% below men's in 1976. By 2015 women's schooling was 5.5% *more* than men's. Women's attainment appears to have really taken off after the military were removed from power in the late 1990's. The contrast between panels A and B indicates that it continues to accelerate to this day; the educational gap between Brazilian women and men continues to widen. By 2015 the proportion of women between the ages of 22 and 26 with at least 11 years of school education was more than 10 percentage points higher than for men. In fact, the gains in women's mean years of schooling came largely from women at the top of the education distribution staying in school longer.

Together, Tables (1) and (2) reveal the classic Kuznets-Lewis pattern of economic transformation for men. But they also reveal an equally compelling if the less commonly told story that gender plays a leading role in the allocation of labor and in the distribution of human capital during economic development. Women's education initially lags men's and then grows to surpass it.

This pattern is bound to affect labor force participation gender ratios. Table (1) shows that men worked for a wage in 1976 and women were in charge of the home. The share of men who reported domestic tasks as their main activity was negligible (0.05%). In the same year 57% of prime-age women devoted *all* of their time to the home. By 2015 the proportion of women who stayed home full-time had fallen to 44%, marking a trend towards the "marketing" of women's labor, a trend that is even more strongly evident in the younger, Panel B cohort. The decline from 53% to 36% in the share that stays home full time is evidence that young women are shaping a new role for themselves in the Brazilian economy.

3 SCHOOL QUALITY

3.1 Earnings Differentials by Place of Study

The literature on Brazilian inequality has always recognized that education plays a central role. Yap (1976) was among the first to point out that Brazilian rural-urban migrants were poor as a group because of their lower educational attainment, not because of anti-migrant discrimination in the cities. Lam and Levinson (1992) demonstrated a nearly tautological link between the distribution of educational attainment and income distribution and therefore between changes in the distribution of years of schooling and changes in the distribution of income in Brazil. Almeida dos Reis and Paes de Barros (1991) found that regional differences in the distribution of attainment are not able to explain large differences in wage inequality across Brazil's major metropolitan areas. They point out sharp differences in the wage-education profiles slopes across cities –differences that could be caused by labor allocation inefficiencies or (more likely) by cross city differences in what it means to have one more year of schooling. Brotherhood, Cavalcanti-Ferreira, and Santos (2019) use 2010 census data to infer variations in Brazilian education quality based on the earnings of men —ages 24 to 65— who migrated to the city of São Paulo after age 24. The large size of the census data sample allows them to discern statistically significant differences in earnings by place of origin while holding market conditions constant. They ascribe differences in what men with the same level of schooling earn in São Paulo to differences in the quality of the schools provided by their places of origin.

As household surveys go, the PNAD samples used in this paper are large. Unfortunately, they are not large enough to capture enough observations in São Paulo to fully replicate the Brotherhood, Cavalcanti-Ferreira, and Santos (2019) method. Instead, I follow Almeida dos Reis and Paes de Barro and examine the earnings of men who are working in one of Brazil's "big 10" cities. All ten cities had populations of more than 1.5 million for the full period under study. Though smaller than the census sample, the PNAD data have the advantage of being more granular. Whereas census data used by Brotherhood, Cavalcanti-Ferreira, and Santos (2019) only report 3 schooling intervals, the PNAD data used in this paper contain actual years of school attainment. I am also able to determine whether the location of prior residence was urban or rural.

In order to replicate the Brotherhood, Cavalcanti-Ferreira, and Santos (2019), I selected the sample of prime-aged men who were living in one of Brazil's "big 10" cities and then estimated "Mincer" log wage regressions of the form

$$y_i = \mu_0 + \mu_v v_i + \mu_{2v} v_{vv}^2 + \mu_s S_i + \mu_{m_i} M_i + \epsilon_i,, \tag{1}$$

where for person i, y_i is the log hourly wage rate, ν_i is labor market experience, S_i is years of formal schooling, and M_i is the location where i attended school (São Paulo City omitted, as the basis for comparison), and the μ_j are parameters to be estimated, j$\in$ $\{0,\ \nu,\ \nu\nu,\ \text{s},\ m_i\}$.

I estimated Equation (1) for men between the ages of 16 and 65 who reported working at least 29 hours in the week prior to the survey, for each of the 1979 [1995 [2005, and 2015 samples.[1] Only schooling locations that have at least 30 observations currently working in a big 10 city are included as valid places of origin. In order to save space, the full estimates are not shown here, but are available upon request. The estimated coefficients on schooling and experience were all significant at the 99% level in all samples. The rate of return to schooling was close to 20%, in 1976, and then declined gradually and monotonically over the years to 12%. Coefficient estimates on experience show the standard positive, but decreasing pattern for all years.

Appendix Table (A1) shows estimates of earnings differentials by rural and urban municipalities of each state that contributed at least 30 men to the large city labor market. The coefficients correspond to μ_{m_i} from Equation (1); they contrast the earnings of a man who was educated in M_i to the earnings of one who was educated in São Paulo City, *ceteris paribus.* Since most migration was from rural to urban areas, more rural origins satisfy the minimum, 30-observation criterion for inclusion. Table (A1) reveals tremendous differences across the country. This reinforces the Brotherhood, Cavalcanti-Ferreira, and Santos (2019) evidence of large variations in the quality of education across Brazil. *Where* a man went to school matters as much as how long he was in school. The poorest areas in the sugar and cotton growing areas of the rural Northeast –Maranhão, Piauí, Ceará, Pernambuco, and Bahia– are among the states with the lowest education quality in the early periods. The 1976 survey also shows a larger quality differential for rural areas in 15 out of the 18 states for which both rural and urban measures are available. Compared to a man educated in the city of São Paulo, a man who attained the same number of years of schooling and had the same years of labor market experience, but went to school in a rural municipality of São Paulo State earned 13% less. Bad as this is, comparable men who went to school in other rural areas of Brazil earned far less. From 29% less if they were educated in Sergipe state to 60% less in if they had the misfortune of being born in Ceará.

The silver lining is that differences in education quality appear to have fallen in many places over the last half century. In rural Maranhão the differential with São Paulo fell from 48 to 16%, in Piauí from 41% to statistically zero, in Ceará from 60 to 28%, in Paraiba from 39 to 9.1%, in Pernambuco from 47 to 21%, and so on for nearly every state that had sufficient data. Though there are fewer states with enough information to estimate differentials for people schooled in urban areas, most of the urban estimates show a similar pattern. The differential for urban Ceará fell from -60 to -41%, urban Pernambuco from -52 to -39%, Minas Gerais from -31 to -11%, Rio from -35 to -11, Paraná from -21 to zero, Rio Grade do Sul from -30 to -3.7, Goiás from -31 to -5.5, and the Federal District (Brasilia), from -28 to 3.1%. This is good news. But the inequalities in Brazil are still enormous.

This section developed a PNAD-based measure of education quality differentials that is based on the Brotherhood, Cavalcanti-Ferreira, and Santos (2019) method. After controlling for the labor market, earnings differences among otherwise identical men appear to be a function of where they went to school. The results suggest that school quality dispersion has fallen since 1976, but that large differences in school quality persist. Poor, rural areas continue to have the biggest quality deficits.

This method is an ingenious way of establishing that there are major differences in education quality, but it does not tell us *why* returns are different for different states. It also suffers from being a static measure. The school quality ascribed to a 65 year old man and a 25 year old man are forced to be equal if they studied in the same place –even though they studied 40 years apart. Also, the method is based on *relative* earnings

[1] The 1985 survey does not report place of birth or any other migration data.

–compared to urban São Paulo. We cannot rule out that relative performance in the rest of the country improved because school quality in São Paulo City got worse.

3.2 New Data on Local Education Expenditures as a Dynamic Measure of School Quality

Brazil has a long history of strong local government. The municipalities of the Federative Republic of Brazil roughly correspond to U.S. counties, but they are more powerful actors. They have an autonomous status within the constitution, a status legally equal to that of the states. In fact, every municipality not only has its own elected mayor it also has a legislative assembly. Municipalities have the power to "institute and collect taxes within their jurisdiction, as well as to apply their revenues" (Constitution, Article 30, III, author translation). One of their main responsibilities is to "maintain programs of pre-school and elementary education" (Article 30, VI). Municipalities are not entirely self-sufficient, by any means; federal and state governments transfers to the municipalities are substantial. Yet, significant dependence on local resources combined with the mandate to finance local schools has to lead to wide variations in the level of educational expenditures and therefore to a wide variation in the quality of education that children receive.

I introduce a new dataset on educational expenditures in this section and present it as a partial explanation for the observed large regional variations in the effects of Brazilian education. The data consist of a yearly municipal expenditures on education in Brazil for the period 1941 to 2004. These data were developed by the author and to his knowledge have never been used before. An important advantage of these data is that they provide a dynamic measure of school quality that varies with when as well as where people went to school.

I follow a well established tradition of using per capita education spending as a proxy for school quality (for instance Altonji and Dunn [1996], and Nicoletti and Birgitta [2018]). I am able to match people to municipality school spending at the time when each PNAD respondent was in school and in the place where he went to school. If he is a non-migrant, or if he migrated to his current place of residence before the age of 11, I assume that he was educated where he currently lives. If he migrated to his current location after age 10, he is categorized as educated in his place of origin. My measure of school quality is a 3-year moving average of per capita municipal education spending centered on the year when respondents turned 10.

The education expenditure data summarized in Table (3) come mainly from Finanças do Brasil (FoB), an irregular series of limited-run publications put out by the Brazilian Ministry of Finance (Ministério da Fazenda). I mostly hand-collected these in Rio de Janeiro, São Paulo, and Brasilia over a number of years and then transcribed them into electronic format using the double-entry method to avoid entry errors.[2] The data are adjusted for currency reforms (there were 8 of these between 1942 and 1994) and corrected for inflation using Brazil's Institute of Applied Economic Research deflator (IPEA), setting June [1979=100. The data are then converted into per-capita municipal expenditures using yearly municipal population estimates based on IPEA population figures over the period.

[INSERT TABLE 3]

[2]The double entry method consists of entering the data twice, independently, and then subtracting one matrix from the other one. Non-zero entries in the resulting matrix are then corrected.

The PNAD sample frame is representative at the state level and can be further partitioned into representative rural/urban sub-samples within each state. PNAD respondent location is recorded down to the municipal level, as is place of birth. In order to merge municipal school expenditure and PNAD respondent data, I aggregated the FoB data to the urban/rural municipality level. For each of the 64 years between 1941 and 2004 there are therefore 2 observations on per capita school expenditures per state, one for the municipality of the capital city, and another for the remaining, rural municipalities of each state.

Summary statistics on per capita educational expenditures in Table (3) are reported in constant [1979 Cruzeiros, by state and by urban/rural area for the entire 64-year period. The number of observations varies across states because some states were unpopulated at the outset[3].

Three features stand out in the school expenditure data of Table (3). First, is the enormous cross-state variation in mean education spending. Even considering only the states that existed over the full 64 years, there is a 10-fold difference in local per capita school spending among the municipalities of state capitals. Constant currency figures range from from Cr$.37 in Santa Catarina to Cr$ 3.54 in Rio de Janeiro. In rural areas spending ranges from Cr$ 0.43 per capita in the poor Northeastern state of Maranhão to more than Cr$ 80 in the wealthy southeastern industrial state of São Paulo[4].

The second pattern that stands out is the high degree of variation over time. Finally, there is a large discrepancy between rural and urban per capita expenditures. Somewhat surprisingly, the discrepancy is in favor of rural areas. This undoubtedly reflects a higher per-pupil cost of building, staffing, equipping, and maintaining schools in sparsely populated areas rather than better schools. This serves as a warning to not interpret higher spending as indicating higher quality without accounting for population density. Summary statistics in Table (3) show enormous dispersion. In most states the standard deviation of per capita education spending was about 1.5 times the mean level.

In order to save space, I will illustrate the dispersion of Brazilian education quality by contrasting the cases of just two states: Bahia, in the Northeast and São Paulo in the Southeast. Salvador, the capital of Bahia, is one of the oldest cities in the Americas. It was the imperial capital of Brazil and the center of Brazil's colonial sugar and cotton slave economy (Graham [1968]). It is now one of the poorest states in the country. São Paulo stands in defining contrast. São Paulo was undeveloped and not feted at all in the early years of the Brazilian history. Yet São Paulo became the center of the country's coffee boom in the mid 1800's and later used coffee wealth to finance the continent's largest industrial center (Baer [2013]). It is now one of the wealthiest states in the Brazil.

In the years after World War II, education expenditures were higher in São Paulo state than they were in Bahia. This was especially true in São Paulo City. Figures (1) and (2) show municipal expenditures on education remained low in both capital cities between 1941 and the end of the military era in 1985. The military had other priorities than educating the populace. When education spending in Bahia and São Paulo appear on separate graphs their spending patterns look similar. But the discrepancy between the rich and the poor states are glaring when they appear together, as in Figure (3). Data for rural municipalities contrasted in Figure (4), shows that inequality of opportunity between rural areas of these two states was even greater.

Per capita education spending remained fairly constant in Brazil, but then exploded when 20 years of

[3] Acre, Amapá, Rondônia, Roraima, and Mato Grosso do Sul, did not exist in 1941, and the state of Tocantins was founded in 1988 and its capital was built in 1989.

[4] The USD/Cruzeiro exchange rate in 1979 was Cr$ 22.58 per US$ (Department of the Treasury [1979).

military dictatorship transitioned to democracy in the 1980's. When the military junta was finally ousted in 1985, the return to democratically determined public goods priorities apparently fueled a surge in education spending. This change was especially evident in rural municipalities.

3.3 The impact of education quality on income distribution

Welch (1966) drew attention to the role of school quality in determining earnings. Behrman and Birdsall ([1983] and [1985]) used the average educational attainment of teachers as a measure of school quality in Brazil and concluded that "The estimated internal social rate of return to quality is at least as large as is that to investment in quantity" (Behrman and Birdsall [1985], 1203). Card and Krueger (1992) established that school quality is nearly as important as years of school attainment in explaining earnings differentials in the United States too, and thirty years later, Jackson and Mackevicius (2024) showed that quality differentials continue to be major determinants of educational outcomes in the United States.

Related work has shown that education quality differentials are important determinants of income within other countries (Psacharopoulos and Velez [1993], Lee and Barro [2001], versus Barro and Lee [1996], Bedi and Edwards [2002], Brotherhood, Cavalcanti-Ferreira, and Santos [2019]) and that only by accounting for the measurement error induced by omitting educational quality we can reconcile empirical micro returns with macroeconomic growth estimates (Krueger and Lindahl [2001]) and make sense of the link between years of school attainment and income differences between countries (Hanushek and Woessmann [2012]).

The Brazilian economics literature is well aware of education quality inequities in that country. For instance, Correia de Souza and Menezes Filho (2020) estimate a frontier production model and find that states that are farthest from the frontier grow more when they invest more in basic education, while states that are near the frontier do better when they increase spending at higher levels of schooling. Nakabashi and Salvato (2007) develop a state-level school quality index based on the proportion of teachers with a college degree, student pass rates, and classroom crowding large. They find that the large differences in education quality across Brazilian states significantly explain the disparity in income levels across states. In summary, the literature shows that years of schooling at best provides a rough approximation to measuring the skills that determine income. We need to account for quality.

I have shown that returns to schooling vary widely by *where* schooling took place and I have shown that local school expenditures in Brazil have varied wildly over time and place. I have not yet shown that these expenditures affect education quality. In particular, I have not shown that school funding differences are reflected in the earnings of adults, many years later.

In order to do so, consider the traditional Mincer equation, modified to incorporate school quality as in Equation (2).

$$y_i = \mu_0 + \mu_v v_i + \mu_{v2} v_i^2 + \mu_s S_i + \mu_{m_i} q_{m_i} S_i + \epsilon_i, \tag{2}$$

Here earnings y_i, depend on her labor market experience v_i, on how much time she spent in school S_i, and on local education quality q_{m_i}, augmented by her time in school. If we define age cohorts narrowly enough, the terms that account for potential labor market experience are subsumed in the intercept. Income differences within a given cohort are then a function of school quality and years of schooling and income inequality is a function of disparity in the quality of schooling and inequality in the distribution of school attainment.

$$\mathrm{Var}(y) = \beta^2\mathrm{Var}(S) + \gamma^2\mathrm{Var}(qS) + 2\beta\gamma\mathrm{Cov}(S, qS) + \Theta. \tag{3}$$

The first term in Equation (3) is the portion of income inequality that is explained by differences in school attainment. It consists of the Mincerian return to schooling β, squared, times inequality in the distribution of school attainment. This first term was reported in Lam and Levinson (1992). The second and third terms are new to the literature. They make up the portion of earnings inequality that comes from differences in the quality of education that current workers received. This quality effect has two parts. Term 2 expresses the the direct inequality-inducing effect of disparity in quality, magnified by the rate of return to quality, γ, squared. The third term accounts for the covariance between school quality and school quantity. This is likely to be positive if economic actors and their parents are rational in choosing a person's optimal school attainment. The last term, Θ, comprises the unexplained portion of earnings variations.

I matched school spending to the location where respondents lived at the time when they were 10 years old, in each of the 1976 [1995 [2005, and 2015 PNAD surveys. I then estimated modified Mincer regressions for 3-year cohorts of FTE men between the ages of 21 and 45, the common age range made possible by the 1941 to 2004 school spending data.[5] Full estimation results are presented in Table (4). In all of the estimates for all of the cohorts and in all four samples, the Mincerian "rate of return" to schooling is positive and significant at the 99% level, ranging from 14 to 16% in 1976 and then falling to 8-9% in 2015. Most importantly, the school quality measure is also positive and strongly significant in all of the samples for all cohorts.[6] The rate of return to quality ranges from 3.4% per Cr$1,000, **per year of schooling completed** in 1976 and falls to 0.009 to 0.02% per year of schooling in 2015. These results clearly establish that local expenditures on education affect school quality in a way that later becomes reflected in the earnings of prime-aged adult men. Since this group has nearly complete LFP, selection bias will be low (Heckman [1974, Mulligan and Rubinstein [2008).

Table (5) reports estimates of all 3 summands of Equation (3). The first result that stands out is the monotonic decline in earnings inequality. Within each PNAD panel in Column (2) shows a clear decline in earnings inequality from the oldest to the youngest cohorts. This finding is reinforced by picking an age group and then following it across the four panels. Brazil experienced a major drop in labor earnings inequality among prime aged males over the past half century.

Column (3) shows that the falling rate of return to schooling was critical in the decline of Brazilian income income inequality. Between 1976 and 2015 Mincer's β fell by about half. Inequality in years of school attainment also fell as overall educational attainment rose, leading the full effect of attainment dispersion –the first term in Equation (3)– to fall by more than two-thirds (Column (5)). Why the Mincer β fell is a bit beyond the scope of this paper, but it is tempting to surmise that when school coverage expanded (more people in school) and school attainment rose (more years of school per person) as documented in Table (2), returns to education were depressed by saturating the market, and/or by pulling underprivileged students into a system that offered them lower quality schools. Note that though the rate of return to quality, γ, fell, the portion of income inequality explained directly by school quality in Column (6) and the portion explained by the interaction between quantity and quality in Column (7) remained surprisingly constant over the last half century.

[5] A 45 year old man in the 1976 survey was 10 years old in 1941, the first year for which I have school spending data. A 21 year old man in the 2015 survey would have been 10 years old in 2004, the last year of school spending data.

[6] Estimates for the full 25 to 45 year old samples appear in Table (A2).

These results suggest that Brazil's period of educational expansion reproduced an educational system that consisted of schools of vastly different quality. Brazilian human capital differentials account for roughly $\frac{1}{3}$ to $\frac{1}{2}$ of income inequality among Brazil's main labor income earners (Column (8)). Between 5 and 19% of this inequality came from differences the quality of education afforded to its workers (Column (9)).

This section has shown the importance of education quantity and quality differences in explaining income inequality and validated the school spending series as a measure of human capital inputs. Since there is much more variation across time and geography in school spending than in years of schooling attained, the spending series provides a much more granular measure of human capital differences.

4 Economic Development and Women's Labor Force Participation Decision

A particularly important shortcoming of the literature is that empirical tests of women's adaptation to economic development are scarce. In fact the bulk of empirical evidence on economic development, labor force participation, and earnings inequality is based on the employment and earnings of men.

Brazil is certainly no exception to this pattern, and gendered roles remain strong in Brazil (Sotomayor [2009], Agenor and Canuto [2015], Berniel et al. [2023]). But the PNAD survey series shows that women's roles are changing rapidly. Table (6) is based on the sub-sample of households where a prime-aged head was cohabiting with a spouse of the opposite sex. It reports headship and labor force participation statistics by gender from 1976 to 2015. When PNAD surveyors enter a home, they are instructed to ask, "*What is the name of the person who is (mainly) responsible for this home*?"[7] In the first year, *all* of the households identified the man as the head of household. There was no exception. But this proportion declined rapidly over the next half century. By 2015, 25% of the households pointed to the woman as head. Table (6) also shows significant changes in women's relation to the labor force. In 1976 only 17.1% worked full time, starkly contrasting the vast majority of men who did so (94.7%). Full time employment (FTE) —defined as working at least 35 hours per week— then declined steadily among men, dropping to 80.3% in 2015. FTE moved in the opposite direction among women. Though the majority of women are still mainly engaged in domestic production, the proportion of FTE women more than doubled, from 17.1% in 1976 to 39.3% in 2015.

Mulligan and Rubinstein (2008) show that a switch from negative to positive selection into the labor force was induced by within-gender wage inequality and that this switch partly explains FLFP growth in the United States. This paper will show enormous differences in within-gender inequality in Brazil wrought by differences in education quality. And, even though full Heckmanian FLFP selection did not change signs in Brazil, the effect of some of the drivers of that selection –like husband's income and human capital itself– *did* change signs. Most importantly, human capital differences induced by the level and distribution of education quality played an important role in the evolution of women's relation to formal labor markets.

[7] Author translation of PNAD [2005], p. 94.

4.1 The FLFP decision

I frame analysis in this section with a stylized model of work decisions that largely fits the gendered nature of developing country labor forces. A country has 2N potential labor force participants, N of whom are men and N women. Every person is endowed with one unit of labor and with a pre-determined amount of human capital. Human capital endowments vary across people in a manner that will be specified below. Utility for person " i" depends on consumption of a home-produced good d and on consumption of a good q that must be purchased on the market. All men work full-time for a wage per unit of time. Men's wages depend on their human capital endowment.

Whereas men work full time for a wage and do not produce d. Women *must* produce d, however they can also choose to allocate some of their time to remunerated labor in the formal labor market in order to augment the consumption of q. Thomas (1990) rejects the unitary or dictatorial household welfare function for Brazil and finds that women and men largely determine the uses of their own resources, while taking contributions to the household by other members as given. In keeping with these findings, I assume that women maximize

$$u\left(d_i, q_i\right) = d_i^{\beta} q_i^{(1-\beta)}, \tag{4}$$

subject to $d_i = 1 - z_i \; and \; q_i = x_i + wh_i z_i. \; Here,$ d_i is time devoted to home production, time that is transformed one-to-one into the home-good. In the utility function, $0 \leq \beta \leq 1$ is the Cobb-Douglas preference parameter, z_i is time spent working in the formal labor market, x_i is the (partially-exogenous) spousal income, h_i is her human capital, and w is the market remuneration per unit of female human capital. As in Heckman (1974), when the solution is interior, her labor supply $z_i^{\text{sup}} > 0$ is observed

$$z_i^{\text{sup}} = \; 1 - \beta\left(1 + \frac{x_i}{wh_i}\right). \tag{5}$$

Given the wage rate and utility function parameter, formal labor supply increases with her human capital and falls with spousal income. Setting $z_i^{\text{sup}} = 0$ in Equation (6) defines the greatest lower bound of the set of corner solutions in (h, x) space:

$$x_i^* = \left(\frac{1-\beta}{\beta}\right) wh_i. \tag{6}$$

For values $x_i > x_i^*$, demand for the home-produced good exceeds the woman's labor endowment. Women with spousal incomes greater than or equal to x^* choose to withdraw from the labor market. It follows that if d is normal, then for any h_i^* there exists a x_i^* such that z_i^*=0.

4.2 Human Capital Production and the Potential Wage

People are naturally endowed with one unit of human capital that is enhanced by time spent in formal schooling (quantity), by the quality of that schooling, by personal effort, and labor market experience. Schultz (1961), Mincer (1974) and Becker (1975) argue that schooling increases human capital and that human capital in turn increases labor market earnings, providing the theoretical foundation for Equations (1) and (2). Early empirical literature focused on the role of years of schooling and labor market experience (Mincer [1974], Lam and Levinson [1992], Psacharopoulos [1993]), while other writers have highlighted the importance of school quality in explaining wages (Welch [1966], Behrman and Birdsall [1983], Card and Krueger [1992], Altonji and Dunn [1996], Bedi and Edwards [2002]) and the most recent vein of this literature emphasizes the role of uneven school quality in creating income inequality (Hanushek and Woessman [2012], Brotherhood, Cavalcanti-Ferreira, and Santos [2019]).

In order to incorporate quality empirically, I assume that the human capital production function has the following form

$$h_i = e^{\gamma_e e_i + \gamma_v v_i + \gamma_{vv} v_i^2 + \gamma_s S_i + \gamma_m q_{m_i} S_i}. \tag{7}$$

The human capital of person *i,* who was educated in location m_i , depends on her grit –the effort e_i she exerts in attaining human capital, on her labor market experience v_i, on how much time she spent in school S_i, and on local education quality q_{m_i}, augmented by her time in school. The γ_j are human capital production function parameters, for j∈ {e, m, v, vv, s}. While many other forms are possible (see the discussion in Bedi & Edwards [2002), Equation (7) is simple and facilitates comparison with earlier literature. It incorporates interactions between school quantity and quality, and automatically sets the impact of school quality on human capital to zero if a person never attended school (Welch [1966], Behrman and Birdsall [1983], Card and Krueger [1992], Bedi and Edwards [2002]). That is, q_{m_i}affects the slope of h_i in (S, h) space, but not its intercept.

Men

Labor allocation is gendered; men devote their labor endowment to wage work. Researchers observe local education quality, q_{m_i}, labor market experience v_i, years of school attendance S_i, and a man's income x_i as the gross returns to his human capital: $x_i = wh_i \times 1$. Given Equation (7), the economic model of men's income can be written as

$$y_i := \text{Log}\,(x_i) = \text{Log}(w) + \gamma_e e_i + \gamma_v v_i + \gamma_{v2} v_i^2 + \gamma_s S_i + \gamma_m q_{m_i} S_i. \tag{8}$$

Since e_i is unobservable, the econometric model for men becomes Equation (2):

$$y_i = \mu_0 + \mu_v v_i + \mu_{v2} v_i^2 + \mu_s S_i + \mu_{m_i} q_{m_i} S_i + \epsilon_i. \tag{9}$$

The error term ϵ_i incorporates effort and other unobservable determinants of income.

Women

Women are responsible for producing the domestic good, but they may also choose to allocate some of their time to wage labor. The alternative use of their time in the production of the home good d, is both essential (in the utility sense) and not marketed. The labor market earnings of woman "i", $\boldsymbol{\psi}_i$, have an economic representation similar to men's, however they face a different labor market than men because of gender discrimination, a higher compensating differential for schedule complementarity with home production (Hall and Mueller [2018], Morchio and Moser [2023]), or non-convex rewards to the length of the work day (Goldin [2014]). This is captured by δ >1 in Equation (10), where women's compensation per unit of time is expressed as

$$\text{Log}\left(\boldsymbol{\psi}_i\right) = \text{Log}\left(\frac{w}{\delta}\right) + \gamma_e e_i + \gamma_v v_i + \gamma_{\text{v}2} v_i^2 \; + \gamma_s S_i + \gamma_m q_{m_i} S_i. \tag{10}$$

4.3 The effect on FLFP of correlation between spousal earnings potentials: A Correlated Spousal Earnings Model

Equations (5) and (6) showed that women supply labor by comparing their potential market earnings to the income of their spouse. This introduces a selectivity problem into the econometric specification of women's earnings. In the context of our data (Table 6), a realistic simplifcation is to assume that all men work at the highest-paying job they can find, but women choose to work only if $h_i > \left(\frac{\beta}{1-\beta}\right)\frac{x_i^*}{\boldsymbol{\psi}_i}$. Only women whose human capital is high in relation to their husband's income will be observed working. The selection bias problem is well known in FLFP econometrics. We cannot use information from working women to predict the earnings of women who do not work (Heckman [1974). We need a reasonable alternative estimate of women's potential earnings $\boldsymbol{\psi}_i$.

One method for getting around the selectivity problem is to use estimated coefficients for men from Equation (8) to predict the potential labor market earnings of women. An index that is highly correlated with women's human capital is

$$\overline{\boldsymbol{\psi}}_i := \hat{\mu_0} + 1(S)\hat{\mu}_{m_i} q_{m_i} + \hat{\mu}_v v_i + \mu_{\text{v}2} v_i^2 + \hat{\mu}_s S_i. \tag{11}$$

This index predicts her earnings as though she were a fully-employed, prime-aged man. $\overline{\boldsymbol{\psi}}_i$ overestimates her earnings by the constant Log(δ). However, this does not stand in the way of using $\overline{\boldsymbol{\psi}}_i$ in contexts where we care about the dispersion of women's earnings. I will use Equation (11) in some of the estimations as a measure of women's potential earnings.

Equation (5) also introduces a censorship problem into the empirical specification since demand for home goods cannot be directly observed. Whereas the time endowment is bounded $d \in$[0,1], demand for home goods is not. In particular, when the husband's income is sufficiently large in relation to a woman's own, it can be true that $d_i > 1$ for $\frac{x_i}{wh_i} > \frac{1-\beta}{\beta}$, even though we only observe $d = 1$.

A third econometric problem arises from the correlation between x_i^* and ψ_i. Assortative mating theory (Becker [1973) and recent empirical evidence (for instance Goldin [2014 and Bratsberg et al. [2023), suggest that x_i^* and ψ_i are jointly determined and that their co-evolution during the process of economic development is the major force behind the changing shape of gendered labor markets.

For the purpose of estimating FLFP, I assume the following correlated spousal earnings ("CSE") framework. Women draw husbands from a lognormal earnings distribution with median income $\tilde{\mu}_{xi}$ that depends on their own potential earnings ψ_i. Specifically, $\tilde{\mu}_{xi}\left(\psi_i\right) = \psi_i^{\alpha+1}$, with $\alpha > 0$. A woman knows the distribution of her husband's income at the time of marriage, including $\tilde{\mu}_{x\text{i}}$, but his actual income x_i^* is only revealed to her after marriage.

$$y_i = \ln\left(x_i\right), \text{ and y} \sim N\left(\psi_i^{\alpha+1}, \sigma_x^2\right). \tag{12}$$

From the definition of the lognormal density, we have that $\mu_y = \tilde{\mu}_x$. So, given the CDF of x, $F_x(x)$, the cumulative density function of d is given by, $F_d(d) = F_x\left(\frac{(d_i-\beta)}{\beta}\psi_i\right) = \Phi\left(\frac{\ln\left(\frac{(d_i-\beta)}{\beta}\psi_i\right)-\mu_y}{\sigma_y}\right) = \Phi\left(\frac{\ln\left(\frac{(d_i-\beta)}{\beta}\right)+\ln(\psi_i)-(1+\alpha)\ln(\psi)}{\sigma_y}\right) =$

$\Phi\left(\frac{\ln\left(\frac{(d_i-\beta)}{\beta}\right)-\alpha\ln(\psi)}{\sigma_y}\right)$.

Finally, we obtain the cummulative density function of d:

$$F_d(d)=\Phi\left(\frac{\ln\left(\frac{(d_i-\beta)}{\beta}\right)-\sum_{j=1}^{n}\omega_j v_j}{\sigma_y}\right) \tag{13}$$

and the pdf, as

$$f_d(d)=\frac{\partial F_d(d)}{\partial d}=\frac{1}{(d-\beta)}\phi\left(\frac{\ln\left(\frac{(d_i-\beta)}{\beta}\right)-\sum_{j=1}^{n}\omega_j v_j}{\sigma_y}\right). \tag{14}$$

5 THREE ESTIMATION METHODS

I employ three estimation methods, to examine the impact of school quality on FLFP in Brazil. The first is maximum likelihood estimation of domestic labor supply from the CSE model developed in the previous section. The second method is Heckman's 2-step selection model, and the third (in the Appendix) is a Tobit model. I find that in all three methods school quality variables contribute significantly to explaining women's labor allocation decisions.

5.1 Maximum Likelihood Estimation of the CSE model

All agents are endowed with one unit of time. So if woman i is observed providing λ_i units of market labor, we know that she is supplying $d_i = 1-\lambda_i$ of domestic labor. But if $\lambda_i = 0$, we only know that $d_i \geq 1$, meaning that such observations in the data are censored. Letting $c_i \in \{0,1\}$ indicate censorship, the likelihood function of observation i is then $l_i = f_d\left(d_i\right)^{(1-c_i)}\left(1-F_d\left(d_i\right)\right)^{c_i}$.

We can now define the likelihood function in terms of this pdf and the survival function $S_i^d(d){:=}1{-}F_i^d(d)$. Defining the log of the likelihood function for the i^{th} observation as, $\ell_i{:=}\ln\left(l_i\right)$, we obtain

$$\ell_i{=}(1-c_i)\ln\left(\frac{\phi\left(\ln\left(\frac{d_i-\beta}{\beta}\right)-\sum_{j=1}^{n}\omega_j v_j\right)}{(d-\beta)}\right)+c_i\ln\Phi\left(\frac{\ln\left(\frac{d_i-\beta}{\beta}\right)-\sum_{j=0}^{n}\omega_j v_j}{\sigma_y}\right), \tag{15}$$

which can be estimated with maximum likelihood methods.

5.2 Heckman 2-step selection model

The Heckman selection model (Heckman [1976) is a 2-step approach. It involves first predicting the instantaneous "hazard function", defined as the ratio of the pdf to the survival function (sometimes called the "inverse of the Mills ratio"). The predicted hazard is then inserted as an ancillary variable in an hours worked equation.

5.3 Tobit

Since hours worked presents a large number of zeros for women, a third reasonable method is to estimate hours worked equations using Tobin's (1956) limited dependent variable model. Observed market labor can be seen as the expression of a two-part decision: (i) work for a wage: yes, or no? Then if "yes", (ii) work for how many hours?

The researcher observes z_i^+, where

$z_i^+ = 0 \quad \textit{if} \quad z_i^{\sup} \leq 0$, and

$z_i^+ = z_i^{\sup}$ if $\quad z_i^{\sup} > 0$.

Or, equivalently

$z_i^+ = 0 \quad$ if $\quad h_i \leq \left(\frac{\beta}{1-\beta}\right)\frac{x_i^*}{\cdot}$, and

$z_i^+ = 1-\beta\left(1+\frac{x_i}{\psi_i}\right)$ if $h_i > \left(\frac{\beta}{1-\beta}\right)\frac{x_i^*}{\psi_i}$.

This two-part function is well-suited for Tobit estimation. I examine Tobin's method for comparison and robustness in the Appendix.

6 Human Capital FLFP Estimation

The previous section introduced a new dataset on school quality and established that these data are reflected in human capital, an important individual characteristic of working men that is rewarded by the marketplace. In this final analytical section I examine how human capital and its dispersion has affected the participation of women in the labor market.

I contrast three methods of estimating how women make their FLFP decisions. The first is a maximum likelihood estimation of the CSE model, the second is a replication of Heckman's 2-step selection model. For robustness, I present Tobit estimates in the Appendix and discuss them briefly here.

I estimated the CSE model using the human capital index of Equation (11). Then, for comparison, I re-estimated it using each of the human capital inputs that form the arguments of Equation (11), in raw form.

I follow a similar procedure for the Heckman and Tobit models. However, I first estimated them without controlling for school quality and then by controlling for it. The quality inclusive versions also first use the human capital index and then the raw human capital input variables. Each of the CSE, Heckman, and Tobit models is also estimated without including marital status, fertility, or husband income, and then again with controls for these pre-determined characteristics, as in Heckman (1974).

6.1 Maximum Likelihood Estimation of the CSE Model

Recalling Equation (15) from the CSE model, the log of the likelihood function for the i^{th} observation is

$$\ell_i = (1-c_i)\ln\left(\frac{\phi\left(\ln\left(\frac{d_i-\beta}{\beta}\right)-\sum_{j=1}^{n}\omega_j v_j\right)}{(d-\beta)}\right) + c_i \ln\Phi\left(\frac{\ln\left(\frac{d_i-\beta}{\beta}\right)-\sum_{j=0}^{n}\omega_j v_j}{\sigma_y}\right) \tag{16}$$

Human capital acquisition affects the allocation of women's time through two avenues in the model. Traditional income and substitution effects result from raising her wage rate. But when her earnings potential rises, it also affects the income density function of the pool of spouses she draws from. The estimating Equation (16) is a reduced form that captures the aggregate effects of her human capital endowment. In other words, an estimate could show a zero net effect if having more human capital made her want to work more, but this was offset by having a husband who earns more and the normality of demand for domestic goods. Similarly, when raw human capital inputs are used to estimate her time allocation, in place of the index, the estimated coefficients on schooling and school quality capture the net effect of the combined and likely opposite impacts of these human capital inputs on her labor force behavior.

Time allocation information in the PNAD surveys is weekly. In the empirical implementation of the model, I therefore assume that women have 168 hours (24x7) to allocate between home production and market work. The dependent variable in the estimates is then normalized by 168, i.e., it measures the percentage of time spent on domestic or market activities.

The human capital index incorporates all of the determinants of her human capital. Also, the reduced form statistical structure of the CSE model combines the effects of her human capital on her direct incentives to work as well as the indirect and likely opposite effect that her higher earnings potential leads to a spouse who is drawn from a stochastically dominant earnings density. Higher mean spousal earnings will increase demand for home produced goods if these are normal goods, and therefore also tend to increase the time she spends in domestic production, as derived in Equation (5).

Estimates of the impact of a woman's human capital on the proportion of her time spent in domestic labor are presented for each survey year in the columns marked '(1)' in Table (7). The human capital index has a strong, negative, and significant impact on the proportion of time that women spend on domestic production in all of the years studied. This index has a further advantage which will now become apparent. Recall the human capital index measures the logarithm of the wage that a woman would earn if she were a prime-aged male –call this her 'male wage'. Since the dependent variable is the percentage of weely time devoted to home production, this leads to a straight-forward interpretation of the coefficients on Hk as elasticities. A one percent increase in a woman's male wage in 1976 led to an mean 0.207 decline in the proportion of time she devoted to domestic tasks, to a 0.129 fall in 1995, 0.179 in 2005, and a 0.234 decline in 2015.

Studies of FLFP customarily also include marital status and the number of young children in the home (Heckman [1976) as predetermined variables. Columns marked '(2)' in Table (7) include controls for her marital status and for the number of young children (below age 7) in her home. These variables have the expected sign in all four samples —inducing her to spend more time on domestic work— but they do not have a noteworthy impact on the human capital coefficient.

In sum, the CSE estimates show that women with more human capital have always spent less time at home –regardless of their marital status or the number of young children they have.[8]

The Hk index been been shown to be a significant determinant of income and of FLFP. As a robustness check, Table (8) eschews the index and examines the roles of school attainment, school quality, and other variables that contribute to human capital formation, in raw form. According to these maximum likelihood estimates, one additional year of school attainment is associated with an average decrease in the proportion of time spent in domestic production, ranging from .11% in 1995 to .213% in 2015. Table (8) indicates that attending a higher quality school as a child reduces the time adult women spend on domestic work and conversely increases the time they spend working for a wage. This is a very significant effect given the extremely wide dispersion in spending across Brazil. In the estimating sample, women who had some formal education experienced local per capita school spending rates that ranged from USD$0.003 to USD$236.4.

[INSERT TABLES 7 and 8 ABOUT HERE]

Maximum likelihood estimates in Table (8) are consistent with theoretical predictions. More years of formal schooling reduce time spent at home and all of the coefficients on school quality are statistically

[8]Heckman also included husband earnings as a predetermined variable. I do not do so here because the essence of the CSE model is the joint determination of the husband and wife's earnings potentials. Husbands earnings are included in replications of the Heckman and Tobit models, below.

significant and negative, as is consistent with theoretical predictions. Also, in every year hours spent in domestic production are higher for married women and they are an increasing function of the number of young children in the home. A major contribution of this paper is finding that higher school quality leads to women working fewer hours in domestic production, and to more market work. The rapid rise in per capita expenditures on education in Brazil over the past 80 years has led to a decrease in the marginal effects of additional expenditures.

6.2 Heckman Selection

I now turn to alternative, well-established estimation methods to see if results from the CSE model are maintained. Whereas the reduced form CSE method estimated the determinants of time spent producing the domestic good, the Heckman (1976) model is typically used to estimate the supply of women's labor to the formal wage labor market. The FLFP decision lends itself to two-step estimation procedures, and indeed has been one of the main reasons for developing these methods. The Heckman model has been widely employed and is well-vetted. I therefore begin with a replication of his results with Brazilian data. In keeping with Heckman's original specification, I omit quality measures and treat the number of hours worked by a woman as the result of two, sequential decisions. In the first stage a zero-one labor force participation indicator is regressed on women's years of schooling, potential job experience, a marital status indicator, the number of young children that she has, and her husband's income. In the second stage, hours worked are regressed on years of schooling, potential job experience and the hazard function value ("inverse Mills ratio") from the first stage.

Results in Table (9) validate the Heckman model yet again. First stage estimates of all the samples indicate that women with more years of schooling and more potential labor market experience are more likely to be observed working, whereas being married and having more children below age 7 reduce this probability. The husband's earnings have a U-shaped effect, negative in 1976 and 1995, and positive thereafter. In the top panel, second stage estimates reveal an interesting change that took place about half-way through the period under study. Conditional on having chosen to work, women with more years of schooling worked for fewer hours in 1976, and 1995, but by 2005 the effect was opposite. The impact of higher school attainment was still strongly significant, but it is now measured to be **positive**.

[INSERT TABLE 9]

Table (10) incorporates the human capital index, Hk, as an instrument of selection and determinant of hours worked. Hk includes local municipal education spending as a dynamic measure, adjusted to the place and time when each person was in school. Since it also accounts for years of schooling and labor market experience, these variables are not repeated in specifications which employ the index.

Table (10) shows that results from the original specification are robust to the inclusion of the index which adjusts human capital for school quality. The "flavor" of first stage estimates in Table (9) is largely unaffected. Coefficient estimates also result in an ambiguous human capital effect. Hk increases the selection of women into formal work but, conditional on working, Hk is associated with fewer hours of work in the first two samples and has no significant effect in the last two. The effect of husband's income is also ambiguous. It also changes signs. It has a negative and significant effect on selection in the first two samples, but a positive and significant one in the later samples. The fact that a woman's human capital and the income of her husband are both ambiguous determinants of FLFP is probably not a coincidence. Higher human capital women have husbands that are drawn from dominant earnings densities, leading to confounding effects when they are treated as independent.

When the raw school quality measures are incorporated into the Heckman model, as in Table (11), years of schooling attained and husbands' earnings maintain their significant, but ambiguous, sign-changing effects. School quality adds a new dimension to the observed differences in women's labor market participation decisions. School quality has a mixed, though largely negative effect of its own on women's decision to work, and a positive and strongly significant effect on the intensity of work among working women.

[INSERT TABLE 10]

Sign changes in important estimates are disconcerting and raise the question of whether they are due to a major behavioral change, or whether the model is improperly specified. Binscatter plots provide an insightful and econometrically useful simplification of empirical relationships under study (Cataneo et al [2024). Data for each sample were sorted by Hk into 20 bins. The points in each graph are the mean of a 0/1 work indicator. Binscatter plots of the empirical relationship between human capital and the probability that women work for a wage in Figure (5) are clearly positive and monotonic and close to linear in all 4 samples –unlike first stage Heckman estimates.

Figure (6) shows a similarly derived relation between Hk and the number of hours worked by women who work. This relation is a more complex and clearly *non*-linear. Furthermore, the inverse-U shape in the binscatter plot is clear and leads to a sign flip in the fitted trend when it is forced to be linear. At low levels of human capital, hours worked increase as Hk rises, but the relation becomes negative near the median. If high human capital women work, they work for *fewer* hours. When the relationship is forced to be linear, the estimated slope flips from negative in 1976 to positive, but decreasing in the subsequent years. Contrary to what is implied by the solid trend lines, the empirical relation between Hk and hours worked is actually maintained through more than a half century of data. Change was not dramatic or sudden, but subtle. The inverse-U shape shifted to the right over time and became flatter at the top. An overall simplification of these changes would be to say that at the lowest human capital levels, women stay home. At the median level, most women work part time, and at low levels, having relatively more Hk leads women to work more hours for pay. The proportion who work continues to rise with human capital, but once median human capital levels are reached, the pattern changes. At lower levels of development, women with above average human capital start pulling back from remunerated labor. This is probably because they have been able to marry high income earning men. At higher levels of development in later years, high human capital women tend to work full-time, just like their husbands –regardless of their husbands' incomes.

The problem with the traditional Heckman model specification is that resulting estimates do not reflect these clear data patterns in a straightforward way. This is obvious in the binscatter plots of Figures (5) and (6). The discrepancy between linearly restricted estimates and empirical patterns appears to be especially problematic for hours worked. When models force trend lines to be linear, we are forced to choose whether it is best to say that the relation between Hk and hours worked is negative, positive, or zero-sloped –even though the truth is that all of these relations are true over some range. The slope in (Hk, Hours) space depends on *where* in the Hk distribution women are.

6.3 Tobit Estimates

As a robustness check, I include Tobit model estimates in the Appendix and a brief discussion of the results here. The left-censoring of the hours worked decision generates data with zero hours worked for women at the corner as in Equation (6), but also for women whose demand for domestic goods exceeds their labor endowment, generating "negative" market labor supply. The high frequency of zeros for FLFP in Table (1) suggests using the Tobit method. Table (A3) reports basic Tobit model estimates of hours worked for women between the ages of 20 and 65 who are either the head or the spouse of the head of their household. Explanatory variables include years of schooling but no school quality controls —implicitly treating all schools as equally effective in creating earnings potential. Having more years of schooling is associated with more hours worked in the formal labor market and all the other variables have the expected effects, except for husband income, which changes signs across samples, as it did in the Heckman estimates. Finally, Table (A4) reports Tobit estimates that employ each woman's human capital index, measured as the wage she could earn *ceteris paribus* if she were a man —*ceteris paribus*. Following Heckman (1976), all three tables include controls for pre-determined marital status, number of children below age 6, and husband's income.

[INSERT TABLE A3]

Estimates in Table (A4) reveal that controlling for more nuanced differences in human capital with the Hk index is used instead of years of schooling mainly affects the estimated role of husband's income. The sign flip happens earlier and the positive/negative difference in slopes becomes more significant. Once again, this suggests that there is a relation between women's human capital and the income of their spouses that is not well specified.

[INSERT TABLE A4]

When education quality measures are incorporated directly, in place of Hk (Table (A5)), the basic estimates change very little, but *both* the effect of quality and the effect of higher spousal income flip signs.

[INSERT TABLE A5]

7 Conclusion

I have introduced a new, granular dataset on education quality differentials that spans more than sixty years. I used the data to examine how human capital formation affects income inequality among men and the labor force participation of women during the process of economic development. Differences in education quality are a significant factor explaining income inequality among fully employed men and an important determinant of the labor force participation of women.

The literature has rationed empirical evidence of how women's labor force participation changes with development. The empirical evidence introduced in this paper adds a new dose and the CSE model adds a new way of conceptualizing the net effect of a woman's human capital on her participation in the wage labor market by allowing her husbands' human capital to be correlated with their own. The analysis of much more granular variations in human capital made possible by these data show that the net effect of human capital accumulation on FLFP has been unambiguously positive for a very long time.

This paper also makes a contribution to understanding the Brazilian development experience in particular. Hanushek and Woessmann (2012) posited a Latina American growth puzzle: why has Latin American growth been sluggish, despite its rich resource base and comparatively high school attainment? True to the pattern evinced by the Latin America region as a whole, PNAD data show that Brazil made enormous gains in school coverage and years of attainment during the past half-century. But unfortunately, the school spending data introduced here show that not all educational gains were created equally. Differences across space and time in education spending have been enormous. The differential education quality children were exposed to grew up to become a very important determinant of returns to education, of labor productivity, and of incomes.

The finding that school quality differences are also crucial determinants of how women allocate their time is new to the literature. School quality differences are a major factor in the labor force participation decisions of women. Given the number of years of school attained by a woman, the quality of the school she attended plays a very strong role of its own. It affects the opportunity cost of her time, and if she marries, it also affects the earnings of her spouse. The changing interplay between a woman's human capital and the earnings of her spouse have played a crucial role in women's labor force participation decisions and in how those decisions have evolved. The FLFP model states that the nature of the interplay is economic: in traditional societies what matters is not the years of schooling she has *per se*, but the relation between how much she can earn if she works and how much her husband earns. The data introduced here show that school quality has played and continues to play a major role in this relation. If we are trying to understand how formal education affects women's relation to the labor force, we must also take into account the fact that her human capital does not only affect her earnings potential, it also determines the earnings potential of the husband she matches with.

Above all, the take-away from this paper is that differences in the quality of schools we provide to children lead to differences in how they live their lives as adults.

8 References

Agenor, Pierre-Richard and Otavio Canuto. 2015. Gender Equality and Economic Growth in Brazil: A Long-Run Analysis. Journal of Macroeconomics. 43:155–172

Alesina, Alberto, Paola Giuliano, and Nathan Nunn. 2013. On the Origins of Gender Roles: Women and the Plough. The Quarterly Journal of Economics. 128 (2): 469−530, https://doi.org/10.1093/qje/qjt005

Almeida dos Reis, J. G. and R. Paes de Barros. 1991. Wage Inequality and the Distribution of Education: A study of the evolution of regional differences in inequality in metropolitan Brazil. Journal of Development Economics. 36 (1): 117-143.

Altonji, Joseph G. 1993. The Demand for and Return to Education When Education Outcomes are Uncertain. Journal of Labor Economics, 11(1), 48–83. http://www.jstor.org/stable/2535184

Altonji Joseph G., and Rebecca M. Blank. 1999. Chapter 48 Race and gender in the labor market, Handbook of Labor Economics, Elsevier, Volume 3, Part C [1999, 3143-3259.

Altonji, Joseph G., and Thomas A. Dunn. 1996. Using Siblings to Estimate the Effect of School Quality on Wages. The Review of Economics and Statistics, 78(4), 665-671. https://doi.org/10.2307/2109953

Anand, & Kanbur, S. M. R. 1993. The Kuznets process and the inequality−−development relationship. Journal of Development Economics, 40(1), 25-52.

Autor, David H., and David Dorn. 2013. "The Growth of Low-Skill Service Jobs and the Polarization of the US Labor Market." American Economic Review, 103 (5): 1553-97.

Baer, Werner. 2013. The Brazilian Economy: Growth and Development. Lynne Rienner Publishers, Inc.; 7th edition

Barro, Robert J. and Jong Wha Lee. 1996. International Measures of Schooling Years and Schooling Quality. The American Economic Review. 86 (2): 218-223.

Barros, Ricardo, Mirela de Carvalho, Samuel Franco, & Rosane Mendonca. 2010. Markets, the State, and the Dynamics of Inequality in Brazil. In L. F. López-Calva & N. Lustig (Eds.), Declining Inequality in Latin America: A Decade of Progress? (pp. 134-174). Brookings Institution Press. http://www.jstor.org/stable/10.7864/j.ctt6wpdl

Becker, Gary S. 1964. Human Capital: A Theoretical and Empirical Analysis, with Special Reference to Education. National Bureau of Economic Research, New York. Distributed by Columbia University Press.

——————. 1965. A Theory of the Allocation of Time. Economic Journal 75 (3): 493-517.

——————. 1973. A Theory of Marriage: Part I. Journal of Political Economy, 81 (4): 813-846.

——————. 1975. Human Capital. 2nd ed. NBER, New York, NY.

Bedi, Arjun S. and John H. Y. Edwards. 2002. The Impact of School Quality on Earnings and Educational Returns: Evidence from a Low-Income Country. Journal of Development Economics. 68: 157-185.

Berniell, Inés , Leonardo Gasparini, Mariana Marchionni, and Mariana Viollaz. 2023. Lucky women in unlucky cohorts: Gender differences in the effects of initial labor market conditions in Latin America. Journal of Development Economics,161: 1-8.

Behrman, J.R., and Nancy Birdsall. 1983. The Quality of Schooling: Quantity Alone is Misleading. American Economic Review 73 (5): 928-946.

Birdsall, Nancy, David Ross, and Richard Sabot. 2001. Education, Growth and Inequality, in Pathways to Growth: Comparing East Asia and Latin America, Nancy Birsall, ed. Washington, D.C. : Published by Inter-American Development Bank ; Baltimore, MD : Distributed by The Johns Hopkins University Press

Bratsberg, Bernt, Simen Markussen, Oddbjørn Raaum, Knut Røed, and Ole Røgeberg. 2023. Trends in Assortative Mating and Offspring Outcomes. The Economic Journal, 133(651): 928-950

Bridgman, Benjamin, Georg Duernecker, and Berthold Herrendorf. 2018. Structural Transformation, Marketization, and Household Production Around the World. Journal of Development Economics. 133: 102–126

Bourguignon, F. & C. Morrisson. 2002. Inequality among world citizens: 1820-1992. American Economic Review, 92(4): 727-744.

Brotherhood, Luiz Mário, Pedro Cavalcanti Ferreira, and Cezar Santos. 2019. Education Quality and Returns to Schooling: Evidence from Migrants in Brazil. Economic Development and Cultural Change 2019 67:3, 439-459.

Card, D., Krueger, A. 1992. Does School Quality Matter? Returns to Education and the Characteristics of Public Schools in the United States. Journal of Political Economy 100 (1): 1 -40.

Cataneo, Matias D., Richard K. Crump, Max H. Farrell, and Yingjie Feng. 2024. On Binscatter. American Economic Review. 114(5): 1488–1514. https://doi.org/10.1257/aer.20221576

Nicoletti, Cheti, and Birgitta Rabe. 2018. The Effect of School Spending on Student Achievement: Addressing Biases in Value-Added Models. Journal of the Royal Statistical Society Series A: Statistics in Society. 181(2):487-515, https://doi.org/10.1111/rssa.12304

Correia de Souza, Eduardo and Naercio Menezes Filho. 2020. O impacto do gasto público e da qualidade por nível de ensino sobre o crescimento da renda dos estados brasileiros. Pesquisa e Planejamento Econômico. 50 (2): 31-56.

Cortés, Patricia. 2023. Immigration, household production, and native women's labor market outcomes: A survey of a global phenomenon. European Economic Review, 158.

Danzer, Alexander M., and Lennard Zyska. 2023. "Pensions and Fertility: Microeconomic Evidence." American Economic Journal: Economic Policy, 15 (2): 126-65.

Duryea, Suzanne, Rafael P. Ribas, Breno Sampaio, Gustavo R. Sampaio, Giuseppe Trevisan. 2023. Who benefits from tuition-free, top-quality universities? Evidence from Brazil. Economics of Education Review, 95: 1-14.

Edwards, John H. Y. 2019. Maids, female labor force participation, and economic development in Latin America. Paper presented at the 2019 Latin American Econometric Society meeting, Puebla, Mexico.

Edwards, Sebastián. 2009. Latin America's Decline: A Long Historical View. NBER Working Paper # 15171, https://www.nber.org/papers/w15171

Ministerio da Fazenda. Finanças do Brasil (various issues). Ministerio da Fazenda, Secretaria do Tesouro Nacional. Brasilia, DF.

Engbom, Niklas and Christian Moser. 2022. Earnings Inequality and the Minimum Wage: Evidence from Brazil. American Economic Review. 112(12): 3803–3847.

Fernandez, Manuel and Julián Messina. 2018. Skill premium, labor supply, and changes in the structure of wages in Latin America. Journal of Development Economics. 135: 555–573

Engels, Friederich. 1884. Der Ursprung der Familie, des Privateigenthums und des Staats. Stuttgart, I.H.W. Dietz.

Goldin, Claudia. 1995. The U-Shaped Female Labor Force Function in Economic Development and Economic History, in T. Paul Schultz, ed. Investment in Women's Human Capital. Chicago, University of Chicago Press.

———. 2014. A Grand Gender Convergence: Its Last Chapter. American Economic Review, 104 (4): 1091-1119.

Graham, Richard. 1968. Britain and the Onset of Modernization in Brazil 1850-1914 (Vol. 4). Cambridge University Press.

Hall, Robert, and Andreas I. Mueller. 2018. Wage Dispersion and Search Behavior: The Importance of Nonwage Job Values. Journal of Political Economy. 126(4): 1594-1637.

Hanushek, Eric A & Ludger Woessmann. 2012. Schooling, educational achievement, and the Latin American growth puzzle. Journal of Development Economics. 99, (2): 497-512,

Heckman, James, 1974. Shadow prices, market wages and labor supply. Econometrica 42, 679-694.

———. 1976. The common structure of statistical models of truncation, sample selection and limited dependent variables and a simple estimator for such models. Annals of Economic and Social Measurement 5: 475-492.

———. 1990. Varieties of Selection Bias. American Economic Review Papers and Proceedings 80: 313-318.

Humphries, John Eric, Juanna Schrøter Joensen, and Gregory F. Veramendi. 2024. "The Gender Wage Gap: Skills, Sorting, and Returns." AEA Papers and Proceedings, 114: 259-264.

IBGE. 2018. Pesquisa Nacional por Amostra de Domicílios Contínua Divulgação Especial Mulheres no Mercado de Trabalho. https://biblioteca.ibge.gov.br/index.php/biblioteca-catalogo?view=detalhesid=2101641.

Jackson, C.K. and Mackevicius, C.L. 2024. What Impacts Can We Expect from School Spending Policy? Evidence from Evaluations in the United States. American Economic Journal: Applied Economics. 16(1), pp.412-446.

Klasen, Stephan. 2019. What Explains Uneven Female Labor Force Participation Levels and Trends in Developing Countries? World Bank Research Observer. 34 (2), 161-197.

Krueger, Alan, B., and Mikael Lindahl. 2001. "Education for Growth: Why and for Whom?" Journal of Economic Literature, 39 (4): 1101-1136.

Kuznets, Simon, 1955, Economic growth and income inequality, American Economic Review, 45(1): l-28.

———. 1966. Modern economic growth: rate, structure, and spread. Yale University Press.

Lam, David, Deborah Levison. 1991. Declining inequality in schooling in Brazil and its effects on inequality in earnings. Journal of Development Economics. 37 (1-2): 199-225.

Lee, Jong-Whaand Barro, Robert J. 2001. Schooling Quality in a Cross–Section of Countries. Economica, 68: 465-488. https://doi.org/10.1111/1468-0335.d01-12

Lewis, Arthur. "Development with Unlimited Supplies of Labour," The Manchester School, May 1954, 22, 139-92.

Lopez-Calva, Luis F., Nora Lustig, and Eduardo Ortiz Juarez. 2012. Declining Inequality in Latin America in the 2000s:The Cases of Argentina, Brazil, and Mexico. Tulane Economics Working Paper Series # 1212.

Mammen, Kristin, and Christina Paxson. 2000. Women's Work and Economic Development. Journal of Economic Perspectives. 14 (4):141-164.

Mincer, Jacob. 1974. Schooling, Experience and Earnings. NBER, New York.

Monteiro, Joana. 2015. Gasto Público em Educação e Desempenho Escolar. Revista Brasileira de Economia. 69(4):467–488.

Morchio, Iacopo and Moser, Christian, The Gender Pay Gap: Micro Sources and Macro Consequences (August 18, 2023). Available at SSRN: https://ssrn.com/abstract=3176868 or http://dx.doi.org/10.2139/ssrn.3176868

Mulligan, C. B., and Rubinstein, Y. 2008. Selection, Investment, and Women's Relative Wages over Time. The Quarterly Journal of Economics, 123(3), 1061–1110. http://www.jstor.org/stable/25098924

Nakabashi, L. and Salvato, M. 2007. Human capital quality in the Brazilian states. Economia. 8 (2): 211-229.

OECD. 2021. Education Policy Outlook: Brazil - With a Focus on National and Subnational Policies, available at: https://www.oecd.org/education/policy-outlook/country-profile-Brazil-2021-EN.pdf.

O'Neill, Donal. 1995. Education and Income Growth: Implications for Cross-Country Inequality. Journal of Political Economy 103, no. 6: 1289-1301.

Park, K.H. 1996. Educational expansion and educational inequality on income distribution. (sic.), Economics of Education Review. 15(1): 51-58

Pinkovskiy, Maxim and Xavier Sala-i-Martín. 2009. Parametric Estimations of the World Distribution of Income. NBER Working Paper # 15433. http://www.nber.org/papers/w15433

Psacharopoulos, George. 1993. Returns to Investment in Education: A Global Update. Policy Research Working Paper Series no. 1067. The World Bank, Washington, DC.

Psacharopoulos, G., Velez, E., 1993. Educational quality and labor market outcomes: evidence from Bogota, Colombia. Sociology of Education 66 (2), 130–145.

Ranis, Gustav, and John Fei. "A theory of economic development." The American economic review (1961): 533-565.

Ray, Debraj and Garance Genicot. 2023. Measuring Upward Mobility. American Economic Review, 113(11): 3044-3089.

Schultz, T.W. 1961. Investment in Human Capital. American Economic Review. 51 (1): 1- 17.

Schultz T. Paul. 1990. Women's Changing Perspective in the Labor Force: A World Perspective. Economic Development and Cultural Change. 38, 451-88.

Shah, Anwar. 1991. The new fiscal federalism in Brazil. Vol. 124. World Bank Publications.

Sotomayor, Orlando J. 2009. Changes in the Distribution of Household Income in Brazil: The Role of Male and Female Earnings. World Development. 37 (10): 1706-1715.

Thomas, Duncan. "Intra-Household Resource Allocation: An Inferential Approach." The Journal of Human Resources, vol. 25, no. 4, 1990, pp. 635-64.

Tobin, J. 1956. Estimation of Relationships for Limited Dependent Variables. Econometrica 26: 24-36.

U.S. Department of the Treasury. 1979. Treasury Reporting Rates of Exchange As of March 31, 1979. https://www.govinfo.gov/content/pkg/GOVPUB-T63_100-7369e48987fe20b8b7821a4829901ee2/pdf/GOVPUB-T63_100-7369e48987fe20b8b7821a4829901ee2.pdf

World Bank (https://data.worldbank.org/indicator/SI.POV.GINI?most$_{r}ecent_{v}alue_{d}esc = false$).

Table 1: Labor Force Participation and Labor Characteristics by Sex, 1976 to 2015

	1976		**1985**		**1995**		**2005**		**2015**	
A. Ages 16 to 65	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Worked	86.65	32.77	85.60	41.63	83.52	49.95	78.47	47.71	77.43	53.60
Looking for Work	2.82	1.28	4.01	2.43	4.75	4.27	8.36	8.02	13.62	11.76
Domestic Tasks	0.05	56.90	0.30	47.47	6.47	48.73	11.75	45.19	12.84	44.00
Studies	3.70	4.89	3.55	5.80	2.52	0.72	6.36	8.82	4.80	5.17
Pensioned	3.60	2.36	4.29	1.89	4.16	6.64	7.41	4.70	5.49	8.96
Hours worked/week	43.44	14.89	42.14	17.11	38.37	19.07	35.47	19.83	31.50	19.03
(St. Dev.)	(18.71)	(21.77)	(19.58)	(21.90)	(20.08)	(21.12)	(20.17)	(20.00)	(20.17)	(20.23)
Zero work hours (%)	11.5	65.2	13.6	57.7	15.8	46.2	19.3	43.9	23.0	45.5
Migrant (%)	49.4	49.3	..	..	50.5	52.2	47.3	49.5	41.9	43.9
Married (%)	59.1	58.1	60.7	59.9	61.1	52.2	57.7	59.0	58.7	57.8
Urban Resident (%)	65.9	68.9	73.8	76.8	79.4	52.2	82.5	85.2	84.4	86.8
Largest 10 City (%)	22.0	23.0	47.9	50.0	63.6	52.2	64.1	66.1	65.0	66.8
B. Ages 22 to 26										
Worked	89.49	38.87	90.05	45.80	86.47	52.96	82.33	55.86	77.00	54.82
Looking for Work	3.17	1.72	3.40	3.40	2.24	1.56	8.75	12.56	20.28	19.43
Domestic Tasks	0.05	53.02	0.15	45.85	6.51	45.52	8.54	41.49	11.70	36.22
Studies	4.10	4.57	2.31	4.08	3.24	5.82	4.83	7.84	6.43	10.04
Pensioned	0.20	0.16	0.23	0.10	0.38	0.14	0.26	0.43	0.25	0.23
Hours worked/week	46.64	17.2	43.63	19.02	39.59	20.74	36.41	21.7	31.29	20.23
(St. Dev.)	(15.8)	(22.3)	(17.1)	(22.3)	(18.65)	(4.17)	(19.20)	(21.17)	(19.49)	(20.51)
Zero work hours (%)	6.5	59.6	9.3	53.7	13.0	43.8	16.5	41.3	23.05	45.21
Migrant (%)	45.8	47.4	..	..	42.7	45.8	36.7	39.1	30.8	32.7
Married (%)	40.9	57.4	42.0	57.1	41.4	55.7	34.2	53.3	34.0	46.9
Urban Resident (%)	66.8	69.8	74.6	77.6	79.9	81.8	83.8	85.2	86.1	87.1
Largest 10 City (%)	21.7	23.1	48.2	50.1	63.1	64.1	64.3	64.4	65.5	64.2

Notes. This table contains information derived from the five PNAD surveys. It consists of two panels on labor force participation and migration, with separate statistics for men and women. Contrasting statistics for men and women over time illustrates secular changes in the role of women. Panel A data are for everyone between the ages of 16 and 65 and Panel B contains the same data for young adults (ages 22 to 26) who are forming the new labor force. Contrasting Panels A and B provides a information on contemporaneous rates of change.

Table 2: School Attainment by Sex, 1976 to 2015

	1976		1985		1995		2005		2015	
A. Ages 16-65	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
Age	33.48	33.19	33.64	33.73	34.40	34.89	35.21	35.84	37.43	38.37
(St. Dev.)	(13.37)	(13.47)	(13.23)	(13.24)	(13.18)	(13.27)	(13.36)	(13.41)	(13.80)	(13.82)
Schooling (Yrs.)	4.05	3.84	5.01	4.98	5.67	5.91	7.13	7.55	8.49	9.10
(St. Dev.)	(3.81)	(3.67)	(4.22)	(4.17)	(4.27)	(4.30)	(4.37)	(4.44)	(4.34)	(4.34)
Zero Schooling	28.3	32.18	21.61	22.63	17.27	16.1	11.63	10.35	14.27	11.96
4+ Schooling	43.9	41.82	41.05	39.53	41.24	39.26	31.5	28.61	20.81	17.98
8+ Schooling	13.57	13.1	17.96	17.77	18.07	18.13	21.38	20.63	20.35	18.50
11+ Schooling	14.24	12.9	19.37	20.08	23.42	26.51	35.5	40.41	44.57	51.57
# Obs.	104,370	112,057	144,699	154,898	97,595	104,846	128,324	137,527	117,362	125,996
B. Ages 22-26										
Age	23.91	23.92	23.93	23.94	23.96	23.99	23.95	23.97	23.95	23.99
(St. Dev.)	(1.42)	(1.41)	(1.41)	(1.41)	(1.41)	(1.41)	(1.41)	(1.40)	(1.43)	(1.42)
Schooling (Yrs.)	4.86	4.93	6.03	6.36	6.36	7.03	8.44	9.16	10.70	10.73
(St. Dev.)	(4.0)	(4.1)	(4.2)	(4.2)	(3.97)	(4.02)	(3.97)	(3.85)	(3.17)	(3.24)
Zero Schooling	19.46	20.19	12.66	10.48	10.93	7.14	5.84	3.86	5.11	3.29
4+ Schooling	43.82	42.45	39.47	38.26	40.99	38.63	24.76	21.57	14.76	9.85
8+ Schooling	16.05	15.00	22.34	21.92	22.34	21.07	21.01	18.74	20.39	16.38
11+ Schooling	20.66	22.36	25.53	29.35	25.73	33.15	48.39	55.83	59.73	70.48
# Obs.	16,569	18,372	22,795	24,983	13,804	14,627	19,066	19,421	13,276	13,208

Notes. This table is based on the same five PNAD surveys and demographic groups as Table (1). Its panels contain educational attainment statistics for men and women. The four-plus, eight-plus, and eleven-plus attainment categories correspond to the way that the Brazilian educational system was organized when the population studied here was in school.

Table 3: Mean, 1941-2004 Municipal School Spending Per Capita
(*Constant 1979 Cr$'000*)

	Rural Municipalities				Urban Municipalities			
State	**1941**	**1961**	**1981**	**2001**	**1941**	**1961**	**1981**	**2001**
Acre	0.31	0.03	0.92	31.61	0.10	0.10	2.12	8.96
Amazonas	0.10	0.03	2.35	34.90	0.35	0.00	0.85	9.93
Pará	0.20	0.07	1.68	33.71	0.12	0.19	0.73	3.19
Amapa	0.03	0.00	0.59	36.71		1.02	1.87	6.88
Rondônia			2.66	33.43		0.01	0.71	3.94
Roraima		0.09	0.06	11.81		0.17	2.43	12.58
Maranhão	0.04	0.03	1.13	25.36	0.11	0.10	0.95	3.07
Piauí	0.04	0.07	0.57	29.35	0.06	0.04	0.52	4.52
Ceará	0.03	0.07	1.77	40.38	0.13	0.13	1.77	4.95
Rio Grande do Norte	0.02	0.11	3.34	49.17	0.01	0.11	1.21	4.32
Paraíba	0.05	0.16	0.92	43.87	0.03	0.05	0.63	3.22
Pernambuco	0.11	0.27	2.67	45.19	0.25	0.17	0.92	3.63
Alagoas	0.05	0.14	1.91	36.83	0.01	0.09	1.53	3.85
Sergipe	0.05	0.10	3.33	45.14	0.07	0.16	1.79	4.69
Bahía	0.12	0.12	2.48	41.60	0.34	0.11	0.85	1.66
Mato Grosso	0.31	0.58	2.56	74.27	0.06	0.04	0.99	4.56
Mato Grosso do Sul	0.31	0.58	4.50	74.86	0.06	0.04	2.03	6.66
Goias	0.12	0.18	25.45	102.34	0.08	0.12	1.08	4.87
Minas Gerais	0.13	0.21	3.70	78.81	0.03	0.13	0.87	3.40
Espirito Santo	0.05	0.11	4.45	80.41	0.02	0.00	1.09	5.32
Rio de Janeiro	0.20	0.93	17.74	257.40	0.01	0.00	6.95	13.48
São Paulo	0.29	1.27	20.28	312.25	0.43	0.66	2.63	7.91
Paraná	0.06	0.27	5.25	72.06	0.09	0.08	1.18	3.90
Santa Catalina	0.21	0.35	4.56	82.07	0.07	0.03	0.27	1.82
Rio Grande do Sul	0.23	1.03	8.04	89.57	0.03	0.20	0.67	3.96

Notes. The education quality data used in this paper covers the entire 1941 to 2004 period. This table provides an impression of the data by reporting figures for 20-year intervals. The data represent per-capita spending on education by rural and urban municipalities for each Brazilian state in constant, 1979 Cruzeiros. The USD/Cruzeiro exchange rate in 1979 was Cr$ 22.58 per US$ (Department of the Treasury, 1979). Since the data are ascribed to the time when survey respondants were in school at age 10, they cover people born between 1931 and 1994. States that did not exist in 1941 have missing values.

Table 4: ***1976, 1995, 2005, and 2015 Log-Wage Estimates for Seven, 3-Year Male Cohorts***

Sample \Age Cohort	(1) 25-27	(2) 28-30	(3) 31-33	(4) 34-36	(5) 37-39	(6) 40-42	(7) 43-45
1976	—	**1976**	—	**1976**	—	**1976**	—
School	0.140***	0.155***	0.164***	0.164***	0.163***	0.168***	0.177***
	(62.57)	(66.71)	(64.43)	(57.68)	(52.73)	(49.76)	(49.11)
School*Quality	0.0344***	0.0313***	0.0308***	0.0381***	0.0366***	0.0581***	0.0311*
	(8.76)	(7.27)	(6.88)	(6.62)	(3.84)	(4.45)	(2.19)
_cons	1.198***	1.270***	1.348***	1.382***	1.447***	1.468***	1.500***
	(93.05)	(98.07)	(95.22)	(98.48)	(96.58)	(98.73)	(92.87)
N	6504	6538	5450	5574	5081	4896	4210
R^2	0.462	0.497	0.530	0.476	0.450	0.468	0.469
1995	—	**1995**	—	**1995**	—	**1995**	—
School	0.107***	0.126***	0.126***	0.137***	0.133***	0.143***	0.145***
	(43.25)	(51.12)	(52.30)	(55.81)	(50.15)	(50.09)	(45.87)
School*Quality	0.00240***	0.00325***	0.00471***	0.00483***	0.0146***	0.0265***	0.0309***
	(8.66)	(8.66)	(8.16)	(5.63)	(7.05)	(7.39)	(7.06)
_cons	0.954***	0.959***	1.033***	1.081***	1.160***	1.150***	1.193***
	(56.99)	(55.76)	(60.49)	(62.47)	(64.20)	(62.18)	(61.80)
N	5780	5789	5950	5528	5107	4778	4137
R^2	0.331	0.394	0.393	0.422	0.424	0.458	0.468
2005	—	**2005**	—	**2005**	—	**2005**	—
School	0.0909***	0.102***	0.104***	0.112***	0.111***	0.120***	0.125***
	(46.59)	(51.07)	(49.27)	(50.64)	(48.15)	(52.00)	(51.55)
School*Quality	0.000619***	0.00103***	0.00146***	0.00231***	0.00333***	0.00409***	0.00550***
	(9.22)	(8.06)	(9.36)	(11.84)	(11.55)	(9.36)	(7.45)
_cons	0.130***	0.151***	0.220***	0.206***	0.267***	0.263***	0.283***
	(7.72)	(9.14)	(13.38)	(12.32)	(15.46)	(15.00)	(15.55)
N	7133	6742	6524	6254	6047	6045	5423
R^2	0.287	0.337	0.349	0.396	0.377	0.396	0.404
2015	—	**2015**	—	**2015**	—	**2015**	—
School	0.0802***	0.0915***	0.0938***	0.0913***	0.0865***	0.0879***	0.0834***
	(34.66)	(41.79)	(42.89)	(42.02)	(41.67)	(40.33)	(35.15)
School*Quality	0.0000893***	0.000189***	0.000322***	0.000600***	0.00108***	0.00124***	0.00160***
	(5.33)	(8.23)	(8.88)	(10.88)	(10.44)	(7.80)	(8.58)
_cons	2.543***	2.499***	2.546***	2.610***	2.737***	2.776***	2.789***
	(105.78)	(110.72)	(115.75)	(123.86)	(144.90)	(146.90)	(132.75)
N	4868	5392	5644	5677	5452	5167	4899
R^2	0.217	0.278	0.291	0.301	0.318	0.305	0.276

t statistics in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. Each of columns (1) through (7) is for a different, 3-year age cohort of men. Every cell in this table is the estimated coefficient from a "Mincer equation" as in Equation (2), on years of school attained, school attainment interacted with school quality, or the constant for that survey year for the age cohort in that column. However, since age differences within each cohort are insignificant, terms involving potential labor market experience are subsumed in the intercepts. School quality is statistically significant at the 99.9% level in 27 out of 28 cases and significant at the 95% level for the oldest 1976 cohort.

Table 5: ***The Contribution of School Quality to Earnings Differentials among Male Cohorts***

Age Cohort	(1) N	(2) Var(y)	(3) β	(4) γ	(5) $\beta^2 Var(S)$	(6) $\gamma^2 Var(Sq)$	(7) $2\beta\gamma Cov(S, Sq)$	(8) R^2	% of R^2 from Quality
				1976	**Sample**				
25-27	6,504	0.83	0.14	0.034	0.33	0.007	0.042	0.46	10.5%
28-30	6,538	0.98	0.16	0.031	0.43	0.005	0.046	0.50	10.2%
31-33	5,450	1.05	0.16	0.031	0.50	0.006	0.053	0.53	11.1%
34-36	5,574	1.02	0.16	0.038	0.43	0.006	0.050	0.48	11.7%
37-39	5,081	1.02	0.16	0.037	0.42	0.002	0.032	0.45	7.7%
40-42	6,483	1.05	0.17	0.062	0.44	0.004	0.050	0.47	11.6%
43-45	4,210	1.10	0.18	0.031	0.49	0.001	0.025	0.47	5.4%
				1995	**Sample**				
25-27	5,780	0.66	0.11	0.002	0.18	0.007	0.032	0.33	11.8%
28-30	5,789	0.83	0.13	0.003	0.28	0.008	0.040	0.39	12.3%
31-33	5,950	0.88	0.13	0.005	0.30	0.007	0.039	0.39	11.8%
34-36	5,528	0.92	0.14	0.005	0.35	0.004	0.028	0.42	7.6%
37-39	5,107	0.98	0.13	0.015	0.36	0.007	0.048	0.42	13.1%
40-42	4,778	1.09	0.14	0.026	0.42	0.009	0.065	0.46	16.1%
43-45	4,137	1.17	0.15	0.031	0.46	0.011	0.079	0.47	19.1%
				2005	**Sample**				
25-27	7,133	0.54	0.09	0.0006	0.13	0.005	0.017	0.29	7.9%
28-30	6,742	0.62	0.10	0.0010	0.18	0.005	0.021	0.34	7.6%
31-33	6,524	0.66	0.10	0.0015	0.19	0.007	0.031	0.35	10.8%
34-36	6,254	0.75	0.11	0.002	0.23	0.013	0.050	0.40	15.8%
37-39	6,047	0.78	0.11	0.003	0.23	0.013	0.049	0.38	16.5%
40-42	6,045	0.85	0.12	0.004	0.28	0.009	0.044	0.40	13.5%
43-45	5,423	0.91	0.12	0.005	0.32	0.007	0.039	0.40	11.4%
				2015	**Sample**				
25-27	4,868	0.36	0.08	0.00009	0.07	0.0017	0.004	0.22	2.8%
28-30	5,392	0.45	0.09	0.00019	0.11	0.0043	0.010	0.28	5.2%
31-33	5,644	0.51	0.09	0.00032	0.13	0.0054	0.015	0.29	6.9%
34-36	5,677	0.56	0.09	0.00060	0.14	0.0092	0.023	0.30	10.7%
37-39	5,452	0.56	0.09	0.0011	0.14	0.0089	0.027	0.32	11.3%
40-42	5,147	0.58	0.09	0.0012	0.15	0.0056	0.022	0.31	9.2%
43-45	4,899	0.62	0.08	0.0016	0.14	0.0081	0.027	0.28	12.8%

Notes. This table is in 4 sections, one for each of the 1976, 1995, 2005, and 2015 PNAD survey samples. Figures are based on the full set of Mincer equation estimates reported in Table 4. Column (2) reports earnings inequality as the variance of the logarithm of earnings for men in the cohort listed on the left, and column (3) shows the rate of return per year of schooling attainment. Column (4) gives the mean percentage income increase per 1,000 1979 Cruseiros per capita spent on education in the area where that man went to school. Column (5) reports the total variation in log income explained by schooling, while columns (6) and (7) are the portion explained by the returns to school quality, the variance in school quality, and the covariance between school quality and school attainment. Column (8) reports the R^2 for that row. Figures in Column (9) represent what proportion of the R^2 comes from including the school quality information.

Table 6: Headship and Labor Force Participation by Sex
Households with a Head & Spouse Living Together Ages 20-55

	1976	**1985**	**1995**	**2005**	**2015**
Full Sample, Ages 20-55	118,005	165,144	210,220	144,449	129,070
Head & Spouse Cohabit (%)	82.8	89.7	81.5	77.8	74.4
Men (%)					
Identified as Head	100.0	99.8	99.0	92.0	75.9
Worked Full Time	94.7	91.2	86.5	83.7	80.3
Women (%)					
Identified as Head	0.0	0.3	1.0	8.0	24.1
Worked Full Time	17.1	22.5	29.4	35.3	39.3

Notes. This table provides an impressionistic view of the typical role of gender in the organization and labor force participation of Brazilian households. The first row reports the number of people of ages 20 to 55 in that survey year. The second row shows what proportion of those people lived in a household where the head and his or her spouse were living together. The remaining entries report two statistics on the men and the women in these "husband-and-wife" households: the proportion of men (women) who were identified as heads and the proportion of men (women) who worked "full-time", at least 35 hours per week.

Table 7: CSE Model Maximum Likelihood Estimates Estimates of of Domestic Labor Supply
Human Capital Index, Fertility & Marital Status

YEAR	**1976**		**1995**		**2005**		**2015**	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Human Capital	-0.207***	-0.229***	-0.129***	-0.131***	-0.179***	-0.176***	-0.234***	-0.233***
	(-36.02)	(-42.41)	(-47.55)	(-48.12)	(-69.04)	(-67.17)	(-69.03)	(-68.40)
# Children		0.0277***		0.0157***		0.0204***		0.0186***
		(8.60)		(6.82)		(8.29)		(6.05)
Married		0.511***		0.113***		0.0253***		0.0323***
		(70.38)		(32.03)		(8.32)		(10.55)
Constant	1.029***	0.765***	0.699***	0.626***	0.536***	0.512***	1.166***	1.138***
	(80.43)	(63.43)	(125.05)	(106.12)	(163.50)	(135.96)	(94.68)	(91.19)
σ_y	0.765***	0.682***	0.463***	0.459***	0.462***	0.461***	0.431***	0.431***
	(184.97)	(186.20)	(280.70)	(280.87)	(326.58)	(326.57)	(300.42)	(300.44)
N	66693	66693	90396	90396	118157	118157	99696	99696
β	0.410	0.400	0.370	0.370	0.390	0.390	0.390	0.390
Loglikelihood	-24509.6	-20921.9	-18691.0	-18049.4	-20024.7	-19935.4	-13403.7	-13314.1

t statistics in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. These are maximum likelihood estimates of the CSE model of Equation (15) for each survey year. Estimates show a consistently negative and significant relation between the amount of a woman's human capital Hk, and the proportion of time that she spends on domestic production. The first column of each year includes only the woman's human capital measure Hk, while second column estimates also control for differences due to her marital status and the number of children below age 7 in her household. Hk coefficients are robust to the inclusion of the marital status and fertility controls.

Table 8: CSE Model Maximum Likelihood Estimates of Domestic Labor Supply
Using 'Raw' Human Capital Inputs Directly, with and without School Quality

YEAR	**1976**		**1995**		**2005**		**2015**	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Sch. Yrs.	-0.0210***	-0.0165***	-0.0110***	-0.0108***	-0.0166***	-0.0165***	-0.0213***	-0.0211***
	(-24.84)	(-18.48)	(-28.06)	(-26.52)	(-49.23)	(-48.37)	(-61.91)	(-61.03)
Sch. Quality		-0.0165***		-0.000154**		-0.284e-4*		-0.189e-4***
		(-14.32)		(-2.88)		(-2.48)		(-4.62)
Age	-0.0938***	-0.100***	-0.0308***	-0.0312***	-0.0314***	-0.0316***	-0.0286***	-0.0291***
	(-25.91)	(-27.47)	(-62.95)	(-61.20)	(-82.15)	(-80.42)	(-68.30)	(-67.23)
Age^2	0.00134***	0.00143***	0.668e-5***	0.674e-5***	0.638e-5***	0.641e-5***	0.567e-5***	0.573e-5***
	(21.50)	(22.84)	(61.95)	(61.32)	(83.77)	(83.11)	(74.14)	(73.78)
# Children	0.0390***	0.0388***	0.0322***	0.0320***	0.0433***	0.0432***	0.0595***	0.0594***
	(12.08)	(12.04)	(13.61)	(13.52)	(17.38)	(17.32)	(19.16)	(19.10)
Married	0.606***	0.605***	0.157***	0.157***	0.0637***	0.0637***	0.0529***	0.0532***
	(73.08)	(73.18)	(41.46)	(41.51)	(20.16)	(20.17)	(17.04)	(17.14)
constant	1.819***	1.922***	1.090***	1.101***	1.126***	1.133***	1.144***	1.159***
	(38.03)	(39.73)	(97.75)	(93.14)	(116.85)	(113.08)	(100.46)	(97.64)
σ_y	0.677***	0.675***	0.450***	0.450***	0.448***	0.448***	0.416***	0.416***
	(186.30)	(186.34)	(281.51)	(281.52)	(327.74)	(327.74)	(301.71)	(301.72)
N	66693	66693	90396	90396	118157	118157	99696	99696
β	0.400	0.400	0.370	0.370	0.390	0.390	0.390	0.390
Loglikelihood	-20645.0	-20543.4	-16478.3	-16474.2	-16776.6	-16773.5	-10027.2	-10016.5

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. In these CSE model estimates, the Hk index of Table (7) is eschewed in favor of the "raw" human capital inputs of Equation (7). Having more years of schooling reduces the amount of labor women supply to domestic production, but attending a better quality school further reduces it significantly in all four samples. The equality of the first two coefficients in 1976 Column (2) is not a typo, it is a coincidence.

Table 9: Basic Heckman 2-step Model of Women's Hours Worked

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
School	-0.692***	-0.129***	0.0510*	0.116***
	(-24.08)	(-3.54)	(2.03)	(5.35)
Experience	-0.0803**	-0.395***	-0.269***	-0.161***
	(-3.11)	(-8.82)	(-9.13)	(-6.28)
Experience^2	0.000892*	0.00658***	0.00472***	0.00302***
	(2.37)	(8.88)	(8.93)	(6.07)
_cons	52.53***	58.74***	51.76***	45.07***
	(77.31)	(41.36)	(55.39)	(57.86)
Selection				
School	0.0621***	0.0232***	0.0169***	0.0227***
	(38.64)	(17.90)	(15.69)	(21.25)
Experience	0.0270***	0.0369***	0.0271***	0.0280***
	(20.00)	(29.97)	(27.36)	(27.28)
Experience^2	-0.000606***	-0.000784***	-0.000705***	-0.000794***
	(-31.70)	(-47.06)	(-52.57)	(-55.01)
Married	-0.793***	-0.293***	-0.230***	-0.253***
	(-38.43)	(-18.73)	(-18.73)	(-22.83)
Husband's Income	-0.0195***	-0.0111***	0.00367*	0.0105***
	(-7.70)	(-4.78)	(2.19)	(7.81)
# Children	-0.0689***	-0.112***	-0.218***	-0.277***
	(-13.13)	(-17.08)	(-32.12)	(-32.13)
_cons	-0.217***	0.0265	0.260***	0.200***
	(-7.55)	(0.94)	(11.01)	(8.33)
/mills				
lambda	-6.180***	-26.60***	-22.41***	-14.87***
	(-19.55)	(-23.59)	(-26.41)	(-21.72)
N	79374	78660	105098	104011

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. This table reports replications of Heckman (1976) using the Brazilian PNAD data for 1976, 1995, 2005, and 2015. Unlike U.S. results reported by Mulligan and Rubinstein (2008), first stage "selection" estimates shown in the bottom panel indicate consistently negative selection of women into wage work in Brazil over the period. Though significant throughout, there is a "sign-flip" on husband's income in the selection equation and another, corresponding one on years of school attainment in the hours worked estimates.

Table 10: Heckman 2-step Model of Women's Hours Worked with Human Capital Index

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
Human Capital	-4.532***	-0.844**	-0.460	0.393
	(-22.61)	(-3.16)	(-1.31)	(0.96)
_cons	57.60***	52.97***	51.94***	43.76***
	(80.87)	(39.58)	(34.98)	(19.23)
Selection				
Human Capital	0.539***	0.262***	0.284***	0.476***
	(43.86)	(30.34)	(35.67)	(45.57)
Married	-0.718***	-0.345***	-0.266***	-0.247***
	(-19.40)	(-18.67)	(-19.32)	(-19.63)
# Children	-0.0745***	-0.0320***	-0.0704***	-0.0609***
	(-12.40)	(-5.12)	(-10.68)	(-7.05)
Husband's Income	-0.0696***	-0.00716**	0.0206***	0.0252***
	(-16.13)	(-2.85)	(11.62)	(17.22)
_cons	-0.505***	-0.0572**	0.0515***	-1.394***
	(-15.72)	(-2.66)	(3.89)	(-36.07)
/mills				
lambda	-6.399***	-24.11***	-25.42***	-15.42***
	(-18.35)	(-19.48)	(-15.11)	(-12.18)
N	44348	63893	84143	76334

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. The Hk index replaces the raw human capital inputs in these first and second stage estimates. A major difference with the standard Heckman model specification is that it contains more information. Since Hk controls for school quality at the time and place where a women went to school, it is a dynamic measure that is not the same for all women of the same age with the same years of schooling. Nevertheless, these estimates evince sign-flips analogous to the ones produced by the standard Heckman specifications reported in Table (9). The first-stage effect of higher husband income is negative and significant in 1976 and 1995, and it is positive and significant in 2005 and 2015. Hk effects are strongly negative and significant in the first of the second-stage estimates, but the effect shrinks and eventually becomes positive but insignificant by the fourth survey year.

Table 11: Heckman 2-step Model of Women's Hours Worked with 'Raw' School Quality Measures

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
School	-0.816***	-0.192***	0.0132	0.138***
	(-20.88)	(-5.17)	(0.49)	(6.34)
Sch. Quality	3.124***	0.210***	0.0672***	0.0156***
	(8.00)	(4.65)	(5.42)	(4.59)
Experience	0.00500	-0.357***	-0.267***	-0.123***
	(0.07)	(-7.20)	(-7.93)	(-4.66)
Experience2	-0.00171	0.00573***	0.00488***	0.00260***
	(-1.03)	(6.89)	(8.21)	(5.13)
_cons	52.05***	56.07***	51.94***	43.98***
	(53.38)	(42.47)	(52.63)	(56.65)
select				
School	0.108***	0.0343***	0.0192***	0.0205***
	(47.34)	(23.97)	(16.98)	(18.64)
Sch. Quality	0.0817***	-0.0129***	-0.00337***	-0.000923***
	(3.95)	(-6.80)	(-5.99)	(-4.55)
Experience	0.0432***	0.0427***	0.0307***	0.0248***
	(10.60)	(22.90)	(25.15)	(21.37)
Experience2	-0.000534***	-0.000867***	-0.000777***	-0.000772***
	(-5.51)	(-28.45)	(-43.38)	(-47.92)
Married	-0.717***	-0.331***	-0.251***	-0.252***
	(-19.72)	(-18.27)	(-19.50)	(-22.38)
Husband's Income	-0.0699***	-0.0219***	0.00270	0.0109***
	(-16.35)	(-8.69)	(1.56)	(7.90)
# Children	-0.0725***	-0.0991***	-0.215***	-0.288***
	(-12.06)	(-14.68)	(-31.06)	(-32.52)
_cons	-0.501***	-0.0161	0.245***	0.300***
	(-10.43)	(-0.46)	(9.24)	(11.22)
/mills				
lambda	-6.662***	-23.44***	-22.92***	-14.66***
	(-19.41)	(-24.24)	(-26.53)	(-21.36)
N	46114	67828	97187	98983

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes. In this table, the Heckman (1976) specification of Table (9) is augmented with a school quality measure. The quality measure is significant at the 99.9% level in all first and second stage estimates, while the remaining coefficients remain largely unchanged in sign, magnitude, and significance. This indicates that the school quality variable is adding new information. However, adding this new information does not change the sign-flips on husband's income in the selection equation and on years of school attainment in the hours worked estimates.

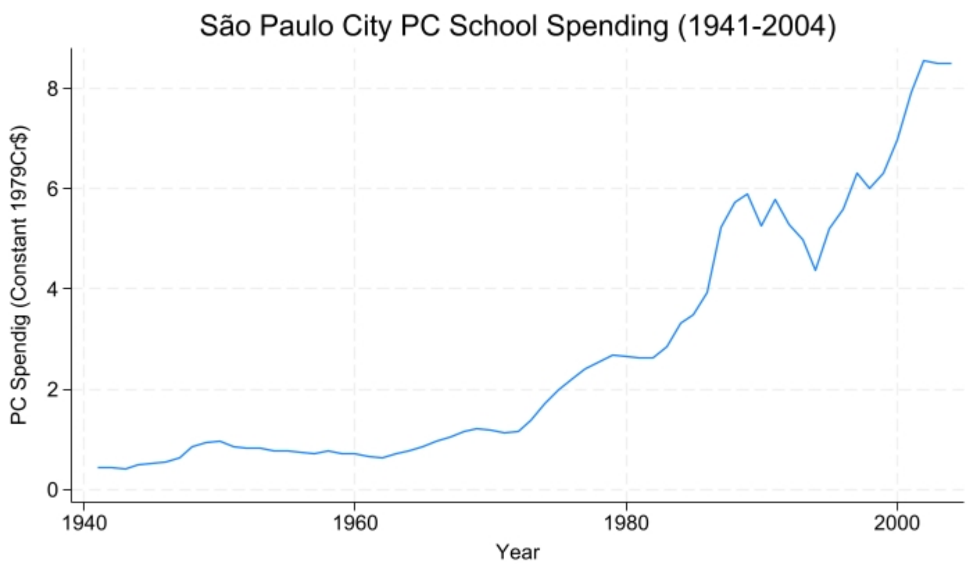
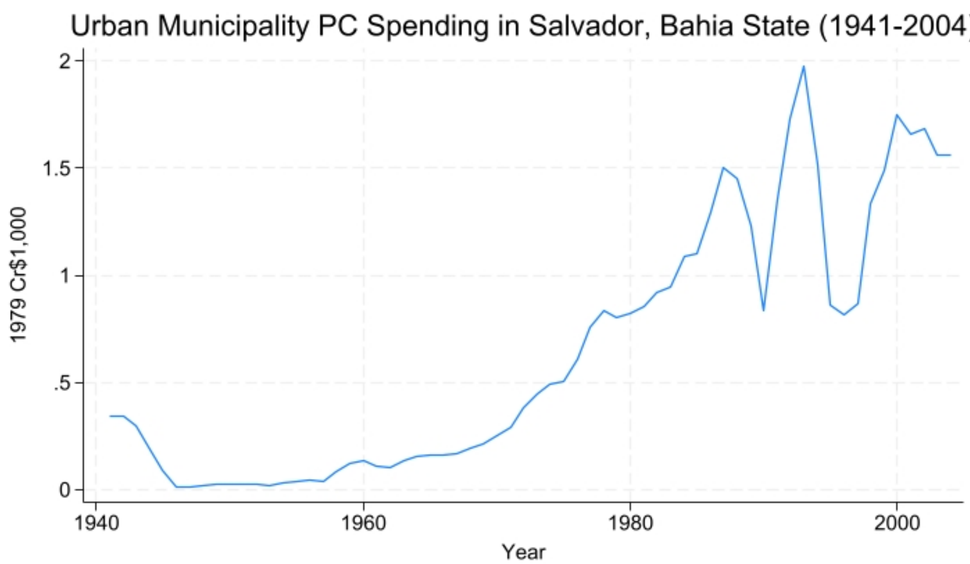


Figure 1: *This figure shows per capita education spending São Paulo City, capital of the Brazilian state of São Paulo, from 1941 to 2004.*



Bahia

Figure 2: *This figure shows per capita education spending in the city of Salvador, capital of the Brazilian state of Bahía, from 1941 to 2004.*

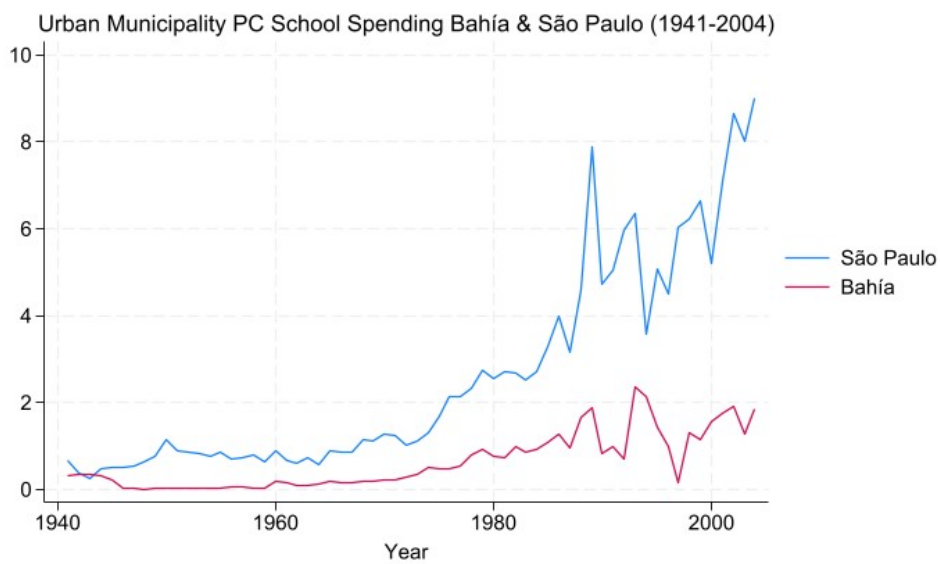


Figure 3: *This figure shows the evolution of real mean per capita education expenditures in the urban municipalities of São Paulo and Bahía, from 1941 to 2004.*

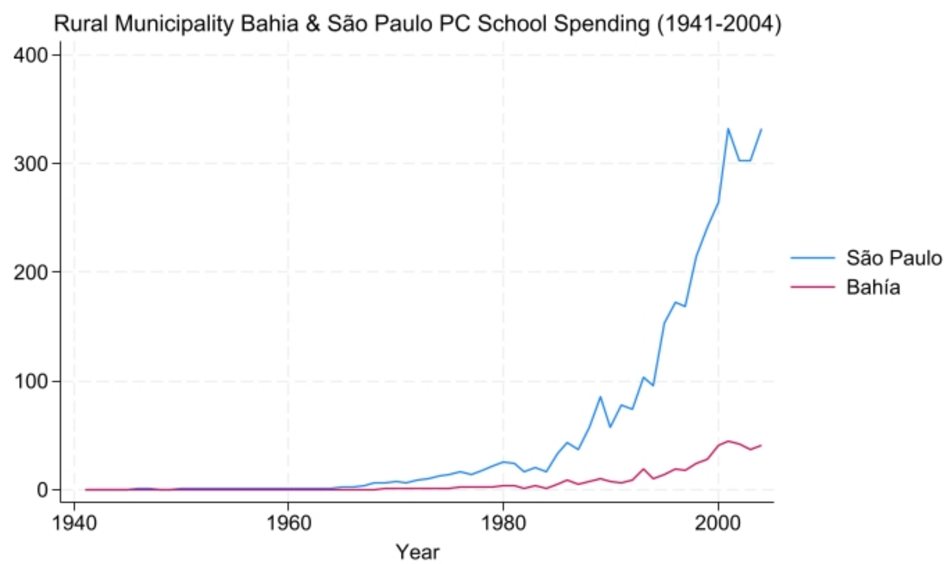
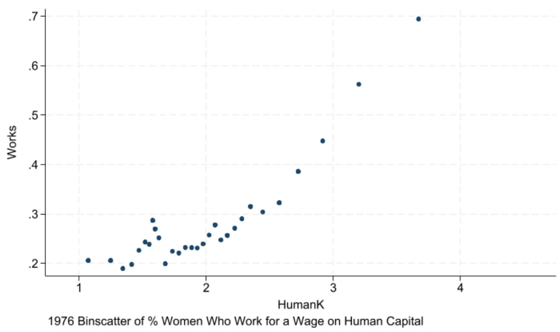
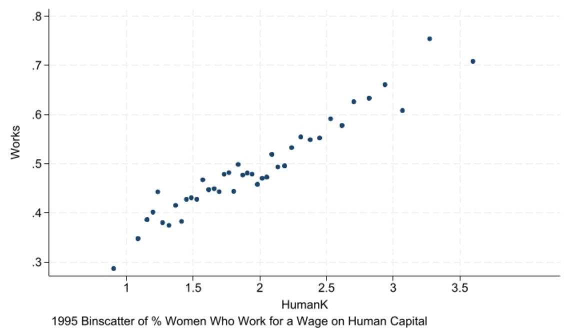


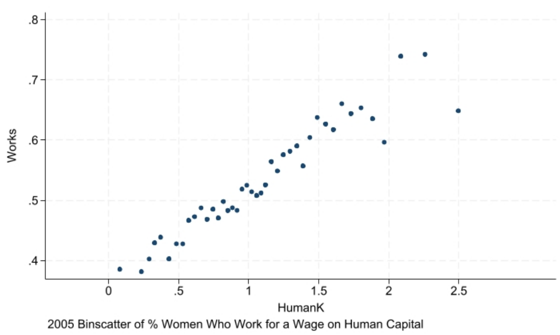
Figure 4: *This figure contrasts the evolution of real mean per capita education expenditures in the rural municipalities of the Brazilian states of Bahía to expenditures in São Paulo and, from 1941 to 2004.*



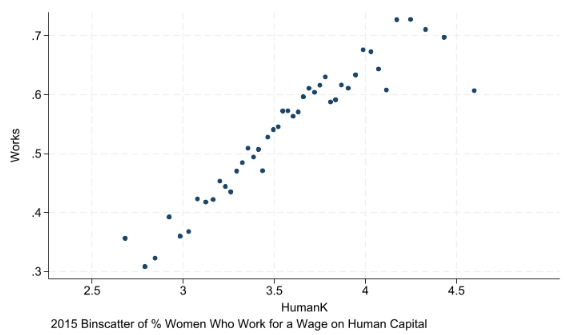
(a) 1976



(b) 1995

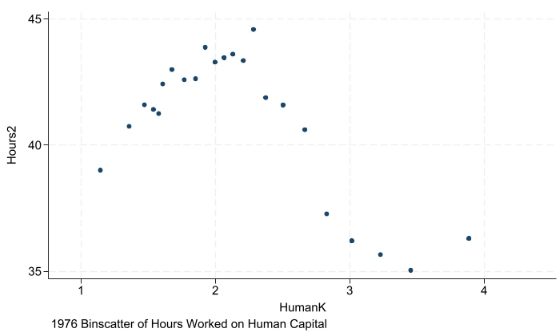


(c) 2005

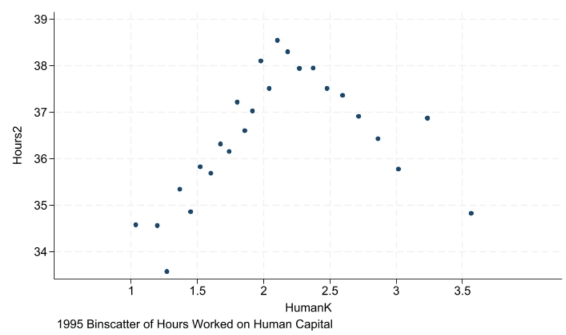


(d) 2015

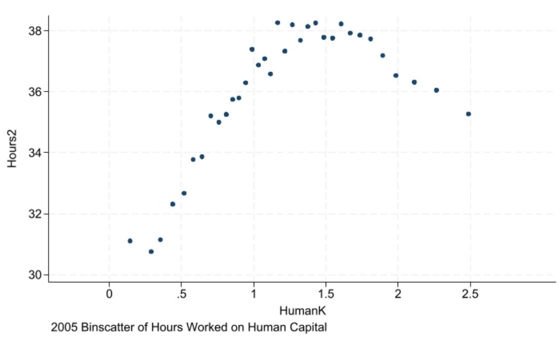
Figure 5: Binscatter plots of the percentage of women who work, by human capital level.



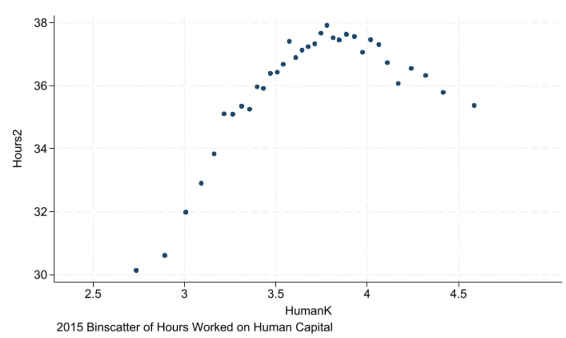
(a) 1976



(b) 1995



(c) 2005



(d) 2015

Figure 6: Binscatter plots of women's weekly work hours, as a function of their human capital level.

9 APPENDIX

Table A1: Large-City Earnings of Fully Employed Men by Location of School Attendance

YEAR	**1976**		**1995**		**2005**		**2015**	
School Location	Rural	Urban	Rural	Urban	Rural	Urban	Rural	Urban
Amazonas			-0.29		-0.36			
(t-stat.)			(-2.29)		(-3.13)			
Pará	-0.20	-0.29	-0.42	-0.48	-0.37	-0.47	-0.35	-0.36
(t-stat.)	(-0.85)	(-2.62)	(-13.23)	(-17.31)	(-17.21)	(-21.80)	(-16.42)	(-21.23)
Tocantins			-0.41		-0.24		-0.14	
(t-stat.)			(-5.00)		(-4.06)		(-2.75)	
Maranhão	-0.48	-0.32	-0.37		-0.26		-0.16	
(t-stat.)	(-4.57)	(-3.74)	(-8.37)		(-8.09)		(-5.89)	
Piauí	-0.41	-0.30	-0.15		-0.10		-0.021	
(t-stat.)	(-3.91)	(-3.08)	(-2.88)		(-2.40)		(-0.55)	
Ceará	-0.60	-0.60	-0.39	-0.60	-0.36	-0.61	-0.28	-0.41
(t-stat.)	(-17.83)	(-22.04)	(-18.09)	(-24.38)	(-19.17)	(-31.21)	(-14.28)	(-22.82)
Rio Grande do Norte	-0.30	-0.29	-0.29		-0.09		-0.0032	
(t-stat.)	(-3.87)	(-3.97)	(-5.94)		(-1.88)		(-0.060)	
Paraíba	-0.39	-0.32	-0.16		-0.17		-0.091	
(t-stat.)	(-9.17)	(-6.19)	(-4.41)		(-5.25)		(-2.82)	
Pernambuco	-0.47	-0.52	-0.36	-0.56	-0.37	-0.56	-0.21	-0.39
(t-stat.)	(-17.59)	(-24.61)	(-17.23)	(-23.99)	(-19.55)	(-29.70)	(-11.03)	(-23.63)
Alagoas	-0.39	-0.23	-0.12		-0.13		0.0051	
(t-stat.)	(-6.37)	(-4.19)	(-2.68)		(-2.89)		(0.11)	
Sergipe	-0.29	-0.08	-0.14		-0.18		-0.10	
(t-stat.)	(-3.84)	(-1.12)	(-2.47)		(-3.18)		(-1.78)	
Bahia	-0.43	-0.39	-0.34	-0.61	-0.28	-0.47	-0.21	-0.39
(t-stat.)	(-17.57)	(-20.20)	(-17.07)	(-28.12)	(-16.47)	(-27.08)	(-12.82)	(-24.64)
Minas Gerais	-0.42	-0.31	-0.24	-0.39	-0.14	-0.26	-0.0066	-0.11
(t-stat.)	(-26.93)	(-23.02)	(-13.75)	(-19.85)	(-8.88)	(-15.98)	(-0.44)	(-7.95)
Espírito Santo	-0.43	-0.23	-0.17		-0.11		-0.04	
(t-stat.)	(-8.69)	(-4.37)	(-3.63)		(-2.32)		(-0.66)	
Rio de Janeiro	-0.44	-0.35	-0.27	-0.36	-0.13	-0.19	-0.046	-0.11
(t-stat.)	(-16.54)	(-26.96)	(-11.02)	(-18.73)	(-5.65)	(-11.40)	(-2.16)	(-7.36)
São Paulo	-0.13	1.00	-0.01	1.00	0.00	1.00	0.05	1.00
(t-stat.)	(-7.84)	..	(-0.31)	..	(-0.07)	..	3.34)	..
Paraná	-0.33	-0.21	-0.13	-0.18	-0.041	-0.13	0.07	0.012
(t-stat.)	(-9.80)	(-8.15)	(-6.05)	(-6.48)	(-2.20)	(-5.87)	(3.85)	(0.68)
Santa Catarina	-0.33	-0.24	-0.0076		-0.01		0.14	
(t-stat.)	(-9.56	(-8.70	(-0.20		(-0.36		3.57	
Rio Grande do Sul	-0.36	-0.30	-0.22	-0.26	-0.060	-0.13	-0.010	-0.037
(t-stat.)	(-15.96)	(-15.31)	(-11.15)	(-12.26)	(-3.65)	(-7.13)	(-0.63)	(-2.35)
Mato Grosso do Sul			0.10		-0.072		0.21	
(t-stat.)			(0.90)		(-0.80)		(2.83)	
Mato Grosso			0.0037		-0.16		-0.038	
(t-stat.)			(0.030)		(-1.99)		(-0.51)	
Goiás	-0.40	-0.31	-0.41	-0.50	-0.18	-0.19	-0.026	-0.055
(t-stat.)	(-14.24)	(-12.99)	(-15.25)	(-15.47)	(-8.00)	(-7.57	(-1.18)	(-2.71)
Distrito Federal				-0.28		-0.21		-0.031
(t-stat.)				(-2.63)		(-3.24)		(-0.62)

Notes. These estimates are based on fully-employed men (ages 20-65) living in a Brazilian city with more than 1.5 million inhabitants during the survey year. Only places of origin with at least 30 observations were included. In the spirit of Brotherhood, Cavalcanti-Ferreira, and Santos (2019), each estimate (t-statistic) is a "Mincer regression" coefficient on a dummy variable that takes on a value of 1 for men who lived in that place when they were 10 years old —and are therefore assumed schooled there. Coefficients measure earnings relative to men schooled in urban São Paulo. For example, *ceteris paribus*, men schooled in rural Pará state, earned 20% less in 1976, 42% less in 1995, 37% less in 2005, and 35% in 2015 than men who went to school in urban São Paulo.

Table A2: Mincer Regressions

Fully-Employed Men Ages 25 to 45

	1976	**1995**	**2005**	**2015**
Yrs. School	0.179***	0.159***	0.136***	0.114***
	(136.38)	(172.26)	(170.16)	(140.54)
School*Quality	0.0336***	0.00317***	0.000897***	0.000165***
	(16.53)	(18.59)	(20.72)	(16.42)
Experience	0.0708***	0.0574***	0.0342***	-0.00270
	(22.64)	(24.01)	(19.80)	(-1.71)
Experience^2	-0.000974***	-0.000615***	-0.000228***	0.000573***
	(-14.72)	(-12.21)	(-5.80)	(14.73)
BigCity	0.185***	0.138***	0.0274***	0.0640***
	(24.41)	(19.93)	(4.87)	(12.16)
_cons	-0.0503	-0.259***	-0.709***	2.088***
	(-1.28)	(-8.51)	(-33.61)	(111.00)
N	44815	58299	70753	64357
R^2	0.442	0.403	0.346	0.284

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A3: Basic Tobit Estimates of Hours Worked –Women Age 20 to 65

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
Yrs. School	2.863***	1.083***	1.154***	1.549***
	(38.05)	(31.52)	(42.93)	(59.41)
Age	4.816***	1.730***	1.322***	1.468***
	(30.86)	(36.20)	(32.87)	(35.51)
Age^2	-0.0672***	-0.000399***	-0.000329***	-0.000340***
	(-34.80)	(-48.41)	(-48.32)	(-50.76)
Married	-43.31***	-12.49***	-8.861***	-7.939***
	(-62.43)	(-33.80)	(-31.02)	(-29.31)
Children of Age $\leq$ 7	-4.599***	-3.987***	-6.114***	-7.042***
	(-17.29)	(-20.19)	(-30.97)	(-29.87)
Husband's Income	-0.000920***	-0.00235***	-0.000253***	0.000102*
	(-15.09)	(-12.29)	(-3.44)	(2.46)
_cons	-75.18***	-22.89***	-13.70***	-25.06***
	(-24.05)	(-16.18)	(-11.39)	(-19.98)
/				
var(e.Hours)	2655.9***	1064.3***	959.9***	903.0***
	(80.22)	(120.94)	(141.73)	(139.06)
N	70272	68297	85578	84432

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A4: Tobit Estimates of Hours Worked with Human Capital Index –Women Age 20 to 65

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
Hk	19.56***	8.844***	9.720***	14.80***
	(32.25)	(33.54)	(41.72)	(50.83)
Married	-48.83***	-11.15***	-5.821***	-4.943***
	(-50.61)	(-27.69)	(-18.63)	(-16.30)
Children of Age$\leq$ 7	-4.396***	-1.758***	-1.836***	-1.381***
	(-15.30)	(-9.40)	(-9.48)	(-5.70)
Husband's Income	-0.000909***	-0.00192***	-0.0000640	0.000261***
	(-11.97)	(-9.37)	(-0.80)	(5.73)
_cons	-14.52***	1.333*	4.928***	-38.53***
	(-9.30)	(2.05)	(12.59)	(-35.35)
/				
var(e.Hours)	2469.3***	1110.6***	1050.9***	996.2***
	(62.90)	(113.29)	(131.99)	(127.26)
N	40053	58829	75005	70946

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A5: Tobit Estimates of Hours Worked Controlling for School Quality –Women Age 20 to 65

	(1) 1976	(2) 1995	(3) 2005	(4) 2015
% Hours in Formal Market Work				
School	3.256***	1.107***	1.134***	1.529***
	(33.45)	(30.07)	(40.05)	(56.45)
School×*Quality*	-0.00343***	0.000101	0.0000780***	0.0000261***
	(-4.05)	(1.23)	(3.68)	(4.70)
Age	6.077***	1.871***	1.353***	1.530***
	(11.20)	(31.45)	(31.36)	(34.76)
Age^2	-0.0856***	-0.000430***	-0.000332***	-0.000348***
	(-9.88)	(-39.25)	(-46.41)	(-49.76)
Married	-48.81***	-12.95***	-8.872***	-8.000***
	(-51.94)	(-33.21)	(-30.12)	(-28.70)
Children of Age$\leq$ 7	-4.725***	-4.003***	-6.126***	-6.946***
	(-16.39)	(-19.65)	(-30.23)	(-28.71)
Husband's Income	-0.000935***	-0.00246***	-0.000251***	0.000107*
	(-12.68)	(-12.38)	(-3.33)	(2.50)
_cons	-89.29***	-25.99***	-14.75***	-26.88***
	(-10.85)	(-15.45)	(-11.40)	(-20.00)
/				
var(e.Hours)	2441.4***	1057.7***	966.9***	903.3***
	(64.03)	(116.47)	(137.91)	(135.47)
N	41456	61912	81358	80009

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$