The impact of public and private transfers on rural poverty in Brazil[1]

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Abstract

In the last 25 years the Brazilian economy has undergone impressive transformations. More individuals gained access to the consumer market, the informal sector has shrunk, and the real value of minimum wage has increased. There has also been a significant decline in inequality due to government efforts to provide income to the most needed, such as the Bolsa Família program (BF), the Benefício de Prestação Continuada program (BPC), and subsidized credit to family agriculture and housing. Some recent studies suggest that public transfers and the dynamics of the job market were the leading causes of poverty and inequality decline in Brazil in the last 15 years. No study, as far as we know, take a comprehensive econometric analysis of the impact of transfers in the rural areas in contemporary Brazil. Combining different data sources to create a balanced panel of state-rural units of analysis, we estimate the impact of the major public (pensions) and private (inter-household) transfers on the dynamics of rural poverty in Brazil between 1996 and 2011. We combine data from the Brazilian National Household Survey (PNAD) and administrative data from State Statistics Bureaus, compiled and distributed by the Brazilian Institute of Geography and Statistics (IBGE), in order to estimate an Generalized Method of Moments-System (GMM-System) dynamic panel model for poverty (FGT measures). Controlling for demographic composition, GSP (Gross State Product) agricultural share, GSP share to GNP (Gross National Product), educational attainment, unemployment rate, land concentration, and age structure, we focus on how public (pensions) and private (inter-household) transfers, as well as their interaction, affected the dynamics of poverty and inequality in the rural contemporary Brazil. Our results show a strong significant impact of both transfers on poverty dynamics, with scale dominance for the retirement income. Despite controls used, poverty persistence is still significant in contemporary rural Brazil.

Keywords: Poverty, Dynamic Panel Model, Public Transfer, Private Transfer, Rural Brazil, GMM-System

1. Introduction

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In the last 20 years, the Brazilian economy has undergone impressive economic and social transformations, leading to significant reduction in poverty and inequality. Hyperinflation was eliminated, more individuals gained access to the consumer market (Rocha, 1996), the informal sector has shrunk (Corseuil et al., 2011), and the real value of the minimum wage has increased (Saboia, 2007). There has also been a significant decline in inequality due to government efforts to provide income to the most needed, such as the Bolsa Família program (BF), the Benefício de Prestação Continuada program (BPC), and subsidized credit to family agriculture and housing. The expansion of the Social Security to the rural areas, and the right of rural elderly women to access non-contributory retirement were also an important social instrument of sectorial income redistribution in the last two decades (Kreter and Bacha, 2006). Empirical evidence suggests that public transfers and the dynamics of the job market were the leading causes of poverty and inequality decline in Brazil in the last 15 years (Soares, 2006; Hoffman, 2010). Although some qualitative and local studies for rural areas of Brazil suggest that the impact of public transfers on poverty and inequality is apparent at both the household and municipality level (Albuquerque et al., 1999; Augusto and Ribeiro, 2006), the long-term dynamics for the country as a whole is largely unknown. The only exception is the recent study conduted by Marinho and Araújo (2010). Their study, however, comprises a shorter period of time (1995 to 2004), and lacks some important predictors of poverty dynamics in the econometric specification, such as the contribution of the agricultural sector to the gross national product, land concentration, and the percent contribution of the regional gross product to the gross national product. Thus, contemporary analysis of the impact of transfers on rural population is an important empirical question not fully addressed.

Rural population is an especially vulnerable group, as household members tend to be involved in vulnerable occupational activities (such as sharecropping, temporary employment, and family agriculture) and household income is usually unstable. This explains the low contributory capacity of rural areas and, consequently, the low coverage rates in most developing countries (Barrientos, 2003; Mesa-Lago, 1994). In response to that, many countries developed formal arrangements to support elderly people in rural areas (Manson and Lee, 2006; Turra and Queiroz, 2005). Although not unique among developing countries (Mesa-Lago 1994), the Brazilian rural pension system is the most comprehensive among them in regards to coverage and targeting of the poor (Afonso and Fernandes, 2005), although this is not intentional as entitlement requirements are not based on income (Schwarzer, 2000). Furthermore, the rural retirement system has been seen as instrumental in the reduction of both social unrest and opposition to the restructuring of the agricultural sector during the import substitution industrialization model in Brazil, as well as a mitigation mechanism for rural-urban migration. This institutional and political environment, fueled by the universalization principles brought about by the 1988 Brazilian Constitution after 20 years of authoritative rule, set the base for the expansion of benefits to all rural households, increasing the political viability and long-term sustainability of the cash transfer scheme (Barrientos, 2003).

International literature on the distributive effect of remittances on rural populations has been long established (Barham and Boucher, 1998; Stark et al., 1986; Taylor et al., 2005), but has yielded conflicting results about the impact of remittances on income inequality. The most convincing theoretical argument and empirical evidence in the literature suggests that remittances, as a responsive benefit of migration, are higher among origin-areas with short-term tradition to outmigration due to the high cost and risk of migration activity. As outmigration becomes more prevalent, risk declines due to social network returns; origin-household incomes and remittances become then less positively or even negatively correlated (Stark et al., 1986; VanWey, 2004). Literature on remittances and poverty is less established and few empirical studies can be found. Taylor et al. (2005) argues that remittances may influence poverty in two possible ways. Remittances might reduce poverty in origin areas by shifting population from low-income rural sectors to higher-income economic sectors through migration. Conversely, remittances may be inefficient in reducing poverty if migration is risky and costly, which prevents poor households from accessing migrant labor markets. Evidence supports the optimistic view that remittances are efficient in reducing poverty but increase their impact when social networks diffuse, reducing cost or risk of migration among the poor (Taylor et al., 2005). Research on remittances and land use also suggests that remittances might work as a combination of altruistic and contractual arrangements to mitigate risk and improve investment capacity of origin households (VanWey, 2004; Adams, 1996).

The aim of this study is to estimate the impact of public and private transfers on rural poverty in Brazil. We analyze the evolution of rural-state level wellbeing indicators from 1996 to 2011 using a GMM-system (Generalized Method of Moments) dynamic panel regression model. Our focus on understanding the distributional wellbeing effects of different sources of transfers (public or private) mirrors international research on the positive impact of transfers on rural wellbeing and investment capacity in developing countries (Barrientos, 2003; Taylor et al., 2005). The rural retirement income itself provides an opportunity to analyze an income shock that is tightly connected to the Brazilian economic growth. This aspect is particularly relevant to redistributive effects of the rural retirement program since its recent contributory requirement is tied to the minimum salary, which experienced a consistent appreciation in real terms from 1995 to the present.

Some studies suggest that the real appreciation of the minimum salary explained a substantial amount of the inequality decline in Brazil in the last decade (Soares, 2006; Hoffmann, 2010). From 1998 to 2008, for example, the minimum salary increased by 58% (Hoffmann, 2010). Barros et al. (2001) however emphasize that a higher value for the minimum salary may have a negative impact on poverty due to a reduction in employment opportunities. This result is modified when considering the impact of the increase in the purchase power of the minimum salary over the retirement income, that shows a strong positive multiplier effect on the economy, with an undoubtedly reduction in poverty (Barros et al., 2001) and inequality (Barros et al., 2007). This simultaneous relation between the value of the minimum salary and poverty/inequality indices is thus clearly mediated by the non-contributory portion of the retirement system, which does not depend on the labor market. This is the main explanation why, even in a system of indirect effects (in a general equilibrium scenario), the social security system functions as a powerful redistributive mechanism, with even stronger effects in its rural segment (Barros et al., 2001; Schwarzer, 2001; Afonso and Fernandes, 2005). The effect of the rural retirement income in metropolitan areas of Brazil is found to be regressive (increase inequality), and this is mainly explained by higher concentration of benefits over one minimum salary and by its connection to the labor market (contributory scheme) (Hoffmann, 2010). The non-contributory character of the rural system is the key that explains its potential as a redistributive mechanism from urban to rural areas, from richer to poorer regions (Afonso and Fernandes, 2005), and from the labor market to the poor elderly (França, 2004).

The literature seems to agree about the positive effect of the rural retirement on inequality decline (Soares et al., 2006; Hoffmann, 2010), although there is mixed evidence about its impact on poverty (Schwarzer, 2000; Barrientos, 2003; Hoffmann, 2006; Marinho and Araujo, 2010). Following França (2004), we argue that part of this mixed evidence regards the limited scope of the studies when defining wellbeing. França (2004), for instance, shows that the social security system is highly relevant for the local economy across Brazilian municipalities, reaching a higher share of the GNP (Gross National Product) than the "Fundo de Participação dos Municípios" in 92 of the 100 municipalities with the highest HDI (Human Development Index). The relevance of the social security system to the local economies will further increase with the rapid population aging under course in Brazil (Ansiliero and Paiva, 2008). Although not directly captured by the traditional poverty and inequality indices, case studies reveal important improvement in non-monetary dimensions of wellbeing among households with elderly receiving the rural retirement income. Schwarzer (2000) found that the rural retirement income is being invested in not only food, but also in house improvement and private health products and services. Augusto and Ribeiro (2006) argue that the rural benefit increased the ability to acquire credit, facilitating the acquisition of durable goods and services, in addition to fostering small businesses. Albuquerque et al. (1999) suggest that the rural retirement prevents rural-urban migration by improving the wellbeing in rural areas and is an important mechanism to reduce extreme poverty and prostitution, as well as to increase elderly's longevity.

Some authors argue that not only the source of income but also its stability may influence the expenditure behavior of households, especially among rural populations where credit and insurance markets are underdeveloped or absent (Rosenzweig, 1988; VanWey, 2004) and property rights not fully established (Ludewigs et al., 2009). Variable income sources (such as remittances) are considered to be generally spent on immediate needs (such as nutrition and non-durable goods), while stable and lasting income sources (such as retirement income) might be channeled towards productive and permanent investments (such as human capital, physical capital, and market-oriented land use systems) (Brown, 2006). We argue that because a certain level of income is needed to trigger productive investment, both sources of income may interact in order to allow reduction in investment poverty. In addition, because rural families with children under 17 years old enrolled in public schools are eligible for the Bolsa-Familia conditional cash transfers, they might use it as a mechanism to improve welfare well-being (Duarte et al., 2007), releasing remittances and retirement income to be invested in productive activities.

2. Data

To estimate the impact of public and private transfers on poverty indicators, we combined microdata from the Brazilian National Household Survey (PNAD), covering the years from 1996 to 2011, and state-level data derived from State Statistics Bureaus, Government State Offices, and SUFRAMA (Superintendência da Zona Franca de Manaus, in Portuguese).

These state-level data are compiled and made publicly available by the Brazilian Institute of Geography and Statistics (IBGE). PNAD is also collected and distributed by IBGE.

We started by using PNAD data at the individual level and generating a series of drivers of poverty, using both individual and household-level information. Then, we collapsed all individual-level data among the rural population at the state level, creating a panel of rural-state data from 1996 to 2011. Because PNAD is not collected in years when the Demographic Census is conducted, we would lose two years over the period analyzed. Thus, we interpolated the missing years (2000 and 2010) using the values of each variable from years right before and after, creating a balanced panel dataset with N = 27 states[2] and T = 16 years. We opted for constraining the analyses to include all rural-state units, but the 7 units belonging to the rural North. This was necessary since PNAD is not representative of the rural areas of states from the North Region in Brazil before 2004. From PNAD microdata we estimated the following variables: poverty indices, inequality indices, retirement income, private transfers, educational attainment, Land Concentration indices, and the unemployment rate. From IBGE aggregate data we estimated the state contribution to Brazilian GNP and the agricultural share of the Gross State Product (GSP). Details of variable construction are given in the next section.

3. Variables Construction

3.1. Dependent Variables

Our dependent variable is represented by poverty indices of the Foster-Greer-Thorebecke family (Foster et al., 1984). Poverty then will be measured by the following three FGT measures: the headcount ratio (P_0), the poverty gap (P_1) and the squared poverty gap (P_2). Each of these measures requires a previously established poverty line, z. With the poverty lines correctly specified, the poverty headcount ratio (P_0) is defined by:

$$P_0 = \frac{h}{n} \quad (1)$$

where h is the number of poor individuals in a population with n persons, with restriction $0 \leq P_0 \leq 1$. This is a measure of incidence or extension, not taking into account poverty intensity. Thus, P_0 is insensitive to decline of a poor's income (Hoffmann, 2000; Simão, 2004). If income insufficiency is considered as the difference $z - x_i$, with $i \leq h$, where z is the poverty line and x_i the income from the x-ith poor, the income poverty insufficiency ratio, I, can be defined as:

[2] Distrito Federal (DF) is included as a typical state. Because it is mainly urban, we performed analyses excluding it but results did not change. So, to avoid reduction in sample size we decided for keeping DF in the final results.

$$I = \frac{1}{hz}\sum_{i=1}^{h}(z - x_i) \quad (2)$$

where hz is the maximum value for income insufficiency if all h poor persons had no income. Thus, the higher the value I, the lower the average income of the poor relative to z. If one calls m as the average income of the poor, given by:

$$m = \left(\sum_{i=1}^{h} x_i\right)^{-h} \quad (3)$$

it can be shown that:

$$I = 1 - \frac{m}{z} \quad (4)$$

Equation (4) shows that, for given values of z and m, I is insensitive to the number of poor persons (h). Measures P_0 and I are complementary ones, the former being insensitive to povety intensity, the latter to poverty incidence (Hoffmann, 1998). Foster, Greer, and Thorbecke (1984) proposed a class of poverty measures, given by the general formula:

$$\varphi(\alpha) = \frac{1}{nz^{\alpha}}\sum_{i=1}^{h}(z - x_i)^{\alpha} \quad (5)$$

Where $\alpha \geq 0$. It can be shown that $0 \leq \varphi(\alpha) \leq 1$, with the following extreme cases: when $\varphi(\alpha) = 0$, all individuals have $x_i > z$; when $\varphi(\alpha) = 1$ $\varphi(\alpha) = 1$, all individuals have $x_i = 0$. Class measure (5) summarizes all above measures, P_0 and I. When $\alpha = 0$, Equation (5) becomes P_0, while $\alpha = 1$ represents $P_0 I$. The latter measure is called poverty gap (P_1). When $\alpha = 2$, FGT represents the severity of poverty (P_2). The measure P_2 is a function of both P_0 and P_1, and of a coefficient of variation for the income of poor individuals, as shown in Hoffmann (1998). Therefore, P_2 is sensitive to the number of poor individuals, how poor they are, and how unequal they are among them (Hoffmann, 2000).

FGT poverty indices are decomposable, that is, their values at the aggregate level may be reconciled by averaging out lower-level indices (such as state or municipality levels), with weights being given by the lower-level share to the aggregate level. They also meet the focal axiom, since they are all insensitive to variation in non-poor income (Hoffmann, 1998; Expert, 2006). Other desirable properties for axiomatic indices are not met by all the three measures. For instance, P_0 does not meet two properties: 1) monotonicity, because it is insensitive to variation of the income among the poor individuals, 2) focal axiom, because it does not respond to within-poor income redistribution. P_1, while satisfying the monotonicity axiom, does not respond to the focal axiom. The only FGT satisfying all the axiomatic properties is P_2, but it is the less intuitive to interpret (Expert, 2006). The different dimensions of income poverty covered by the three FGT indices described above justify their separate use for the analyses given in this study. In this study we present results for P_0 only, but all regression results using P_1 and P_2 are available upon authors' request.

Because the cost of living is heterogeneous in different parts of Brazil (including rural-urban differences), we used the regional poverty lines estimated by the Brazilian Institute for Applied Economics[3] (IPEA, 2014). The original IPEA data on regional poverty lines are available for the period 1976-2009, although years when PNAD is not collected have missing information for the series. Thus, we need to estimate values for 2000, 2010, and 2011. For 2000 we used the average value from 1999 and 2001 for each regional value. For the years 2010 and 2011, we used forecasted values from an ARIMA (0,1,1) regression model. Because all series showed unit root, requesting correction for changing averages over time, deterministic projections would be naïve estimates. Estimates of the regional poverty lines used in our poverty measures are available upon authors' request.

3.2. Independent Variables

Our two state variables are the rural retirement income and income received from other households. These are our proxies for public and private transfers. Because our models are estimated at the rural-state level over time, we tested different proxy specifications, such as the proportion of individuals in the rural area not covered by the rural retirement system, as well as the proportion of individuals not receiving any private income from non-coresidents. Because results with both type of measures did not change results significantly, we opted for the income measures since it is more intuitive to interpret, as well as a direct component of the FGT poverty measures used.

Following the procedure proposed by Marinho and Araujo (2010), we estimated the rural retirement income by first identifying all individuals who declared in the PNADs to be receiving one minimum salary as retirement income, who were living in the rural area, and who were at least 60 years old if men and at least 55 years old if women. Although we acknowledge that this proxy may underestimate the number of individuals covered by the Rural Retirement System, since there can be individuals living in rural areas and contributing to the General Social Security System or receiving more than one minimum salary, it is impossible to identify the exact persons receiving the specific income using PNAD data. Then, we multiplied the numbers of beneficiaries above identified by the nominal value of the Brazilian minimum salary for each year from 1996 to 2011. This gives us a proxy for the total amount of money provided by the Rural Retirement System in rural Brazil. Finally, we divided this total amount by the number of individuals in the rural area of each state for each year, resulting in a per capita rural retirement income, as suggested by Marinho and Araujo (2010). Different from the authors, however, we acknowledge that this measure would not

[3] Although the regional poverty lines are currently available on IPEA website, in the beginning of the paper writing the data was available upon request only. We would like to thank Emerson Marinho and Eduardo Araújo, from the Economics Department at Ceará Federal University (CAEN/UFC), who sent us the data for use.

capture the impact of public transfers on poverty in the econometric models, because it is highly contaminated by the effect of age structure. Therefore, in the regression models, we controlled for age structure to standardize demographic structures across states. For age structure, we defined two variables: proportion of individuals aged 15 to 64 years old and proportion of individuals aged at least 65 years old per state and year.

To proxy private transfers we used individual income received from other households. This is non-coresident private income transfers as appeared in PNAD, with no transformation. Although a direct measure of income private transfer, it is clearly under declared since other datasets, such as the Budget Family Survey (POF, in Portuguese), also collected by IBGE, shows significant higher levels of transfers (Diniz et al., 2007). We could not use POF, however, since it is not representative for rural areas, in addition to having only three points in time available for the period here studied. Because of the known downward bias in the level of private transfers, we tried to use the proportion of rural residents receiving any money from other households, but results did not changed significantly. An interaction term between public and private transfers was created to capture the triggering effect pointed out above. The interaction is expected to be negative, powering the effect of public income on poverty for those receiving more volatile, non-public income. Because we are dealing with nominal values of income in a time series, current values had to be deflated to account for inflation over the period analyzed. We used the Courseil and Foguel (2002) implicit deflator for PNAD. The index was adapted to reflect real prices at 1996 values (baseline). We also tried the implicit deflator for the Gross National Product, estimated by IBGE, with no difference in trends. Because the deflator suggested by Courseil and Foguel (2002) is specific for PNAD, we decided to use that one. All transfers were transformed to Neperian logarithm to correct heavily positively skewed distributions across states for every year.

For inequality indices, we tried different specifications: Gini, Mehran, and Piesch. Because Mehran is more sensitive to pro-poor redistribution, Piesch to pro-rich redistribution, and Gini is an average of both indices; we tested models with each one of those (Hoffmann, 2004). We decided for Mehran because of its sensitivity to change in poor income, since we are interested in the dynamics of poverty over time. In addition, because inequality indices vary little from year to year and cross-sectionally in absolute terms, we took the Neperian logarithm to reveal hidden scale heterogeneity. To control for regional economic factors, we used the proportion of Gross State Product due to the agricultural sector, as well as the proportion of GSP to the Gross National Product. These variables were available from aggregate data at IBGE website. We also controlled for differences in education attainment of the rural population across state and over time. Because average years of education completed showed little variance cross-sectionally and temporally, we estimated a proxy as the proportion of individuals in the rural area with at least 8 years of education completed. This strategy avoids lack of explanatory power of education on poverty due to lack of data variability. We also used a land concentration index. The Brazilian Agricultural Census could be used, although it would be available for 3 years only. Thus, we created a proxy using the PNAD microdata. We first summed all land owned by rural employers and autonomous farmers in hectares (first, second, and third parcels as informed by PNAD questionnaire). To avoid bias in the calculation of the land concentration index due to influential cases (extreme outliers), we excluded those farm owners with land area above 3 standard deviations. Then, the Merhan index was used to estimate how unequal land areas were distributed across states and over time. As for the income inequality measure, we took the Neperian logarithm of the land concentration index to reveal hidden heterogeneity. Finally, unemployment rates were used as traditionally defined: the proportion of unoccupied individuals in the rural area divided by the number of economically active individuals in rural areas.

4. Methodology

To estimate the impact of public and private transfers on poverty dynamics, we use a first order linear dynamic panel model of the form:

$$y_{it} = \rho y_{it-1} + x_{it}\beta + u_i + \varepsilon_{it} \qquad (6)$$
$$for\ i = \{1, \cdots, N\}\ and\ t = \{1, \cdots, T\}$$

where u_i represents the individual heterogeneity, and captures the non-observed and time invariant effects which affect the dependent variable. This individual effect includes a wide range of factors, such as geographic characteristics and cultural factors. The ε_{it} term represents the idiosyncratic errors vector, identically and independently distributed. In a dynamic panel of this sort y_{it} exhibits *state dependence*, that is, the current y level depends on its level in the last period, even after the individual heterogeneity (u_i) and other control variables (x_{it}) are included in the model. The lagged y in Equation (6) is, by construction, correlated with the individual effects, since y_{it-1} contains u_i. However, usual methods used to eliminate individual effects, such as the *within transformation* (Fixed Effects estimation), and the first difference transformation, still yield inconsistent parameter estimators. The inconsistency persists because such transformations induce correlation between the transformed error terms and the transformed lagged variable, y_{it-1}.

The usual dynamic panel estimation consists of transforming variables in first difference, or forward orthogonal deviations, in order to eliminate the individual effects[4]. Then, it uses Two Stage Least Square (2sls) or Generalized Method of the Moments (GMM) estimation with appropriate selection of instruments to reduce the correlation of the first difference of the lagged dependent variable (Δy_{it-1}) and the transformed error terms ($\Delta\varepsilon_{it}$). The use of the GMM method for dynamic panels was first introduced by Holtz-Eakin, Newey and Rosen (1998), latter developed by Arellano and Bond (1991), Arellano and Bover (1995), and Bludell and Bond (1998). The GMM for panel data allows simultaneously control for individual and temporal effects, at the same time attenuating endogeneity created by the inclusion of the lagged dependent variable in the model as an explanatory variable. There are at least two main variants of GMM estimators for dynamic panel: the first difference GMM estimators (Arellano and Bond, 1991) and the GMM-system (Blundell and Bond, 1998). The GMM estimator in first differences consists in estimating the regression equation with all variables as the first difference of the original variables in level, using lags of the lagged

[4] The within transformation can be used if the available instruments are strictly exogenous; for models in which the strict exonegeneity is violated, instead holding sequential exogeneity only, first difference is a better strategy (Wooldridge, 2010). For unbalanced panels a common strategy is to perform forward orthogonal deviations, minimizing loss of cases (Arellano and Bover, 1995). Forward orthogonal transformations consist in subtracting the average of future values of the variable from its current value.

term, y_{it-1} ($t \geq 3$) and the lagged exogenous variables as instruments so that endogeneity induced by the correlation between the lagged differenced endogenous variable and the differenced errors are attenuated. Arellano and Bond (1991) suggest using the lagged explanatory variables in level as instruments for the equation in first difference. Blundell and Bond (1998) developed a GMM-system estimator, which combines in the parameters equation the equations in first difference with the equations in level. The former are instrumented by the lagged variables in level, while the latter are instrumented by the variables in first difference. This empirical strategy is the solution for variables with unitary root.

In this paper, we test both approaches to test which specification yields more robust results, using the Hansen test for instruments, as in the Arellano/Bond approach, and the Sargan test for extra instruments used in the GMM-system approach (Blundell and Bond, 1998). We also performed the Arellano-Bond test for error autocorrelation, since GMM-system estimators are consistent under two conditions: validity of extra instruments used and absence of serial autocorrelation of residuals (Bludell and Bond, 1998).

4.1. Empirical Model

Our equation in level for the FGT poverty measures is defined as:

$$P_{\alpha,it} = \beta_0 + \beta_1 P_{\alpha,it-1} + \beta_2 ln[PubInc_{it}] + \beta_3 ln[PrivInc_{it}] + \beta_4 ln[PubInc\ x\ PrivInc_{it}] + \beta_5 ln[Mehran_{it}] + \beta_6 ln[\%AgrGSP_{it}] + \beta_7 ln\left[\%\frac{GSP}{GNP}_{it}\right] + \beta_8 ln[MehranLand_{it}] + \beta_9 UnempRate_{it} + \beta_{10}\%Persons_{it}^{15-64} + \beta_{10}\%Persons_{it}^{65+} + u_i + \varepsilon_{it} \qquad (7)$$

Where:

P_α = FGT poverty index ($\alpha = 0,1,2$)
$PubInc$ = Per capita rural retirement income (deflated)
$PriVinc$ = Income received from non-coresidents (deflated)
$Mehran$ = Mehran income inequality index
$\%AgrGSP$ = Agricultural share of the Gross State Product (%)
$\%GSP/GNP$ = State share of the Gross National Product (%)
$MehranLand$ = Mehran land inequality index
$UnempRate$ = Unemployment rate (%)
$\%Persons^{15-64}$ = Proportion of individuals aged 15 to 64
$\%Persons^{65}$ = Proportion of individuals aged at least 65

Our equation in first difference is given by:

$$\Delta\{P_{\alpha,it}\} = \beta_1\Delta\{P_{\alpha,it-1}\} + \beta_2\Delta\{ln[PubInc_{it}]\} + \beta_3\Delta\{ln[PrivInc_{it}]\} + \beta_4\Delta\{ln[PubInc\ x\ PrivInc_{it}]\} + \beta_5\Delta\{ln[Mehran_{it}]\} + \beta_6\Delta\{ln[\%AgrGSP_{it}]\} + \beta_7\Delta\left\{ln\left[\%\frac{GSP}{GNP_{it}}\right]\right\} + \beta_8\Delta\{ln[MehranLand_{it}]\} + \beta_9\Delta\{UnempRate_{it}\} + \beta_{10}\Delta\{\%Persons_{it}^{15-64}\} + \beta_{10}\Delta\{\%Persons_{it}^{65+}\} + \Delta\{\varepsilon_{it}\} \quad (8)$$

where $\Delta\{P_{\alpha,it}\} = P_{\alpha,it} - P_{\alpha,it-1}$. Because $E\left\{\Delta\{P_{\alpha,it-1}\}, \Delta\{\varepsilon_{it}\}\right\} \neq 0$, Ordinary Least Square estimators would be biased and inconsistent. Thus, instruments for $\Delta\{P_{\alpha,it-1}\}$ must be used. Assuming the moment conditions $E\left\{\Delta\{P_{\alpha,it-s}\}, \Delta\{\varepsilon_{it}\}\right\} = 0$ for $t = 3,4, \ldots, T$ and $s \geq 2$, good instruments for Equation (7) would be $\Delta\{P_{\alpha,it-s}\}$ for $t = 3,4, \ldots, T$ and $s \geq 2$, as suggested by Arellano and Bond (1991). This empirical strategy eliminates weak endogeneity only. In the presence of strong time persistence of poverty, a strategy of estimation in system, combining level and difference as instruments (see discussion above) would yield consistent estimators under endogeneity, where errors are correlated in the past, present, and future (Arellano and Bover, 1995; Bludell and Bond, 1998). The inconsistency in estimators for Equation (7) with differenced instruments is asymptotically irrelevant on T. We believe the consistency gain from the GMM-system over the GMM-difference is relatively small for our panel, since we have a relatively large time window (from 1996 to 2011). To assure robust results, we test both strategies.

Additional endogeneity had to be considered in our empirical model. We assumed that Mehran, % GSP/GNP, and %AgricGSP are endogenous to poverty. In addition, the argument that retirement income is exogenous to poverty is only valid until 2005; from 2006 on, contributory rules became effective, varying by type of benefit. Thus, contributory capacity and retirement income become endogenous starting at 2006, with individuals in 2006 who could retire without any contribution suddenly being forced to contribute. This could raise poverty in the year following the change in contribution rule. Thus, we instrumented retirement income in 2006 with a lagged value for 2006, 2 lags for 2007, and so on. As shown in Tables 3, we used 176 instruments in the GMM-system regressions (all Sargan tests for over identification not significant at 5%). For all estimated models, we weighted the covariate matrix with the individual variance to produce robust standard errors of parameters. To compare gain in parameter consistency, we show three models for P_0 FGT poverty measure with increasing consistency: Ordinary Least Square, Fixed Effect, and GMM-System with forward orthogonal deviations.

5. Results

5.1. *Descriptive Results*

As discussed in section 4, our analytical panel sample comprises 320 observations for the level dataset and 280 for the instrumented first difference dataset. In total, 20 rural state units over 16 years (1996 to 2011) were used in our analysis of poverty dynamics. Table 1 shows that the grand average poverty level over the period (and states) was 49.49%, ranging from values as low as 7.5% (São Paulo in 2011) and as high as 86.9% (Ceará in 1997). Most of the variation observed comes from differences in poverty incidence across states, although poverty decline over time for each state is considerably high, especially after 2003 (line trends available upon authors' request). As previously pointed, the per capita retirement income (expressed in 1996 R$) is contaminated by differences in state-year age structure. This being said, variation is balanced from between states and within state over time variation. With an average of R$11.57 per individuals, values range from R$0.78 (Distrito Federal in 1996) to R$30.79 (Rio Grande do Sul in 2011). Average monthly income from non-coresidents shows a low value, as large as R$98.31, ranging from R$14.68 (Espírito Santo in 2008) to R$2,065.24 (Espírito Santo in 2006). The range of extreme values concentrated in the same state in such a short period of time reflects the volatile nature of this income source, with the largest share of variation being accounted by the within state differences over time.

Table 1: Descriptive statistics for variables used in the dynamic panel models of rural poverty in Brazil - 1996 to 2011

Variable		***Mean***	***Std. Dev.***	***Min***	***Max***
Headcount ratio	overall	49.49	22.74	7.49	86.88
	between		20.77	24.01	75.29
	within		10.28	24.03	68.35
Poverty gap	overall	23.72	14.75	2.00	54.95
	between		13.56	8.20	42.08
	within		6.50	2.12	36.88
Squared poverty gap	overall	14.61	10.45	0.88	39.62
	between		9.59	3.91	28.61
	within		4.63	-2.49	26.09
Per capita rural retirement income	overall	11.57	5.57	0.78	30.79
	between		3.72	2.50	18.40
	within		4.22	3.36	23.96
Income received from non-coresidents	overall	98.31	149.66	14.68	2065.24
	between		46.87	50.00	230.59
	within		142.49	-117.60	1932.96
Mehran income inequality index	overall	0.603	0.047	0.485	0.799
	between		0.035	0.553	0.708
	within		0.033	0.522	0.738
Mehran land	overall	0.953	0.057	0.600	1.000

inequality index	between		0.032	0.876	0.993
	within		0.048	0.641	1.078
Agricultural share of the Gross State Product	overall	9.23	6.44	0.21	35.35
	between		6.25	0.32	26.20
	within		2.07	-0.16	18.37
State share of the Gross National Product	overall	4.76	7.45	0.50	36.72
	between		7.62	0.53	34.53
	within		0.35	2.80	6.95
Unemployment rate	overall	3.09	2.52	0.20	12.97
	between		2.35	0.90	9.73
	within		1.03	-1.21	6.56
Proportion of 15 to 64 years old individuals	overall	0.70	0.02	0.63	0.77
	between		0.02	0.67	0.73
	within		0.02	0.65	0.74
Proportion of 65 and over years old individuals	overall	0.10	0.02	0.03	0.19
	between		0.02	0.05	0.14
	within		0.02	0.07	0.15

Source: PNAD 1996-2011 (IBGE); State-level Statistics (IBGE)
Note: N = 320 / n = 20 / T = 16. North states excluded.

The grand mean for Mehran income inequality index was estimated of as 0.603, ranging from 0.485 (Santa Catarina in 2009) to 0.799 (Distrito Federal in 1998). Within and between variations are balanced, as for the Mehran land concentration index. For the latter, an estimated grand mean of 0.953 hides strong differences from 0.600 (Distrito Federal in 2001) to 1.000 (Bahia in 2007). The occurrence of an index as high as 1.000 is a clear sign of declaration bias from values reported in PNAD, but its relative distribution across states mirrors the history of land concentration in rural Brazil. Both sectorial and regional economic indicators have their grand mean variation mostly accounted by regional (between) differences, as expected. In contemporary rural Brazil, the importance of the agricultural sector to the GSP in our sample ranged from 0.2% (Distrito Federal in 2006) to 35.5% (Mato Grosso in 2004). The regional contribution to the GNP ranged from 0.5% (Piauí in 2001) to 36.7% (São Paulo in 1997), mirroring the strong regional concentration of economic activities in Brazil. Finally, unemployment rates in rural areas also showed strong between-state variation, ranging from as low as 0.2% (Mato Grosso do Sul in 1996) to as high as 13.0% of the economically active population (Distrito Federal in 2007).

5.1. *Regression Results*

We turn now to our regression results. Table 2 shows the estimated effects of public and private transfers on rural poverty in Brazil. Using the headcount ratio as the dependent variable, we compare three estimation procedures: OLS, Fixed-Effect, and GMM-System. As shown in the table, we found a significant persistence effect of poverty over time, even in the GMM-System. As expected, OLS usually overestimates the lagged coefficient, due to strong bias in the estimation caused by endogeneity. Also, as expected by the econometric theory, GMM-System standard errors are larger, leading to more conservative hypothesis tests for coefficients. We found that both rural retirement and private income have a significant statistical impact on poverty in rural Brazil with scale dominance for public transfer. Also as discussed by the economic literature on the influence of income volatility on investment decisions and poverty reduction, our interaction effect is highly significant, meaning that the ability of public income to reduce poverty is powered by the additional levels of income from private transfers. It is worth nothing that endogeneity in both income sources were taken into account in the GMM-System estimation, with additional tests required by the estimation procedure being validated by the non-significant value for the Sargan test (Prob > chi^2 = 0.5890). The Arellano-Bond test for zero autocorrelation in first-differenced errors shows that autocorrelation is present for the first order difference, justifying the use of GMM-System strategy. Results are consistent for all poverty measures used (results for the other two FGT measures available upon request).

Table 2: Determinants of Poverty Dynamics in Rural Brazil from 1996 to 2011 - Headcount Ratio

Variable	***OLS***	***Fixed Effect***	***GMM-System***
Lagged headcount ratio	0.905***	0.542***	0.544***
	(0.0358)	(0.0494)	(0.0454)
Ln of public retirement income	-1.779	-13.59***	-16.65**
	(4.062)	(4.271)	(7.095)
Ln of Mehran inequality index	5.908	12.63***	8.391
	(3.617)	(4.364)	(6.643)
Ln of private transfer income	-2.095	-6.133***	-8.293**
	(2.136)	(2.056)	(3.607)
Interaction (public x private income)	-0.572	-2.256**	-2.816*
	(0.870)	(0.935)	(1.456)
Ln of GSP agricultural share	-0.348	4.101***	5.206***
	(0.356)	(0.987)	(1.088)
Ln of GNP state share	0.108	-7.884**	-8.050**
	(0.270)	(2.847)	(3.541)
% of individuals with 8 years + of education	-16.50***	-34.86***	-41.74***
	(5.988)	(11.43)	(12.74)
Ln of Mehran land concentration index	0.421	2.875	4.791
	(3.117)	(3.267)	(4.875)
Unemployment rate	0.0778	-0.183	-0.199
	(0.170)	(0.138)	(0.198)
% of 15-64 years old individuals	-47.95**	-22.65	-18.10
	(20.50)	(16.90)	(20.81)
% of 65+ years old individuals	-43.14*	-56.01**	-35.06
	(24.73)	(20.15)	(27.08)
Constant	54.18***		

	(18.31)		
Global Test (F or Wald)	F(12,287) = 1703.48	F(12,19) = 329.16	Wald chi2(12) = 3258.34
Sargan test of overidentifying restrictions (Prob > chi2 = 0.5890)			S chi2(164) =159.30
Arellano-Bond test for zero autocorrelation in first-differenced errors	First Order	P[z(-10.13)>z]=0.0000	
	Second Order	P[z(1.37)>z]=0.1697	
Observations	320	320	280
Number of groups (States)		20	20
Instruments (#)			176
R-squared	0.980	0.928	

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1
Source: PNAD - 1996/2011 – North states excluded.

For the control variables, effects go in the direction expected by the theory and empirical literature. Although the Mehran income inequality index was not significant in the GMM-System regression for P_0, its effect is statistically valid for the other two FGT indices (not shown). This is expected since P_1 and P_2 are more sensitive to income distribution, with the latter even more sensitive to change in relative positions of the poor among them. As shown in Table 2, the sectorial and regional economic proxies show strong impacts on rural poverty. For instance, a 1-point increase in the ln of the agricultural share to GSP for a state in Brazil (approximately 2.71% increase in the original scale) would raise poverty by 5%. The same increase in the state contribution to GNP would decrease poverty by 8%. Education is also quite powerful in reducing poverty; with an increase in 1% of individuals in the rural area with at least 8 years of education reducing rural poverty by almost 42%. The high impact of education is explained by the variable used in our model, which emphasizes the top part of the educational distribution; using average years of education would have a smaller effect. Scale here, however, is irrelevant.

6. Final Remarks

Although the Brazilian Social Security System was created in 1923, it was just in 1971 that the government established a non-contributory retirement program to cover the rural population, incorporated to the general retirement system after 1988 Brazilian Constitution was promulgated (Schwarzer, 2000). Different from other international experiences, the Brazilian rural retirement system was universal and non-contributory until mid-2006, depending on age and affiliation of the elderly to agricultural activities (Schwarzer, 2000; Kreter and Bacha 2006). There is also no means test. Thus, the presence of an elderly in the household represents an income shock to the family with likely impacts on investment capacity of smallholders and powerful sectorial redistribution (Carvalho-Filho, 2008; Ramos and Arend, 2012). To date, local level studies suggest that the retirement income has important impact on the local economy, as well as on the well-being of beneficiary households, increasing family expenditures on food and home improvement (Schwarzer, 2000; Barrientos 2003).

There is a vast literature on poverty and inequality decomposition in Brazil showing empirical evidence of the importance of retirement income for poverty and inequality dynamics in the last 15 years. Hoffmann (2006), for instance, shows that pensions and retirement income have a concentration effect on inequality. This pro non-poor effect of the social security system in Brazil is likely to be reflecting the high values of benefits from government bureaucrats. In addition, before recent changes in the implementation of an upper limit for benefits, beneficiaries with higher ability to contribute still receive substantial income from the system. In rural areas little is known about the impact of the non-contributory rural retirement program on poverty and inequality decline. A recent study (Marinho and Araujo, 2010) found no effect of retirement income on rural poverty for Brazil between 1995 and 2004. Their study caries a detailed econometric analysis to account for endogeneity of variables and the dynamic nature of poverty over time.

From 2004 to 2011 a dramatic change in monetary well-being indicators took place in Brazil. In the last 20 years, many studies found a significant reduction in poverty and inequality (Hoffmann, 2010, Corseuil et al., 2011). Afonso and Fernandes (2005), for instance, argue that the rural retirement income has a very strong impact on poverty, since the highest estimated Internal Return Rate for benefits is found in areas where the Rural Retirement is more important, such as the rural areas of the North and Northeast regions of Brazil. Ansiliero and Paiva (2008) also show the consistent increase in the social security coverage in Brazil, with rural areas accounting for the highest percentage increase. Thus, the recent combination of increase in the minimum salary, which the rural retirement benefit is tied, and coverage expansion of the rural system produced a different context for the role played by the retirement income on rural poverty. Building on Marinho and Araujo (2010) study, we estimated a dynamic panel model of rural poverty combining different datasets and including a wider time horizon, from 1996 to 2011. Different from the authors, we found a strongly and consistent significant impact of the rural retirement program on poverty dynamics for contemporary Brazil, confirming empirical evidence as in Barrientos (2003) and evidence from institutional analyses (França, 2004; Kreter and Bacha, 2006).

Our findings, compared to Marinho and Araujo (2010), may reflect a variety of differences: 1) a larger panel dataset (important for the asymptotic properties of estimators and hypothesis testing), 2) differences in the transformation of variables (we did try logarithm transformations of the dependent variable, but empirical distribution suggests that FGT measures should be used as estimated), 3) inclusion of additional variables, such as land concentration index, sectorial importance of agriculture, and regional contribution to GNP, and, finally, 4) the inclusion of age structure in the models. The last difference is key, since by construction the per capita rural retirement benefit in the authors' study was jointly capturing monetary impact and age structure influence on poverty. Our econometric results confirm the importance of the rural retirement benefit for rural wellbeing in Brazil, as suggested by many previous qualitative studies (Albuquerque et al., 1999; Augusto and Ribeiro, 2006).

7. References

Adams, R. H. (1996). "Theories of International Migration: A Review and Appraisal", in *FCND Discussion Paper 17*, International Food Policy Research Institute (IFPRI).

Afonso, L. E.; Fernandes, R. (2005). “Uma estimativa dos aspectos distributivos da previdência social no Brasil”, in *Revista Brasileira de Economia*, Vol. 59, N. 3, p. 295-334.

Ansiliero, G.; Paiva, L. H. (2008). “The recent evolution of social security coverage in Brazil”, in *International Social Security Review*, Vol. 61, N. 3, p. 1-28.

Albuquerque, F. J. B.; Lobo, A. L.; Raimundo, J. S. (1999). “Análise das repercussões psicossociais decorrentes da concessão de benefício rurais”, in *Psicologia: Reflexão e Crítica*, Vol. 12, p. 503-519.

Arellano, M.; Bond, S. (1991). “Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations”, in *Review of Economic Studies*, Vol. 58, p. 277-297.

Arellano, M.; Bover, O. (1995). “Another look at the instrumental variable estimation of error-components models”, in *Journal of Econometrics*, Vol. 68, p. 29-51.

Augusto, H. A.; Ribeiro, E. M. (2006). “O idoso rural e os efeitos das aposentadorias rurais no domicílio e no comércio local: O caso de Medina, nordeste de Minas”, in *Anais do Encontro Nacional de Estudos Populacionais 15*, Caxambú. ABEP.

Barham, B.; Boucher, S. (1998). “Migration, Remittances, and Inequality: Estimating the Net Effects of Migration on Income Distribution”, in *Journal of Development Economics*, Vol. 55, N. 2, p. 307-331.

Barrientos, A. (2003). “What is the impact of non-contributory pensions on poverty? Estimates from Brazil and South Africa”, in *CPRC/DEPP Working Paper* 33, Manchester: IDPM, University of Manchester.

Barros, R. P. D.; Corseuil, C. H.; Cury, S. (2001). “Salário mínimo e pobreza no Brasil: estimativas que consideram efeitos de equilíbrio geral”, in *IPEA Working Paper*, N. 779.

Barros, R. P.; Carvalho, M.; Franco, S.; Mendonça, R. (2007). “A importância da queda recente da desigualdade para a pobreza”, in *IPEA Working Paper*, N. 1258.

Blundell, R.; Bond, S. (1998). “Initial conditions and moment restrictions in dynamic panel data models”, in *Journal of Econometrics*, Vol. 87, p. 115-143.

Brown, S. S. (2006). “Can Remittances Spur Development? A Critical Survey”, in *International Studies Review*, Vol. 8, p. 55-75.

Carvalho Filho, I. (2008). “Household income as determinant of child labor and school enrollment in Brazil: evidence from Social Security Reform”, in *IMF working paper*, N. 08241. Disponível em: <http://www.imf.org/external/pubs/ft/wp/2008/wp08241.pdf>. Acessed in: Jan, 20, 2014.

Corseuil, C.; Foguel, M. (2002). “Uma sugestão de deflatores para rendas obtidas a partir de algumas pesquisas domiciliares do IBGE”, in *IPEA Working Paper*, N. 897.

Corseuil, C. H.; Moura, R.; Ramos, L. (2011) “Determinantes da expansão do emprego formal: o que explica o aumento do tamanho médio dos estabelecimentos?” São Paulo: *Revista Economia Aplicada*, v. 15, n. 1.

Diniz, B. P. C.; Silveira, F. G.; Bertasso, B.; Magalhes, L. C. F.; Servo, L. M. S.; Silveira, F. G., ...; Piola, S. F. (2007). “As pesquisas de orçamentos familiares no Brasil”. in *Gasto e consumo das famílias brasileiras contemporâneas*, Vol. 2, p. 17- 74.

Duarte, G. B.; Sampaio, B.; Sampaio, Y. (2007). “Impactos do programa Bolsa Família sobre os gastos com alimentos de famílias rurais”, in *Fórum Banco do Nordeste*.

EXPERT GROUP ON POVERTY STATSITICS. Rio Group. (2006). *Compendium of best practices in poverty measurement*. Santiago de Chile: CEPAL. 156pp. Disponível em: <http://www.eclac.cl/publicaciones/xml/3/26593/rio_group_compendium.pdf>

França, A. S. (2004). *Previdência social e a Economia dos Municípios*. Brasília, DF: ANFIP. 17pp.

Foster, J.; Greer, J.; Thorbecke, E. (1984). “A class of decomposable poverty measures”, in *Econometrica: Journal of the Econometric Society*, p. 761-766.

Hoffmann, R. (1998). *Distribuição de Renda: Medidas de Desigualdade e Pobreza.* São Paulo: Edusp. 275pp.

Hoffmann, R. (2000). “Mensuração da desigualdade e da pobreza no Brasil”, in Henriques, R. (Org.). *Desigualdade e pobreza no Brasil.* Rio de Janeiro: IPEA. Pp. 81-107.

Hoffmann, R. (2004). “Decomposition of Mehran and Piesch inequality measures by factor components and their application to the distribution of per capita household income in Brazil”, in *Brazilian Review of Econometrics*, Vol. 24, N. 1, p. 149- 171.

Hoffmann, R. (2006). “Transferências de renda e a redução da desigualdade no Brasil e cinco regiões entre 1997 e 2004”, in *Econômica*, Vol. 8, N. 1, p. 55-81.

Hoffmann, R. (2010). “The Evolution of Income Distribution in Brazil: What Promotes and What Restricts the Decline in Inequality”, in *Annals of the conference A COMPARATIVE ANALYSIS OF GROWTH AND DEVELOPMENT: ARGENTINA AND BRAZIL*. University of Illinois, Chicago.

Holtz-Eakin, D.; Newey, W. K.; Rosen, H. S. (1988). “Estimating Vector AutoregressionsWith Panel Data”, in *Econometrica*, Vol. 56, p. 1371-1396.

IPEA (Brazilian Research Institute for Applied Economics). (2014). “Linhas de Pobreza Estimadas para Cada Região” – IPEADATA, Rio de Janeiro. Available at: <www.ipeadata.gov.br/>. Accessed in: Jan, 20, 2014.

Kreter, A. C.; Bacha, C. J. C. (2006) “Avaliação da eqüidade da Previdência no meio rural do Brasil.”, in *Rev. Econ. Sociol. Rural*, Vol. 44, N. 3, p. 467-502.

Manson, A.; Lee, R. (2006). “Reform and support systems for the elderly in developing countries: capturing the second demographic dividend”, in *Genus*, Vol. LXII, N. 2, p. 11-35.

Marinho, E.; Araujo, J. (2010). “Pobreza e o Sistema de Seguridade Social Rural no Brasil”, in *Revista Brasileira de Economia*, Vol. 64, N. 2, p. 161-174.

Mesa-Lago, C. (1994). *Changing Social Security in Latin America: Toward Alleviating the Social Costs of Economic Reform*. Lynne Rienner Publishers Inc. 210 pp.

Ludewigs, T.; D'Antona, A. O.; Brondízio, E. S.; Hetrick, S. (2009). "Agrarian Structure and Land Use Change along the Lifespan of Three Colonization Areas in the Brazilian Amazon", in *World Development*, Vol. 37, N. 10, p. 1348-1359.

Ramos, M. P.; Arend, S. C. (2012). "O impacto da reforma da previdência social rural brasileira nos arranjos familiares: uma análise para entender a composição dos domicílios dado o aumento da renda dos idosos", in *Rev. Bras. Estud. Popul*, Vol. 29, N. 1, p. 67-86.

Rosenzweig, M. R. (1988). "Labor markets in low-income countries". In: Chenery H., Srinivasan T. N. (eds) *Handbook of development economics* (Vol. I). North Holland, Chapter 15, p. 713-762.

Schwarzer, H. (2000). "Impactos Socioeconômicos do Sistema de Aposentadorias Rurais no Brasil – Evidências Empíricas de um Estudo de Caso no Estado do Pará.", in *IPEA Working Paper*, N. 729.

Schwarzer, H. (2001). "Paradigmas de Previdência Social Rural: Um Panorama da Experiência Internacional", in *Planejamento e Políticas Públicas*, Vol. 23, p. 119-163.

Simão, R. C. S. (2004). *Distribuição de renda e pobreza no estado de Minas Gerais.* Masters in Economics – Escola Superior de Agricultura Luiz de Queiroz, Universidade de São Paulo, Piracicaba, São Paulo.

Soares, S. S. D. (2006). "Análise de bem-estar e decomposição por fatores da queda na desigualdade entre 1995 e 2004", in *Econômica*, Vol. 8, N. 1, p. 83-115.

Stark, O.; Taylor, E.; Yitzhaki, S. (1986). "Remittances and Inequality", in *The Economic Journal*, Vol. 96, p. 722-740.

Taylor, J. E.; Moran, J.; Adams, R.; Lopez-Feldman, A. (2005). "Remittances, Inequality and Poverty: Evidence from Rural Mexico", in *Agriculture and Resources Working Papers*, N. 05-003.

Turra, C. M.; Queiroz, B. L. (2005). "Transferências intergeracionais: uma análise internacional', in *Revista Brasileira de Estudos de População*, Vol. 22, N. 1, p. 191-194.

VanWey, L. K. (2004). "Altruistic and Contractual Remittances between Male and Female Migrants and Households in Rural Thailand", in *Demography*, Vol. 41, p. 739-756.

Windmeijer, F. (2005). "A finite sample correction for the variance of linear efficient two-step GMM estimators", in *Journal of econometrics*, Vol. 126, N. 1, p. 25-51.

Wooldridge, J. M. (2010). *Econometric analysis of cross section and panel data*. MIT press.

Yang, D.; Choi, H. (2007). "Are remittances insurance? Evidence from rainfall shocks in the Philippines", in *World Bank Economic Review*, Vol. 21, N. 2, p. 219-248.