Age and Education in the Course of Development: Does Composition Matter?

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Abstract

In this analysis, we estimate the impact of the changing relative size of the adult male population classified by age and education groups on the earnings of employed males living in 502 Brazilian local labor markets during four time periods between 1970 and 2000. The effects of shifts in the age distribution of the working age population have been studied in relation to the effect of the baby-boom generation on the earnings of different cohorts in the U.S. However the question has received little attention in the context of the countries of Asia and Latin America, which are now experiencing substantial shifts in their age-education distributions. Taking advantage of the huge variation across Brazilian local labor markets, our models suggest that age-education groups are not perfect substitutes, so that own cohort-education size depresses earnings, as expected by the theory. Compositional shifts are influential, attesting that this approach represents a fruitful way of studying this central problem in economic development, going beyond the effects normally analyzed by formal labor market equations.

1. Introduction

This study estimates the impact of the changing relative size of the adult male population classified by age and education groups on the earnings of employed males living in 502 Brazilian local labor markets during four time periods comprising a thirty-year time span. The paper takes advantage of the substantial variation in age and educational structure across and within these local labor markets over the 1970-2000 period to test the impact of the relative size of age-education groups on earnings.

A basic tenet of the demographic dividend framework is the following identity:

y = per capita income = $(Y/EWA)*((\theta*WA)/P)$

where:

Y is total income

EWA is the employed population

WA is the working age population

θ is the employment rate

P is total population.

The first demographic dividend is determined by the independent impact of the age structure (WA/P) on per-capita income. Productivity is defined as (Y/EWA), which is basically affected by the production function. This productivity factor can also be affected by a shift in the population's age structure, causing the so called second demographic dividend, which is the impact of population aging on capital accumulation via capital deepening. Another component of this productivity factor is the average earnings of the employed population. It is precisely this component that can be exogenously affected by shifts in the age and educational structures, according to the shape of the labor demand curve for each age and educational labor factor.

We contribute to the demographic dividend literature by inserting the cohort size and supply-demand frameworks in the determination of the productivity component that appears in the basic identity described above. This is done in the context of local labor markets and longerrun variation in the age-education structure of the labor force. By incorporating this structure into a formal model of labor demand, we identify an additional indirect effect of age structure on per-capita income. This effect is ignored in the literature that centers its analysis only on the first demographic dividend, namely, on the direct impact of age structure on per capita income.

2. Literature Review

The decline in the dependency ratio caused by rapid fertility decline has been shown to have influenced economic development in the countries of East and Southeast Asia (e.g., Bloom, Canning and Sevilla 2003; Bloom, Canning and Malaney 2000; Bloom and Freeman 1986; Mason 2005; Mason and Feng 2005; Williamson 2003). The focus of those studies was on the shifting age distribution of the populations rather than the rate of population growth. The dependency ratio first increased after the mortality decline at the beginning of the demographic transition, but then fell after fertility began to fall precipitously in those countries. This process has been called a "demographic dividend" whereby a changing age distribution allows for fewer investments in the youngest cohorts, enabling resources to be allocated for investments in economic development and family welfare. The higher proportion of people in adult age groups is a temporary effect since, after some decades, this population will age and the dependency ratio will again increase. Because of the temporary nature of the dependency ratio decrease, this process has also been called a "window of opportunity" for the implementation of specific policies to generate economic growth.

The significance of fertility swings and a shifting age distribution on economic development was also analyzed in studies of the influence of the "baby boom" on labor market outcomes in the United States (Freeman 1979; Welch 1979). Freeman (1979) analyzed the effect of changes in the age structure of the workforce on age-earnings profiles in the U.S. Because of the "baby boom" that followed World War II and peaked between 1955 to1960, there was an especially significant change in the age structure of the U.S. workforce in the late 1960s and early

1970s, when the number of young persons increased very rapidly. The principal finding was that the age-earnings profile of male workers appears to be significantly influenced by the age composition of the workforce. In previous studies, the age-earnings profile was usually viewed as a stable economic relationship determined by human capital investment decisions, assuming that earnings rise only with age and experience as a result of individual investment behavior. Freeman indicates that from the late 1960s through the mid-1970s, when the number of young workers increased rapidly, the earnings of young male workers fell relative to the earnings of older male workers, altering male age-earnings profiles, particularly for college graduates.

Welch (1979) analyzed the 1968-1976 March Current Population Surveys (CPS) in order assess the impact of the change in age composition experienced by the U.S. due to the entrance of the post-World War II baby boom cohorts in the job market. The main hypothesis was again that the changing age composition of the workforce affected earnings patterns. The key finding was that the pressure of a workforce whose average age is rapidly declining reduces wages of new entrants. Moreover, cohort size depresses earnings, and most of the effect is felt early in one's career. Welch also suggests that cohort size-effects increase with the level of schooling. Berger (1985) suggest that the cohort-size effect persists over much of the working lives of large cohorts.

`The cohort-size studies suggested that shifts in factor supply (the baby boom) led to a decline in wages, so that demand shifts did not explain all the wage variation. By the same token, an increase in the supply of skilled labor should lead to a relative decline in the wage of skilled relative to unskilled labor. In the context of a production function with a constant elasticity of substitution (CES) and downward-sloping demand for relative skill, an increase in the provision of skilled labor will lead to a decline in the skill premium (defined as the wage of skilled workers divided by that of unskilled). In developed countries, in contrast, the skill premium increased while the supply of educated workers has risen steadily. Katz and Murphy (1992) found that the relative supply of skilled labor combined with smoothly rising demand explain US relative wage trends between 1967 and 1987. Autor, Katz, and Kruger (1998) used a longer time series to test the smooth rising demand hypothesis, and found some evidence that accelerating demand explained the US wage premium shifts. An alternative explanation for the rising wage premium is the role of trade, the US engagement with countries in which skills are relatively scarce. An institutional explanation may also be suggested to the extent that the real minimum wage and the bargaining power of unions declined during this period.

Triest, Sapozhnikov and Sass (2006) have conducted the most recent analysis of population aging and the structure of wages in the United States. Their analysis explores the effect of labor market experience, relative cohort size and real wage growth on real wage by level of education using the March Current Population Survey (CPS) from 1964 and 2004. Their models indicate that: (1) increases in relative cohort size are associated with decreases in wages; (2) although real wages initially increase with labor market experience, there is a significant decrease in the rate of growth as experience increases; (3) there was a general increase in the economic return to educational attainment; (4) changes in the age and experience composition of the labor force will continue to have an important influence on the structure of wages; (5) the initial increase in the experience premium generated by the baby boom's entry into the labor market is now being reversed as the baby boom progresses through middle age and approaches retirement. More specifically, Triest, Sapozhnikov and Sass (2006) emphasize that baby boomers born in 1950 were a large fraction of the college educated labor force when they entered the labor market. At that time, their wages would have been 18 percent higher if their relative cohort size was the same as that of the 1970 cohort when entering to the labor force. Large cohorts depress their own wages relative to those of other cohorts in the labor force at the same time.

While these studies all refer to the US case, they illustrate the power of the supply-demand framework and the richness of combining age and schooling as basic labor inputs driving wage variations.

3. Methods

In Brazil, fertility decline was first apparent in the metropolitan cities of Rio de Janeiro, São Paulo, and Porto Alegre in the early 1960s, which had total fertility rates below 5 at the time. From there the decline spread to the interior of the states of the Southeast, and to the capital cities of states in the Central-West, North and Northeast, finally reaching the interior and rural areas of those regions in the 1980s. At the municipal level in the year 2000, there were a substantial number of entities with total fertility rates above 4, while there were also many in which fertility had fallen below the replacement level. The difference in the timing and speed of the fertility transition led to substantial differences in age distribution across states and municipalities at different points in time (Potter, Schmertmann and Cavenaghi 2002).

Since there is a very pronounced trend in the age distribution that has substantial variation across regions, states and municipalities, this paper seeks to take advantage of this variation at the local level. The data available for Brazil permit an analysis of the phenomenon at the municipality or county (*município*), microregion (agglomeration of municipalities), or state level.

Using a smaller unit of analysis such as the county or state, of course, poses the question of internal migration, which has not been incorporated in most previous analyses undertaken at the national level. The migration component could be important factor in this context since the main population streams have been moving from areas with higher fertility to those with lower fertility. While we fully intend to incorporate migration in future analyses, we do not include it in the models presented below.

3.1. Data

The longest series of data on age, education and earnings available to researchers come from the Brazilian censuses taken in 1960, 1970, 1980, 1991, and 2000. Microdata from these censuses are available from long-form questionnaires administered to every fourth household

(25% samples) in 1960, 1970 and 1980, and to every fifth (20%) or every tenth (10%) household in 1991 and 2000 depending on the size of the municipality. In all cases, there are records for every individual in the selected households that contain information on that person's age, sex, marital status, educational attainment, enrollment in school, and, if employed, occupation and earnings. There are also questions on migration, including state of birth, previous residence, and residence five years before the census.

The lowest level of geographic identifier on these records is the *município*. In previous work, Potter, Schmertmann and Cavenaghi (2002) have established minimum comparable areas that account for the changing definitions and division of municipalities through the years, as the absolute number of municipalities has increased from approximately 2,300 in 1960 to 5,280 in 2000. They have also found it convenient to aggregate municipalities to microregions, yielding 502 comparable areas across the five censuses. Thus, it is possible to calculate various parameters of the age distribution as well as the labor force outcomes, education indicators, and migration rates of these 502 areas at each of the five points in time.

3.2. Creating aggregate-level data

More specifically, for each microregion there are 48 observations, since age was categorized into four groups, education into three groups, and four different census years were used (1970, 1980, 1991, and 2000), as we will explain below. A new age-education variable with twelve categories was generated. For each microregion, age-education, and year cells, the mean income was calculated, correcting for currency changes and inflation. The proportion of males age 15-64 in each age-education group was also calculated by microregion in each census year. .

The decision to generate age and education groups in the analysis of earnings was based on previous labor market studies. Welch (1979) found that workers in adjacent experience cells are more likely to influence each other's labor market opportunities (within educational attainment) than workers in different experience groups (between educational attainment groups).

Taking into account Welch's findings, Borjas (2003) and Triest, Sapozhnikov and Sass (2006) classified information on education attainment in different groups for the estimation of labor outcomes models. Borjas (2003) analyses the impact of immigrant share on labor market outcomes in the United States by different education groups. He estimated separate models for four education groups: high-school dropouts, high-school graduates, persons who have some college (between thirteen and fifteen years of schooling), and college graduates or more. Using the March Current Population Survey (CPS) from 1964 and 2004, Triest, Sapozhnikov and Sass (2006) estimate the impact of labor market experience and relative cohort size on level of real wage by five different levels of education: (1) less than high school (high school dropouts), (2) high school graduates, (3) individuals with some college, (4) college graduates, (5) individuals with post-college education (graduate education).

Borjas (2003) and Triest, Sapozhnikov and Sass (2006) created variables for labor market experience based on information on age and educational attainment of workers. Borjas "assume(d) that the age of entry into the labor market is 17 for the typical high school dropout, 19 for the typical high school graduate, 21 for the typical person with some college, and 23 for the typical college graduate. Let A_T be the assumed entry age for workers in a particular schooling group. The measure of work experience is then given by (Age – A_T). I restrict the analysis to persons who have between 1 and 40 years of experience" (Borjas 2003, p.1341). Triest, Sapozhnikov and Sass (2006) also constructed different groups of work experience based on age and educational attainment: (Age – 17) for high school dropouts; (Age – 18) for high school graduates; (Age – 20) for people with some college; (Age – 22) for college graduates; and (Age – 24) for people with graduate education. However, instead of creating a variable for work experience, we use indicators for different age-education groups as independent variables. Age is categorized in four groups: (1) youth (15-24) , young adults (25-34), experienced adults (35-49), and older adults (50-64).

Our educational attainment classification was based on findings suggested by Riani (2005). She noted that although by the year 2000, the majority of Brazilians between ages 7 and are in school, and large fractions were completing elementary school (between one and eight years of schooling). Moreover, she indicates that there was a decrease in regional, race and rural-urban differentials in elementary school attainment. On the other hand, although more people are attending secondary school, the proportion of people with between nine and twelve years of schooling is still low, and regional differences are still significant. Riani emphasizes that elementary education is spreading to the whole population. However, since completion of high school is still low, the differentials tend to increase because people with better socioeconomic status are the first ones to obtain more education. Taking into account the specifics of the Brazilian population, education attainment will be classified in three main groups: zero to four years of schooling; five to eight years of schooling; and at least nine years of schooling. The first group includes illiterate people and those in the first phase of elementary school (one to four years of schooling). The second group contains people in the second phase of elementary school. The third and final group is comprised of people in high school (nine to eleven years of schooling), and people with some college education (at least twelve years of schooling).

The dependent variable is the logarithm of the mean real income in a group defined by microregion, age-education, and year. Since there were changes in Brazilian currency across time, as well as dramatic inflation during the period, the nominal wage was converted to base one in January 2002. To correct for currency changes, wages in 1970 and 1980 were divided by 2,750,000,000,000; and in 1991, they were divided by 2,750,000. To correct for inflation, wages were divided by 0.000000000000264 in 1970; 0.000000000005778 in 1980; 0.000067602304350 in 1991; and 0.902716061809642 in 2000, as suggested by Corseuil and Foguel (2002).

Time refers to four different censuses used in the analysis: 1970, 1980, 1991, and 2000. The 1960 census was not included in this analysis because information for earnings is categorized in the microdata, and not continuous as in the other censuses. We have not yet finished developing a procedure to "smooth" these categorical frequencies so as to be able to calculate mean earnings for each microregion-age-education group in 1960.

This analysis was done using models in which there was a fixed-effect for each microregion. These areas were first homogenized by Potter, Schmertmann and Cavenaghi (2002) in 518 areas, in order to have comparable areas across the 1960, 1970, 1980, and 1991 censuses. To incorporate the 2000 census, microregions were redefined by the same authors into 502 comparable areas across the country.

The appropriate weights to use in the regression would be the square-root of the number of people (observations) in each microregion, age, education, and year cell. However, the results that will be shown in this paper were not generated using weights because STATA does not allow the use of weights in fixed effects models. We are considering ways to work around this restriction, but are still in the process of devising a solution. Finally, cells with fewer than 25 people receiving income in a specific microregion, age-education, and year group were not included in the regression.

3.3. Estimation of models

After assembling the aggregate data by microregion, age-education, and year cells, as well as getting information on mean income and number of people in each cell, and proportion of people in each age-education group by year and microregion, fixed-effects models were generated using the following formulations. Let W be the logarithm of wages, the dependent variable (and it could be any other outcome); let X be an independent variable or vector of independent variables. Let i denote an area (microregion), t denote time (census year), and c denote a cell (in our case, age-education group). Duplicating Triest, Sapozhnikov and Sass (2006) would involve estimating:

(1) $W_{itc} = \beta_0 + \beta_1 X_{itc} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1 \ldots K; t = 1 \ldots T$,

where each observation is a time period, an area and a demographic cell, υ_i is a vector of area fixed effects, and θ_t is a vector of dummies for each year (time fixed effects). This formulation implies that there are unobservable time and area effects, but that the substitution parameter (and this is essentially an elasticity of factor prices) is identical for all cells. Note that the variation in this model arises solely from variation among cells within an area over time.

Equation (1) generates twelve different regressions, one for each combination of the age and education groups cited above. The pooled model in (1) can be estimated in one single regression, including twelve proportions of people in each one of the age-education groups, eleven dummies for age-education groups, and three dummies for census years. The reference groups are people between 15 and 24 years of age, and with zero to four years of schooling, observed in the 1970 census (see Table 3).

An approach that would be the same econometrically but would allow for cross-effects, and thus accord more closely with theory by explicitly allowing labor-labor substitution (see Hamermesh, 1993), is:

(2) $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_2 X_{itc'} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1 \ldots K; t = 1 \ldots T$,

where c' refers to the other combinations of age-education. This formulation allows for substitution parameters that indicate how a change in the fraction of the population in one cell alters the wage of people in another cell. There will be 10 terms in $X_{itc'}$ in Equation (2), since presumably the X variable measures the fraction of the work force in area i and time t that is in the cell, so that the X's sum to one and only 10 are independent.

Equation (2) generates twelve different regressions. The pooled model of Equation (2) includes in one single regression all cross-proportions of people for each one of the twelve age-education groups from each one of the other eleven age-education groups, eleven dummies for age-education groups, and three time dummies.

In a first step, one could estimate both Equations (1) and (2) more simply by dropping the υ_i and θ_t terms and ignoring area and time fixed effects. Thus there would be four formulations to estimate before going into still weaker assumptions. The first of those assumptions is that the production parameters vary over time. In the case of the model in (1) this leads to:

(1') $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_3 \theta_t X_{itc} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1 \ldots K; t = 1 \ldots T$.

This equation interacts the X variable(s) with T-1 time dummies. It allows testing for the constancy of the effects over time. Equation (1') nests Equation (1). The pooled model of Equation (1') can be estimated in one single regression (see Table 4).

We can estimate a model analogous to Equation (2) that allows all the substitution parameters to vary over time by adding to (2) both the same own-quantity interactions with time dummies that were added to Equation (1) to create Equation (1') and also interactions of the cross-quantity terms with the set of T-1 time dummies:

(2') $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_2 X_{itc'} + \beta_3 \theta_t X_{itc} + \beta_4 \theta_t X_{itc'} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1 \ldots K; t = 1 \ldots T$.

Note that Equation (2') nests each of Equations (1'), (2) and (1). The pooled form of (2') can be generated in one single regression, including all cross-proportions of people for each one of the twelve age-education groups.

Less general formulations of Equations (1') and (2') would simply take a continuous time indicator, TIME, going from 1 to T, and interact it instead of each of the dummies θ_t with the X_{itc} in Equation (1) and with X_{itc} and $X_{itc'}$ in Equation (2). Those formulations implicitly allow for linear trends in the production parameters.

The most general formulations that make sense would take Equations (1') and (2') and allow for the possibility of area-specific trends in the production parameters. Thus one might generalize still further and estimate:

(1") $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_3 TIME_t X_{itc} + \beta_5 \upsilon_i TIME_t X_{itc} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1 \ldots K; t = 1 \ldots T$.

Here one might even like to allow for time-specific and area-specific production parameters by interacting the X_{itc} with the vector θ_t instead of with the continuous variable TIME in Equation (1"). That could be done; but such an extensive formulation means estimating separate production parameters for each area in each time period, something that is not likely to be very productive.

As it is, the formulation in Equation (1”) implies that there are separate production parameters for each area, but that in each area the production parameter is characterized by a linear trend. Of course, one would use this and (1’) to test for the significance of the β_5, the area-specific trends in the production parameters.

Finally, one can take the same tack with the more general versions in Equations (2) and (2’) and estimate:

(2”) $W_{itc} = \beta_0 + \beta_1 X_{itc} + \beta_2 X_{itc'} + \beta_3 TIME_t X_{itc} + \beta_4 TIME_t X_{itc'} + \beta_5 \upsilon_i TIME_t X_{itc} +$

$\beta_6 \upsilon_i TIME_t X_{itc'} + \upsilon_i + \theta_t + \varepsilon_{itc}$, $i = 1\ldots K; t = 1\ldots T$.

This formulation allows for time trends in area-specific production parameters describing both own- and cross-substitution effects. Throughout all of this the trick is to move from the simplest formulation, Equation (1) without area or time fixed effects, to the increasingly general formulations. In each case one would test the validity of the additional specifications using the appropriate F-statistics.

4. Results

4.1. Descriptive analysis

As discussed in previous sections, the age distribution of the population of Brazil has been changing. Figure 1, based on UN projections, shows that the child dependency ratio will continue to decrease significantly in the next decades. Moreover, the old-age dependency ratio has been increasing since 2000, and is going to increase even more in coming years. These patterns are related to the decline in the total fertility rate since the 1960s (Table 1). Since fertility declined so abruptly, the proportional share of younger groups also declined.

<<< Figure 1 and Table 1 >>>

However, the fertility decline began in different parts of the country at different times. The difference in the timing and speed of the fertility transition led to substantial differences in the age distribution across regions, states and municipalities at different points in time. On one hand, Figure 2 illustrates the percent of young adults (25-34 years of age) with at least nine years of schooling in all 502 Brazilian microregions for 1970-2000 censuses. There is a clear increase over time in the proportion of young adults with high educational attainment . At the same time, differences among microregions are pronounced and persistent. Higher proportions of this age-education group are observed in Southeastern (SE), Southern (SO), and Center-Western areas (CW), comparing to Northern (NO) and Northeastern (NE) areas. On the other hand, Figure 3 shows that the percent of adults (35-49) with low levels of education (zero to four years of schooling) has been decreasing through time for all microregions. However, areas in the Southeast and South of Brazil indicate a greater decrease in the proportion of men in low-educated group, compared to the North and Northeast of the country.

<<< **Figures 2 and 3** >>>

In Figure 4, age distributions for four selected microregions are shown for 1970 and 2000. (Data are shown only for these two points in time to allow a clearer picture of the changes.) The curves for the Northeastern microregions (in Piauí, and Ceará) indicate that the age distributions in 1970 and 2000 were similar. These microregions do not have significantly different patterns between 1970 and 2000, unlike the Southeastern (in Rio de Janeiro) and the Southern (in Rio Grande do Sul) microregions. In the Southeast and South of Brazil, there were significant changes in age distribution, with a growth in the proportion in older ages from 1970 to 2000.

<<< **Figure 4** >>>

Figure 5 illustrates the distribution of the male population by education for 1970 and 2000 for the same microregions shown in Figure 4. In general, there was a growth in the percentage of people with higher levels of schooling across the years. Furthermore, education curves indicate that Northeastern microregions have lower levels of education than the microregions in the South and Southeast.

<<< **Figure 5** >>>

The changes in education distribution were substantial between the 1970 and 2000 Brazilian censuses in all regions. Changes in age distribution were also observed in the same period for at least areas in the Southeast and South of the country. Moreover, differences among regions suggest the need to use models that take into account the specificities of these localities, through the use of fixed effects for microregions, and to focus the analysis on changes through time.

4.2. Influence of changing age distribution on labor market outcomes

Since there are 502 microregions, 12 age-education groups, and four censuses, the maximum number of possible observations in the regressions would be 24,096. However, the minimum cell-size requirement changes the maximum number of possible observations dropped to 19,704.

Table 2 shows the percent of male population by year and age-education groups in Brazil. In general, the numbers indicate that the proportion of people with zero to four years of schooling fell from 1960 to 2000. For example, the proportion of people between 15-24 years of age, and 0-4 years of schooling dropped considerably from 30.84% in 1960 to 9.04% in 2000, a decrease of more than three times. Moreover, proportions of people with five to eight years of schooling, as well as those with at least nine years of schooling increased during the period. The highest increases in proportion of people with at least nine years of education were the ones for individuals with 15-24 years, from 1.08% in 1960 to 10.24% in 2000, an increase of almost 9.5 times; and for those with 35-49 years of age, from 0.91% in 1960 to 8.46% in 2000, an increase of more than nine times in forty years.

<<< Table 2 >>>

Table 3 illustrates estimates of the pooled form of Equation 1. The indicator variables for age-education groups show that within each age category the incomes are higher for those people with more schooling. For instance, people between 25 and 34 years of age and with zero to four years of schooling earn 1.52 more times [exp(0.42)] than people between 15 and 24 years of age with same education (reference category). At the same time, young adults (25-34) with at least nine years of schooling earn 6.05 more times [exp(1.80)] than the reference group.

<<< Table 3 >>>

The coefficients of the proportions of people in age-education groups indicate that greater negative impacts on income exist for people with more years of education. In order to interpret these coefficients, it is necessary to calculate the elasticity for each one of them, because the proportions of people across age-education groups vary over time. For instance, the negative impact of –16.23 for the oldest age group (50-64) with five to eight years of schooling is not greater than the impact of –3.32 for the youngest age group (15-19) with same education, if we consider the mean national age-education distribution in each one of the censuses. This is clarified by the columns of elasticity in Table 3, which were estimated taking into account the age-education distribution by year from Table 2. These are estimates of the time-varying elasticities of complementarity.

The estimated elasticities show that negative impacts of proportions of males by age-education groups are greater in those groups with higher education (five to eight years of schooling, and at least nine years of schooling). Moreover, negative impacts increase over time, since Brazil has been experiencing an increase in the proportion of people in groups of higher education. On the other hand, coefficients for groups with lower education show a decrease of the negative impact on income over time, due to the fact that these groups have been experiencing a decrease in their proportional share in the whole population.

As can be seen in Table 3, an increase of ten percent in the number of people with five to eight years of schooling, between 15 and 24 years of age, reduces their income by 1.8 percent (-0.179) in 1970, and 4.1 percent in 2000. For young adults (25-34) in the same education group, the impact also increases over time, from an income reduction of 1.2 percent in 1970 to 4.6 percent in 2000 with an increase of ten percent in the number of people in this group. The same happens for adults (35-49) and older adults (50-64) with five to eight years of schooling. Coefficients in the highest education groups also show a significant negative impact on income. However, older adults (50-64) with at least nine years of schooling, and even adults (35-49) in this education group, have a smaller negative impact on income. This result is partly due to the fact that they represent smaller proportional shares in the population compared to younger groups, generating lower elasticities. In summary, the increase in the proportion of people with higher years of schooling generates higher negative impacts on their income compared with groups with lower education, and this impact has been increasing over time.

These estimates permit the comparison of the predicted mean monthly real income among different age-education groups over the range of the actual proportions of people in these groups in Brazilian microregions. Figure 6 illustrates predicted income for young adults (25-34) with at least nine years of schooling, and adults (35-49) with zero to four years of schooling. The intention is to verify the pattern of income in different years between young adults with higher education, and adults with low education. Usually, older people are the ones receiving higher incomes, as can be seen with the dummies for age-education groups for adults (35-49) and mature adults (50-64) in Table 3. However, Figure 6 shows that when those adults have little education, their income is very low, even compared with young adults. Further, these graphs illustrate lower outcomes for individuals who live in microregions with higher proportions of people in their age-education group. In this case, the negative impact of a higher proportion of people in one's own group on earnings is greater for young adults, as can be seen by their steeper curves compared to the flatter ones for adults with lower education.

<<< **Figure 6** >>>

Table 4 shows regression results that allow the own-quantity effects (proportion of people in age-education groups) to vary over time directly rather than because the underlying stocks of workers change (pooled form of Equation 1'). Interactions with year indicate that the negative impact of proportions of people in the microregion have been decreasing across the years. This is observed mainly for 1991 and 2000, in which positive coefficients offset the negative impacts of proportion of people in 1970 (reference category). Those coefficients have the highest values for the last age-education group (50-64 years, 9+ years of education) with estimates equal to 18.07 in 1991, and 19.41 in 2000. Other age-education groups also present significant positive coefficients principally for the last two years analyzed.

<<< **Table 4** >>>

We also estimated the pooled form of Equation 2 in order to predict earnings that take into account cross-effects. The results are not presented in a table format, because the number of coefficients (147 in total) makes it hard to interpret the results. A way to analyze the results is through the ratios presented in Figure 7 for males with 35-49 years of age, and between five and eight years of schooling from 1970 to 2000. For each one of the four selected microregions (the same as those in Figures 4 and 5), there is a ratio between the predicted mean income of the own-effect model (pooled form of Equation 1) and the predicted mean income from a regression equation with only the age-education group and year dummies, that is to say with no age-education group proportions included (See Table 5). The intention is to show the difference in estimated earnings between a model that takes into account age and education structure (proportions of people by age-education groups) and the usual model of labor demand, which considers only the direct impact of age on per capita income (through the age-education group dummies). A second set of ratios in Figure 7 compares the predicted income from the model with cross-effects (pooled form of Equation 2) and the usual model of labor demand. The horizontal line shows the baseline predicted values as unitary, and should be compared to the other curves which show the ratio of the predicted income from the comparison model to those of the baseline model. The dashed line indicates the ratio between observed and predicted income by the model with only age-education indicators and provides an indication of how well the predicted points fit the data in the particular case of the selected area.

The curves in Figure 7 indicate that the slope of predicted earnings from the own-effects and cross-effects models accord fairly well with the slope actually found in the data, all in relation to predicted earnings based on a model without group size effects. Note, however, that this slope is less in the two Northeastern areas than it is in the two areas from the South and Southeast, due to the greater shift in the proportions in this group found in those in those areas. The proportion of adults (35-49) between five and eight years of schooling increased from 4.9 percent in 1970 to 8.9 percent in 2000 in the Southern microregion. The selected Southeastern microregion had an even greater increase in this group proportion, going from 2.8 percent in 1970 to 10.2 percent in 2000. However, the Northeastern areas did not present as much improvement, changing from 0.5 percent (1970) to 2.8 percent (2000) in the microregion in Piauí, and from 0.7 percent (1970) to 4.3 percent (2000) in the selected area in Ceará (these proportions are not shown in the table).

Finally, as Figure 7 indicates, the observed variation through time has a pattern more similar to the predicted income from the cross-effects model than from the own-effects model. This finding is clearest for the microregions in the Southeastern and Southern regions, and, of course, these are just four selected areas from the 502 in the data set.

<<< **Figure 7** >>>

Another way to view the results from the own-effects model is to report how the changing national distribution of males in age-education groups from 1970 to 2000 affects the predicted group earnings (Figure 8). In order to accomplish this exercise, the national proportions of males by age-education groups and census years presented in Table 2 were used to calculate predicted earnings, applying the coefficients from Table 3. The figure shows that groups presenting a decline in their proportion over time will experience gains in their income, and vice-versa. As in Figure 7, the ratio of the predicted earnings from the own-effects model to the predicted income from the classic labor market model is plotted in the graph, as well as the flat baseline. Figure 8 illustrates the curves for all three education groups, but only for adult males age 35-49. Comparing the curves to the baseline, one can see that the low education group (zero to four years of schooling) has predicted earnings from the own-effects models increasing over time. This pattern results from this group decreasing from 22.66 percent in 1970 to 13.32 percent in 2000 (Table 2) as a proportion of the total male labor force. On the other hand, the two education groups with increasing shares (5-8, and 9+ years of schooling) present an own effect that affects income negatively in comparison with the baseline.

<<< **Figure 8** >>>

5. Final Remarks

In this paper, we have tackled an old question, but in a very different context, and with a new way of extracting lessons from the data. Interesting and important results concerning the effects of shifts in the age distribution of the working age population have been obtained by a series of authors by looking at this question in relation to the effect of the baby-boom generation on the earnings of different cohorts in the U.S. But the question has received little attention in the context of the countries of Asia and Latin America, which are now experiencing substantial shifts in their age distributions due to large and rapid declines in fertility. In these countries, these shifts in the age distribution have also been accompanied by fairly dramatic increases in educational attainment.

One important difference between the U.S. case and Latin American countries such as Brazil concerns the magnitude of regional differences in the timing of both the educational and demographic transitions. These changes were fairly homogeneous across the U.S., but varied enormously geographically in Brazil. It is this heterogeneity that both motivates and enables the regional approach to the problem that we have undertaken in this analysis.

The first and most important result from our models is that relative group size matters. The coefficients of the proportions of people in age-education groups tend to have a negative impact on income, with the greatest negative impacts on income occurring for groups with more years of education. This is completely in line with what one would expect from theory. Age-education groups are not perfect substitutes, so that own cohort-education size depresses earnings. That the effects increase with education is consistent with the observation (Hamermesh, 1993, Chapter 3) of lower own-wage elasticities as education increases. While there may have been demand shifts over the thirty year time span, they do not appear to have been large enough to compensate for the supply variation, leading to the negative coefficients in our own-group effects model.

The effects and magnitude of biased technological change and/or institutional changes are suggested by the positive interaction terms in Table 4. Although the interaction terms were positive, just a few of them were strong enough to exceed the negative effects obtained for the baseline period (1970). By 2000, however, the interaction term for the unskilled labor (0-4 years of schooling) was positive and stronger than the negative effect in the baseline for all but one age group. This could be indicating either the operation of a somewhat surprising unskilled bias in technological change or the operation of institutional shifts, perhaps in minimum wage laws, over the period. Both possibilities offer fruitful areas for future research.

Table 4 also showed that demand and institutional shifts were insufficient to compensate for the downward pressure of supply shifts in the case of middle and high levels of schooling, a result that is valid for all age groups. The only exception is the case of high schooling in the oldest age group, where we find a case of positive elasticity—the more people with higher education, the higher their earnings.

Both demographic and educational shifts are likely to have important redistributive effects for those found in groups that are either growing or shrinking rapidly, or whose "neighbors" (in terms of both age and education) are either shrinking or growing. And, as can be seen, some very large shifts have taken place in Brazil over the last four decades, and will continue into the foreseeable future.

The results presented here are preliminary, and we still have a lot of work to do in deciding which of the many possible models best fits, and captures the most important features of the data. We also have yet to fully come to terms with the ways that migration between areas or changes in female labor force participation might be influencing the results. Nevertheless, we believe the results are encouraging, and indicate that there is considerable benefit to be derived from analyzing the influence of change in age-education distributions for "local" labor markets in developing country contexts such as Brazil.

With respect to the subject of this PAA session, "The Demographic Dividend", it has been clear from the outset that we have been addressing a related but different question. The main focus in the dividend literature has been on the dependency ratio--the ratio of those of labor force age to those both younger and older. In countries such as Brazil, this ratio is undergoing dramatic change and will continue to do so in coming years, no doubt with many important consequences. It is also the case that the composition of the Brazilian labor force, in terms of both age and educational attainment, is undergoing dramatic shifts. What we have tried to investigate here is whether those compositional shifts will have an effect beyond those normally analyzed in the Mincer earnings equation, so that we need to study their role in the context of a formal theory of labor demand. The first indications are that these compositional shifts are, indeed, influential, and that this approach represents a fruitful way of studying this central problem in economic development.

Acknowledgements

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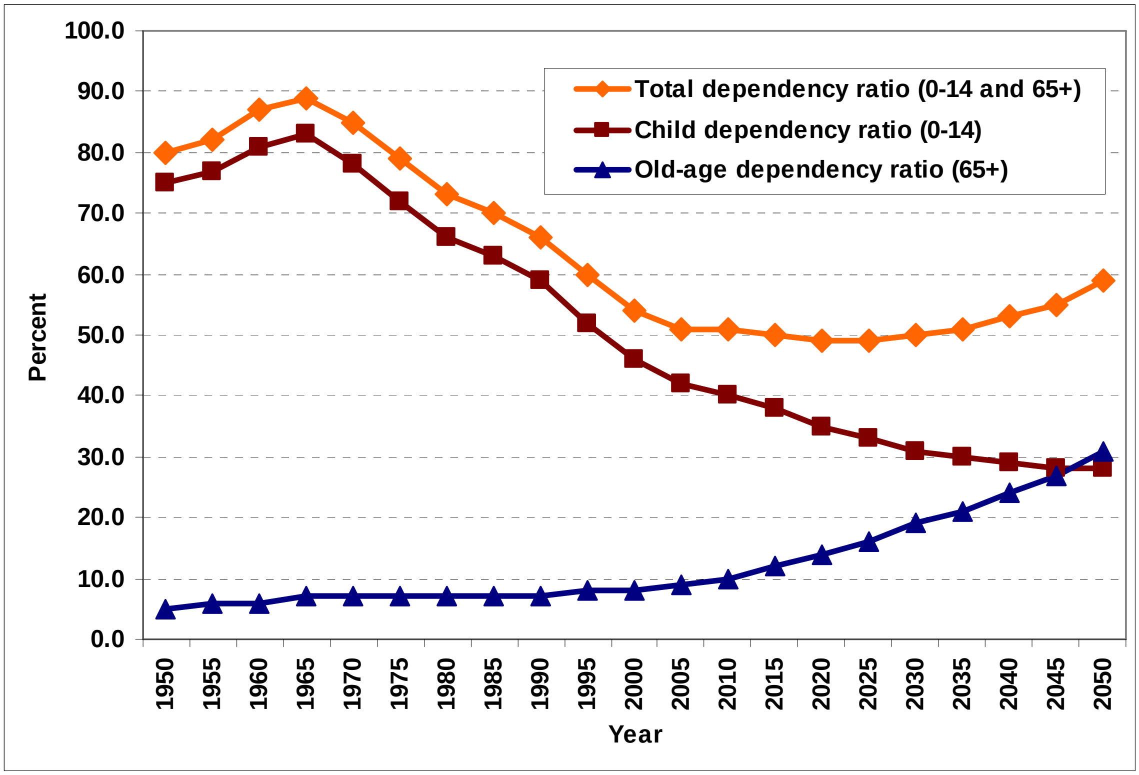
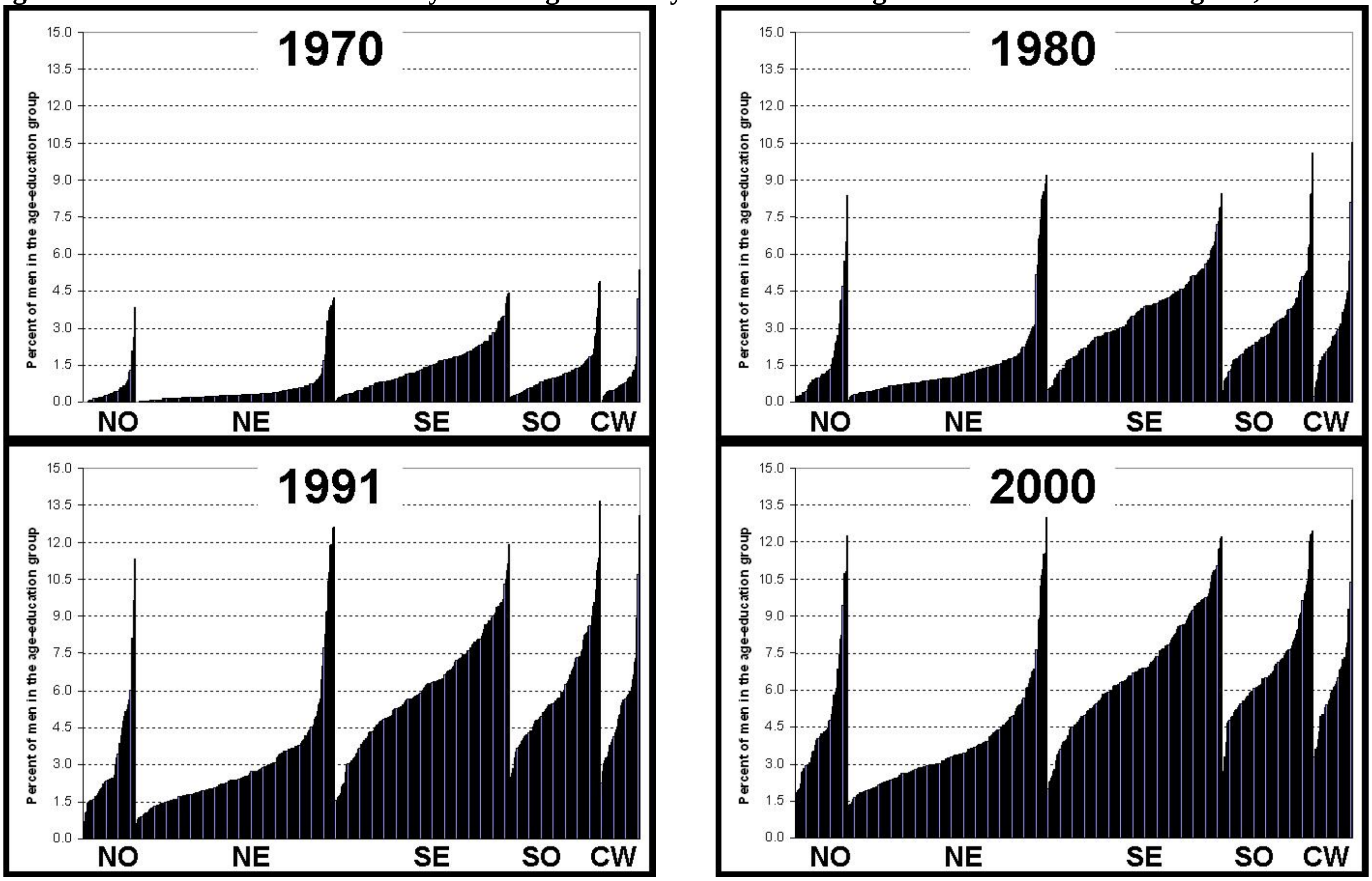
Figure 1. Dependency ratios in Brazil, 1950-2050.



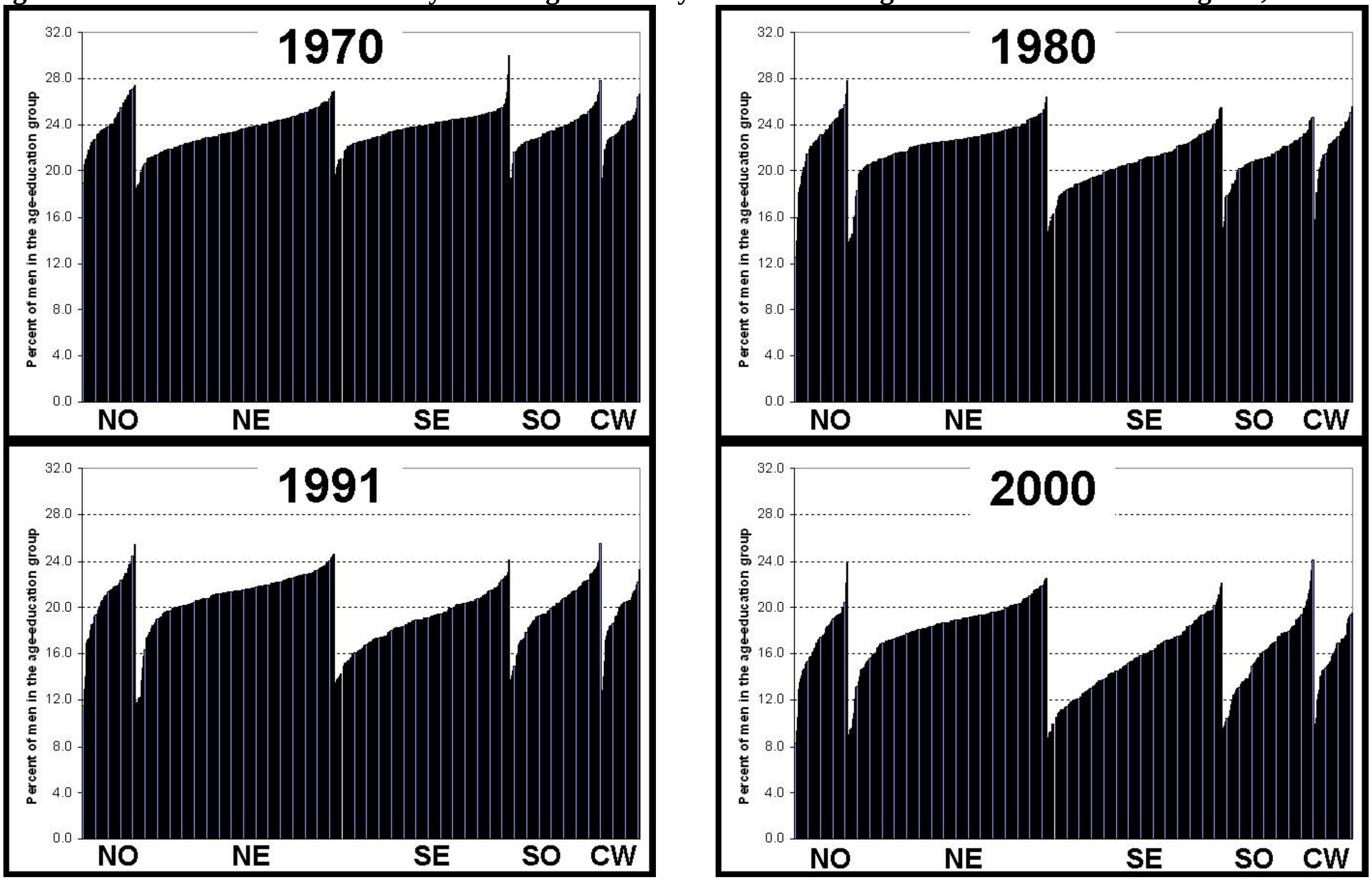
Source: United Nations - http://esa.un.org/unpp (in August 16, 2006 - medium variant).

Figure 2. Percent of men with 25-34 years of age and 9+ years of schooling in 502 Brazilian microregions, 1970-2000.



Source: 1970-2000 Brazilian Censuses.

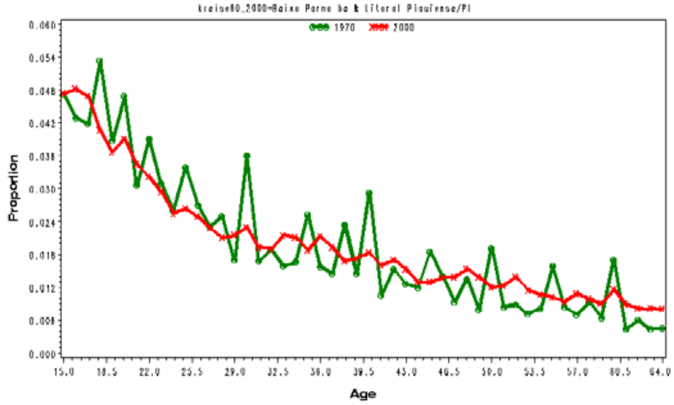
Figure 3. Percent of men with 35-49 years of age and 0-4 years of schooling in 502 Brazilian microregions, 1970-2000



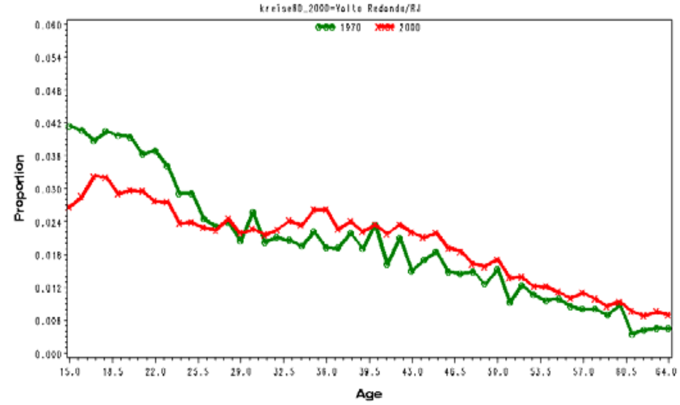
Source: 1970-2000 Brazilian Censuses.

Figure 4. Changes in the age distribution in selected Brazilian microregions, 1970 and 2000.

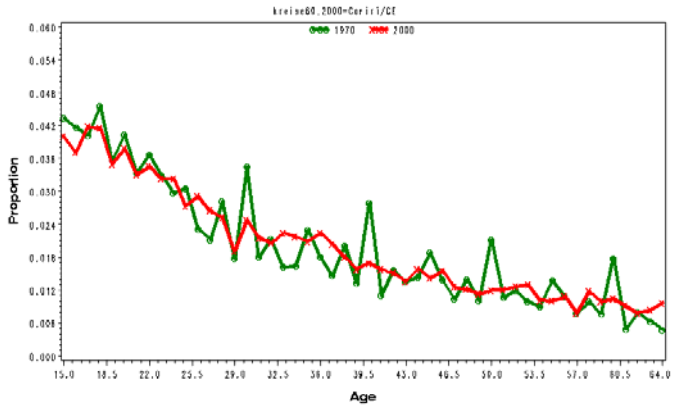
Baixo Parnaíba & Litoral Piauiense - Piauí - Northeast



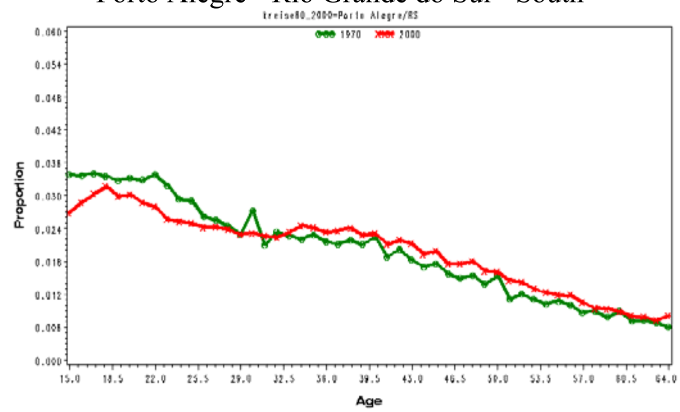
Volta Redonda - Rio de Janeiro - Southeast



Cariri - Ceará - Northeast

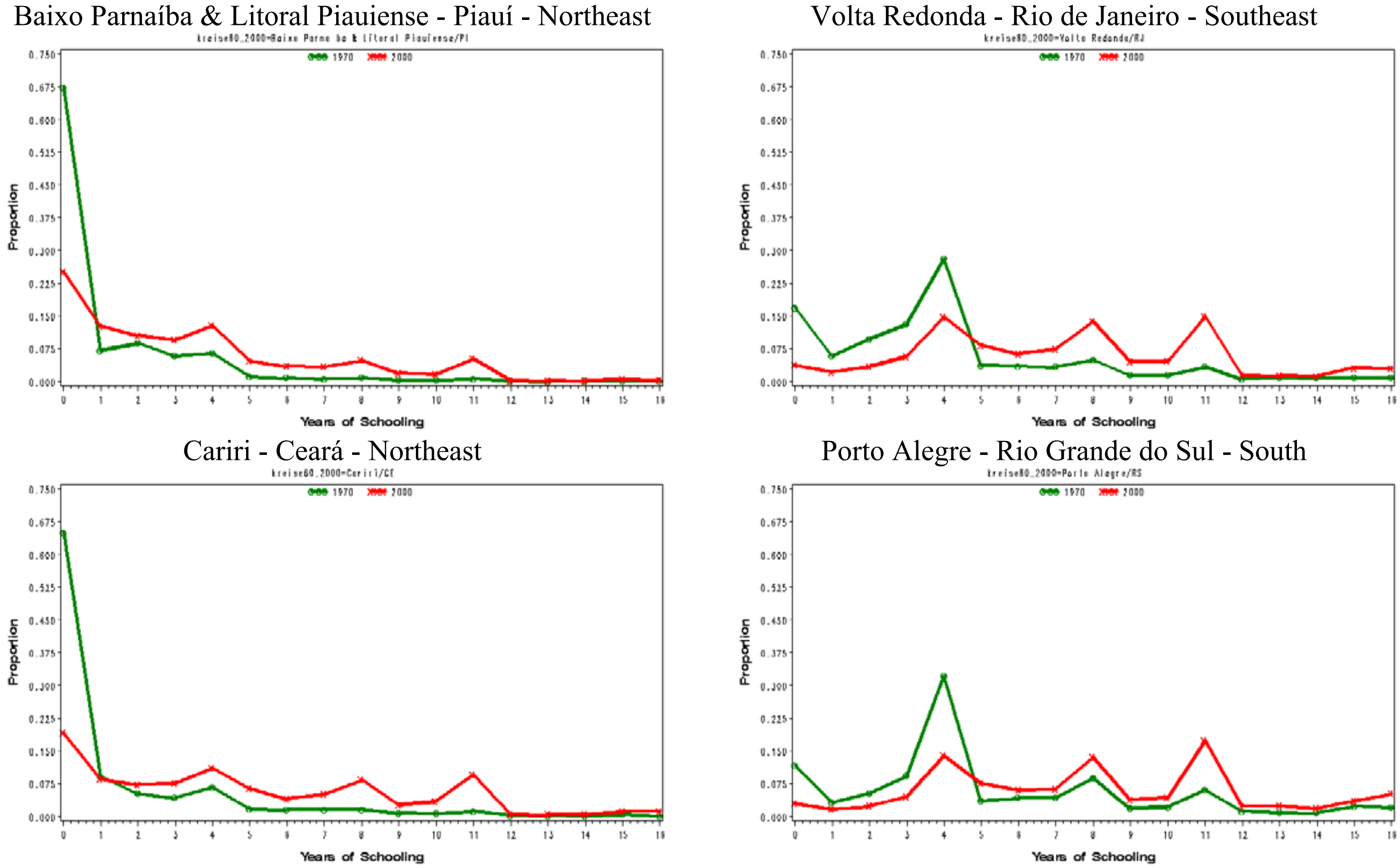


Porto Alegre - Rio Grande do Sul - South



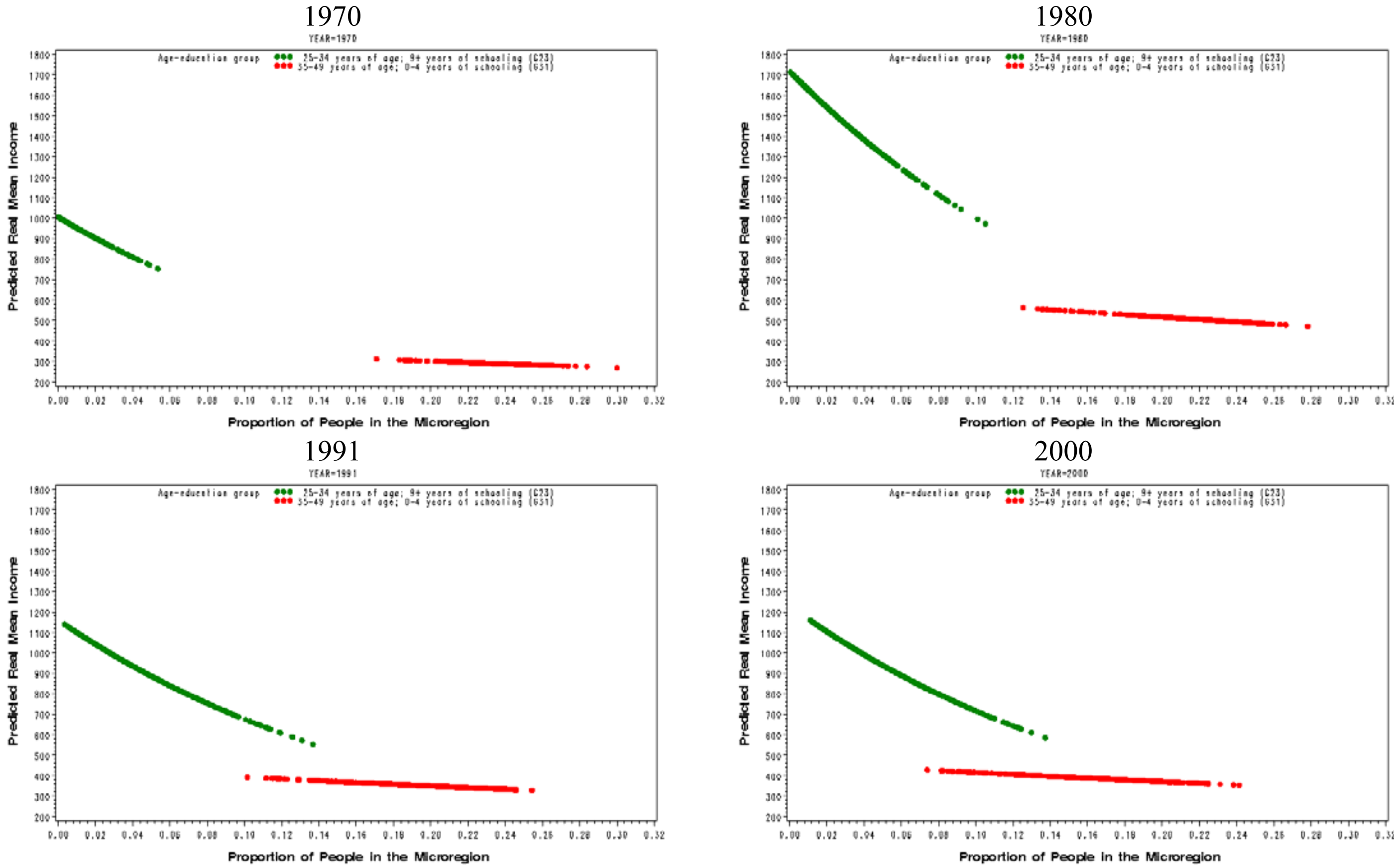
Source: 1970 and 2000 Brazilian Censuses.

Figure 5. Changes in the education distribution in selected Brazilian microregions, 1970 and 2000.



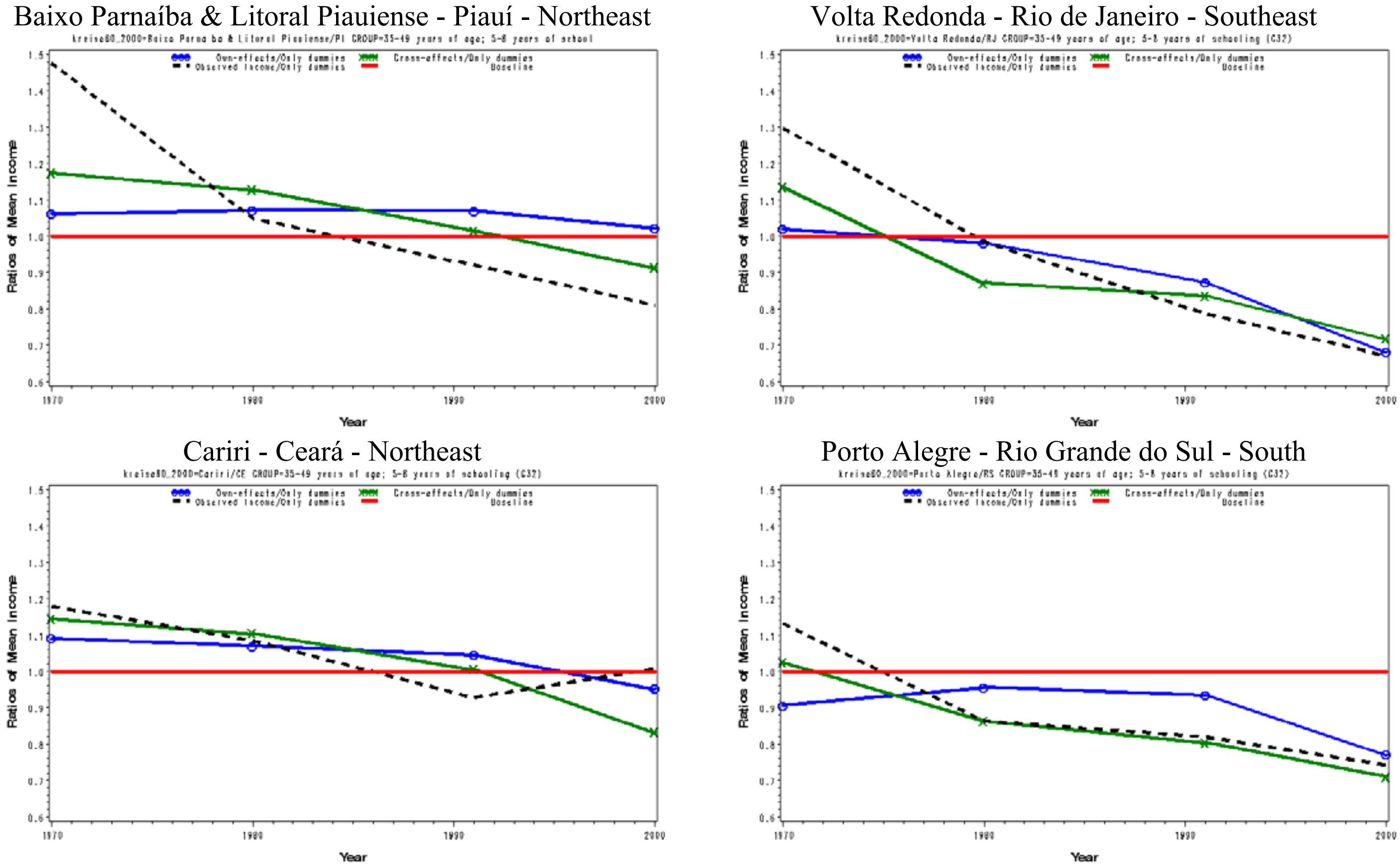
Source: 1970 and 2000 Brazilian Censuses.

Figure 6. Predicted real mean monthly income[+] by proportion of people in Brazilian microregions, estimated from regression results shown in Table 3[++], for young adults (25-34) with at least nine years of schooling (9+), and adults (35-49) with zero to four years of schooling, 1970-2000.



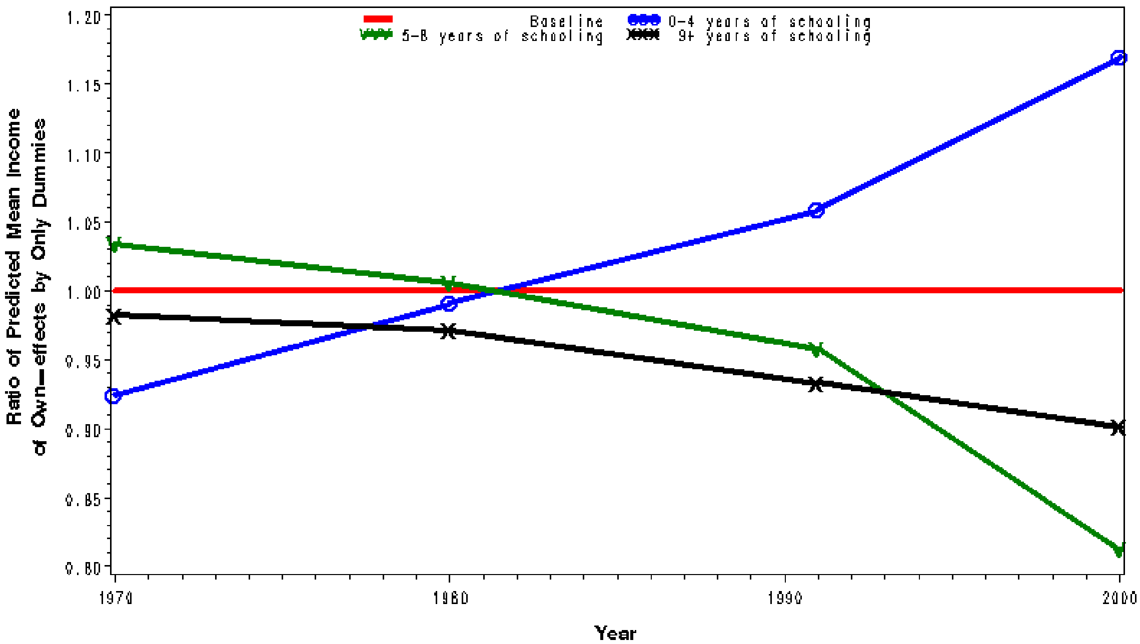
[+] Nominal income was converted to base 1 in January 2002, taking into account changes in currency, and inflation.
[++] This model is the pooled form of Equation 1.
Source: 1970-2000 Brazilian Censuses.

Figure 7. Ratios of Predicted Mean Income of Selected Regressions and Observed Income by Predicted Mean Income of Regression Only with Dummies for Males with 35-49 Years of Age and 5-8 Years of Schooling by Year, 1970-2000.



Source: 1970-2000 Brazilian Censuses.

Figure 8. Ratio of Predicted Mean Income of Own-effects Model (Table 3) by Predicted Mean Income of Regression Only with Dummies using National Age-Education Distribution for Males with 35-49 Years of Age by Year and Education Group, 1970-2000.



Source: 1970-2000 Brazilian Censuses.

Table 1. Total fertility rate, infant mortality rate, and life expectancy at birth in Brazil, 1950-2050.

Period	Total fertility rate	Infant mortality rate (per 1,000 births)	Life expectancy at birth (years)
1950-1955	6.15	134.7	50.9
1955-1960	6.15	121.9	53.3
1960-1965	6.15	109.4	55.7
1965-1970	5.38	100.1	57.6
1970-1975	4.72	90.5	59.5
1975-1980	4.31	78.8	61.5
1980-1985	3.8	63.3	63.1
1985-1990	3.1	52.4	64.9
1990-1995	2.6	42.5	66.6
1995-2000	2.45	34.1	68.8
2000-2005	2.35	27.4	70.3
2005-2010	2.25	23.6	71.9
2010-2015	2.15	20.3	72.9
2015-2020	2.06	17.1	74.2
2020-2025	1.98	14.3	75.2
2025-2030	1.92	12.1	76.2
2030-2035	1.86	10.3	77
2035-2040	1.85	9.1	77.8
2040-2045	1.85	8.1	78.5
2045-2050	1.85	7.5	79.2

Source: United Nations - http://esa.un.org/unpp (in August 16, 2006 - medium variant).

Table 2. Percent of male population by year and age-education group in Brazil, 1960-2000.

Age-education Group	1960	1970	1980	1991	2000
15-24 years 0-4 years of schooling	30.84	28.19	20.59	14.61	9.04
15-24 years 5-8 years of schooling	2.63	5.38	10.53	12.09	12.46
15-24 years 9+ years of schooling	1.08	2.74	5.87	5.97	10.24
25-34 years 0-4 years of schooling	22.66	19.71	16.39	12.41	8.82
25-34 years 5-8 years of schooling	1.18	1.98	3.90	6.82	7.63
25-34 years 9+ years of schooling	1.19	2.00	4.77	7.40	8.12
35-49 years 0-4 years of schooling	24.47	22.66	19.02	17.11	13.32
35-49 years 5-8 years of schooling	0.98	1.62	2.39	3.67	6.73
35-49 years 9+ years of schooling	0.91	1.59	2.84	5.54	8.46
50-64 years 0-4 years of schooling	13.21	12.84	11.72	11.49	10.36
50-64 years 5-8 years of schooling	0.43	0.65	0.94	1.16	1.99
50-64 years 9+ years of schooling	0.40	0.62	1.05	1.72	2.84
Total	**4,039,104**	**25,760,594**	**32,613,947**	**43,434,534**	**53,177,964**

Source: 1960-2000 Brazilian Censuses.

Table 3. Fixed effects for Brazilian microregions of year dummies, age-education group dummies, and proportions of people in age-education groups in the logarithm of the monthly real income[+], 1970-2000.[++]

Variables	Coefficients	Elasticity taking into account age-education distribution in Table 2: 1970	1980	1991	2000
Constant	5.11***				
1970	---				
1980	0.54***				
1991	0.14***				
2000	0.20***				
Dummies for age-education groups:					
15-24 years; 0-4 years of schooling (G11)	---				
15-24 years; 5-8 years of schooling (G12)	0.60***				
15-24 years; 9+ years of schooling (G13)	0.98***				
25-34 years; 0-4 years of schooling (G21)	0.42***				
25-34 years; 5-8 years of schooling (G22)	1.22***				
25-34 years; 9+ years of schooling (G23)	1.80***				
35-49 years; 0-4 years of schooling (G31)	0.83***				
35-49 years; 5-8 years of schooling (G32)	1.59***				
35-49 years; 9+ years of schooling (G33)	2.17***				
50-64 years; 0-4 years of schooling (G41)	0.82***				
50-64 years; 5-8 years of schooling (G42)	1.71***				
50-64 years; 9+ years of schooling (G43)	2.24***				
Proportions of people in age-education groups:		**1970**	**1980**	**1991**	**2000**
Proportion with 15-24 years; 0-4 years of schooling (G11)	-0.07	-0.020	-0.014	-0.010	-0.006
Proportion with 15-24 years; 5-8 years of schooling (G12)	-3.32***	-0.179	-0.350	-0.401	-0.414
Proportion with 15-24 years; 9+ years of schooling (G13)	-4.85***	-0.133	-0.285	-0.290	-0.496
Proportion with 25-34 years; 0-4 years of schooling (G21)	-0.35**	-0.069	-0.057	-0.043	-0.031
Proportion with 25-34 years; 5-8 years of schooling (G22)	-5.99***	-0.119	-0.233	-0.409	-0.457
Proportion with 25-34 years; 9+ years of schooling (G23)	-5.41***	-0.108	-0.258	-0.401	-0.439
Proportion with 35-49 years; 0-4 years of schooling (G31)	-1.15***	-0.261	-0.219	-0.197	-0.153
Proportion with 35-49 years; 5-8 years of schooling (G32)	-7.22***	-0.117	-0.172	-0.265	-0.486
Proportion with 35-49 years; 9+ years of schooling (G33)	-3.12***	-0.050	-0.089	-0.173	-0.264
Proportion with 50-64 years; 0-4 years of schooling (G41)	-1.52***	-0.195	-0.178	-0.175	-0.157
Proportion with 50-64 years; 5-8 years of schooling (G42)	-16.23***	-0.106	-0.152	-0.189	-0.324
Proportion with 50-64 years; 9+ years of schooling (G43)	-0.25	-0.002	-0.003	-0.004	-0.007
Other regression statistics:					
Number of observations	19,704				
Number of groups	502				
Sigma u	0.34				
Sigma e	0.20				
Rho	0.73				
F (26; 19,176)	8,446.56***				
F (501; 19,176)	57.32***				

[*] Significant at $p<.05$; [**] Significant at $p<.01$; [***] Significant at $p<.001$.
[+] Nominal income was converted to base 1 in January 2002, taking into account changes in currency and inflation.
[++] This model is the pooled form of Equation 1.
Source: 1970-2000 Brazilian Censuses.

Table 4. Fixed effects for Brazilian microregions of year dummies, age-education group dummies, proportions of people in age-education groups, and interactions of those proportions with year in the logarithm of the monthly real income[+], 1970-2000.[++]

Variables	Coefficients			
Constant	5.30***			
1970	---			
1980	0.44***			
1991	-0.07***			
2000	-0.05***			
Dummies for age-education groups:				
15-24 years; 0-4 years of schooling (G11)	---			
15-24 years; 5-8 years of schooling (G12)	0.52***			
15-24 years; 9+ years of schooling (G13)	0.90***			
25-34 years; 0-4 years of schooling (G21)	0.44***			
25-34 years; 5-8 years of schooling (G22)	1.12***			
25-34 years; 9+ years of schooling (G23)	1.68***			
35-49 years; 0-4 years of schooling (G31)	0.75***			
35-49 years; 5-8 years of schooling (G32)	1.51***			
35-49 years; 9+ years of schooling (G33)	2.12***			
50-64 years; 0-4 years of schooling (G41)	0.75***			
50-64 years; 5-8 years of schooling (G42)	1.61***			
50-64 years; 9+ years of schooling (G43)	2.23***			
		Interactions with year:		
Proportions of people in age-education groups:		**1980**	**1991**	**2000**
Proportion with 15-24 years; 0-4 years of schooling (G11)	-0.76***	0.34***	0.92***	1.31***
Proportion with 15-24 years; 5-8 years of schooling (G12)	-5.10***	0.72*	3.28***	3.06***
Proportion with 15-24 years; 9+ years of schooling (G13)	-4.94***	-1.09	2.17***	1.75**
Proportion with 25-34 years; 0-4 years of schooling (G21)	-1.59***	0.97***	1.34***	1.62***
Proportion with 25-34 years; 5-8 years of schooling (G22)	-6.81***	0.17	2.97**	3.19***
Proportion with 25-34 years; 9+ years of schooling (G23)	-1.58	-2.30**	-0.81	-1.85*
Proportion with 35-49 years; 0-4 years of schooling (G31)	-1.96***	1.00***	1.59***	1.65***
Proportion with 35-49 years; 5-8 years of schooling (G32)	-8.60***	0.95	2.84*	3.85**
Proportion with 35-49 years; 9+ years of schooling (G33)	-4.77***	-1.04	3.78**	3.44**
Proportion with 50-64 years; 0-4 years of schooling (G41)	-3.33***	1.67***	2.72***	3.54***
Proportion with 50-64 years; 5-8 years of schooling (G42)	-8.73**	-2.60	-0.74	1.40
Proportion with 50-64 years; 9+ years of schooling (G43)	-16.05***	1.50	18.07***	19.41***
Other regression statistics:				
Number of observations	19,704			
Number of groups	502			
Sigma u	0.33			
Sigma e	0.19			
Rho	0.74			
F (62; 19,140)	3,930.88***			
F (501; 19,140)	53.44***			

[*] Significant at p<.05; [**] Significant at p<.01; [***] Significant at p<.001.
[+] Nominal income was converted to base 1 in January 2002, taking into account changes in currency and inflation.
[++] This model is the pooled form of Equation 1'.
Source: 1970-2000 Brazilian Censuses.

Table 5. Fixed effects for Brazilian microregions of year dummies, and age-education group dummies in the logarithm of the monthly real income+, 1970-2000.++

Variables	Coefficients
Constant	5.15***
1970	---
1980	0.51***
1991	0.07***
2000	0.08***
Dummies for age-education groups:	
15-24 years; 0-4 years of schooling (G11)	---
15-24 years; 5-8 years of schooling (G12)	0.31***
15-24 years; 9+ years of schooling (G13)	0.76***
25-34 years; 0-4 years of schooling (G21)	0.38***
25-34 years; 5-8 years of schooling (G22)	0.99***
25-34 years; 9+ years of schooling (G23)	1.62***
35-49 years; 0-4 years of schooling (G31)	0.61***
35-49 years; 5-8 years of schooling (G32)	1.40***
35-49 years; 9+ years of schooling (G33)	2.10***
50-64 years; 0-4 years of schooling (G41)	0.63***
50-64 years; 5-8 years of schooling (G42)	1.58***
50-64 years; 9+ years of schooling (G43)	2.29***
Other regression statistics:	
Number of observations	19,704
Number of groups	502
Sigma u	0.31
Sigma e	0.22
Rho	0.66
F (14; 19,188)	12,866.38***
F (501; 19,188)	65.46***

* Significant at $p<.05$; ** Significant at $p<.01$; *** Significant at $p<.001$.
+ Nominal income was converted to base 1 in January 2002, taking into account changes in currency and inflation.
++ This is the classic labor market model.
Source: 1970-2000 Brazilian Censuses.