Growth with redistribution: Brazil shows the way?

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Abstract

The factors that drove one of the more remarkable spells of redistribution in recent memory are assessed through a comprehensive decomposition approach. Results show that the key ingredients of a rapid and sharp reduction in inequality witnessed in Brazil during the first decade of the new century had been in place for long time. These included a falling dependency ratio, steadily rising schooling levels and importance of women in the labor market. Singular about the 2000s was economic growth associated with falling returns to college and positive changes in the distribution of earnings accruing to men. Whereas in the past the latter served to increase household income dispersion and counterbalance positive structural developments, their impact on distribution was reversed during the decade, thus generating the sharp observed fall in inequality. Conditional and other transfers placed secondary to minor roles.

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1 Introduction

The dawn of new century witnessed one of the more remarkable episodes of redistribution in recent memory in a country known more for being one of the most unequal in the world. After soaring during two decades of scorching growth, fluctuating in the macro-economically turbulent 1980s, and going nowhere in an ensuing period of meager but stable growth, Brazilian inequality diminished sharply and consistently for the first time in recorded history. In a span of less than a decade inequality fell by a fifth and the poverty headcount ratio by an even larger extent. Measures sensitive to developments at the very bottom of the distribution suggest that income changes left very few people behind.

Such impressive developments coincided with implementation of social policies designed to directly address the state of affairs, as conditional and unconditional transfers grew in scope and economic importance. They captured academic, national, international attention, served to equalize the distribution of income and lower poverty (Soares et al., 2009; Ferreira, Leite and Ravallion, 2010), but the degree to which they accounted for aggregate trends may be overstated or far from well understood. Changes also coincided with a long period of sustained economic growth, continued improvement in schooling attainment, secular growth in importance of women in the labor market, and influential demographic shifts, most notably falling dependency rates.

An assessment of the forces behind such remarkable distributive shifts has policy implications that go beyond the case of Brazil. It can shed light on the factors that foster inclusive growth and the overall determinants of distribution. Declining fertility is a global phenomenon. Female labor force participation is following different trends across the world, rising in some areas like Latin America and stalled or declining in others (Elborgh-Woytek et al., 2013), with potentially large impacts on distribution. Education, while vital for augmenting individual and social capabilities, does not possess a solid reputation as an equality-enhancing policy tool.

Results show that ingredients behind the sharp reduction in inequality had been in place for long time. A gradual shift towards households with fewer children relative to adults generated an important demographic dividend in helping reduce income differentials and raise household incomes at the very bottom of the distribution. Similarly, increasing importance of women in the labor market served to reduce income disparities across households and lower poverty, even its more intense manifestation. Steadily rising schooling levels were associated with declining education-income differentials and allowed for an acceleration in college attendance during the 2000s.

During the decade, the college premium began falling across an ample range of the earnings distribution, but the most singular aspect of the period operated through its positive effects on earnings accruing to men. Whereas in the past, male earnings distribution changes raised poverty, inequality, and countered positive structural trends, the economic expansion resulted in a reversion of influence. Therefore, the decade represented a period when changes in earnings accruing to men, earnings accruing to women, education, and demographics operated in the same poverty and inequality-reducing direction. The end result was the large distributive results observed during the period, with 80% of the fall in the headcount ratio and close to half of the fall in intense poverty accounted by demographics, changes in household work structure, schooling, and returns to skill.

The study begins with a survey of the literature on the determinants of distribution in Brazil and a description of the data source and metrics in Sections 2 and 3. Trends in poverty and inequality are laid out in the following section, with the role played by demographics evaluated in Section 5. Section 6 extends the methodological approach to incorporate the potential influence of changes in pension distribution, the introduction of *Bolsa Escola/Familia*, and the role of the labor market and gender. In light of results pinning earnings to the bulk of the positive shifts in distribution, Section 7 examines their source, with special focus on education, skill prices, and changes in household work structure such as the rise in two-income couples. The last section concludes with a summary and discussion of possible lessons for development and distribution.

2 Survey

Scorching growth in the 1960s accompanied by soaring inequality and little impact on poverty fostered great interest in Brazil and intense debates on the relationship between growth and distribution (Fishlow, 1972, 1980; Fields, 1980; Beckerman and Coes, 1980). Langoni (1973) placed education at the heart of the problem, but despite impressive advances over ensuing decades, distribution remained stubbornly unmoved (Barros, Henriques and Mendonça, 2001). Higher attainment reduced dispersion through compression of earnings differentials, but also raised it through population composition effects (Sotomayor, 2004; Menezes, Fernandes and Pichetti, 2006). Through forces that raised residual dispersion, wage inequality rose even in periods when schooling returns fell and education became more equally distributed (Lam and Levison, 1992).

One of these influences was the macroeconomic environment. Income differences increased during the early 1980s Fed recession and shot up even further in the ensuing recovery, with price level changes bordering on hyperinflation also resulting in lower real wages (Hoffmann and Kageyama, 1986; Bonelli and Ramos, 1995; McIntyre and Pencavel, 2004). The advent of a prolonged period of price stability brought forth by the 1994 *Real* Plan was therefore a welcome event in many respects (Neri, 2006). Poverty fell and distribution swings associated with inflationary bursts and stabilization programs became a thing of the past.

Other reforms that accompanied or preceded the plan had additional positive impacts on distribution. Trade liberalization resulted in reduced skill differentials (Gonzaga, Menezes and Terra, 2006). Lower prices associated with freer trade and other market-oriented reforms reduced poverty and income differentials, as their impact was larger at the bottom of the distribution (de Carvalho and Chamon, 2012).

Policies more directly aimed at affecting income distribution had mixed effects. Labor market policies in the form of a higher minimum wage have not been associated with rising incomes at the bottom of the family income distribution (Neumark, Cunningham and Siga, 2005). However, income transfers have had more success. Between the mid-1980s and mid-2000s, when the economy grew little, poverty declined due partly to them (Ferreira, Leite and Ravallion, 2010). Public policy innovations such as conditional transfers have also served to reduce inequality (Soares et al., 2009).[1]

Important questions remain, especially regarding the unusual and impressive progress in distribution experienced over the decade of the 2000s. The relative importance of transfers is still unclear. Pension payments rose markedly during the period but their effectiveness as poverty and inequality-fighting policy tools is unknown. If economic growth was an important part of the explanation, what were the mechanisms that brought about a reduction in poverty and inequality? Did education play a part and are there conditions that foster a positive role in distribution? What does demographics add to the picture and given its structural nature, what lies in store for the future? These are among the questions this study aims to address.

3 Data

Data are derived from the 1977 to 2009 National Household Surveys (Pnads), conducted annually by the Brazilian Institute of Geography and Statistics (Ibge) with interruptions in 1994 for budgetary reasons and in 1980, 1991,

[1]Brazilian data have also shed light on related issues such as the measurement of capabilities and the role of geography and the firm in wage determination (see Bourguignon, Ferreira and Menéndez, 2007; Menezes, Mueller and Ramey, 2008; Fally, Paillacar and Terra, 2010).

and 2000 when the census was fielded. The survey is national in character, but rural areas of the Brazilian Midwest and the Amazon region were not surveyed until 1979 and 2004, respectively. For comparability sake, observations from the latter region are excluded from the analysis. Focus is placed on the post-*Real* plan period and therefore on observations drawn from the 1995-2009 Pnads.

Distribution trends are estimated using equivalent household income constructed by dividing household income by an equivalence scale of the form $(\mathrm{A}_i + kC_i)^s$. A and C refer to the number of adults and children in a household, k to the resource cost of children relative to adults, and s is a parameter reflecting scale economies in the production of household goods. K and s are set to 0.4 and 0.9, respectively, where the former choice is the upper bound of the cost of children relative to adults as estimated by Deaton and Muellbauer (1986) using data from poor countries. The larger share of food costs relative to other expenses in developing economies limits the extent of scale economies, and the choice of s attempts to reflect the limitation. In comparison, the equivalence scale of the UK implies parameter values k=0.53 and s=0.77 (Jenkins and Cowell, 1994) while an alternative US poverty measure explicitly choses k=0.7 and s=0.65–0.75 (Citro and Michael, 1995).

Equivalent income is weighted by the sum of the sample weights of each household member. In doing so, a personal distribution of income is derived under the assumption that each member receives an equal share of the household's equivalent income. Since the assumption may be untenable for certain individuals such as boarders, domestic employees and their relatives, they are excluded from the analysis. There is no official poverty line in Brazil and in place the study compares equivalent income to thresholds developed by Rocha (1997). Incomes are deflated using the National Consumer Price Index (INPC) corresponding to the month when the year's Pnad was carried out. The base period is August 1994, date when the precursor of the current monetary unit was introduced.

4 Trends in inequality and poverty

Until recently, never in recorded Pnad or Census history had Brazilian inequality fallen consistently for an extended period of time. A brief decline in the latter part of the 1970s was followed by rising levels over the 1980s (Figure 1). Macroeconomic stability brought forth by the currency anchor known as the 1994 *Real* Plan had limited effects on inequality, until it began falling in earnest in the 2000s. Bottom and top-sensitive Atkinson inequality indices place the 2001-09 fall at around 20%.[2]

[Figure 1]

In the right panels of Figure 1 poverty follows a volatile pattern, especially when viewed through measures that take into account intensity and distribution aspects of the condition. These reflect substantial sensitivity to the macroeconomic environment, reacting strongly to episodes of recession and high inflation, particularly during the early 1990s when monthly inflation reached 80%. Macroeconomic stability generated a one-time effect on the headcount ratio, until 2003 when it began falling precipitously. Between 2003 and 2009, the headcount ratio fell by 42% while measures that place emphasis on the bottom of the distribution establish declines in the order of 46%.

[Figure 2]

The timing discrepancy between the start of the inequality and the poverty-declining period is explained by the income trends depicted in the top left panel of Figure 2. While incomes at the bottom half of the distribution remained generally unchanged between 1995 and 2003, those at the top eroded steadily, and especially so between 2001 and 2003. The compression generated by the relative income movements is valued by inequality measures as reflecting a fall in dispersion between 2001 and 2003, but is given no weight by poverty indices that are concerned only about developments occurring much further down the distribution. The pickup in economic activity starting in 2003 lifted incomes at all segments of the distribution in such a manner as to produce both a decline in poverty and inequality.

Therefore, it could be argued that the post-*Real* Plan period can be divided into what could be termed as consolidation and growth sub-periods that coincide with the economy's performance in distribution. Growth between 1995 and 2003 was slow and characterized by fits and starts associated with the Asian and Russian financial crises, the recession of the early 21th century, and high interest rates associated with a currency anchor—eventually dropped for an inflation target. Inequality and poverty hardly budged, and Lorenz curve and poverty dominance checks cannot establish unambiguous changes during the period. Growth picked up markedly after 2003 and comparisons establish a fall in inequality according to all Lorenz-consistent measures. Similar checks establish that poverty results are independent of where the threshold is set and the index used to aggregate the condition. That is, the 2009 distribution both Lorenz and first order-dominates that of 2003.

[2] Atkinson measures with higher parameter values place more emphasis on changes at the bottom of the distribution.

The two sub-period distinction is reinforced by an examination of household income distributions, described through kernel density methods. Displayed in the top right panel of Figure 2, they show that there are no large or obvious differences between those corresponding to 1995 and 2003, with the exception of a small increase in density mass at income levels below the poverty line. A comparison of 2003 and 2009 densities is another matter, reflecting a substantial shift in mass to the right. Specifically, differences depicted in the bottom panels of the figure show a marked decline in households with incomes up to twice the poverty line and a concurrent increase in those with incomes between two and six times the threshold.

5 Demographics

Distribution changes can be driven by factors associated with the business cycle, but may also be influenced by structural trends that can magnify or counter shorter-term effects. Demographics can be one such source. As in many other countries, the Brazilian population has over time become older, more urbanized, and constituted in smaller households that are also more likely to be headed by women. Just between 1995 and 2009, the share of the population living in female-headed households rose from 16% to 31%, the urbanization rate increased from 79% to 85%, and the household ratio of children to adults fell by a third.[3]

An assessment of the impact of the phenomena requires a comparison of observed changes in distribution with those that would have occurred in absence of demographic shifts. The latter is a hypothetical proposition but DiNardo, Fortin and Lemieux (1996) offers a counter-factual approach involving re-weighing of sample weights. If households are distinguished by only one trait such as rural or urban setting, generating the distribution that would have resulted if the urbanization ratio had not risen over time calls for giving less weight to urban households in the current distribution and more weight to rural ones. The procedure resembles the Oaxaca-Blinder counterfactual with the important feature of allowing assessment of changes across the whole range of the distribution rather than restricting itself to the means. Akin to it, general equilibrium effects are not taken into account and results may depend on which distribution is used as reference.

Specifically, the procedure can be adapted to the objectives of this study in the following manner, denoting $F(y, z, t)$ as a joint distribution of household income y, characteristics z, and date t. The density of income at a point

[3]The ratio fell from 0.44 to 0.30 between 1995 and 2009.

in time $f_t(y)$ can be expressed as the integral of the density of income conditional on a set of household characteristics on a date t_y over the distribution of characteristics at a date t_z:

$$f_t(y) = \int_{z\in\Omega_z} dF(y, z \mid t_{y,z} = t) = \int_{z\in\Omega_z} f(y \mid z, t_y = t) dF(z \mid t_z = t) \ (1)$$

where Ω_z refers to the domain of attributes. Assuming that the conditionals are not affected by the distribution of characteristics, the year t density that would result if the distribution of characteristics had been as in a period t' is then:

$$\int_{z\in\Omega_z} f(y \mid z, t_y = t) dF(z \mid t_z = t')$$
$$= \int_{z\in\Omega_z} f(y \mid z, t_y = t)\Psi_z(z) dF(z \mid t_z = t) \ (2)$$

$$\text{where } \Psi_z(z) \equiv dF(z \mid t_z = t')/dF(z \mid t_z = t) \ (3)$$

and constitutes the only difference between the expressions for the observed and the counter-factual densities. Thus, a counter-factual density can be obtained from an observed one through use of a re-weighting mechanism with weights $\Psi_z(z)$. When households are characterized by more than one attribute, one may be interested in effects of changes in one trait given that nothing else changes. In this case, the income density can be re-written in the following manner, where the vector z is partitioned into z_1 and z_2, the former containing one household attribute and the latter all others. Suppressing domains:

$$\int\int f(y \mid z_1, z_2, t_y = t) dF(z_1 \mid z_2, t_{z_1|z_2} = t') dF(z_2 \mid t_{z_2} = t)$$

$$= \int\int f(y \mid z_1, z_2, t_y = t)\Psi_{z_1|z_2}(z_1, z_2) dF(z_1 \mid z_2, t_{z_1|z_2} = t) dF(z_2 \mid t_{z_2} = t) \ (4)$$

$$\text{where } \Psi_{z_1|z_2}(z_1, z_2) \equiv dF(z_1 \mid z_2, t_{z_1|z_2} = t')/dF(z_1 \mid z_2, t_{z_1|z_2} = t)(5).$$

$\Psi_{z_1|z_2}(z_1, z_2)$ can be arrived at by estimating $Pr(z_1|z_2, t_{z_1|z_2})$ for dates t and t' through parametric means. Re-weighting the period t distribution is then performed as detailed in the one-attribute case.

[Table 1]

Logits are used to estimate $\Psi_z(z)$ and generate counter-factual trends that hold constant the urbanization ratio, gender and age of the head of household, and households' dependency rates. Results in Table 1 establish that all demographic shifts were either neutral or equalizing in nature, with urbanization and the dependency ratio playing the more prominent roles. Had dependency rates remained unchanged over time, inequality would have declined by 16-17% rather than by the observed 17-19%. The phenomenon therefore accounted for 7% to 9% of the 1995-2009 fall. With the exception of the rise in female-headship, demographic trends also served to raise income and lower poverty. Strongest of all was that associated with falling dependency, which accounted for 18% of the fall in the headcount ratio and 30% of that of the FGT(3) measure.

[Figure 3]

In addition to summary measures, entire distributions can be generated and counter-factual changes differenced from observed ones to arrive at demographic effects. Depictions in Figure 3 shed light on conclusions derived from distribution indices. Falling dependency effects were large and strongest at the bottom of the distribution. Specifically, they were centered at income levels equivalent to one-half of the poverty line, hence the larger effects reflected by income and distribution-sensitive poverty measures. Rising female headship raised the frequency of incomes under the poverty line, but the impact was more than offset by the trends towards households with fewer dependents.

All years under examination were used in constructing a reference distribution, with the end result of producing counter-factuals with the population characteristics of the 1995-2009 time period. However, Table 1 provides in brackets estimates that use the first five and last five distributions as reference. These establish that conclusions are not highly dependent on the choice. For example, the share of the decline in the headcount ratio attributable to falling dependency ranges from 17% to 19%, depending on the reference. In the same manner, the share of the fall in the A(.5) index ranges from 7.2% to 7.8%.

Altogether, results point to the fact that distribution trends depicted in Section 4 operated under substantial tail winds originating from favorable demographic shifts. Counter-factuals that hold constant all factors at once establish that effects accounted for up to 16% and 25% of the observed falls in inequality and poverty, respectively. Their influence is far from trivial and when evaluating the forces behind the balance of the progress in distribution, demographic effects need to be held constant. The DFL methodology provides a straightforward solution, through use of the re-weighted distributions when accounting for additional sources of change.

6 Income sources

Conditional transfers, pensions, and earnings appear as potentially important factors and their assessment requires a methodological approach that can accommodate examination of the impact of changes in income source distribution. To that end, the study combines the re-weighting approach with a rank-preserving income exchange mechanism. Adapted to the household setting by Cancian and Reed (2001), implementation involves a first stage that consists of summarizing a distribution of (say) year t pensions by 1000 points. To do so, each individual i in a given age range or other agreed upon population characteristic is associated with a ranking R_{ti} that gives the person's millicile in that distribution. Coupled with the rankings are the corresponding pension income levels calculated as the mean of all individuals sharing the same millicile in year t: $\mu_t(r)$ with $r \in 1, 2, \ldots, 1000$. The second stage involves the construction of a household income HY_{tf} distribution as follows, where q_{tf} is a list of the pension income rankings in household f, OY_{tf} represents the sum of household income from sources other than pensions, and $e(A_{tf}, C_{tf})$ is an equivalence scale that uses information on the number of adults and children in a household.

$$\mathrm{HY}_{1f} = [\Sigma_{j\in[q_{1f}]}\mu_1(R_{1j}) + OY_{1f}] \times e(A_{1f}, C_{1f})$$

Total income of household f in year t=1 is the sum of pension income of household members in an agreed upon age range or characteristic—where each is assigned a pension corresponding to his ranking in the pension distribution—plus household income from all sources other than retirement income. In practice, the 1000-point approximation of the distribution of pension income turns out to be more than adequate, as Atkinson inequality measurements of actual and "summarized" household income distributions are identical up to three decimal places. However, the utility of the procedure is not reproducing distributions, but simulating the distribution that would have obtained if only the dispersion of an income source had remained as in another period. That is, the procedure consists of constructing a year 2 counter-factual distribution

$$\mathrm{HY}_{2f} = [\Sigma_{j\in[q_{2f}]}\mu_1(R_{2j}) + OY_{2f}] \times e(A_{2f}, C_{2f}).$$

Total household income in year t=2 is the sum of pension income of all pensionable age household members—estimated with year 2 rankings but the corresponding pension levels of the distribution in year 1—plus year 2 household income from all sources other than pensions. Any discrepancy between an observed and a counter-factual change in inequality can be attributed to changes in the distribution of the income source being held constant. For example, if a counter-factual associated with pension distribution produces an inequality trend that is 10% smaller than the observed change, then pensions account for 10% of the aggregate change in dispersion.

Conclusions can however depend on the choice of examination period and little light is shed on processes that could have developed, diminished, or changed course in between. Visual evidence, often a useful complement to numerical results, is also lacking. To address these limitations, the approach is extended to allow for the construction of counter-factual distribution trends. To do so, rather than using beginning, end, or other period distributions as reference, all are employed in its construction. The counter-factual definition then changes from how inequality would have behaved if the distribution of an income source had remained as in a given year, to how inequality would have behaved if the distribution of the income source had remained unchanged at its mean over time. Counter-factuals are estimated for each period and used in constructing trends. Throughout, demographics is held constant using re-weighted distributions, where the the aim is that of adding explanatory power beyond that already established.

Some drawbacks remain, the first being that the approach does not provide for exact decompositions of a change in distribution. Also, while the reference distribution is an average over all possibilities, it is still a choice and results can differ from those drawn from other references such as beginning or end-period distributions. Last, as with any decomposition procedure, possible interactions within the household or general equilibrium effects are not taken into account.[4]

Yet, the procedure presents a number of advantages. It is capable of distinguishing between changes in the distribution of income sources and changes in household structure such as that related to marriage, simply replacing male (female) earnings distributions with those prevailing in another year. In contrast, in inequality index decompositions by income source, households headed by unmarried females are assigned zero for male earnings and hence increased frequency of those types of units translates to rising male earnings inequality. Similarly, an increase in the share of households without children results in increased conditional transfer inequality. A second advantage of the approach is that simulation of entire distributions allows examination of impact on income densities and summary measures, thus permitting a more nuanced and robust understanding of the nature of the changes in dispersion. For the same reason, poverty analysis can be incorporated within the framework in a straightforward manner. Third, the method it is geared towards understanding changes rather than levels as is the case of more commonly used source decompositions.[5] Fourth, it is independent of measurement unit, hence allowing use of per capita or equivalent household income definitions rather than the total household income measure required by most income source decomposition methods. It is also non-parametric in

[4]For example, links have been established between male unemployment and female labor force participation and between household wealth and child labor (Fernandes and de Felício, 2005; Soares et al., 2012).

[5]Soares et al. (2009) is a recent example of a Gini decomposition by income source (Fei, Ranis and Kuo, 1978).

nature, requiring few assumptions about income-generating mechanisms and last, it is amenable to use in conjunction with the DFL framework, hence allowing for a more comprehensive accounting of the forces behind distributive change.

6.1 Retirement income

One such source is pension income, whose distribution is a major contributor to the high levels of household income inequality observed in Brazil relative to other countries such as the US (Bourguignon, Ferreira and Leite, 2008). In being linked to a minimum wage that rose markedly through the decade of the 2000s, higher pensions or changes in their distribution could have had important effects on both poverty and inequality—serving to increase or decrease it in the latter case. To evaluate the role they played, counter-factual trends are constructed to depict how inequality and poverty would have behaved in absence of changes in their distribution.[6]

[Figure 4]

Results in the top panels of Figure 4 establish that pensions played an equalizing role throughout the post-*Real* Plan period. Observed inequality falls faster than counter-factual trends and as a result, counter-factual to observed ratios depicted in the figure rise over 1995-2009. Without changes in pension distribution and demographics, inequality would have declined by 12% rather than by the 15-16% that obtains when only demographics is held constant.[7] The source therefore accounted for 17% to 22% of the long-term fall, with the higher-range estimate corresponding to the more top-sensitive A(0.25) measure (Table 2).

[Table 2]

The source's effect on poverty was more substantial, but only when viewed through the headcount ratio. That is, retirement income accounted for 17% of the fall in the headcount ratio, but only 7% of the fall in poverty viewed through a measure that takes into account intensity and distributive aspects of the condition. Large pension changes do not appear to have accrued to individuals at the very bottom of the income distribution. The main ex-

[6] Retirement income refers to old age (*aposentadoria*) and other regular payments (*pensão*) but both sources are referred to inter-changeably as pensions or retirement income. Pensions can be drawn relatively early in life in Brazil and the distribution of pension income is therefore held constant among the population of at least 45 years of age.

[7] To avoid clutter and repetition, "observed" and "counter-factual" changes refer henceforth to observed and counter-factual changes that net out demographic effects.

planation for the remarkable fall in poverty, and especially its more intense manifestation, lies somewhere else.

6.2 Transfers

In 2001 Brazil introduced at the federal level a program offering a small income supplement on a number of conditions. *Bolsa escola* paid a monthly allowance of about US$5 per child to low-income households that agreed to register, ensure minors' school attendance and scheduled vaccinations. Two years later, the program and other initiatives providing small sums intended for food and cooking gas purchases were folded into a new one named *Bolsa famĩlia* that also provided an unconditional income supplement to very low-income households, regardless of family composition. Over time, emphasis has shifted from conditions to transfers and its evolution has not been devoid of controversy due to the program's potential for generating dependency and political patronage (Hall, 2008), and its high costs relative to efficiency benefits (Glewwe and Kassouf, 2012).

Evaluation of its distributive impact is hindered by the fact that there is no dedicated entry for *Bolsa escola/famĩlia* income in the Pnad survey. Rather, it is classified as "other" income, along with interest and dividends. The situation is not hopeless, since program rules can be exploited for getting a better measure of the income source. A threshold approach could re-code to zero "other" income from households not meeting the program's qualification rules. A stipend approach could exploit the fact that benefits are standard sums, for example, *Bolsa escola* and basic family allowances being R$15 and R$50, respectively. Both thresholds and stipends have been revised since 2006, but are published and straightforward to identify.

An approach combining threshold and stipend rules is used for identifying *Bolsa escola/familía* income and generating ratios of counter-factual to observed distribution trends presented in the bottom panels of Figure 4. These show that conditional transfers had positive but limited effects on inequality. That is, observed inequality declines more quickly than the counter-factual trend, but only slightly more so. The income source also served to reduce poverty, with effects concentrated at the very bottom of the distribution. That is, counter-factual to observed poverty ratios rise by 6% with respect to the headcount ratio, but by 13% and 15% with respect to the FGT(2) and FGT(3) indices, respectively.

The program's more substantial effects on poverty are not unexpected as it is targeted to low-income households, defined in 2009 as those with per capita income under R$100. The figure is close to the poverty line used in this study, and the income supplement can lift a household above the threshold. Yet, the fact that its impact is centered at the very bottom of the distribution is indicative of its targeting efficacy. In all, 9% to 13% of the fall in inequality and 11% of the decline in the headcount ratio can be attributed to the program. Its contribution towards alleviating intense poverty is far larger and equivalent to 30% to 42% of the change in the condition.

6.3 Earnings

Earnings accruing to male and female household members are important sources of income that have followed different paths over time. While the share of the former in household income has fallen from 60% in 1995 to 50% in 2009, that of the latter has risen from 22% to 26%. At the same time, earnings from women have become less unequally distributed and those of men less correlated with household income (Table 3).

[Table 3]

The top panels of Figure 5 demonstrate that earnings distribution changes among men raised poverty and inequality up to the early 2000s. Effects on poverty were of much greater magnitude and reflected developments that were especially detrimental when evaluated through measures sensitive to changes at the very bottom of the distribution. Counter-factual to observed poverty ratios fall by 21% in the case of the headcount measure, but by 35% and 39% in the case of the FGT(2) and FGT(3) indices, respectively. Without variation in the distribution of male earnings, poverty would have fallen substantially over 1995-2003.

[Figure 5]

Much changed in the early 2000s. From that moment on, counter-factual to observed distribution ratios begin rising, signifying that changes in male earnings became both equalizing and poverty-reducing, although the ground lost in the earlier period is made up only as reflected by the headcount ratio. That is, they accounted for 22% of the decline in the headcount ratio over 1995-2009, but -9% and -34% of the drop in the FGT(2) and FGT(3) indices, respectively.

In contrast, the role of women in distribution was consistently positive. Inequality trends shown in the bottom panels of Figure 5 rise through the entire post-*Real* plan period, and poverty ones more markedly so after 2003. Therefore, while the 2000s represented a turn of influence for men, for women they represented an accentuation of their positive role on distribution. In all, more than a third of the fall in inequality experienced over 1995-2009 can be accounted by changes in their earnings distribution, followed in order of importance by male earnings, pensions, and conditional transfers—the latter two income sources responsible in combination for another third of the fall. On the poverty front, female earnings effects were far more pronounced. Associated with about half of the fall in the headcount ratio, they accounted for the bulk of the fall in intense poverty. Along with transfers, the income source helped offset long-term poverty and inequality-enhancing developments in the distribution of male earnings.

Table 2 also presents in brackets decomposition results that use as reference beginning and end-period distributions rather than the mean over all periods. Alternative results generally center on the stated ones. For example, pensions' contribution to the 1995-2009 fall in the headcount ratio range from 15% to 17%, and from 19% to 22% with respect to the change in the A(.50) index. Corresponding numbers for shares accounted by conditional transfers are 7-11% and 6-10%. Bands are wider for results regarding contributions of female earnings, showing somewhat stronger ones when using more current distributions as reference. However, rankings and basic conclusions remain unchanged. The bulk of the positive movements in short and long-term distribution originated in the labor market. Depending on the reference distribution, female earnings were responsible for 41% to 52% of the long-term decline in the headcount ratio and a much larger share of the fall in intense poverty. Earnings changes among men turned into a positive distributive force during the 2000s, and along those of women accounted for about two-thirds to three-quarters of the impressive advances in distribution witnessed during the decade.

7 Education, skill prices, and household work structure

In light of the central role of the labor market in diminishing income disparities and even reducing intense poverty, possible pathways are explored, with education being a prime suspect. Attainment has been rising for a long time (Figure 6), and over the post-*Real* Plan period mean schooling increased from 5.6 to 7.7 years among the the population aged 25 to 64. There is evidence of an acceleration in high school completion and college achievement, though the latter is restricted to the 2000s. Other things equal, returns can be expected to decline.

[Figure 6]

In Figure 7 skill price trends are depicted and established through earnings function that regress log monthly earnings on education and age polynomials as well as household status and metropolitan area residence binaries. Prices are allowed to vary across the earnings distribution through use of quantile regressions that estimate returns at the median and at the top and bottom deciles and quartiles. Since the objective is observing the behavior of skill prices and evaluating their potential effects on distribution, rather establishing returns to education, strictly speaking, samples are unrestricted in including all individuals with positive earnings.

[Figure 7]

The top left panel shows returns to an additional year of schooling fluctuating with the macroeconomic environment but trending downward, then falling more consistently after price stability was achieved. They also "fan out," in the sense that with the exception of those at the bottom decile, prices fall more at the lower than upper part of the earnings distribution. For example, between 1995 and 2009 the return to an additional year of school falls by 20% at the 25th percentile, but by 8% at the 90th.

Disaggregation by degree establishes dramatic declines in returns to intermediate and high school, and more so during the 2000s. At the median of the earnings distribution, the intermediate school premium falls from 45% in the mid-1990s and 38% in the beginning of the 2000s, to 29% by the end of the decade. At the bottom decile of the distribution, however, returns rise markedly—a phenomenon possibly associated with changes in the value of the minimum wage. Very similar numbers obtain for the return to high school completion, with the exception that the rise in prices at the bottom is much less pronounced.

The college premium is another matter, in showing a slight upward trend until the 2000s. Thereafter, it begins falling everywhere but the top of the distribution where it remains rather unchanged. Further down, it falls by 3% at the top quartile, 10% at the median, and 35% at the bottom decile. The decade witnessed a pickup in college completion and Fortin (2006) finds a close link between labor supply and the college premium. However, an explanation for the fanning out of returns is harder to come by. It could reflect increased skill heterogeneity associated with a rapidly expanding tertiary sector, or increased discrimination of ability differences on the part of employers. Whatever the explanation, it is clear that Brazilian returns to college begin to decline for the first time at an ample range of the distribution, and to rise less rapidly or remain constant at the top.

[Table 4]

Besides education, another potential labor market mechanism capable of generating greater equality and lower poverty is labor force participation, with changes in the number and composition of workers in a household unit capable of affecting absolute and relative incomes. To evaluate their potential for impact, units are classified into eight categories using work status and gender of the head of household and spouse, if there is one. The first four refer to households headed by unmarried individuals (or without spouses present) and classified as headed by working or non-working males or females. The remaining four categories refer to households with couples: where both individuals work, both do not work, where only the male works, or only the female works. Work status and gender of other members is not taken into account in order to focus on the more likely main income-earners and keep classifications at a manageable size.

Table 4 presents evidence of what appears to be a structural decline in units with couples where the male works outside the home and the female does not. While 42% of the population lived in such a household in 1995, 30% did so less than a decade and a half later. Some share of the population transitions to units headed by single women, a trend that is slow and structural, but another share transitions to households where both the husband and wife work. The latter phenomenon appears cyclical and one of the defining characteristics of the 2000s, as it is largely exclusive to the period. Between 1995 and 2001 the share of the population living in such units held steady at around 25%, but rose to 30% by 2009.

7.1 Sequential decomposition

The rank-preserving income exchange methodology took household work structure as given, but its influence on distribution can be evaluated through the approach described in Section 5, where equation 3 is estimated with household type binaries as covariates. Noting Ψ_D and Ψ_S as the re-weighting functions for demographics and household work structure, respectively, $\Psi_D * \Psi_S$ holds both demographics and household work structure to the levels of a reference distribution. Marginal impacts can be arrived at through a sequential decomposition described in Table 5.

[Table 5]

Its first and last rows state that initial and final-period income densities and summary statistics can be computed using data and household weights provided by the surveys, with their percentage difference constituting the observed change in distribution. To account for the role of demographics, the initial-period data is weighted by an Ψ_D that sets demographics at their final-year level. The percentage difference between this counter-factual and the initial distribution is calculated and if equivalent to (say) 10% of the observed change, demographics is responsible for that share of the trend.

Similarly, the impact of household work structure is assessed re-weighing the initial data through use of $\Psi_D * \Psi_S$, where household work structure is also held at its final year levels. Following the same numerical example, if the percentage difference between the counter-factual and the initial distribution is equivalent to 12% of the observed change, household work structure drives 2% of the trend.

Evaluation of education's impact could be performed in a similar manner, but assigning the schooling attributes of another period would not account for important correlations with variables such as age, gender, or geographic setting, whose effects can be confounded with schooling. Sample re-weighting can be conditioned on the relevant covariates, with the pertinent re-weighing function described in equation 5. Estimated through an ordered logit, Ψ_E sets schooling levels of heads of household and spouses, conditioned on z_2, at their final-year levels. The impact of education on distribution can then be ascertained as in the case of demographics and household work structure.

Finally, getting a grip on the influence of skill prices requires their estimation. If returns at the mean were an adequate reflection of those elsewhere in the distribution, OLS would suffice as estimator. However, Brazilian skill prices appear increasingly dependent on position in the distribution and to accommodate this feature of the country's labor market, earnings distributions are constructed using a quantile regression approach (see Machado and Mata, 2001; 2005 and Appendix). Skill prices are arrived at using the final-year sample, with the dependent variable being earned income of the household head and spouse, if there is one. Covariates include quadratics in mean schooling and age of working heads of household and spouses, fraction that is female fully interacted with age and schooling, as well as binaries for metropolitan area residence and working couples. Observed earnings of heads of household and their spouses are then replaced with estimated earnings arrived at using initial period covariates but final-year skill prices. Estimated earnings and re-weights $\Psi_D * \Psi_S * \Psi_E$ are then used to generate counterfactuals.

[Table 6]

Decomposition results in Table 6 show that both education and skill price changes proved equalizing, with the latter playing the far stronger role. Their combined influence operated most prominently at the higher end of the distribution, in accounting for close to half of the fall of the more top-sensitive A(.25) index versus a third with respect to the A(.75). Skill price and education effects on distribution appear even more impressive in light of the fact that they only reflect impact through working heads of household and spouses rather than all income-earners and pertain to the global distribution of income and not one restricted to economically active households. The flip side of generally falling skill prices is that over 1995-2009 they put pressure on incomes and poverty. However, increased schooling attainment more than offset the negative price effects and was the largest contributor to the marked fall in poverty evident during the period under examination. Close to half of it can be attributed to increased schooling among working heads of household and their spouses.

Work structure shifts were a mixed bag over 1995-2009, with increased frequency of both unmarried heads of households and two-income couples. Over the 2000s however, the latter trend dominated and was a small contributor to the poverty decline evident during the decade. Its limited impact on distribution suggests that most positive changes occurred within rather than across household types. That is, earnings changes associated with schooling raised incomes of one and two-income couple households, lifting many from poverty. Changes occurring within these households also resulted in sharply reduced income disparities. Demographic effects were not far behind in importance, as decomposition results confirm their role, second to education in generating the impressive poverty decline.

However, the stated results are contingent on the order in which effects are estimated. The decomposition is sequential, unlike that employed in Section 6. It could be performed in reverse order, but many other possibilities would remain unexplored. To account for as much information as possible, decompositions are performed for all possible sequences, with the average over all 4! decompositions presented in the second panel of Table 6. These confirm that schooling was by far the most important poverty-reducing factor and falling skill prices the most important dispersion-reducing phenomenon. Degrees of effects change, with demographics losing and skill prices gaining explanatory power.[8] Averaged over all possible decomposition sequences, a remarkable 80% of the fall in the headcount ratio and close to half of the fall in intense poverty occurring during the 2000s can be accounted by demographics, changes in household work structure, schooling, and returns to skill. Demographics, education, and returns to labor explain 40% to 60% of the drop in inequality, with more top-sensitive measures reflecting the stronger effects.

[Figures 8-9]

Ratios of counter-factual to observed distribution trends presented in Figures 8 and 9 also demonstrate the factors' impact on distribution when each is evaluated first in the sequential decomposition order. Observed trends fall faster than most counter-factuals, and ratios rise steadily over time. Such is the case with education's consistently positive effect on poverty, skill price ef-

[8]The price component picks up explanatory power from demographics when it precedes the latter in sequential order.

fects on inequality, and demographics positive influence on both. In contrast, the impact of skill prices on poverty changes direction during the 2000s, when rising prices begin to lower poverty and account for a third of the fall in the headcount ratio over the decade. At first glance the phenomenon appears in contradiction to declining skill returns established earlier, but Figure 7 shows that these did not fall everywhere in the earnings distribution. High school and especially intermediate schooling premia rose at the bottom decile. Although more markedly after the mid-2000s, the minimum wage has risen throughout the post-*Real* period and could have played a role (Figure 10).

[Figure 10]

A binding wage floor is evident through returns to primary school that fall dramatically since the mid-1990s, but not at the 10th and 25th percentiles (Figure 10). Further down the skill distribution, the base pay for labor as reflected by the value of the intercept term of the earnings function begins to rise in the early part of the 2000s. Consistent with a minimum wage that becomes increasingly more binding in covered and compliant sectors, its variability across the distribution is also reduced.

Additional evidence in the bottom right panel shows log earnings by schooling that begin to rise in 2003 and more so at the bottom of the skill distribution. That is, while earnings among high school graduates rise by 12%, those of individuals with primary of no formal schooling increase by 17% and 20% respectively, hence the positive effect of skill prices on both absolute and relative incomes during the decade. The development coincides with the beginning of the impressive economic expansion of the 2000s and it is difficult to assess the relative importance of the wage floor versus economic growth. Both likely played a role, with the latter certainly allowing for a more rapid expansion of the former.

8 Conclusion

After soaring during decades of formidable growth, fluctuating in the macroeconomically turbulent 1980s, and stagnating in an ensuing stable but low-growth environment, Brazilian inequality fell consistently in the 2000s for the first time in recorded history. Millions left poverty and measures that place emphasis on the bottom of the distribution suggest that few households were left behind. These impressive changes occurred during a period of rising pensions and conditional transfers, but also under the influence of an expanding economy. They took place in the context of long-standing developments related to demographics, schooling, and growing importance of women in the labor market. The study incorporated a comprehensive decomposition approach to isolate effects on distribution and ascertain potential lessons for development and distribution.

Results demonstrate that the impressive trends in distribution manifested over the decade of the 2000s benefited from developments that had been brewing for some time, the first being substantial tailwinds provided by favorable demographics. Most prominent was falling dependency, with a shift towards households with fewer children relative to adults generating an important demographic dividend in helping alleviate intense poverty as well as inequality. A second structural trend involved the rising importance of women in the labor market. Throughout the post-*Real* Plan period, changes in the distribution of earnings accruing to women served to reduce income disparities across households and lower poverty, even its more intense manifestation.

Markedly rising schooling attainment also exerted an equalizing influence, especially in a context of a stable macroeconomic environment. Returns to intermediate and high school fell and even the college premium began to fall during the 2000s at an ample range of the earnings distribution. McIntyre and Pencavel (2004) posit that in periods of high inflation real wages may be bargained in relative terms, hence the lack of relationship between inflation and inequality the authors establish in the Brazilian case. If in fact the relationship holds, as inflation recedes, wages may gradually cease to be set in multiples of a reference such as the minimum-wage. Alongside rapidly rising education levels, this de-linking could be a factor behind the marked decline in education earnings differentials observed since price stability was achieved.

Yet, during the 2000s the most notable distributive development concerned the relationship between the male labor market and the distribution of household income. Whereas in the past changes in the distribution of earnings accruing to men had substantial and negative effects on poverty and household income disparities, the economic expansion resulted in a reversion of influence. The decade then represented a period when changes in earnings accruing to men, earnings accruing to women, education, and demographics operated in the same poverty and inequality-reducing direction. The source of the turn of events for men is a subject beyond the scope of this study, but the strong economic upturn was characterized by expansions of male-dominated sectors such as manufacturing, construction, and a rise in male labor force participation that until then was in secular decline.

The elements behind equalizing growth appear then to emerge. They include secular trends in the form of demographics and continued rising importance of women in labor market, but also factors that generated expansions across economic sectors. Increased schooling and macroeconomic stability reduced inequalizing pressures in the form of greater education-earnings differentials. In all, demographics, education, and falling returns to labor explain 40% to 60% of the drop in inequality, with more top-sensitive measures reflecting the stronger effects. A remarkable 80% of the fall in the headcount ratio and close to half of the fall in intense poverty accounted by demographics, changes in household work structure, schooling, and returns to skill. Pensions and conditional transfers lowered poverty and inequality but in relative terms played secondary roles.

Of concern is the replicability of economic growth experienced during the 2000s and driven to a large extent by Asian demand. Moreover, the fact that it took such a strong growth spell to produce equalizing and poverty-reducing changes in male earnings suggests that much may be undone in an extended period of low growth. In such circumstances, pressure for continued progress in distribution through transfers should be weighted against potential disincentive effects.[9] Presented evidence suggests that policies facilitating education and female labor force participation can have large impacts on distribution and achieve equity as well as efficiency goals.

9 Bibliography

Barros, R., Henriques, R., Mendonça, R., 2001. A estabilidade inaceitável: Desigualdade e pobreza no Brasil. Instituto de Pesquisa Econômica Aplicada Discussion Paper No. 800. Rio de Janeiro: Instituto de Pesquisa Econômica Aplicada.

Beckerman, P., Coes, D. Who benefits from economic development: Comment. American Economic Review 70(1), 246–49.

Bonelli, R., Ramos, L., 1995. Distribuição de renda no Brasil: Avaliação das tendências de longo prazo e mudanças na desigualdade desde meados dos anos 70. Revista Brasileira de Economia 49(2), 353–73.

Bourgignon, F., Ferreira, F., Menéndez, M., 2007. Inequality of opportunity in Brazil. Review of Income and Wealth 53(4), 585–618.

Bourgignon, F., Ferreira, F., Leite, P., 2008. Beyond Oaxaca-Blinder: Accounting for differences in household income distributions. Journal of Economic Inequality 6(2), 117–48.

Burtless, G., Sotomayor, O., 2006. Labor supply and public transfers. In: Bosworth, B., Collins, S., Soto Class, M. The economy of Puerto Rico:

[9]See de Carvalho (2008) for an analysis of the effect of pension rules on retirement behavior in Brazil and Burtless and Sotomayor (2006) for an analysis of the effect of transfers on labor force participation in another middle income economy.

Restoring growth. Washington DC: Brookings Institute Press.

de Carvalho, I., 2008. Old age benefits and retirement decisions of rural elderly in Brazil. Journal of Development Economics 86(1), 129-46.

de Carvalho, I., Chamon, M., 2012. The myth of post-reform income stagnation: Evidence from Brazil and Mexico , Journal of Development Economics 97(2), 368-386.

Citro, C., Michael, R., 1995. Measuring poverty: A new approach. Washington DC: National Academy Press.

Deaton, A., Muellbauer, J., 1986. On measuring child costs: With applications to poor countries. Journal of Political Economy 94(4), 720–44.

DiNardo, J., Fortin, N., Lemieux, T., 1996. Labor market institutions and the distribution of wages, 1973–1992: A semi-parametric approach. Econometrica 64(5), 1001–44.

Elborgh-Woytek, K., et al., 2013. Women, work, and the economy. IMF Discussion Note, Washington DC.

Fally, T., Paillacar, R., Terra, C., 2010. Economic geography and wages in Brazil: Evidence from microdata. Journal of Development Economics 91(1), 155–68.

Fei, J., Ranis, G., Kuo, S., 1978. Growth and the family distribution of incomes by factor components. Quarterly Journal of Economics 92 (1), 17–53.

Fernandes, R., de Felício, F., 2005. The entry of the wife into the labor force in response to the husband's unemployment: A study of the added worker effect in Brazilian metropolitan areas. Economic Development and Cultural Change 53(4), 887–911.

Ferreira, F., Leite, P., Ravallion, M., 2010. Poverty reduction without economic growth. Journal of Development Economics 93(1), 20–36.

Fields, G. 1980. Who benefits from economic development, American Economic Review 70(1), 242–45.

Fishlow, A., 1972. Brazilian size distribution of income. American Economic Review 62(2), 391–402.

Fishlow, A., 1980. Brazilian size distribution of income: Comment. American Economic Review 70(1), 250–6.

Fortin, N., 2006. Higher-education policies and the college wage premium: Cross-state evidence from the 1990s, American Economic Review 94(4), 959-87.

Glewwe, P., Kassouf, A., 2012. The impact of the *Bolsa Escola/Familia* conditional cash transfer program on enrollment, dropout rates and grade promotion in Brazil, Journal of Development Economics 97(2), 505-17.

Gonzaga, G., Menezes, N., Terra, C., 2006. Trade liberalization and the evolution of skill earnings differentials in Brazil. Journal of International Economics 68(2), 345–67.

Hall, A., 2008. Brazil's Bolsa Família: A double-edged sword? Development and Change 39(5), 799–822.

Hanushek, E., Woessman, L., 2012. Schooling, educational achievement, and the Latin American growth puzzle , Journal of Development Economics 99(2), 497-512.

Hoffmann, R., Kageyama, A., 1986. Distribuição da renda no Brasil, entre famílias e entre pessoas, em 1970 e 1980. Estudos Econômicos 16(1), 25–51.

Jenkins, S., Cowell, F., 1994. Parametric equivalence scales and scale relativities. The Economic Journal, 104(425), 891–900.

Lam D., Levison, D., 1992. Declining inequality in schooling in Brazil and its effects on inequality in earnings. Journal of Development Economics 37(1/2), 199–225.

Langoni, C., 1973. Distribuição de renda e desenvolvimento econômico do Brasil Rio de Janeiro: Expressão e Cultura.

Machado, J., Mata, J., 2001. Earnings functions in Portugal 1982-1994: Evidence from quintile regressions, Empirical Economics 26(1), 115–34.

Machado, J., Mata, J., 2005. Counterfactual decomposition of changes in wage distributions using quantile regression. Journal of Applied Econometrics 20(4), 445–465.

McIntyre, F., Pencavel, J., 2004. The effect of macroeconomic turbulence on real wage levels and the wage structure. Economic Development and Cultural Change 52(4), 681–717.

Menezes, N., Fernandes, R., Pichetti, P., 2006. Rising human capital but constant inequality: The education composition effect in Brazil. Revista Brasileira de Economia 60(4), 407–24.

Menezes, N., Mueller, M., Ramey, G., 2008. The structure of worker compensation in Brazil, with a comparison to France and the United States. Review of Economics and Statistics 90(2), 324–46.

Neri, M., 2006. Miseria, desigualdade e estabilidade: O segundo Real. Rio de Janeiro: Fundaçao Getúlio Vargas.

Neumark, D., Cunningham, W., Siga, L., 2004. The effects of the minimum wage in Brazil on the distribution of family incomes: 1996–2001. Journal of Development Economics 80(1), 136–59.

Reed, D., Cancian, M., 2001. Sources of inequality: Measuring the contributions of income sources to rising family income inequality. Review of Income and Wealth 47(3), 321–33.

Rocha, S., 1997. Do consumo observado à linha de pobreza. Pequisa e Planejamento Econômico 27(2), 313–52.

Soares, S., Osório, R., Soares, F., Medeiros, M., Zepeda, E., 2009. Conditional cash transfers in Brazil, Chile and Mexico: Impacts upon inequality. Estudios Económicos (Special Issue), 207–24.

Soares, R., Kruger, D., Berthelon, M., 2012. Household Choices of Child Labor and Schooling: A Simple Model with Application to Brazil. Journal of Human Resources 47(1): 1–31.

Sotomayor, O., 2004. Education and changes in Brazilian wage inequality, 1976-2001. Industrial and Labor Relations Research Review 58(1), 94–111.

Table 1—Percentage share of 1995-2009 fall in distribution index accounted by demographics

Factor	Inequality/Poverty Index					
	A(.25)	A(.50)	A(.75)	FGT0	FGT2	FGT3
Age of head	0	1	2	4	7	9
	[0, -1]	[1, 0]	[2, 2]	[5, 3]	[8, 6]	[10, 9]
Gender of head	3	3	2	-5	-8	-9
	[4, 3]	[3, 3]	[2, 3]	[-6, -4]	[-10, -6]	[-12, -7]
Urbanization	7	7	6	3	4	5
	[5, 9]	[5, 8]	[5, 7]	[2, 3]	[4, 5]	[4, 6]
Dependency ratio	7	8	9	18	25	30
	[7, 7]	[8, 7]	[9, 9]	[19, 17]	[24, 26]	[28, 31]
All	15	15	15	13	17	19
	[13, 15]	[13, 15]	[13, 16]	[13, 13]	[14, 21]	[14, 25]

Note—Characteristics are held constant at their 1995-2009 average. Also shown in brackets are counter-factuals that hold them constant at their 1995-99 (left bracket) and 2005-09 means (right bracket).

Table 2—Share of change in distribution index accounted by pensions, conditional transfers, male and female earnings

1995-2009	**Share (%) of fall in index accounted by income source:**					
	A(.25)	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Pensions	22	19.8	17.1	16.4	10.7	7.2
	[24, 22]	[22, 19]	[19, 15]	[17, 16]	[12, 5]	[9, 0]
Transfers	6.2	7.6	10.2	8.9	27.7	39
	[8, 5]	[10, 6]	[13, 8]	[11, 7]	[38, 18]	[55, 26]
Female earnings	36.3	37.1	39.6	46	80.2	96.7
	[26, 39]	[29, 42]	[31, 42]	[41, 52]	[72, 84]	[90, 97]
Male earnings	22.9	24.2	21.4	22.1	-8.8	-33.4
	[26, 32]	[28, 30]	[26, 26]	[22, 19]	[-8,11]	[-29, -35]

2001-2009	**Share (%) of fall in index accounted by income source:**					
	A(.25)	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Pensions	12.3	11.2	9.2	7	1.4	0
	[13, 11]	[12, 9]	[11, 6]	[7, 5]	[3, 0]	[1, 0]
Transfers	5.9	7.4	9.3	6.4	15.4	18.9
	[7, 5]	[9, 6]	[12, 8]	[8, 5]	[20, 11]	[26, 14]
Female earnings	28.2	28.3	28.8	27.4	35.4	37.5
	[19, 31]	[21, 31]	[21, 32]	[26, 30]	[32, 38]	[34, 40]
Male earnings	36.3	39.4	40.9	43.8	46.3	44.3
	[38, 36]	[41, 39]	[42, 40]	[43, 42]	[44, 47]	[42, 45]

Note—Income source distributions are held constant at their 1995-2009 mean. Also shown in brackets are counter-factuals that hold them constant at their 1995-99 (left bracket) and 2005-09 means (right bracket).

Table 3—Source share in household income, inequality in distribution, and correlation with total household income

	Male Earnings	**Female Earnings**	**Pensions**	**'Other' Income**
Income share				
1995	.630	.214	.125	.008
2001	.571	.235	.164	.009
2003	.555	.237	.175	.011
2009	.530	.257	.176	.020
Inequality				
1995	.381	.637	.779	.976
2001	.413	.629	.752	.955
2003	.412	.615	.739	.897
2009	.403	.572	.726	.798
Correlation				
1995	.882	.740	.583	.745
2001	.854	.777	.847	.417
2003	.842	.746	.603	.230
2009	.831	.742	.566	-.136

Note—Inequality is measured through use of the A(.50) index.

Table 4—Distribution of the population by work status and gender of the head of household and spouse

Work status of head of household and spouse	Year 1995	1998	2001	2003	2006	2009
Single non-working male	1.1	1.3	1.4	1.5	1.5	1.7
Single working male	2.8	2.9	3.0	3.0	3.3	3.4
Single non-working female	7.9	8.9	9.5	9.7	9.9	10
Single working female	7.6	7.8	8.5	9.0	9.6	9.9
Working male, non-working female	43.6	41	38.1	36.2	32.3	30.5
Working female, non-working male	3.3	3.7	4	4.3	4.2	4.2
Working male and female	25.1	25.5	25.5	26.3	28.7	29.7
Non-working male and female	8.6	10.2	10.1	10	9.9	10.1

Table 5—Sequential decomposition

Date of survey	Counterfactual holds constant:	Income definition	Re-weighting factor
1995/2001	-	Equivalent Income	none
1995/2001	Demographics	Equivalent Income	Ψ_D
1995/2001	+Household work structure	Equivalent Income	$\Psi_D * \Psi_S$
1995/2001	+ Schooling of householders and spouses	Equivalent Income	$\Psi_D * \Psi_S * \Psi_E$
1995/2001	+ Skill prices of householders and spouses	Estimated EI	$\Psi_D * \Psi_S * \Psi_E$
2009	-	Equivalent Income	none

Note—Ψ_D is estimated through a logit with a binary for survey year 2009 as dependent variable and as independents gender of the head of household, urban residence, and third order polynomials in age of the head of household and per household ratios of children to adults. Ψ_S is arrived at similarly with the exception that the exogenous variables are binaries for 8 household types, with two-income earner households being the base group. Ψ_E is estimated through an ordered logit where the dependent variables are 13 range possibilities for mean schooling of the head of household and spouse, if there is one present. Independent ones include the fraction of householders and spouses in one of eight age categories, the fraction that is female, and a binary for metropolitan area residence. Estimated earnings of working heads of household and spouses reflecting skill prices of 2009 are arrived at through the approach of Machado and Mata (2001; 2005).

Table 6—Share of fall in distribution index accounted by demographics, household work structure, schooling, and returns to labor

	PRIMARY SEQUENTIAL ORDER					
1995-2009	**Share (%) of fall in index accounted by factor:**					
Factor	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Demographics	17.4	17.7	18.9	20.4	31.6	36.4
Work structure	-2.7	-3.2	-5.3	1.7	-7	-14
Education	12.6	5.2	-1	53.6	51	49.8
Returns to labor	36.2	35.9	32.6	4.8	-12.3	-25.2
2001-2009	**Share (%) of fall in index accounted by factor:**					
Factor	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Demographics	13.3	12.7	12.8	12.7	15	16.1
Work structure	0.3	0	-.6	6.7	3.7	2.1
Education	8.6	3.7	0	33.2	26.7	23.8
Returns to labor	38.8	35.7	28.3	26.7	9.6	0
	AVERAGE OVER ALL POSSIBLE SEQUENCES					
1995-2009	**Share (%) of fall in index accounted by factor:**					
Factor	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Demographics	5.3	6.3	8.4	12.1	22.6	27.8
Work structure	-3.3	-4.3	-7.5	4.6	-6.1	-14.8
Education	17.7	10.5	5	54.6	58.1	59.5
Returns to labor	42.7	42.3	39.1	9.1	-10.1	-24.6
2001-2009	**Share (%) of fall in index accounted by factor:**					
Factor	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Demographics	5.3	5.5	6.6	8.2	10.5	11.7
Work structure	0.1	0.4	-0.7	6.8	3.2	1.5
Education	10.6	5.8	2.2	31.4	26.7	24.6
Returns to labor	44.2	40.2	32.1	32.5	13.8	3.4

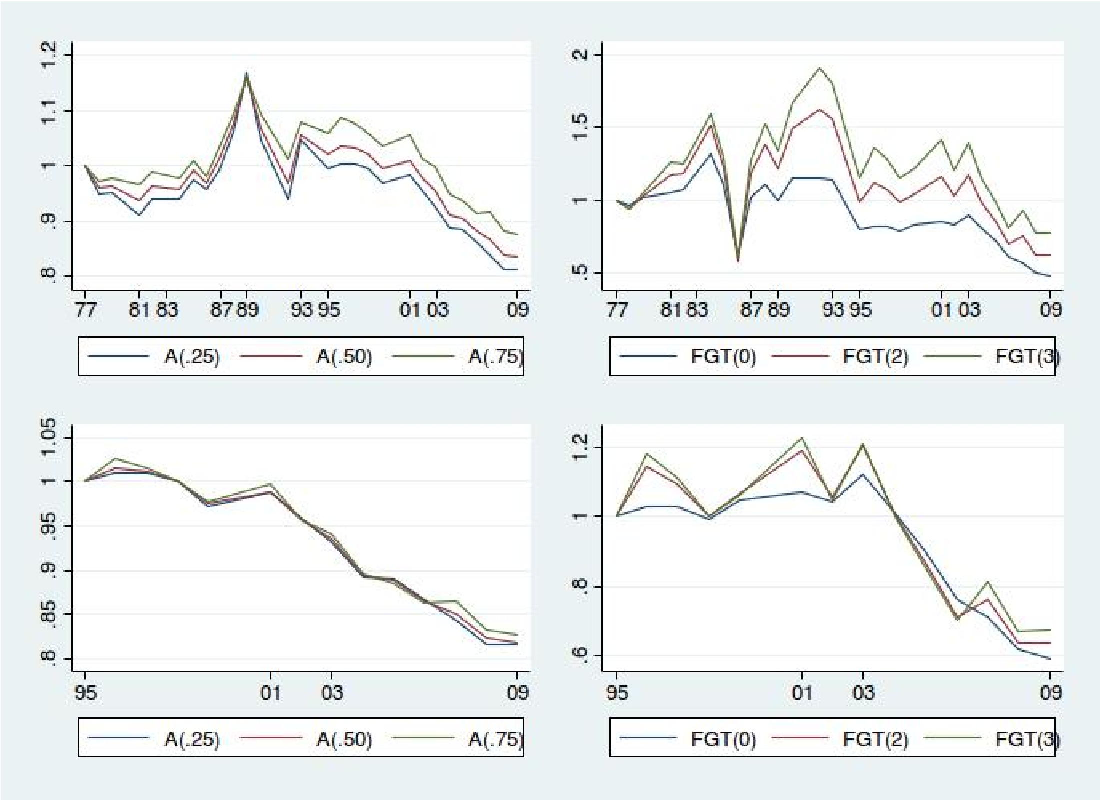

Figure 1: Long-run and post-*Real* Plan trends in poverty and inequality

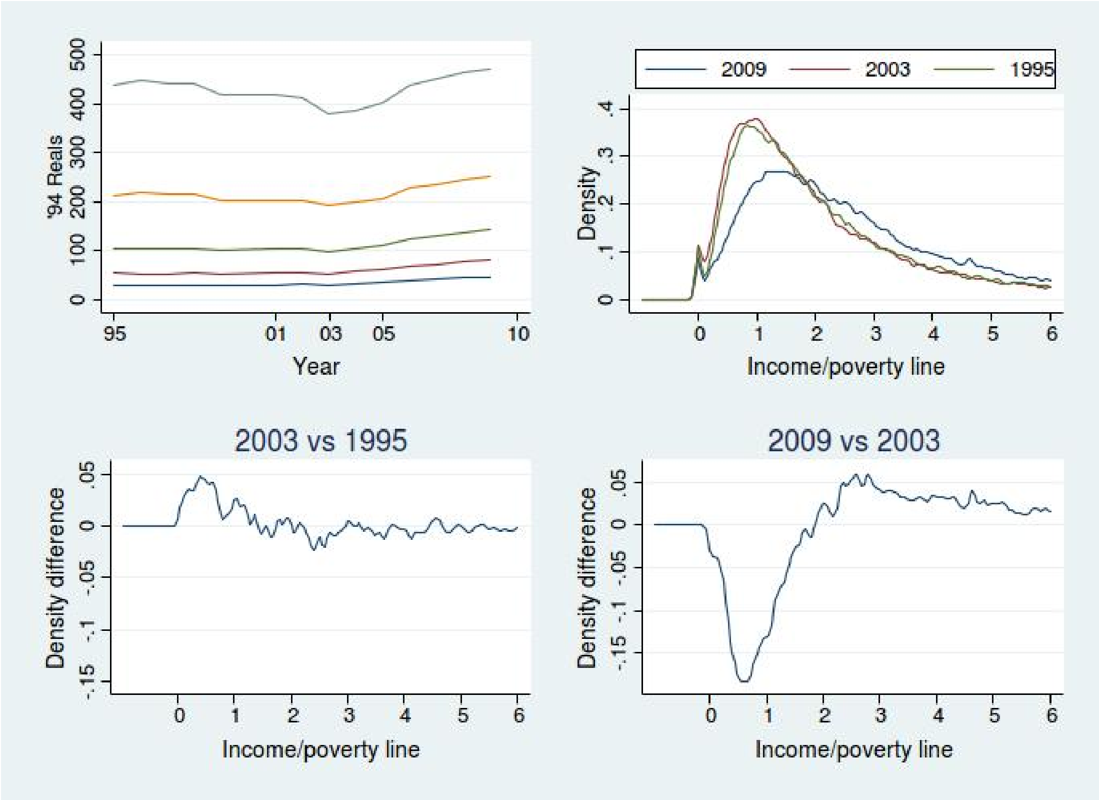

Figure 2: Equivalent income at the 10th, 25th, 50th, 75th, and 90th percentiles of the equivalent income distribution (top left), 1995, 2003, and 2009 income densities (top right), 1995-2003 density difference (bottom left) and 2003-09 density difference (bottom right)

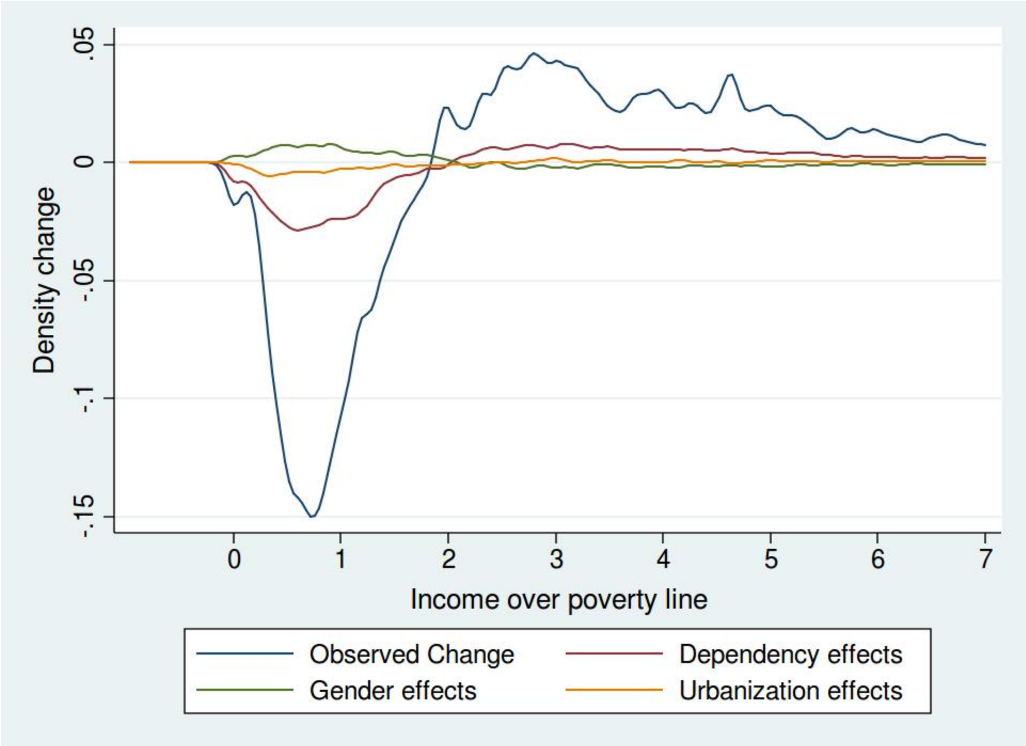

Figure 3: Observed change in income density over 1995-2009 and contributions by demographic phenomena

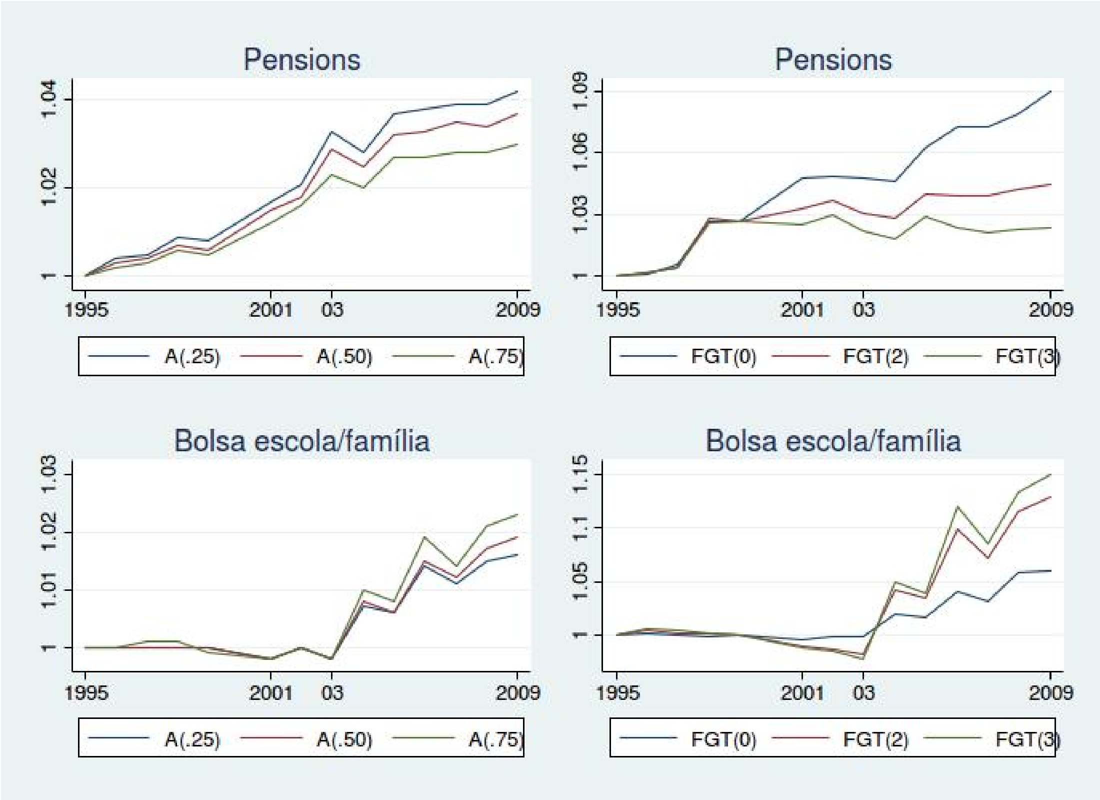

Figure 4: Normalized counter-factual trends in distribution holding constant the distribution of pensions (top panels) and transfers (bottom panels)

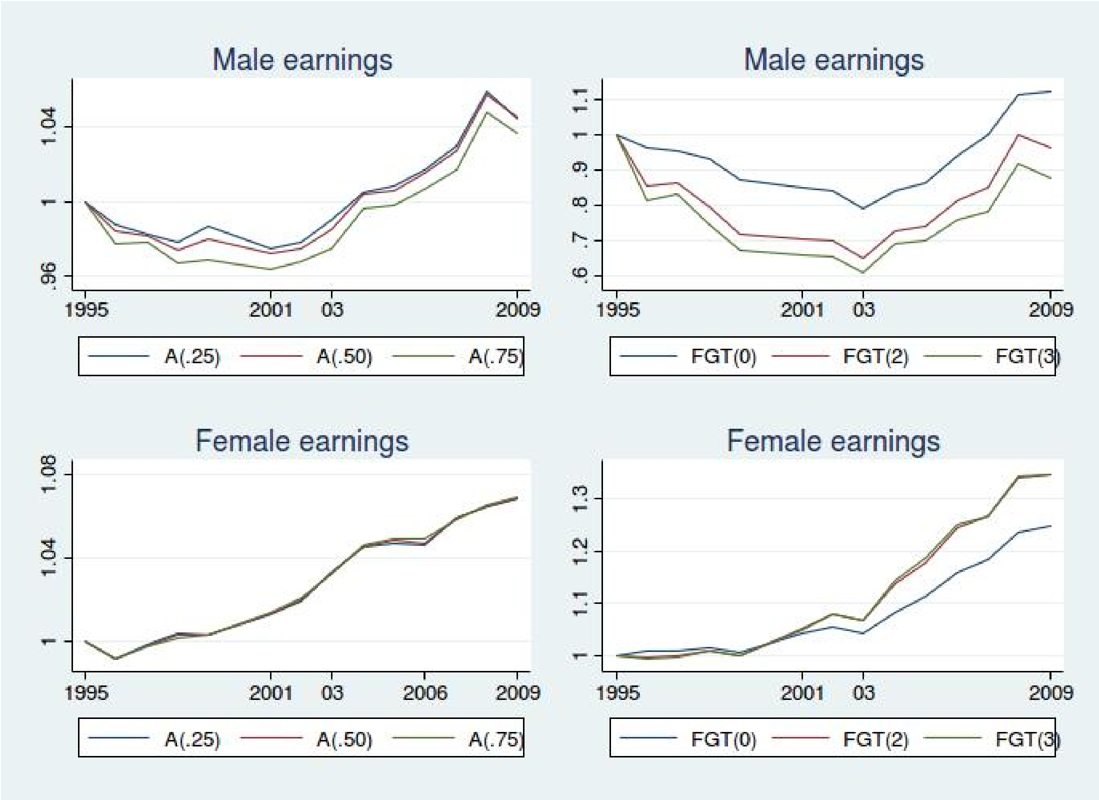

Figure 5: Normalized counter-factual trends in distribution holding constant the distribution of male earnings (top panels) and female earnings (bottom panels)

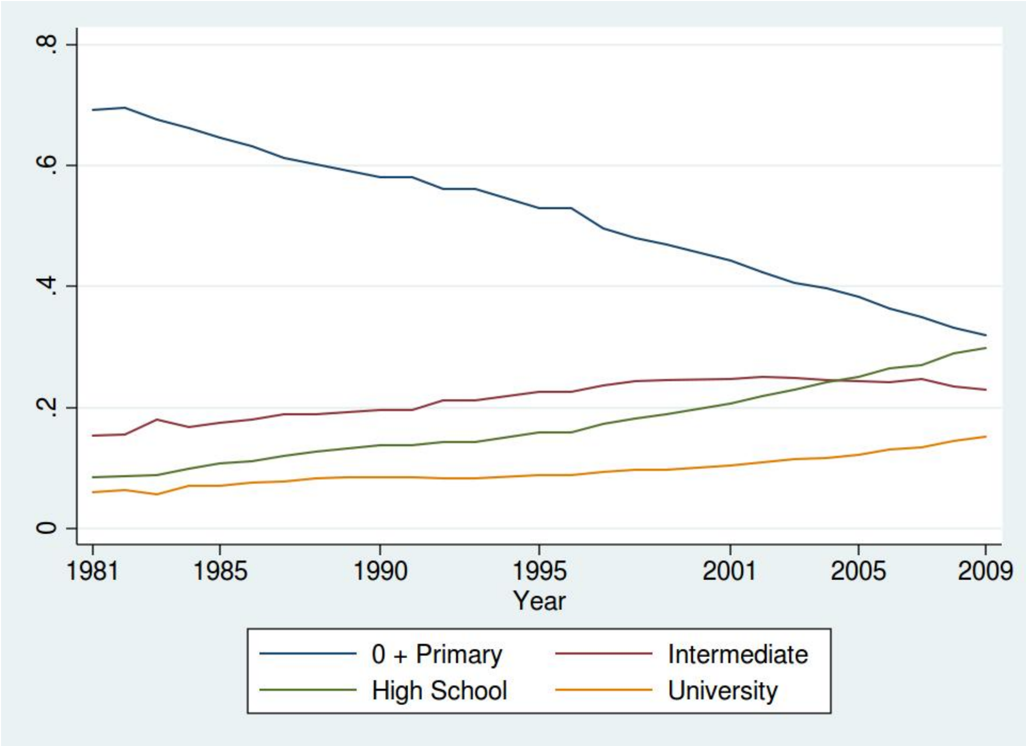

Figure 6: Share of population aged 25-64 by schooling attainment

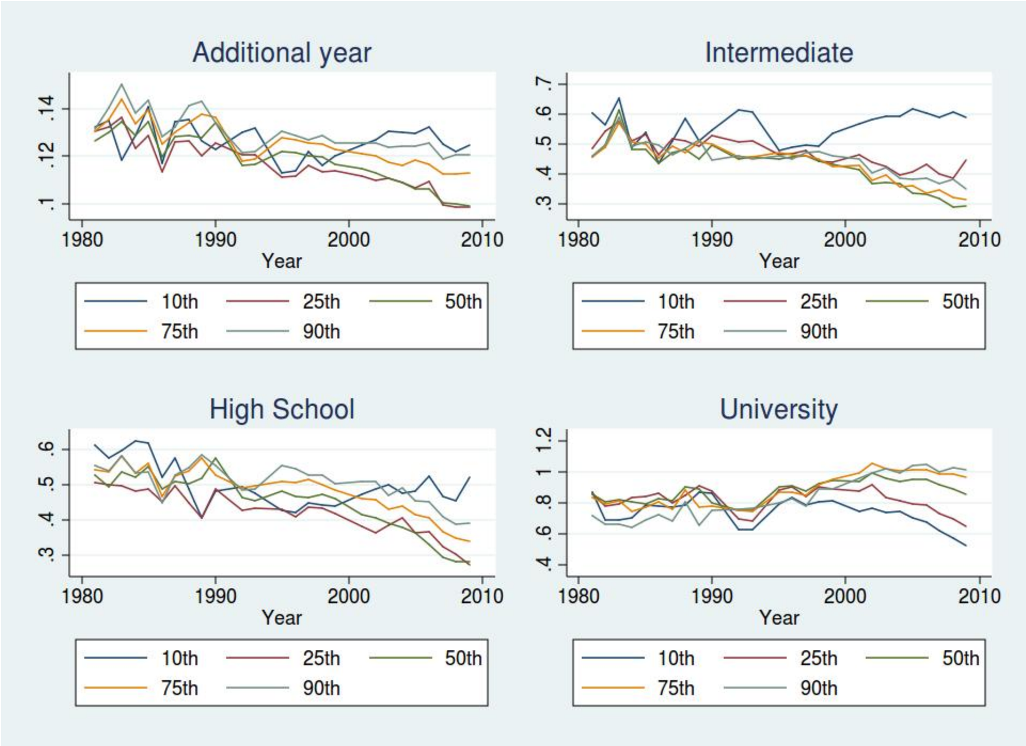

Figure 7: Return (%) to school by percentile in earnings distribution

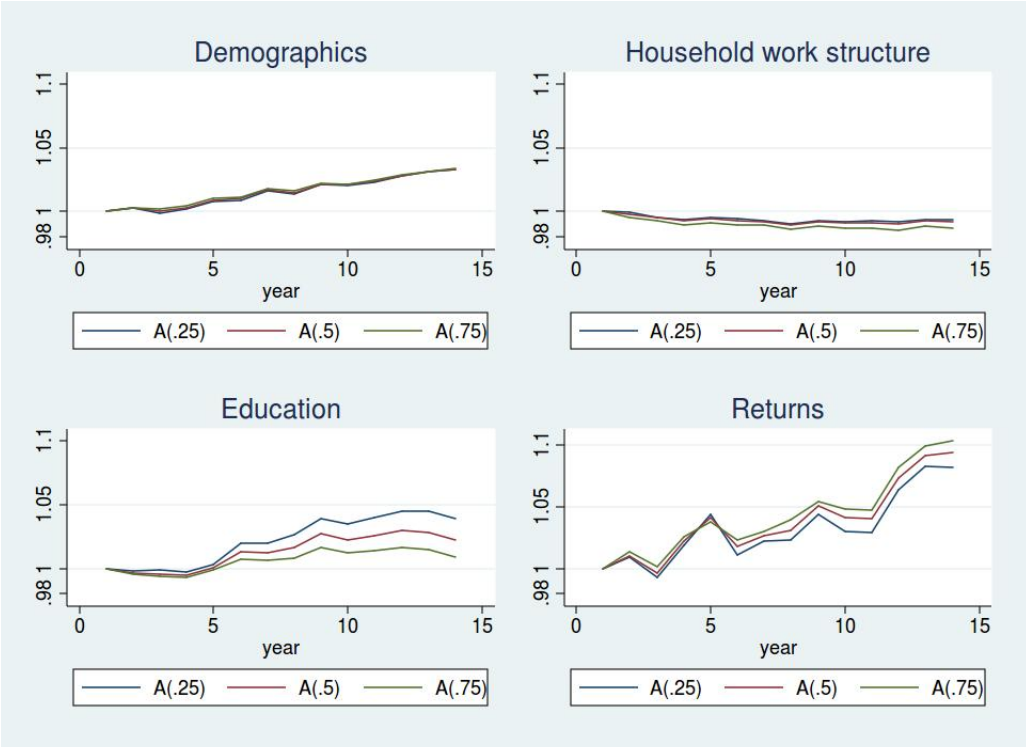

Figure 8: Normalized counter-factual trends in inequality holding constant demographics, household work structure, schooling, and returns to skill

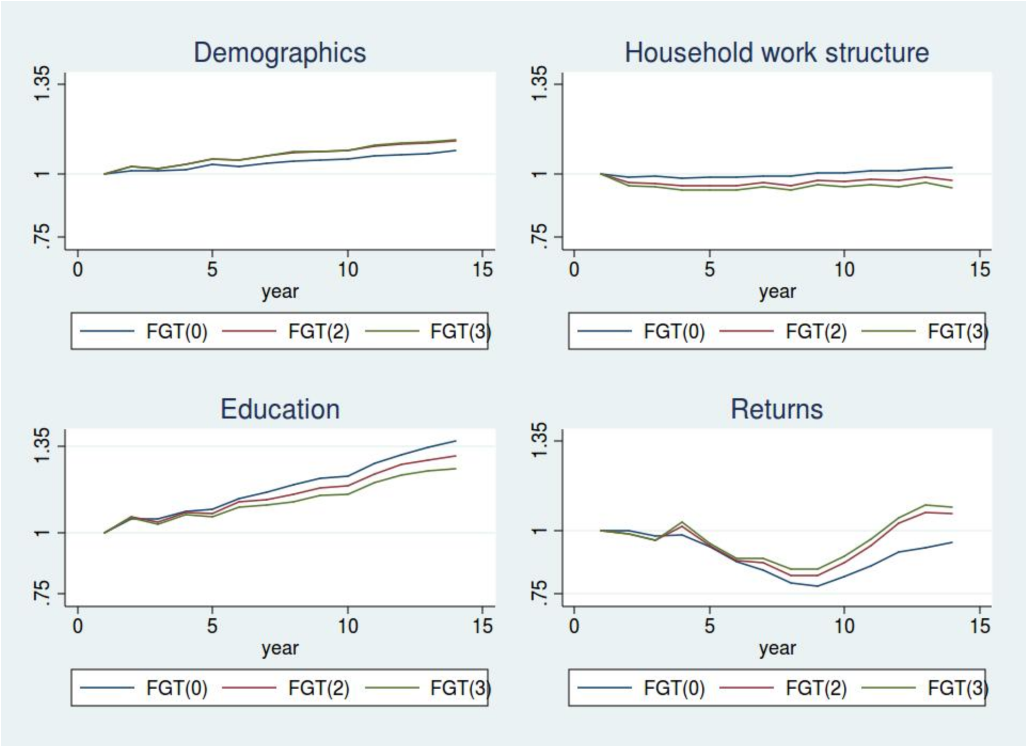

Figure 9: Normalized counter-factual trends in poverty holding constant demographics, household work structure, schooling, and returns to skill

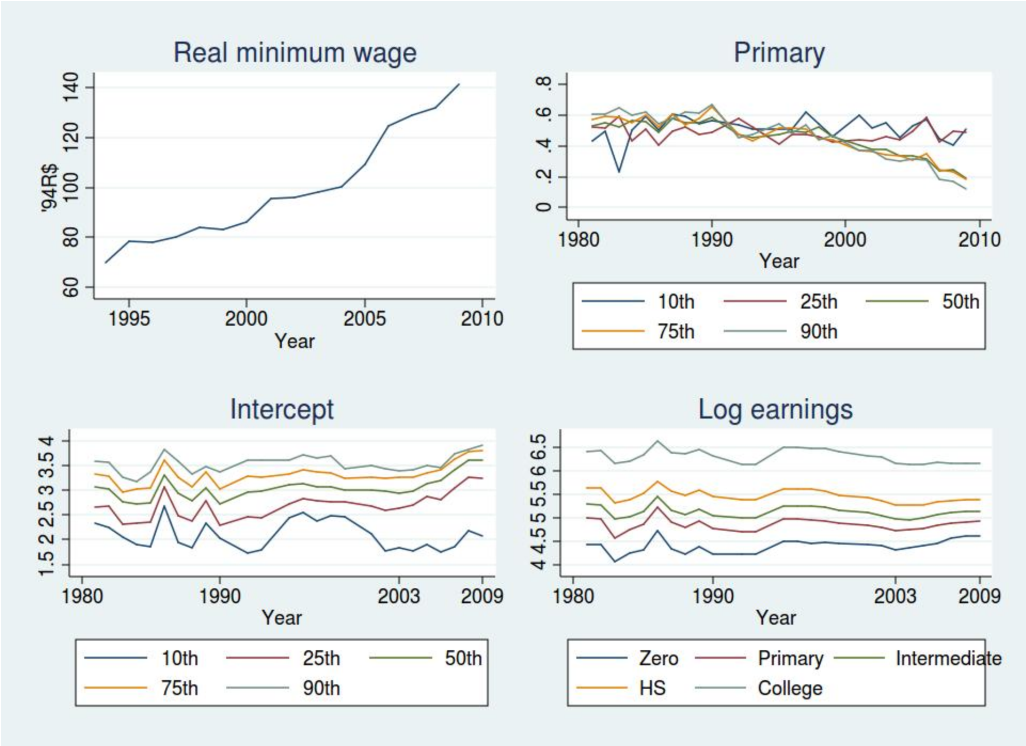

Figure 10: Real minimum wage, returns to primary schooling, earnings functions intercept and log earnings by schooling

10 Appendix A

The choice of parameters for the equivalence scale was based on research of costs of children relative to adults and on economies of scale in the production of households goods that were posited to be lower in Brazil relative to developed economies. Likewise, the poverty line is based on research estimating costs of food baskets supplying minimum caloric requirements. Nevertheless, sensitivity analysis to scale and poverty threshold can shed more light on issues addressed in the paper. To that end, Tables 3 and 6 are re-computed using an alternative equivalence scale and a range of poverty lines. The equivalence scales allows for more economies of scale (s=.77), higher costs of children to adults (k=.53), and reflects choices more akin to a developed economy, in this case the UK. Along with the developed country scale, the poverty line is raised by 10%, 25% and 50%.

Re-computations shown below establish that the alternative equivalence scale has marginal effects on conclusions dealing with inequality. Effects on those related to poverty are not as small due to the fact that the alternative scale results in higher equivalent income and lower poverty. With a poverty line that is re-calibrated up by just 10%, focus is placed on a population further down the income distribution than evaluated in the paper. As the poverty line is adjusted upwards by 25% and then 50%, incomes further up the distribution are taken into account. The importance of pensions in poverty increases somewhat, while that of conditional transfers does the opposite, as could be expected given their means-tested nature. More salient differences pertain to the influence of changes in the distribution of earnings among men. With higher poverty lines, they have a relatively smaller positive impact on the headcount ratio and a less negative one with respect to income and distribution-sensitive measures, reinforcing what poverty measures demonstrate: that male earnings changes had their most negative influence at the very bottom of the distribution. Last, as the line is adjusted upwards, changing prices lose their small positive effect on the headcount ratio, a result consistent with the hypothesis in the main text relating the phenomenon to an increase in the real minimum-wage.

Determinants of decline in measures using alternative equivalence scale and poverty thresholds

1995-2009	Fall in index accounted by source/factor (%):					
Income source	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Pensions	22.9	20.5	17.7	17.5	9.1	4.4
Transfers	6.5	7.9	10.4	11.3	37.1	51.7
Female earnings	37.4	38.1	40.5	52.8	100	119
Male earnings	24	25.1	22	25.6	-22.3	-56.4
	Threshold up by 25%			**Threshold up by 50%**		
Income source	**FGT0**	**FGT2**	**FGT3**	**FGT0**	**FGT2**	**FGT3**
Pensions	18.1	12.6	8.9	22.2	14.6	11.8
Transfers	8.6	26.4	37.3	5.7	20.5	28.6
Female earnings	47.5	81.2	98.3	47.2	71	85
Male earnings	22.3	-2.7	-25.6	16.5	6.2	-8.9

1995-2009	Fall in index accounted by source/factor (%):					
Factor	**A(.25)**	**A(.50)**	**A(.75)**	**FGT0**	**FGT2**	**FGT3**
Demographics	22.2	19.9	17.2	16.5	10.5	6.9
HH work structure	8.6	10.1	13.2	11.0	29.8	42.1
Education	36.5	37.2	39.9	45.5	80.8	97.7
Returns to skill	23.4	24.6	21.7	22.0	-9.1	-33.8
	Threhold up by 25%			**Threshold up by 50%**		
Factor	**FGT0**	**FGT2**	**FGT3**	**FGT0**	**FGT2**	**FGT3**
Demographics	22.2	19.9	17.2	16.5	10.5	6.9
HH work structure	8.6	10.1	13.2	11.0	29.8	42.1
Education	36.5	37.2	39.9	45.5	80.8	97.7
Returns to skill	23.4	24.6	21.7	22.0	-9.1	-33.8

11 Appendix B

Machado and Mata (2001, 2005) provide a strategy for decoupling prices from attributes in distributions through an approach that may be viewed as an adaptation of the Oaxaca-Blinder decomposition to quintile regression. Denoting $Q_{\Theta}(y|\ x)$ for $\Theta \in(0,1)$ as the Θth percentile of the distribution of y given x, and $\beta(\Theta)$ as a vector of quintile regression coefficients, the approach is based on two foundations. First, in the same manner that percentiles fully characterize the marginal distribution of a variable y, conditional percentiles do the same to the distribution of y given x. Second, by the probability integral transformation theorem, if $\Theta_1, \Theta_2, \Theta_3 ... \Theta_m$ are drawn from a distribution U(0, 1), the corresponding m estimates of the conditional quantiles at x $\{x' \widehat{\beta}(\Theta_i)\}_{i=1}^{m}$ are distributed as the estimated distribution of y given x. Hence, a sample from the marginal distribution of y can be generated by drawing a random sample of covariates x from an appropriate distribution.

Specifically, defining y(t) as log earnings and x(t) as a vector of personal characteristics at time t, the construction of a sample from an earnings distribution is carried out in four steps. The first involves the generation of a random sample of size m, $u_1, u_2, u_3 \ldots u_m$, from a distribution U(0, 1). Secondly, for each $u_1, u_2, u_3 \ldots u_m$ a quintile regression is estimated to arrive at the corresponding $\widehat{\beta}^t(u_i)$ using covariates from a sample X at time t. Third, a random sample of size m is generated with replacement from the rows of X at time t, represented by $x_i^*(t)$, i=1, 2, 3 ... m. Finally, $\{y_i^*(t) \equiv x_i^*(t)' \widehat{\beta}^t(u_i)\}_{i=1}^{m}$ constitutes a random sample of size m of the desired distribution.

However, the point of the exercise is not reproducing distributions. Instead, counter-factuals can be constructed to reflect how earnings would have evolved if only skill levels had changed over time or if only skill prices had done the same. From the preceding discussion, these can be readily constructed by estimating coefficients and by drawing covariates from appropriate distributions. That is, to generate the earnings density that would have obtained in year t if individual characteristics had remained as in year t' the above-mentioned steps are carried out with the exception that while regression coefficients are derived using a sample drawn from year's t distribution, samples are integrated out using covariates from year t'. Likewise, to construct the counter-factual of what would have been the wage density in t' if characteristics had been as in t, quantile regressions coefficients are arrived at using the t' distribution and the sample integrated out using covariates from year's t distribution.