Growth with redistribution:
Did Brazil show the way?

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Abstract

A comprehensive decomposition approach is applied to identify the factors that generated one of the most remarkable episodes of redistribution in recent memory, whereby in a span of a decade Brazilian inequality fell by one-fifth and the incidence of poverty declined by two-thirds. It is argued that important elements related to these developments had been in place for some time. These included macroeconomic stability, long-standing increases in schooling attainment, and favorable demographics in the form of a declining dependency ratio that proved to be both poverty- and inequality-reducing. Robust economic growth during the 2000s and the early 2010s resulted in across-the-board income gains that were widely shared, owing to mechanisms that favored advances at the lower end of the distribution of earnings. Demographics, as well as changes in educational attainment, labor force participation, and lower-skill prices explain most of the significant drop in the headcount ratio, with education and interaction between economic growth and a rising wage floor accounting for the bulk of the change. The same factors explain up to 85% of the decline in income dispersion that was especially driven by markedly compressed earnings differentials. Human capital accumulation and strong labor market institutions thus stand out as key mechanisms linking economic growth to redistribution.

1 Introduction

During the first decade of the 21st century, one of the most remarkable episodes of redistribution unfolded in a country known as one of the most unequal in the world. After rising through decades of formidable growth, fluctuating in the macroeconomically turbulent 1980s, and stagnating in an ensuing period of meager growth, Brazilian inequality began to fall sharply and consistently for the first time in recorded history. Income growth was strongest at the bottom half of the distribution, lifted millions from poverty, and left almost no one behind. These developments coincided with the implementation of social policies intended to generate such results, as conditional and unconditional transfers grew in scope and economic importance. However, they also occurred in the context of con-
continued improvement in educational attainment, macroeconomic stability, demographic
shifts that included plunging dependency rates, and perhaps most notably, during a long
period of robust economic growth. Even though the question of what caused distribu-
tion to change is a controversial subject that is still far from well understood, popular
sentiment and academic research lean toward an explanation with a strong social policy
component (Barros et al., 2006, 2010). This is also supported by evidence regarding the
poverty-reducing role that public transfers played in the years leading to the economic
upturn (Ferreira, Leite, and Ravallion, 2010).\footnote{See also The Economist, A new form of radical centrist politics is needed to tackle inequality without hurting economic growth, October 13, 2012.}

An assessment of the forces that generated such shifts in distribution has policy im-
lications that go beyond the case of Brazil. Although there is an extensive body of
literature on the causal relationship between inequality and growth, and to a lesser ex-
tent on how they move together, evidence on the mechanisms that link them is much less
abundant, with answers also important for reconciling differences in short- and long-run
relationships between inequality and growth (Halter, Oechslin, and Zweimüller, 2014).
Similarly, whereas economic growth has been established as a major instrument of poverty
reduction (Dollar and Kraay, 2002, Kraay, 2006), with Brazilian time series evidence be-
ing consistent with the result (Sotomayor, 2006), unanswered questions remain. Kraay
(2006) decomposes poverty alleviation into growth and redistribution components, and
finds little cross-country evidence on the determinants of the latter mechanism, includ-
ing institutional quality, trade openness, primary education, inflation, and government
consumption. Ravallion (2012) stresses the importance of a better understanding of the
factors that hinder or favor poverty reduction in developing economies.

The remarkable run of redistribution in Brazil represents an exceptional opportunity
to gather evidence on the mechanisms linking growth to distribution, with theory pointing
to human and physical capital accumulation, technology, and interactions with institu-
tions and geography (García-Peñalosa, 2010). Barros et al. (2010) provide an assessment
that highlights public transfers as an important source of redistribution in Brazil, but
the conclusions hinge on a long series of counterfactuals in a sole arbitrary sequence
and on assessments of population attribute effects that are derived from factor income counterfactuals. Population and factor income explanations can be effectively combined, but through use of a regression-based approach (Cowell and Fiorio, 2011). The strategy proposed in this study is well tested, adapted to this end, and extended to evaluate skill price effects across the earnings distribution, incorporating a quantile regression approach that has proved to be useful for examining Brazilian gender and racial wage disparities (Salardi, 2012). The empirical approach also generates results that are not dependent on sequential order, and it is capable of assessing impact on poverty as well as on inequality.

The evidence presented in this paper indicates that Brazilian distribution trends benefited from elements that had been in place for some time. These included human capital accumulation and demographic changes that had consistently positive effects on distribution. Specific to the economic expansion were changes in skill prices that served to diminish inequality and poverty through stronger earnings gains at the lower end of the income distribution, a phenomenon produced to a large extent by interaction between robust economic growth and a rising wage floor. Pensions and conditional transfers played secondary to minor roles.

These developments stand in contrast to past experience, when limited progress in fighting poverty between 1985 and 2004 was associated with lackluster macroeconomic performance, and with a low elasticity of poverty alleviation with respect to growth (Ferreira, Leite, and Ravallion, 2010). Institutional factors can reconcile markedly different growth consequences on distribution. That is, in the context of an economic expansion capable of mitigating its unemployment costs, a rising wage floor fostered a wider distribution of the benefits of growth and a reduction in poverty and inequality that no other policy instrument had been able to achieve in the past.

The research that supports these assertions begins with an assessment, in Section 2, of distribution trends and potential drivers. Section 3 lays out an empirical strategy designed to ascertain the impact of education, demographics, labor force participation, and skill prices, including those associated with a rising minimum wage. Section 4 provides a description of the data source, and Section 5 discusses results that are compared, in
Section 6, to those emerging from a thesis rooted in public transfers or geography. Section 7 concludes with a summary and a discussion of policy implications.

2 Distribution, the labor market, and demographics

Interest in the relationship between Brazilian growth and distribution goes back to the 1960s, when a long period of rapid growth was accompanied by increased inequality and limited reduction in poverty, a development that begged the questions of why the benefits of growth were so thinly shared, and if welfare had actually increased during the decade (Fishlow, 1972, Fields, 1977). Continued expansion in the 1970s and macroeconomic instability in the 1980s and early 1990s had further negative impacts on distribution. Inequality rose during the 1980s Federal Reserve recession and increased even further in the ensuing recovery, with near hyperinflation resulting in increased inequality and lower real wages (Hoffmann and Kageyama, 1986, Hoffmann, 1995, Bonelli and Ramos, 1995, McIntyre and Pencavel, 2004).

Successful price stabilization brought forth by the Real Plan in 1994 was therefore a welcome event in many respects (Neri, 2006), as poverty fell and distribution swings associated with inflationary bursts and stabilization programs became a thing of the past. However, the high interest rates supporting the currency anchor adopted to tame hyperinflation reduced the prospects for economic expansion that were further curtailed by the effects of the Asian and Russian financial crises. The negative consequences on distribution were countered by price stability and increased public transfers (Ferreira, Leite, and Ravallion, 2010), changes in the distribution of earnings among women (Sotomayor, 2009), and potentially by increased female and child labor force participation, phenomena that have been tied to male unemployment and low family income (Fernandes and de Felício, 2005, Soares, Kruger, and Berthelon, 2012).

[Table 1]

\^ Other market-oriented reforms that accompanied or shortly preceded the plan had additional positive effects on distribution (de Carvalho and Chamon, 2012).
A sustained economic upturn began in the 2000s when global commodity demand helped the Brazilian economy grow vigorously between 2004 and 2011, except for 2009, when the economy contracted slightly. The expansionary period extended to 2012 and 2013, although growth was under 3%. Estimated measures of household income in Table 1 establish that relative to their levels at the beginning of the decade, income rose by 45%, inequality fell by a fifth, and the share of population living in poverty dropped by two-thirds. What is more, all distributions between 2009 and 2013 can be shown to first-order and Lorenz dominate that of 2003, and as a result, the implications of falling inequality and poverty hold for all Lorenz-consistent inequality measures, and for all thresholds and indices used to identify and aggregate poverty.³

These remarkable changes in distribution could have been influenced by factors associated with the business cycle, but also by structural trends capable of magnifying or countering shorter-term effects. Demographics can be one such source. As in many countries, the Brazilian population has over time become older, more urbanized, and dispersed in smaller households that are also more likely to be headed by women. Between 1995 and 2013, the share of the population living in female-headed households rose from 16% to 36%, that residing in census-classified urban areas grew from 79% to 86%, and the children-to-adults household ratio fell by over a third (Table 1).

Education can represent another structural driver, especially considering steadily increasing schooling achievements, which translated to increases of 54% and 45% in years of schooling among heads of household and working individuals, respectively. All things being equal, returns attributed to skill can be expected to decline. To determine whether this is indeed the case, they are estimated with earnings functions that regress log monthly earnings on education and age polynomials, head of household status, and metropolitan area residence binary variables. Quantile estimators determine returns at the median and at the top and bottom deciles and quartiles, using samples that are derived by retaining information on household members engaged in salaried employment and between the ages of 18 and 64.⁴

³Lorenz curves and cumulative distribution functions are available from the author on request.
⁴As is customary when estimating returns to education, earnings refer to those associated with primary
Results in Figure 1 show steadily declining returns to an additional year of school that also “fan out,” in the sense that they fall less at the upper than the lower end of the earnings distribution—for example by 12% at the 90th percentile, but 24% at the 25th. Disaggregation by degree establishes dramatic declines in returns to intermediate and high school, especially during the 2000s and the early 2010s. At the median, the wage gap between completed high school and intermediate education drops from 47% in the mid-1990s to 39% in the early 2000s, and to 21% in 2013.

Figure 2 demonstrates that falling returns during the economic upturn do not necessarily mean lower wages, as they resulted from wage growth that was negatively correlated with schooling. That is, whereas earnings among high school graduates rose by 21%, those of individuals with intermediate or primary schooling increased by 38% and 42%, respectively. Within these schooling classifications, earnings gains were greatest at the lower end of the distributions, as manifested in Figures 1 and 2 by returns to intermediate school that rose at the bottom decile, and returns to primary school that fell less at the bottom quartile and decile. Further down the pay scale, the wage received by individuals without formal schooling and experience—reflected by the intercept term of the earnings function—increased, and its variability across the distribution declined. Altogether, these developments are consistent with a progressively more binding minimum wage in covered and compliant sectors.

The college wage premium follows a different path that manifests a slight upward trend until the 2000s, when it begins to fall everywhere but at the top decile. From 2003 to 2013, the wage gap between completed college and high school fell by 8% at the top quartile, 16% at the median, and 35% at the bottom decile. The decade witnessed a pickup in university completion, and Fortin (2006) finds a close link between labor supply and college premium, but an explanation for the fanning out of returns is harder
to come by. The expansion in a college loan program in the 2000s could have resulted in increased college graduate heterogeneity, but the speed of development suggests additional mechanisms at work. Regardless, it is clear that for the first time in recorded history, the return to a college education declined in Brazil at an ample range of the distribution and rose less rapidly or remained constant at the top.\(^5\)

[Table 2]

Labor market developments can also influence the distribution of income through changes in labor force participation. Table 2 classifies households into eight categories using work status and gender of the head of household and spouse, if applicable. The first four refer to households headed by an unmarried individual (or without a spouse present) and classified as headed by a working or non-working male or female. The remaining four categories refer to households headed by a couple where both individuals work, both do not work, only the male works, or only the female works.

A decomposition of the population along the stated categories points to a secular decline in units with couples where the male works outside the home but the female does not. That is, whereas 44% of the population lived in such a household in 1995, only 28% did so about two decades later. Part of the population transitions to units headed by a single woman, a trend that is slow and structural, whereas another part transitions to households where both the husband and wife work. The latter phenomenon appears cyclical and is one of the defining characteristics of the 2000s, as it is largely exclusive to the period. Between 1995 and 2001, the share of population living in such units held steady at around 25% but rose to 30% by 2009.

3 Empirical strategy

An assessment of the impact of changes in demographics, household work structure, education, and skill prices requires estimation of distribution trends that would have

\(^5\)Annual survey data, available since 1976, reflect returns to college that rise during the late 1970s, the 1980s, and the 1990s (Sotomayor, 2004, Menezes, Fernandes, and Picchetti, 2006).
occurred in their absence. This is a hypothetical proposition, but DiNardo, Fortin, and Lemieux (1996) offer a counterfactual approach involving reweighing of sample weights. Adapting their approach to the objectives of this study and denoting $F(y, x, t)$ as a joint distribution of household income $y$, household characteristics $x$, and date $t$, the density of income in year $t$, $f_t(y)$, can be expressed as the integral of the density of income conditional on $x$ and on a date $t_y$, over the distribution of $x$ at date $t_x$: $f_t(y) = \int_{x \in \Omega_x} f(y \mid x, t_y = t) dF(x \mid t_x = t)$, where $\Omega_x$ is the domain of attributes. Assuming that the distribution of $x$ does not affect the conditionals, the year $t$ density, based on a distribution of characteristics as in year $t'$, is

$$f_t(y) = \int_{x \in \Omega_x} f(y \mid x, t_y = t) dF(x \mid t_x = t) \Psi_x(x) dF(x \mid t_x = t),$$

where $\Psi_x(x) \equiv dF(x \mid t_x = t')/dF(x \mid t_x = t)$, and constitutes the only difference between the two expressions. A counterfactual density can therefore be obtained from an observed one through use of a reweighting mechanism $\Psi_x(x)$. That is, given a distribution of income in year $t$ and a variable $x$ indicating (say) residence in rural or urban areas, generating the distribution of income based on an urbanization composition observed in year $t'$ involves giving more (less) weight to observations in rural settings and less (more) weight to observations in urban ones, in such a manner as to reflect the distribution of $x$ in year $t'$.

The decomposition sequence and estimating equations are laid out in Table 3, where the 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 the distribution. Controlling for demographics involves reweighting initial-period data by a function $\Psi_D$ that sets household demographic characteristics at final-year levels (row 2). If the percentage difference between this counterfactual and the initial distribution is equivalent to (say) 10% of the observed change, demographic changes are responsible for that share of the trend. With the reweighting function for household work structure denoted as $\Psi_S$, $\Psi_D \times \Psi_S$ sets the demographic and household work structure to the reference distribution levels (row 3). In the same numerical example, if the percentage difference between the
initial distribution and the counterfactual is equivalent to 12% of the observed change, work structure drives 2% of the trend.

[Table 3]

The impact of education could also be evaluated by employing a reweighting function that sets schooling to final-period levels, but such an approach would not account for important correlations with other variables such as age, gender, or geographic setting, whose effects can be confounded with schooling. Alternatively, the reweighting function can be arrived at by conditioning on the relevant covariates through use of an ordered logit, with the resulting $\Psi_E$ setting schooling levels of labor force participants at final-year levels. The impact of education on distribution is then ascertained as in the case of demographics and household work structure.

Understanding the influence of skill prices requires their estimation, and if returns at the mean were an adequate reflection of those elsewhere in the distribution, OLS would suffice as an estimator. However, the effects of Brazilian skill prices appear increasingly dependent on their position in the distribution, and to accommodate this feature of the country’s labor market, earnings distributions are constructed using a quantile regression approach (Machado and Mata, 2005). If $Q_\Theta(w|x)$ for $\Theta \in (0,1)$ denotes the $\Theta$th percentile of the distribution of $w$ given $x$, and $\beta(\Theta)$ a vector of quantile regression coefficients, the formulation is based on two foundations. First, conditional percentiles characterize the distribution of $w$ given $x$ as percentiles do the marginal distribution of the variable $w$. 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 $i=1, 2, 3, \ldots, m$ estimates of the conditional quantiles of $w$ at $x$, $\{x'\hat{\beta}(\Theta_i)\}_{i=1}^m$, are distributed as the estimated distribution of $w$ given $x$. Given estimates of conditional quantiles, integration over $x$ produces a sample from the marginal distribution.

In practical terms, and with $w(t)$ defined as earnings and $x(t)$ as a vector of personal attributes at time $t$, a sample is constructed from a marginal distribution of earnings in four steps. The first involves generation of a random sample of size $m$, $u_1, u_2, u_3,...,u_m$, from a distribution $U(0, 1)$. Second, for each of $u_1, u_2, u_3,...,u_m$, a quantile
regression is estimated to arrive at the corresponding \( \hat{\beta}_t(\gamma_i) \) 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, \( \{w^*_i(t) \equiv x^*_i(t)\hat{\beta}_t(\gamma_i)\}_{i=1}^m \) constitutes a random sample of size \( m \) from the applicable distribution.

However, the point of the exercise is not to reproduce distributions but to construct counterfactuals reflecting how earnings would have evolved if only skill levels or prices had changed over time. From the preceding discussion, these can be readily constructed by estimating conditional quantiles and by drawing covariates from appropriate distributions. For the decomposition in Table 3, the final-year sample is used to arrive at skill prices by regressing earned income of households with working individuals on third-order polynomials in mean schooling and age of economically active household members, share of female labor force participants, binaries for metropolitan area residence and number of workers in the household, and interaction terms between education, age, and gender. Marginal earnings distributions are constructed with conditional quantiles integrated over beginning-period covariates, and resulting income distributions, calculated with estimated household earnings, are reweighted with \( \Psi_D \times \Psi_S \times \Psi_E \) (row 5).

Considering a minimum wage that grew markedly during the upturn that began in the 2000s, it would be of particular interest to evaluate its contribution to the substantial decline in inequality and poverty observed during the period. The DFL approach would appear to be well suited to the task, but assessment is challenged by the fact that this study’s unit of observation is the household, rather than the individual. Although it would be simple to replace the lower tail of a distribution with one corresponding to a distribution with a different minimum wage as the DFL approach allows, this procedure cannot be applied in the case of a distribution of household earnings that itself is product of multiple distributions of individual earnings. Even so, estimation of skill price effects can still be performed in two steps, much in the same way that DFL distinguishes between wage densities below and above minimum wage levels. That is, the first step involves estimation of skill price effects among households where the maximum of earnings of the head of household or spouse (if one is present) is the minimum wage. This set does not
exhaust all households where someone draws the wage floor but does represent cases most clearly bounded and a conservative estimate of impact. Higher-skill price effects can be evaluated through a second step that measures impact among all households, which in a sequential decomposition corresponds to estimating influence above that already established. Minimum wage effects can be particularly sensitive to macroeconomic conditions that can temper or augment its unemployment costs, so that consequences can be more accurately described as resulting from the interaction between the wage floor and the state of the business cycle.

Finally, it should be noted that Ghosh (2014) proposes a semi-parametric extension to the Machado and Mata decomposition, which performs better when (or where) the conditional quantile function is non-linear. The results, based on US data, show that the latter decomposition underestimates the effects at the top and overestimates those at the lower end of the wage distribution. However, Chernozhukov, Fernández-Val, and Melly (2013) find similar results with parametric models specified in a flexible form when comparing quantile and distribution regression estimators to obtain counterfactual distributions. Salardi (2012) also finds small differences associated with decompositions based on Machado and Mata (2005), Melly (2005, 2006), and Firpo, Fortin, and Lemieux (2009) when examining Brazilian racial and gender earnings differentials. Nevertheless, in addition to incorporating a flexible specification for household earnings, the study adopts a range of distribution measures to determine whether results are robust in terms of sensitivity at the middle versus the extremes of the income distribution.

4 Data

Data are drawn from the National Household Surveys (Pnads), conducted by the Brazilian Institute of Geography and Statistics between 1995 and 2013. The annual survey is national in character, but rural areas of the Amazon region are excluded from analysis for the sake of comparability, as data collection in these locations began only in 2004. All surveys were undertaken in the context of price stability achieved in 1994.
The unit of observation is “equivalent household income,” constructed by dividing household income by an equivalence scale of the form \((A_i+kC_i)^s\). \(A_i\) and \(C_i\) refer to the numbers of adults and children, respectively, in the i-th 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 limits the extent of scale economies in developing countries, and the choice of \(s\) attempts to reflect the restriction. In comparison, the equivalence scale of the UK implies the parameter values \(k=0.53\) and \(s=0.77\) (Jenkins and Cowell, 1994) while an alternative US poverty measure explicitly chose \(k=0.7\) and \(s=0.65–0.75\) (Citro and Michael, 1995).

Equivalent income is weighted by the sum of the sample weights of household members, thus deriving a personal distribution of income under the premise 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 and domestic employees and their relatives, they are excluded from analysis. 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, the month in which the precursor of the current monetary unit was introduced.

There are no official poverty thresholds in Brazil, and in their place the study adopts those developed by Rocha (1997, 2003), based on per capita expenditures required to comply with a nutritional norm of 2,100 calories per person per day and augmented to incorporate non-food consumption. They take into account cost of living differences across Brazil’s main metropolitan areas and between metropolitan, urban, and rural areas. Metropolitan São Paulo has the highest poverty line at R$95.7, and Belém the lowest at R$64.2 (August 1994 currency values); the latter figure is comparable to the national minimum wage at the price index base period. Urban non-metropolitan and rural poverty lines average the equivalent of 73% and 43% of corresponding metropolitan
area thresholds, respectively.\textsuperscript{6}

Income distributions are summarized by Atkinson and Foster-Greer-Thorbecke-class distribution measures. As well as satisfying a number of useful properties, both are defined by a parameter that generates sensitivity to income changes at different segments of the distribution. With \( z \) as the poverty threshold, \( n \) the number of observations, \( n_p \) the number of observations classified as poor, \( y_i \) the \( i \)-th income of the poor, and \( \alpha \) a parameter value, the FGT index is formulated as

\[
\text{FGT}(\alpha) = \frac{1}{n} \sum_{i=1}^{n_p} \frac{(z-y_i)^\alpha}{z^\alpha}.
\]

Measures defined by parameter values over 0 satisfy the monotonicity axiom, and those defined by values over 1 result in indices that also respect the transfer principle (Foster, Greer, and Thorbecke, 1984). In practice, the most common choices for \( \alpha \) have been 0, 1, and 2 (Foster, Greer, and Thorbecke, 2010), producing measures that have been termed to reflect the incidence, depth, and severity of poverty, respectively (Ravallion, 1994).

Atkinson-class inequality indices, which satisfy the principle of transfers and scale and population independence, are based on social welfare theory and defined by a parameter that explicitly incorporates aversion to inequality, with higher values generating measures that are more sensitive to income changes at the lower end of the income distribution. The class is formulated as

\[
A(\alpha) = 1 - \frac{\xi}{\mu}, \quad \text{where} \quad \xi = \left[ \frac{1}{n} \sum_{i=1}^{n} y_i^{1-\alpha} \right]^{\frac{1}{1-\alpha}}, \quad y_i \text{ refers to the income of the } \text{i-th observation}, \mu \text{ indicates average income, and } \alpha \text{ and } n \text{ are as already defined. Amiel, Creedy, and Hurn (1999)} \text{ find that a parameter value of 0.25 is consistent with elicited attitudes towards inequality; therefore, it is adopted for use in this study. Nonetheless, two other indices reflecting greater aversion to inequality and defined by parameters 0.5 and 0.75 are also incorporated to evaluate robustness of results. Finally, inequality measurements can be particularly sensitive to extreme values of potentially contaminated data resulting from coding problems, income under-reporting, or other issues (Cowell and Victoria-Feser, 2002), and the concern is addressed by data trimming, specifically by removing from analysis the bottom and top 0.5\% of observations of the distribution of equivalent household income.}

\textsuperscript{6}States without large metropolitan areas are assigned thresholds equivalent to the mean threshold of their region. Timmins (2006) finds that it is reasonable to extrapolate in this manner, as is commonly done given the absence of price data outside the major metropolitan areas.
5 Results

The proposed empirical strategy presents a number of advantages, but one drawback of any sequential approach is that conclusions can depend on the order in which explanatory factors enter in the decomposition sequence. The order could be reversed and results compared, but many other possibilities would remain unexplored. Thus, to account for as much information as possible and avoid determinations that hinge on arbitrary orderings, decompositions are performed in all $4!$ possible sequences; the average results are presented in Table 4.

[Table 4]

These establish that the impressive changes in distribution observed in Brazil during the 2000s and the early 2010s benefited from a demographic dividend that served to reduce inequality, the headcount ratio, and especially, the prevalence of severe poverty. The disproportionate consequences at the bottom of the distribution of income can be appreciated through counterfactual densities that are differenced from the observed ones to arrive at demographic effects. Depicted in Figure 3, these show that the rise in female headship had small negative impacts on income that were counterbalanced by also small consequences of urbanization and population aging. However, by far the most important positive element was declining dependency, whose influence was much greater and centered on an income level equivalent to one-half of the poverty line. Altogether, demographics accounted for up to 16% and 22% of the long-term decline in inequality and poverty, respectively.

[Figure 3]

Changes in labor force participation associated with the business cycle or with structural factors such as rising female participation, could have provided additional support to the trends in distribution, but results show that this was not the case, at least over the long term. The explanatory power of the work structure variable is small, likely reflecting countervailing effects associated with the fact that the frequency of both zero-income and
two-income households rose over 1995-2013. During the 2000s, however, the rise in two-income households was the dominant phenomenon, a trend that was a minor contributor to the headcount ratio decline evident during the decade. The limited influence of the variable even during the upturn suggests that positive income changes occurred within rather than across household types. That is, rather than labor force participation effects, changes in earnings among participants were the main force behind the movements in distribution.

Increased educational attainment, with associated earnings growth, accounted for well over half of the post-Real Plan reduction in poverty and up to 38% of that taking place during the upturn. Household incomes also grew as a result of changes in skill prices that explain about a quarter of the remarkable fall in the headcount ratio that occurred during the 2000s and the early 2010s. The influence of skill prices declines as a function of the sensitivity of the poverty measure to income changes at the bottom of the distribution, a relationship that is consistent with mechanisms that were more effective in helping households cross the poverty line than in raising incomes of the very poor, who compared to higher-income population subgroups, are more likely to be economically inactive.7

Equality was also enhanced by increased educational attainment and changes in skill prices. Skill price changes were the far more important element, in accounting for up to two-thirds of the long-term decline in inequality and over half of the shorter-term trend. Of particular interest is the portion associated with a minimum wage, which grew by 70% during the economic upturn, and results show that changes in lower-skill prices were responsible for over 80% of the equalizing skill price effect.

Impact on poverty was even stronger, as the entire poverty-reducing effect of skill prices was related to the interaction between economic growth and a rising wage floor, an assessment that stands in contrast to pre-economic expansion data that reflect limited minimum wage consequences on family income or poverty (Ipea, 2000, Barros et al., 2001, Neumark, Cunningham, and Siga, 2006). Growth and accompanying increases in

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7Averaged over 1995-2013, households with incomes under 0.5 times the poverty line are three times more likely to be economically inactive than households with incomes between 0.5 and 1 times the poverty threshold.
labor demand can counter the unemployment costs of the wage floor, leaving benefits primarily in the form of higher wages among workers directly affected by the wage floor, and possibly among those indirectly affected through spillover effects higher up in the wage distribution. In order to attempt to reconcile past and current findings, counterfactual trends are constructed to determine how poverty would have behaved throughout 1995-2013 in the absence of changes in lower-skill prices. These show that the positive association between a rising wage floor and lower poverty is observed only during the period encompassed by the economic upturn.\(^8\)

[Figure 4]

Figure 4 sheds further light on the mechanisms at work. Household earnings trends displayed in the top panels suggest that falling inequality was associated with growth in earnings, which was strongest at the bottom half of the distribution. Positive trends in earnings accruing to women at the lower end of the distribution predate the expansion, in contrast to those of men, whose earnings fell at most segments of the distribution in the preceding years. The economic upturn was thus an exceptional period of across-the-board earnings growth that also diminished inequality through stronger gains at the lower end of the income distribution.

Household earnings growth rates shown in the bottom panels confirm the earnings decline among men in the years leading to the expansion. In contrast, growth during the upturn was broad-based, robust, statistically significant, and negatively correlated with position in the distribution, peaking in the bottom quintile, where male and female household earnings grew by 77% and 106%, respectively. Moreover, changes were equalizing in nature, as reflected by statistically significant Lorenz-dominance of distributions of male and female earnings in 2013 over the corresponding distributions in 2003.\(^9\)

Demographics, education, skill prices, and changes in the work structure of households together account for 77% to 85% of the impressive decline in inequality experienced in Brazil during the 2000s and the early 2010s, with education and skill differentials

\(^8\)See the online appendix.
\(^9\)Statistical significance is based on bootstrapped standard errors and on Lorenz curves estimated using Stata’s svylorenz module developed by Stephen Jenkins.
explaining the bulk of the change. Demographics, work structure, education, and changes in lower-skill prices explain 85% of the fall in the headcount ratio and a somewhat smaller share of the fall in the FGT(1) and FGT(2) indices. Education comes out as the dominant element followed by favorable changes in skill prices, which appear closely tied to a combination of economic growth and wage floor effects.\textsuperscript{10}

6 Public transfers and geography

A thesis pointing to the labor market as the primary source of striking advances in Brazilian distribution would benefit from direct estimation and confirmation of small roles attributable to pension and conditional transfer income. Such an evaluation, however, cannot be confidently accommodated by the strategy laid out in Section 3, as there are no obvious models for incorporating public transfers through a regression approach, contrary to labor income determination, informed by the theory of human capital. Instead, a rank-preserving income exchange mechanism is adopted (Reed and Cancian, 2001), and extended to allow estimation of counterfactual trends in distribution.

Implementation involves a first step that consists of summarizing a distribution of (say) year $t$ pensions by 1000 points, where each individual in a given age range or other agreed-upon population characteristic is associated with a ranking that gives the person’s millicile in that distribution. Coupled with the rankings, the function $\mu_t(\cdot)$ returns the corresponding pension income level, calculated as the mean of all individuals sharing the same millicile in year $t$. The equivalent income in household $i$ in year $t$, $y_{it}$, is then arrived at as follows:

$$y_{it} = \left[ \sum_{j \in [q_{it}]} \mu_t(r_j) + o_{it} \right] \times e(a_{it}, c_{it}),$$

where $q_{it}$ is a list of pension income rankings of all pensionable-age individuals in household $i$ at time $t$, $r_j$ is the j-th pension income ranking in the household, $o_{it}$ represents income from sources other than pensions, and $e(a_{it}, c_{it})$ is an equivalence scale that uses\textsuperscript{10}.

The online appendix establishes that results are not materially sensitive to the choice of equivalence scale, poverty line, or decomposition period.

\textsuperscript{10}The online appendix establishes that results are not materially sensitive to the choice of equivalence scale, poverty line, or decomposition period.
information on the number of adults and children in a household. That is, household income is the sum of pension income of all pensionable-age household members—where each is assigned a pension corresponding to his or her ranking in the pension distribution—plus all other income accruing to the household. Similarly, the income in a future year, $t'$, which would obtain only if the distribution of pension income had remained as in year $t$, can be expressed as

$$y_{it'} = \left[ \sum_{j \in q_{it'}} \mu_t(r_j) + o_{it'} \right] \times e(a_{it'}, c_{it'}).$$

Income in year $t'$ is the sum of pension income of all pensionable-age household members—estimated with year $t'$ rankings, but with the corresponding pension levels of the distribution in year $t$—plus year $t'$ household income from all sources other than pensions. Any discrepancy between an observed and a counterfactual change in inequality can then be attributed to changes in the distribution of the income source being held constant. That is, if in comparison to an observed fall of 20%, inequality falls by 10% in a counterfactual that holds pension distribution constant, changes in the income source can be identified as equalizing in nature and associated with 50% of the observed change.

The procedure presents a number of advantages over other approaches such as inequality index decompositions by income source. It is independent of the distribution measure, can estimate contributions to changes over time, and can distinguish between changes in distribution and changes in household structure.\textsuperscript{11}

### 6.1 Retirement income

Pension 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). Linked to a minimum wage that rose markedly through the 2000s and early 2010s, pensions could have had important effects on both poverty and inequality, serving to either raise or lower the latter. To evaluate their impact, counterfactuals are

\textsuperscript{11}For more information, see Sotomayor (2009), where the exchange mechanism is used to ascertain the influence of rising labor market importance of women on the distribution of income in Brazil.
constructed to arrive at an estimate of how inequality and poverty would have behaved in the absence of changes in pension distribution.\textsuperscript{12}

Results in Table 5 establish that the income source played an equalizing role associated with 13\% to 15\% of the post-2001 decline in inequality and with 20\% to 22\% of the long-term fall. Its influence on poverty was less substantial, responsible for just 11\% of the long-term drop in the headcount ratio and 9\% of the decline in the FGT(2) measure. Pension increases therefore generated only modest benefits to low-income individuals, but observed declines in poverty could still be related to policies placing greater focus on the bottom of the income distribution, notable examples being conditional and unconditional transfers.

\section*{6.2 Bolsa Escola/Família}

In 2001, Brazil introduced at the national level a program that offered an income supplement, based on a number of conditions. \textit{Bolsa Escola} paid a monthly allowance of about US$5 per child to low-income households that agreed to register and ensure school attendance and scheduled vaccinations for minors. Two years later the program, along with other initiatives that provided small sums of money for food and cooking gas purchases, were folded into a new policy instrument named \textit{Bolsa Família}. This new program also provided an unconditional income supplement to very low-income households, regardless of family composition. Over time, the focus of \textit{Bolsa Família} has evolved from conditions to transfers, generating controversy over its potential for encouraging dependency and political patronage (Hall, 2008), as well as its high costs relative to efficiency benefits (Glewwe and Kassouf, 2012).

The assessment of the distributive impact of the program is hindered by the absence of a dedicated entry for \textit{Bolsa Escola/Família} income in the Pnad survey. Rather, it

\textsuperscript{12}Retirement income comes in the form of old age (aposentadoria) and other regular payments (pensão) that can be drawn relatively early in life in Brazil. The distribution of retirement income is therefore held constant among the population aged at least 45 years. Demographic characteristics are also held constant through use of the DFL method.
is classified as “other” income, along with interest and dividends. The situation is not hopeless, since program rules can be used to get a better measure of the income source. “Other” income from households not meeting the qualification rules of the program could be recoded to zero in a threshold approach. A stipend approach could exploit the fact that benefits are standard sums—for example, *Bolsa Escola* and basic family allowances are R$15 and R$50, respectively. Both thresholds and stipends have been revised since 2006, but are published, and are straightforward to identify.

An approach combining the threshold and stipend rules is used to identify *Bolsa Escola/Família* income and construct counterfactuals that establish that the policy instrument had positive, but limited, effects on inequality associated with just 3% to 5% of the 2001-2013 fall. Small but beneficial effects on poverty are also evident, with somewhat larger consequences reflected by income- and distribution-sensitive poverty measures. Up to 6% of the remarkable fall in poverty that took place in the 2000s and the early 2010s can be associated with the income source.

### 6.3 Geography

Lastly, geographical factors in a continent-sized country could have played a role in distribution trends. Thresholds used to identify poverty take into account regional differences in the cost of acquiring basic necessities, but adjusting for inflation through a national price index assumes that they did not change over time. This need not be the case. Regional cost of living differences can fall as input and output markets become more integrated across a country, but can also become more pronounced to the extent that policies such as trade reforms have greater impact on some areas than others. This has been shown to be the case in Brazil (Kovak, 2013).

Potential impact on poverty trends can be evaluated using thresholds revised with region-specific price indices, rather than a single national one. Implementation establishes limited consequences on poverty estimates (Table 6), with an assessment of a long-term
decline that is 2% greater than that based on a national price index. Behind the result is the fact that inflation was higher in some areas than in others, though not significantly. That is, whereas the price index rose by a compounded annual average of 6.5% at the national level between 1995 and 2013, regional changes were bounded at the upper end by an increase of 7% in Rio de Janeiro and at the lower end by an increase of 6.3% in Fortaleza in the poorer Northeast.

Cost of living differences could also affect the assessment of both levels and trends in inequality, and Timmins (2006) develops utility-based deflators that can be used to estimate real income inequality. These measures, which include state-level estimates, are disaggregated at the Brazilian equivalent of a county, a variable unavailable in annual survey data. Their application to the data used in this study lowers inequality levels by some 3%, with little impact on trends. Almeida and Azzoni (2013) also take on the issue of establishing regional income deflators, with the added task of ascertaining cost of living changes over time. Costs are determined to be highest, at 14% above average, in Brasília; and lowest, at 16% under average, in Fortaleza, with changes over time generally reflecting regression on the mean. Use of deflators lowers inequality estimates by some 3% and changes the trend estimates, though not to a large degree, revising declines upward by 1% relative to amounts not accounting for changes in cost of living differences.

Geography appears to play a more important role in distribution through regional initial conditions or characteristics than through prices. That is, whereas declining inequality was a general phenomenon across Brazil, it was below average in the country’s main metropolitan areas. Income growth outside these traditional centers of economic and political power was stronger, and inequality fell 9% to 13% over the national trend. It fell even more, by 37% to 45% over the national trend in the Brazilian South, a region with the highest levels of income, and equality, and arguably the country’s strongest institutions. Thus, favorable initial conditions, associated in past research with increased elasticity of poverty alleviation with respect to growth (Ferreira, Leite, and Ravallion, 2010), also appear to have the same effect on inequality.
Conclusion

Results show that the remarkable period of redistribution observed in Brazil during the 2000s and early 2010s benefited from elements that had been in place for some time. A shift towards households with fewer children relative to adults generated an important demographic dividend that reduced inequality, the headcount ratio, and especially, the prevalence of severe poverty. Long-standing progress in increasing educational attainment exerted downward pressure on skill premia, particularly in the context of a stable macroeconomic environment. A singular feature of the period was a robust upturn in the business cycle associated with across-the-board income gains that were widely shared, owing to mechanisms that favored advances at the lower end of the distribution of earnings.

Demographics, as well as changes in educational attainment, labor force participation, and prices to lower levels of skill explain 85% of the impressive fall in the headcount ratio experienced during the 2000s and early 2010s, with education and interaction between economic growth and a rising wage floor accounting for the bulk of the change. The same factors explain 77% to 85% of the decline in income dispersion that was powered to a large extent by markedly compressed earnings differentials. Human capital and strong labor market institutions thus stand out as key mechanisms linking economic growth to redistribution. Price stability could also have played an equalizing role to the extent that it facilitated a gradual de-indexing of the economy and a phase-out of practices that hard-wired skill differentials and favored groups better placed to cope with inflation.\textsuperscript{13}

These developments are timely for a country that has historically struggled to find effective tools for reduction of poverty and inequality. In terms of policy implications, however, the evidence presented does not provide unqualified support in favor of minimum wage increases as effective instruments of redistribution, since more research is required for isolating effects from other confounding factors such as concurrent changes in labor demand. Results do establish that in the context of a robust economic expan-

\textsuperscript{13}For example, as inflation recedes, wages may gradually cease to be set in multiples of a reference such as the minimum wage.
sion, a rising wage floor can foster a wider distribution of the benefits of growth, and lead to substantial reductions in poverty and income inequality, where the challenge involves improving monetary and fiscal policy management and the policies’ potential for promoting full employment. In addition, results suggest that investment in education and other initiatives that augment individual and social capabilities can lead to income redistribution by enhancing prospects for long-term growth.

8 Bibliography


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Salardi, P.; Wage disparities and occupational intensity by gender and race in Brazil: An empirical analysis using quantile decomposition techniques, University of Sussex (2012).


Table 1  
Descriptive statistics

<table>
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<th></th>
<th></th>
<th></th>
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<th></th>
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<tbody>
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<td><strong>Households:</strong></td>
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<td></td>
<td></td>
<td></td>
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<tr>
<td>Equivalent monthly income ('94 R$)</td>
<td>188</td>
<td>180</td>
<td>168</td>
<td>235</td>
<td>261</td>
</tr>
<tr>
<td>Equivalent adults</td>
<td>3.55</td>
<td>3.36</td>
<td>3.30</td>
<td>3.04</td>
<td>2.98</td>
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<td>Ratio of children to adults</td>
<td>.436</td>
<td>.385</td>
<td>.367</td>
<td>.284</td>
<td>.274</td>
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<td>Share in urban areas</td>
<td>.790</td>
<td>.838</td>
<td>.842</td>
<td>.863</td>
<td>.861</td>
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<tr>
<td>FGT(0)</td>
<td>.239</td>
<td>.257</td>
<td>.269</td>
<td>.115</td>
<td>.091</td>
</tr>
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<td>FGT(1)</td>
<td>.088</td>
<td>.100</td>
<td>.104</td>
<td>.044</td>
<td>.037</td>
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<td>FGT(2)</td>
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<td>.057</td>
<td>.058</td>
<td>.027</td>
<td>.024</td>
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<td>Atkinson(.25)</td>
<td>.130</td>
<td>.127</td>
<td>.121</td>
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<td>.241</td>
<td>.230</td>
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<td>Atkinson(.75)</td>
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<td>.349</td>
<td>.330</td>
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<td>.276</td>
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<td>Number of records (1000s)</td>
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<td>101</td>
<td>105</td>
<td>104</td>
<td>109</td>
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<td><strong>Householder:</strong></td>
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<td></td>
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<tr>
<td>Age</td>
<td>45.2</td>
<td>45.5</td>
<td>46.1</td>
<td>46.9</td>
<td>47.6</td>
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<td>Years of schooling</td>
<td>4.64</td>
<td>5.34</td>
<td>5.65</td>
<td>6.82</td>
<td>7.16</td>
</tr>
<tr>
<td>Share that is female</td>
<td>.162</td>
<td>.206</td>
<td>.222</td>
<td>.341</td>
<td>.361</td>
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<td><strong>Working individuals:</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Age</td>
<td>35.6</td>
<td>35.9</td>
<td>36.2</td>
<td>37.3</td>
<td>37.8</td>
</tr>
<tr>
<td>Years of schooling</td>
<td>6.30</td>
<td>7.21</td>
<td>7.61</td>
<td>8.80</td>
<td>9.13</td>
</tr>
<tr>
<td>Share that is female</td>
<td>.369</td>
<td>.389</td>
<td>.397</td>
<td>.416</td>
<td>.422</td>
</tr>
<tr>
<td>Monthly income from work ('94 R$)</td>
<td>347</td>
<td>314</td>
<td>279</td>
<td>362</td>
<td>396</td>
</tr>
<tr>
<td>Number of records (1000s)</td>
<td>117</td>
<td>137</td>
<td>143</td>
<td>144</td>
<td>147</td>
</tr>
</tbody>
</table>

Note: Household- or householder-based observations are weighted by the sum of the sample weights of household members to arrive at statistics presented in the top two panels. Descriptive statistics referring to working individuals, shown in the bottom panel, are weighted by the corresponding person weights.

Source: Author’s calculations from household- and person-based Pnad data.
Table 2
Share (%) of the population by work status and gender of the head of household and spouse

<table>
<thead>
<tr>
<th></th>
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<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Lone non-working male</td>
<td>1.1</td>
<td>1.3</td>
<td>1.3</td>
<td>1.5</td>
<td>1.6</td>
<td>2.0</td>
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<tr>
<td>Lone working male</td>
<td>2.7</td>
<td>2.8</td>
<td>2.9</td>
<td>3.2</td>
<td>3.4</td>
<td>3.9</td>
</tr>
<tr>
<td>Lone non-working female</td>
<td>7.9</td>
<td>8.9</td>
<td>9.4</td>
<td>9.7</td>
<td>10.3</td>
<td>11.0</td>
</tr>
<tr>
<td>Lone working female</td>
<td>7.7</td>
<td>7.9</td>
<td>8.6</td>
<td>9.5</td>
<td>10.0</td>
<td>10.1</td>
</tr>
<tr>
<td>Working male &amp; non-working female</td>
<td>43.9</td>
<td>41.3</td>
<td>38.4</td>
<td>34.6</td>
<td>30.7</td>
<td>28.2</td>
</tr>
<tr>
<td>Working female &amp; non-working male</td>
<td>3.3</td>
<td>3.7</td>
<td>4.0</td>
<td>4.1</td>
<td>4.2</td>
<td>4.0</td>
</tr>
<tr>
<td>Working male &amp; female</td>
<td>25.1</td>
<td>24.3</td>
<td>25.5</td>
<td>28.0</td>
<td>29.8</td>
<td>29.5</td>
</tr>
<tr>
<td>Non-working male &amp; female</td>
<td>8.3</td>
<td>9.9</td>
<td>9.8</td>
<td>9.5</td>
<td>9.9</td>
<td>11.2</td>
</tr>
</tbody>
</table>

Source: Author’s calculations from household-based Pnad data.
## Table 3
### Sequential decomposition

<table>
<thead>
<tr>
<th>Date of survey</th>
<th>Counterfactual holds constant:</th>
<th>Income definition</th>
<th>Reweighting factor</th>
</tr>
</thead>
<tbody>
<tr>
<td>1995/2001</td>
<td>-</td>
<td>Equivalent Income</td>
<td>none</td>
</tr>
<tr>
<td>1995/2001</td>
<td>Demographics</td>
<td>Equivalent Income</td>
<td>$\Psi_D$</td>
</tr>
<tr>
<td>1995/2001</td>
<td>+ Household work structure</td>
<td>Equivalent Income</td>
<td>$\Psi_D \times \Psi_S$</td>
</tr>
<tr>
<td>1995/2001</td>
<td>+ Schooling of economically active</td>
<td>Equivalent Income</td>
<td>$\Psi_D \times \Psi_S \times \Psi_E$</td>
</tr>
<tr>
<td>1995/2001</td>
<td>+ Lower &amp; higher-skill prices</td>
<td>Estimated EI</td>
<td>$\Psi_D \times \Psi_S \times \Psi_E$</td>
</tr>
<tr>
<td>2013</td>
<td>-</td>
<td>Equivalent Income</td>
<td>none</td>
</tr>
</tbody>
</table>

Note: $\Psi_D$ is estimated through a logit with a binary for survey year 2013 as dependent variable and as independents gender of the head of household, urban area residence, and third-order polynomials in age of the head of household and per household ratio of children to adults. $\Psi_S$ 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. $\Psi_E$ comes from an ordered logit where the dependent variables are 13 range possibilities for mean schooling of working household members. Independent ones include a third-order polynomial in mean age of economically active household members, share of labor force participants that is female, and a binary for metropolitan area residence. The price effect is estimated in two steps, first among households that would be constrained by the minimum wage in the reference distribution, and then for all other households.
Table 4
Share of fall in distribution index accounted by demographics, household work structure, schooling, and returns to education

<table>
<thead>
<tr>
<th>Factor</th>
<th>1995-2013 Share (%) of fall in index accounted by factor:</th>
<th>2001-2013 Share (%) of fall in index accounted by factor:</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>A(.25) A(.50) A(.75) FGT0 FGT1 FGT2</td>
<td>A(.25) A(.50) A(.75) FGT0 FGT1 FGT2</td>
</tr>
<tr>
<td>Demographics</td>
<td>11.1 12.4 15.5 9.2 14.7 21.8</td>
<td>10.3 10.9 12.7 7.4 9.9 12.7</td>
</tr>
<tr>
<td>Work structure</td>
<td>-0.1 -1.3 -4.6 2.8 0 -4.9</td>
<td>1.5 0.4 -2.1 3.9 1.8 -0.9</td>
</tr>
<tr>
<td>Education</td>
<td>25.3 20.1 16.0 55.2 61.2 70.5</td>
<td>16.3 12.0 8.4 35.7 36.4 37.6</td>
</tr>
<tr>
<td>Lower-skill prices</td>
<td>39.3 40.1 41.1 34.3 31.3 27.5</td>
<td>50.4 50.3 48.9 37.7 35.0 31.4</td>
</tr>
<tr>
<td>Higher-skill prices</td>
<td>19.4 22.8 25.8 -24.5 -28.9 -33.9</td>
<td>6.6 8.6 9.4 -11.5 -13.0 -13.8</td>
</tr>
</tbody>
</table>

Source: Author’s calculations from household-based Pnad data.

Table 5
Share of fall in distribution index accounted by pensions and by conditional transfers

<table>
<thead>
<tr>
<th>Income</th>
<th>1995-2013 Share (%) of fall in index accounted by income source:</th>
<th>2001-2013 Share (%) of fall in index accounted by income source:</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>A(.25) A(.50) A(.75) FGT0 FGT1 FGT2</td>
<td>A(.25) A(.50) A(.75) FGT0 FGT1 FGT2</td>
</tr>
<tr>
<td>Pensions</td>
<td>22.1 21.1 20.1 10.9 9.4 8.8</td>
<td>15.2 14.3 12.7 5.2 3.0 1.6</td>
</tr>
<tr>
<td>Transfers</td>
<td>2.8 3.5 5.0 1.9 4.9 9.0</td>
<td>3.0 3.9 5.2 1.6 3.8 6.2</td>
</tr>
</tbody>
</table>

Source: Author’s calculations from household-based Pnad data.
<table>
<thead>
<tr>
<th>Cost of living adjustment/regional trend</th>
<th>A(.25)</th>
<th>A(.75)</th>
<th>FGT0</th>
<th>FGT2</th>
</tr>
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<tbody>
<tr>
<td><strong>Cost of living (CL)</strong></td>
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<td></td>
<td></td>
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<tr>
<td>Unadjusted initial (1995) values</td>
<td>.130</td>
<td>.352</td>
<td>.239</td>
<td>.047</td>
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<tr>
<td>Adjustment for regional CL differences$^{a,b}$</td>
<td>.127</td>
<td>.343</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Adjustment for regional CL differences$^{a,c}$</td>
<td>.126</td>
<td>.341</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Unadjusted final (2013) values</td>
<td>.097</td>
<td>.276</td>
<td>.091</td>
<td>.024</td>
</tr>
<tr>
<td>Adjustment for regional CL differences$^{c,d}$</td>
<td>.094</td>
<td>.267</td>
<td>.089</td>
<td>.023</td>
</tr>
</tbody>
</table>

**Unadjusted % change in distribution as ratio of national % change (1995-2013)**

<table>
<thead>
<tr>
<th>Region</th>
<th>A(.25)</th>
<th>A(.75)</th>
<th>FGT0</th>
<th>FGT2</th>
</tr>
</thead>
<tbody>
<tr>
<td>North</td>
<td>.732</td>
<td>.587</td>
<td>.866</td>
<td>.638</td>
</tr>
<tr>
<td>Northeast</td>
<td>1.01</td>
<td>.917</td>
<td>.984</td>
<td>1.15</td>
</tr>
<tr>
<td>Southeast</td>
<td>.860</td>
<td>.884</td>
<td>.956</td>
<td>.732</td>
</tr>
<tr>
<td>Midwest</td>
<td>1.05</td>
<td>1.05</td>
<td>1.19</td>
<td>1.26</td>
</tr>
<tr>
<td>South</td>
<td>1.37</td>
<td>1.45</td>
<td>1.24</td>
<td>1.09</td>
</tr>
</tbody>
</table>

Note: $^a$Poverty thresholds developed by Rocha (2003) already account for regional differences in cost of living. $^b$Timmins (2006) provides deflators for arriving at initial period estimates of real income inequality. $^c$Almeida and Azzoni (2013) provide deflators for initial and final periods. $^d$Final period poverty thresholds are revised using region-specific inflation rates. Results associated with the FGT(1) and A(.50) measures fall between those shed by the indices in the table, and are excluded to reduce clutter. The full table is available from the author on request.

Source: Author’s calculations from household-based Pnad data.
Figure captions

Figure 1: Rate of return to an additional year of school, intermediate education, high school, and university, by percentile in the earnings distribution

Source: Author’s calculations from person-based Pnad data.

Figure 2: Minimum wage, rate of return to primary education, earnings function intercept, and log earnings by schooling attainment

Source: Portalbrasil.net/salariominimo (top left panel) and author’s calculations from person-based Pnad data (remaining panels).

Figure 3: Observed change in the density of household income over 1995-2013 and contributions associated with population aging, a declining dependency ratio, increased urbanization and prevalence of female-headed households

Source: Author’s calculations from household-based Pnad data.

Figure 4: Household earnings at the median and at the top and bottom deciles and quartiles (top panels) and household earnings growth by gender of the income earner and percentile in the distribution (bottom panels)

Note: aIn households with working men. bIn households with working women.
Source: Author’s calculations from household-based Pnad data.
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