The Impact of Pensions and Private Transfers on Rural Poverty in Brazil*

El impacto de las pensiones y las transferencias privadas en la pobreza rural en Brasil

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Abstract

Combining different data sources to create a balanced panel of rural state units of analysis, we estimate the impact of pensions (public) and inter-household (private) monetary transfers on the dynamics of rural poverty in Brazil between 1996 and 2015. We combine data from the Brazilian National Household Survey and administrative data from State Statistics Bureaus, in order to estimate a Generalized Method of Moments-System dynamic panel model for poverty. Controlling for demographic composition, GSP (Gross State Product) agricultural share, GSP share to GNP (Gross National Product), educational attainment, unemployment rate, and land concentration, we focus on how pensions and inter-household transfers, as well as their interaction, affected the dynamics of poverty in the rural contemporary Brazil through an increase in the investment capacity

Keywords

Poverty
Dynamic GMM-system panel model
Public transfer
Private transfer
Rural Brazil

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of households. Our results show a significant and positive 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, suggesting that both transfers, even when combined, are limited to fight the structural component of poverty.

Resumen

En este artículo estimamos el impacto de las pensiones (transferencias públicas) y las transferencias monetarias (privadas) entre hogares en la dinámica de la pobreza rural en Brasil entre 1996 y 2015. Para ello, combinamos los datos de la Encuesta Nacional de Hogares y datos administrativos de la Oficina Estatal de Estadística para estimar un modelo de panel dinámico de la pobreza utilizando el método generalizado de los momentos, y controlando por la composición demográfica, la participación agrícola en el producto bruto estadal (GSP), la participación del GSP en el producto nacional bruto (PNB), el nivel educativo, la tasa de desempleo y la concentración de la tierra. Nuestros resultados muestran que hay un impacto significativo y positivo de ambas transferencias en la dinámica de la pobreza, con predominancia de los ingresos por pensión. A pesar de los controles utilizados, la persistencia de la pobreza sigue siendo un problema importante en el Brasil rural contemporáneo, lo que sugiere que los ingresos combinados son limitados para fomentar la capacidad de inversión de los hogares.

Introduction

In the last 25 years the Brazilian economy has undergone impressive economic and social transformations, leading to a significant improvement in well-being. Hyperinflation was eliminated, more individuals gained access to the consumer market (Rocha, 1996), the informal sector has shrunk (Corseuil, Moura y Ramos, 2011), and the real value of the minimum wage has increased (Saboia, 2007). There has also been a 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 the subsidized credit to family agriculture and housing (Araujo & Flores, 2017; Guedes & Araújo, 2009; Januzzi, 2016). The expansion of the Social Security to the rural areas and the right of the rural elderly women to access non-contributory retirement were also an important social instrument of sectorial income redistribution in the last two decades (Alcantara, 2016; Kreter & Bacha, 2006; Valadares & Galiza, 2016). There was an increase in poverty and inequality since 2015 (Rocha, 2019) and unemployment is currently affecting 13.4 million Brazilians (IBGE, 2019). Under this scenario, the role played by the social security system is key to assure minimum living conditions, especially among the rural population.

The current rural pension system encompasses three types of eligible beneficiaries: rural workers, individual contributors, and special insured. The first group contributes financially to the system exactly as their urban counterparts. The second group

comprises those who provide temporary labor and are frequently involved in precarious working conditions. The special insured are those involved in family agriculture, including their partners and children above 16 years old who work under this labor regime. This category represents 99 % of total eligible rural workers in Brazil (Valadares & Galiza, 2016). To be able to retire, a rural worker must be at least 60 years, if a man, and 50 years, if a woman. It is also mandatory to prove 15 years of involvement in rural activities. Since mid-2006, the rural worker is also supposed to contribute 2.1 % on the total gross agricultural revenue, waived for those producing for self-consumption (Stivali, 2017). In Brazil, the rural pension system is the most comprehensive among the developing countries in regard to its coverage and the targeting of the poor (Afonso & Fernandes, 2005; Mesa-Lago, 1994; Zuanazzi, Fochezatto, & Júnior 2018). These features are not intentional as entitlement requirements are not exclusively based on income (Schwarzer, 2000; Valadares & Galiza, 2016). 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 (Oliveira & Aquino, 2017). This institutional and political environment, fueled by the universalization principles brought about by the 1988 Brazilian Constitution, set the basis for the expansion of benefits to all rural households (Zuanazzi et al., 2018).

Empirical evidence suggests that public transfers, specially pensions, and the dynamics of the job market (Medeiros, Souza, & Castro, 2015) were the leading causes of poverty and inequality decline in Brazil since 1995 (Barbosa & Constanzi, 2009; Hoffman, 2010; Soares, 2006). Some qualitative and local studies for rural areas suggest that the impact of public transfers on poverty and inequality is apparent at both the household and municipality levels (Albuquerque, Lobo, & Raimundo, 1999; Augusto & Ribeiro, 2006; Oliveira & Aquino, 2017). However, the long-term dynamics for the country as a whole is largely unknown. The only exceptions are the studies conducted by Marinho & Araujo (2010), Caetano & Monasterio (2014) and Valadares & Galiza (2016). These three studies analyze the link between pension and poverty nationally but use different strategies. Valadares & Galiza (2016) use microlevel data to simulate different scenarios of change in eligibility, coverage and above-inflation adjustment of the rural pension benefit. This is a particularly interesting study as it incorporates the recent debate in Brazil on how the public pension system should change to meet its long-term fiscal sustainability. The first simulation is the most extreme, assuming how the headcount poverty ratio would change if rural pensions were eliminated. According to their findings, poverty would increase from 49.5 % to 67.0 % in 2014. The second simulation unpegs the adjustment of the pension value to the real gains of the minimum wage. In this scenario, poverty would increase to 53.2 %. The third simulation considers that only individuals 65 years and above would be eligible for the receipt of the rural benefit. In this case, poverty would jump to 57.5 per cent.

Caetano & Monasterio (2014) and Marinho & Araujo (2010) look at the link between rural pensions and poverty/inequality indices using macrolevel data. Caetano & Monasterio (2014) found that GNP (gross national product) and the public pension transfers are negatively correlated, as are the GNP and the beneficiary/contributor ratio. The authors point to a redistributive flow from richer, more urban municipalities to the poorest and predominantly rural ones. This result is suggestive of a progressive system, contributing to the reduction of the Brazilian regional inequality. Marinho & Araujo (2010) take a slightly different approach, looking at how pensions impact poverty at the rural state level directly. Their study, however, comprises a shorter period of time (1995 to 2004),

when expansion of the consumer and labor market was still in its first steps. It also lacks some important predictors of poverty dynamics in their econometric specification, such as the contribution of the agricultural sector to the GNP, land concentration, and the percent contribution of the regional gross product to the GNP. Thus, contemporary analysis of the impact of transfers on the rural Brazil is an important empirical question not fully addressed. This becomes even more relevant in times where the increasing demographic pressure on the social security system threatens its ability to maintain itself in the absence of a profound reform, despite its form, extension, and inclusion/exclusion criteria (Zuanazzi et al., 2018).

One of the main caveats that prevents the universalization principle to fully insure eligible beneficiaries are the criteria used by the government to classify a person as special insured. To prove the minimum of fifteen years linked to rural activities one has to provide documents in addition to interviews and witnesses. In a setting where informality prevails, judgement of the validity is subjective, yielding 30 % of the benefits received under judicial claims (Valadares & Galiza, 2016). In addition to the high levels of informality, differences in age structure and how families and the government work to smooth consumption in later stages of the life cycle limit the scope and impact of the public pensions on poverty reduction in Brazil. Miranda (2007) shows that a higher proportion of households in the rural areas receives monetary transfers from other households than their urban counterparts. He also found that monetary transfers decline for families that become eligible for the rural retirement income. This crowding out effect suggests that families and the government work complementary. Moreover, because the age structure of the rural population is younger, many households are not eligible to receive the rural pension income. This demographic feature makes them more dependent on other sources of income, including private monetary transfers (Guedes, Queiroz, & VanWey, 2009; Raad & Guedes, 2015; Turra & Queiroz, 2005).

The international literature on the distributive effect of private transfers in rural populations has been long established (Barham & Boucher, 1998; Stark, Taylor, & Yitzhaki, 1986; Taylor, Moran, Adams, & López-Fieldman, 2005), but has yielded conflicting results. The most convincing theoretical argument and empirical evidence in the literature suggests that monetary transfers are mainly a positive benefit of selective migration (VanWey, Hull, & Guedes, 2013). These transfers, also called remittances, are higher among origin-areas with short-term tradition to outmigration. As outmigration becomes more prevalent, risk declines due to social network returns, and origin-household incomes and remittances become then less positively or even negatively correlated (Stark et al., 1986; VanWey, 2004). Literature on private transfers and poverty is less established, and few empirical studies can be found. Taylor and collaborators (2005) argue that remittances may influence poverty in two possible ways. They might reduce poverty in origin areas by shifting population from low-income rural sectors to higher-income economic sectors through migration. Conversely, they may be inefficient in reducing poverty if migration is risky and costly, which prevents poor households from accessing migrant labor markets. Evidence support the optimistic view that private transfers are efficient in reducing poverty and increase their impact when social networks diffuse, reducing the cost or the risk of migration among the poor (Taylor et al., 2005). Whenever these transfers are invested by the households in the origin areas, their dependence on this particular income source is likely to decline in the long run. However, because of their instability and unpredictability, private transfers seem to have limited ability to fight poverty structurally (Adams, 1996; De Sherbinin et al., 2008; VanWey, 2004; VanWey et al., 2013).

The aim of this study is to estimate the impact of public pensions and inter-household monetary transfers on rural poverty in Brazil. We analyze the evolution of rural-state level wellbeing indicators from 1996 to 2015 using a GMM-system (Generalized Method of Moments) dynamic panel regression model.

Data

To estimate the impact of pensions and inter-household monetary transfers on poverty indicators, we combined microdata from the Brazilian National Household Survey (PNAD) from 1996 to 2015 with 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 aggregate data at the rural-state level from 1996 to 2015. 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[1] and T = 19 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, interhousehold monetary transfers, educational attainment, land concentration indices, and the unemployment rate. From IBGE aggregate data we estimated the state contribution to the Brazilian GNP and the agricultural share of the Gross State Product (GSP). Details of variable construction are given in the next section.

Variables construction

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_o), 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_o) 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 $0 \leq P_o \leq 1$. This is a measure of incidence or extension, not taking into account poverty intensity. Thus, P_o is insensitive to decline of a poor's income (Hoffmann, 2000; Simão,

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1   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.

2004). If income insufficiency is considered as the difference $z-x_i$, with $i \le h$, 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:

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

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_o and I are complementary, the former being insensitive to povety intensity and 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 $\varphi(\alpha)=1$, all individuals have $x_i=0$. Class measure (5) summarizes all above measures, P_o and I. When $\alpha=0$, Equation (5) becomes P_o, while $\alpha=1$ represents $P_o I$. 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_o 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 (Expert, 2006; Hoffmann, 1998). Other desirable properties for axiomatic indices are not met by all the three measures. For instance, P_o 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 measure 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. Here we present results for P_o 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[2] (IPEA, 2018). 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 to 2015. For 2000 we used the average value from 1999 and 2001 for each regional value. For the years 2010 and 2011 to 2015, we used forecasted values from an ARIMA (0,1,1) time series 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.

Independent variables

Our two state variables are the rural retirement income and income received from other households. These are our proxies for public and monetary 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 significantly, we opted for the income measures in the final model. They are more intuitive to interpret and are a direct component of the FGT poverty measures used.

Following the procedure proposed by Marinho & 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 a man and at least 55 years old if a woman. Then, we multiplied the numbers of beneficiaries above identified by the nominal value of the Brazilian minimum salary for each year from 1996 to 2015. 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 & Araujo (2010). Different from the authors, we acknowledge that this measure would not 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.

Individual income received from other households was used as a proxy for monetary private transfer. 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, show significant higher levels of transfers (Campolina Diniz, Gaiger Silveira, Freire Bertasso, De Magalhães, & Mendes Santos Servo, 2007). We could not use POF, however, since it is not representative of rural areas, in addition to having only three points in time available for the period here

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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 change 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 & 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 & 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 over time, 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 cannot be used, because it is 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.

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:

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

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[3]. 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, & 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 term, y_{it-1} *(t ≥ 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 & Bond (1991) suggest using the lagged explanatory variables in level as instruments for the equation in first difference. Blundell & 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 study, we use 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,

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[3]   The within transformation can be used if the available instruments are strictly exogenous; for models in which the strict exogeneity 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 & Bover, 1995). Forward orthogonal transformations consist in subtracting the average of future values of the variable from its current value.

since GMM-system estimators are consistent under two conditions: validity of extra instruments used and absence of serial autocorrelation of residuals (Bludell & Bond, 1998).

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 (α=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:

$$\Delta\{P_{\alpha,it}\} = \beta_1\Delta\{P_{\alpha,it-1}\} + \beta_2\Delta\{ln[PubInc_{it}]\} + \beta_3\Delta\{ln[PrivInc_{it}]\} +$$

$$\beta_4\Delta\{ln[PubInc \; x \; PrivInc_{it}]\} + \beta_5\Delta\{ln[Mehran_{it}]\} + \beta_6\Delta\{ln[\%AgrGSP_{it}]\} +$$

$$\beta_7\Delta\left\{ln\left[\% \frac{GSP}{GNP_{it}}\right]\right\} + \beta_8\Delta\{ln[MehranLand_{it}]\} + \beta_9\Delta\{UnempRate_{it}\} +$$

$$\beta_{10}\Delta\{\%Persons_{it}^{15-64}\} + \beta_{10}\Delta\{\%Persons_{it}^{65+}\} + \Delta\{\varepsilon_{it}\}$$

where $\Delta\{P_{\alpha,it}\}=P_{\alpha,it} -P_{\alpha,it-1}$. Because $E\{\Delta\{P_{\alpha,it-1}\},\Delta\{\varepsilon_{it}\}\}\neq0$, Ordinary Least Square estimators would be biased and inconsistent. Thus, instruments for $\Delta\{P_{\alpha,it-1}\}$ must be used. Assuming the moment conditions $E\{\Delta\{P_{\alpha,it-s}\},\Delta\{\varepsilon_{it}\}\}=0$ for $t=3,4,...,T$, and $s\geq2$ good instruments for Equation (7) would be $\Delta\{P_{\alpha,it-s}\}$ for $t=3,4,...,T$ and $s\geq2$, as suggested by Arellano & 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 & Bover, 1995; Blundell & 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 2015). 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 2.1 % on the total gross agricultural revenue. This rule continued to be waived for those producing for self-consumption. 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. Although econometrically sound, this endogeneity is likely to be virtually irrelevant for the following two reasons: firstly, the commercial production from family agriculture is highly informal, and secondly, Brazil lacks an effective enforcement mechanism to guarantee the totality of tax collection. 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_o FGT poverty measure with increasing consistency: Ordinary Least Square, Fixed Effect, and GMM-System with forward orthogonal deviations.
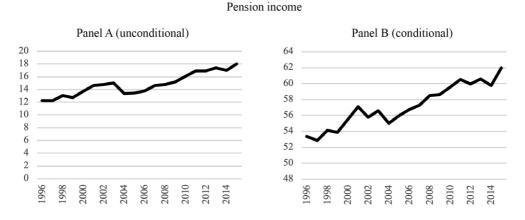
Results

Descriptive results

Pensions are a more important income source for rural than for urban households in Brazil. In 2015, 32 % of rural households received retirement income against 26 % in urban areas. The importance of this income to the overall household income was also higher in rural (18 %) than in urban areas (13 %). If one considers only those households receiving the benefit, the share of pensions represented 62 % and 55 % in rural and urban households, respectively. The proportion of households receiving income from non-coresidents (monetary private transfers), however, is similar across sites (about 2 %), as it is its importance to the overall household income (0.7 % and 0.8 % in rural and urban areas, respectively). Interestingly, the economic importance of monetary private transfers among those receiving any positive value was 35 % for rural and 45 % for urban households, reflecting the crowding out effect already identified in previous studies (Miranda, 2007). Figure 1 presents the evolution of the income share on total household income from pensions and inter-household transfers among rural residents from 1996 to 2015. Panels A and B show the unconditional and conditional share of the income sources, respectively. Pensions increased their share on total income over time, while the non-coresident transfers declined its relative importance. Moreover, the private monetary transfers had a very erratic trend over the years – a pattern commonly reported in other settings.

As discussed in a previous section, our analytical panel sample comprises 380 observations for the level dataset and 340 for the instrumented first difference dataset. In total, 20 rural-state units over 19 years (1996 to 2015) 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 2015) 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

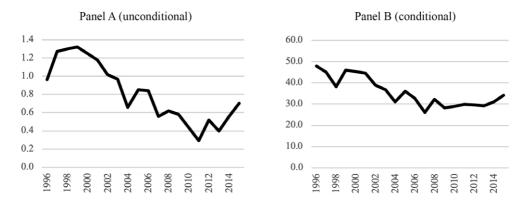considerably high, especially after 2003 (line trends available upon authors' request). As previously pointed, the per capita retirement income (expressed in 1996 R$) is contaminated by differences in state-year age structure. This being said, variation is balanced from between states and within state 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 2015). 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 (also shown in Figure 1), with the largest share of variation being accounted by the within state differences over time.

Figure 1
Evolution of the share represented by pensions and inter-household transfers on total household income among rural residents – Brazil, 1996 to 2015

Pension income

Panel A (unconditional)   Panel B (conditional)

Monetary inter-household income

Panel A (unconditional)   Panel B (conditional)

Source: PNAD 1996-2015 (IBGE); State-level Statistics (IBGE).

Table 1

Descriptive statistics for variables used in the dynamic panel models of rural poverty in Brazil, 1996 to 2015

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	0.60	0.05	0.49	0.80
between		0.04	0.55	0.71
within		0.03	0.52	0.74
Mehran land inequality index				
overall	0.95	0.06	0.60	1.00
between		0.03	0.88	0.99
within		0.05	0.64	1.08
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 individuals 15 to 64 years old				
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 individuals 65 and over years old				
overall	0.10	0.02	0.03	0.19
between		0.02	0.05	0.14
within		0.02	0.07	0.15

Note: N = 380 / n = 20 / T = 19. North states excluded.
Source: PNAD 1996-2015 (IBGE); State-level Statistics (IBGE).

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).

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. 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 and positive 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 the retirement 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 considered 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 > chi² = 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).

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. 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.

Table 2
*Determinants of poverty dynamics in rural Brazil
from 1996 to 2015 - Headcount ratio*

Variable	OLS	Fixed effect	GMM-System
Lagged headcount ratio	0.905***	0.542***	0.544***
	(0.036)	(0.049)	(0.045)
Ln of public retirement income	-1.779	-13.590***	-16.650**
	(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	-16.500***	-34.860***	-41.740***
	(5.988)	(11.430)	(12.740)
Ln of Mehran land concentration index	0.421	2.875	4.791
	(3.117)	(3.267)	(4.875)
Unemployment rate	0.078	-0.183	-0.199
	(0.170)	(0.138)	(0.198)
% of 15-64 years old individuals	-47.950**	-22.650	-18.100
	(20.500)	(16.900)	(20.810)
% of 65+ years old individuals	-43.140*	-56.010**	-35.060
	(24.730)	(20.150)	(27.080)
Constant	54.180***		
	(18.310)		
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	

Note: Robust standard errors in parentheses. North states excluded.
*** p<0.01, ** p<0.05, * p<0.1.*
Source: PNAD - 1996/2015.

Conclusion

This study estimated the impact of public pensions and inter-household monetary transfers on rural poverty in Brazil from 1996 to 2015. We found that poverty declines significantly among rural areas receiving higher levels of aggregate pension income. A 1% increase in the per capita pension income would lead to an average decline in poverty ratio of about 16% if no private transfers were observed. Our results also suggest a relevant role played by the monetary private transfers, even though its effect size is approximately half of the effect estimated for public pensions.

Using the same econometric model and similar data, 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. Our findings, compared to the authors', may reflect a variety of differences: 1) a larger panel dataset, with our data covering a period of intense economic growth (after 2004), 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, 4) inclusion of monetary private transfers and the leverage effect on public transfers, and finally 5) 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 the monetary impact and the age structure influence on poverty.

Existing literature seems to agree about the positive effect of the rural retirement on inequality decline (Hoffmann, 2010; Soares, 2006), although there is mixed evidence about its impact on poverty (Barrientos, 2003; Hoffmann, 2006; Marinho & Araujo, 2010; Schwarzer, 2000). 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 than the "Fundo de Participação dos Municípios" in 92 of the 100 municipalities with the highest Human Development Index. Afonso & Fernandes (2005) also found a very strong impact of public pensions on poverty, with the highest estimated Internal Return Rate for benefits in areas where the Rural Retirement was more important, such as in the rural areas of the North and Northeast regions of Brazil. The relevance of the social security system to the local economies will further increase with the rapid population aging under course in the country (Ansiliero & Paiva, 2008; Cuevas, Karpowicz, Mulas-Granados, & Soto 2017).

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 not only in food, but also on house improvement and private health products and services. Augusto & 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) and Lima & Braga (2016) 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. A more recent study from Oliveira & Aquino (2017)

points to the use of these transfers on the reduction of family debts in rural areas. Collectively, these case studies show how heterogeneous public pensions are used across the rural areas of Brazil.

The rural retirement income itself provides an opportunity to analyze an income shock that is tightly connected to the Brazilian economic growth in the period. 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 since 1995 (Hoffmann, 2010; Soares, 2006). Barros, Corseuil, & Curry (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 its purchase power over the retirement income, reflected by a strong positive multiplier effect on the economy (Barros et al., 2001; Barros, Carvalho, Franco, & Mendoça, 2007). This simultaneous relation between the value of the minimum salary and poverty indexes is thus clearly mediated by the non-contributory portion of the system, which does not depend on the labor market. This is the main explanation why the social security system functions as a powerful redistributive mechanism, with even stronger effects in its rural segment (Afonso & Fernandes, 2005; Barros et al., 2001; Schwarzer, 2001).

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 are not fully established (Ludewigs, D'Antona, Brondízio, & Hetrick, 2009). Variable income sources, such as monetary private transfers, are generally spent on immediate needs, while stable and lasting income sources, that include public pensions, might be channeled towards productive and permanent investments (Brown, 2006; Oliveira & Aquino, 2017). 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. Indeed, we found a reinforcing effect of poverty reduction represented by a significant interaction between public pensions and monetary private income. We also found a persistent effect of poverty over time, even after accounting for both types of transfers. This result is suggestive of more structural components of rural poverty in Brazil.

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 (Kreter & Bacha, 2006; Schwarzer, 2000; Stivali, 2017). 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 & Arend, 2012).

Because of the rapid population aging under course and the reduction in the ability of urban areas to continue to finance the rural pension system, some studies have been investigating alternatives to preserve wellbeing with a more fiscally sustainable scheme in the long run. The criteria to be used, however, is still open to debate.

Some authors give priority to the demographic aspects contributing to a deficit in the pension system. Stivali (2017) argues that a single minimum age at retirement should be applied, regardless of gender, place of residence and other sociodemographic attributes of eligible beneficiaries. Even though life expectancy at birth differs across sociodemographic groups, the conditional life expectancy at the modal age of retirement is less variable. What is missing in this particular claim is that individual and regional socioeconomic heterogeneities lead to asymmetric contributory capacity and different trajectories in the labor market (Valadares & Galiza, 2016).

The rural population is an especially vulnerable group, as household income is usually unstable and their members tend to be involved in vulnerable occupational activities such as sharecropping, temporary employment, and family agriculture. Data from 2014 reveal that more than 70 % of individuals in rural areas started working before completing 15 years of age, while in urban areas this figure is lower than 50 % (Valadares & Galiza, 2016). In addition, two thirds of those working on agriculture activities were involved in family agriculture and only 40 % had formal labor contracts. This explains the low contributory capacity of rural areas and, consequently, the low coverage rates (Valadares & Galiza, 2016). This situation is not exclusive to the Brazilian society as most developing countries face similar challenges related to their rural population (Barrientos, 2003; Mesa-Lago, 1994).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, Moran, Adams, & López-Feldman, 2005).

Despite the strong effects found in this study, we acknowledge that the proxy used for the public pension income may underestimate the number of individuals covered by the Rural Retirement System, since there can be rural residents contributing to the General Social Security System or receiving more than one minimum salary. Therefore, it is impossible to identify the exact number of persons receiving the specific income using PNAD data. Valadares & Galiza (2016) showed that in 2014 PNAD would identify 4.1 million beneficiaries against 9 million according to the official data from the Social Security system. The difference between the survey and the administrative data is explained by two factors. First, the Law 8,212 contemplates individuals living in urban agglomerates but who work in agriculture activities. Second and most important, the eligibility criterion is the link to rural activities and not the place of residence, as used to find those eligible from PNAD data. The second limitation of our study is the lack of non-monetary private transfers on our models. Many studies in rural areas worldwide found that other types of transfers, including visits, food, help with days of work, and medicines, are more often reported as private flows of resources received by rural households (Bartolome & Vosti, 1995; Hull & Guedes, 2013; VanWey, 2004). This evidence is also true for many rural households in Brazil (Guedes et al., 2009; VanWey & Cebulko, 2007). Unfortunately, there is no longitudinal data representative to all rural areas in Brazil available. Because of these limitations we argue that the impacts found represent a lower bound on the ability of pensions and private transfers to fight rural poverty in Brazil.

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