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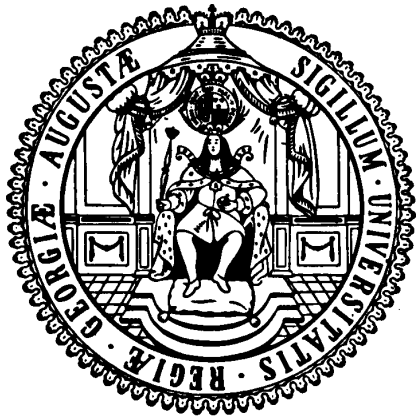

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Macroeconomic Performance and Inequality: Brazil 1983-1994

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Macroeconomic Performance and Inequality: Brazil 1983-1994*

Manoel Bittencourt†

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Abstract

We examine how macroeconomic performance, mainly in the role of high rates of inflation, affected earnings inequality in the 1980s and early 1990s in Brazil. The results—based initially on national time-series, and then on the relatively novel sub-national panel time-series $T \succ N$ data and analysis—show that the extreme inflation, combined with the incomplete indexation coverage seen at the time, had a regressive and significant impact on inequality. Thus, sound macroeconomic policies, which keep inflation low and stable in the long run, are to be a necessary first step of any policy package implemented to alleviate inequality in Brazil.

Keywords: Brazil, inequality, inflation, indexation.

JEL Classification: D31, E31, O11, O54.

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1 Introduction and Motivation

We examine the impact that macroeconomic performance, mainly in the role of high rates of inflation, had on earnings inequality in the 1980s and early 1990s in Brazil. The importance of this subject is, firstly, for the distinguishing features in terms of poor macroeconomic performance and high inequality, which are important not only for Brazil, but also for other developing countries that presented similar poor economic conditions during roughly the same period of time[1]. Secondly, the subject is equally important for emerging developing countries, some of which still present poor macroeconomic performance to this day. Thirdly, the link between macroeconomic performance and inequality in Brazil has been markedly different from the one seen in developed countries.

For the analysis we use a data set that combines a fairly long time series T with a shorter panel N variation, and which presents novel and interesting features in terms of estimation. Firstly, time-series data might well be non-stationary, and therefore the issue of testing for unit roots in panels is theoretically relevant for specification and estimation purposes. Secondly, there is the question of having heterogeneous dynamic panels. The treatment of heterogeneity is one of the central questions in panel time-series $T \succ N$ analysis, for in its presence the estimates might be biased. Thirdly, there is the possible occurrence of between-region dependence in the data. This is an important matter that, if not taken into account, can lead to the situation of getting little gain in using panel estimators instead of different time series for each region.

All in all, the evidence shows that chronic extreme inflation had a regressive impact on inequality. The high-inflationary environment seen in Brazil at the time had a significant and positive effect on the Coefficient of Variation and Gini coefficient, and a negative one on the shares of the first four quintiles of the earnings distribution. However, these results are in contrast with the evidence from developed countries, where the subject has, in fact, attracted consistent attention.

The first wave of studies on, e.g. the US, covers the post-war period until the early 1970s. Schultz (1969), Metcalf (1969), Thurow (1970), Beach (1977) and Blinder and Esaki (1978), employing a range of methods based

[1]For instance, Bolivia, Colombia, Mexico, Peru and Tanzania. See Bulíř (2001).

on aggregate national time-series data, report that inflation had small, but not always statistically significant, progressive effects on inequality[2]. However, Metcalf (1969) and Thurow (1970) also suggest that those groups more reliant on imperfectly-indexed public transfers—families with a female head and poor blacks respectively—are more prone to lose with higher inflation. A second wave of studies that incorporates data from the 1980s includes Blank and Blinder (1986), and Cutler and Katz (1991). Their results confirm the previous studies, but with smaller and less precise inflation effects on inequality. More recently, and with data from the 1990s, Romer and Romer (1999) report that inflation remains progressive on inequality in the US[3]. Thus, it is fair to say that in *developed* countries inflation is believed to be progressive through the debtor and creditor channel, with the poor being the debt holders, and therefore the main beneficiaries of *moderate* rates of inflation that keep their non-indexed debts fixed in the short run.

On the other hand, Brazil has been known for its high inequality, e.g. Gini coefficients of .623 and .601 in 1976 and 1995 respectively, and also for its chronic high rates of inflation, especially in the 1980s and early 1990s. For the latter, this paper covers a particularly traumatic period in which Brazil experienced not only high inflation, but also bursts of hyperinflation in the late 1980s and early 1990s, and again in the middle of the 1990s. The subject of inequality and inflation has been often anecdotally debated, however, given the lack of a constant and reliable stream of data until late 1970s—the literature on Brazil is, not surprisingly—thin and relatively recent.

Kane and Morisett (1993) report that the shares of the four lowest quintiles of the income distribution were regressively affected by inflation in the 1980s in Brazil. Cardoso et al. (1995) also investigate the impact of inflation on inequality in the 1980s. Employing time series from metropolitan regions they find that inflation has significant effects in raising inequality in each region separately. Barros et al. (2000) pool time series with regional information from 1982 to 1998 and consider the existence of fixed effects across regions. Their findings confirm the ones contained in the previous studies, with or without the presence of regional fixed effects. Also using data

[2]Schultz (1969) also makes use of Dutch data covering roughly the same period, and he reports the same sort of qualitative results as for the US.

[3]Complementary to the above, Nolan (1987) uses UK data covering the 1960s and 1970s. He reports that over time the shares of the top two quintiles decrease relative to the shares of the first and third quintiles of the income distribution when inflation rises.

from the 1980s, Ferreira and Litchfield (2001) estimate an aggregate time series. They too report regressive effects of inflation on inequality. Therefore, these studies on Brazil indicate that, differently from what happens in developed countries, inflation presents regressive effects on inequality, with inflation being regressive for its high rates that—among other distortions caused—offset the debtor and creditor channel. More specifically:

- first, in an economy that presents and requires either cash-in-advance constraints or different shopping-time allocations for the consumption of a certain bundle of goods—e.g. Lucas and Stokey (1987), Sturzenegger (1992), Erosa and Ventura (2002), and Cysne et al. (2005)—the existence of inflation acts as a tax on consumption of goods requiring cash, therefore leading people to substitute consumption of cash for goods requiring financial or indexed assets. All the same, with this process of financial adaptation, the rich are able to hold interest-bearing assets, credit and foreign assets, i.e. currencies that are not affected by the inflation tax. On the other hand, the poorer are financial-assets constrained—with little or no access to, e.g. simple but *indexed* bank accounts, not to mention more sophisticated financial assets—and therefore having to hold the highly-taxed cash instead[4].

- second, imperfect wage indexation due to lower bargaining power by the poor, since in the Brazilian formal labour market indexation during the high-inflation period was a function of wage levels, with higher wages being overindexed and the lower ones severely underindexed[5].

- third, Kane and Morisett (1993), Crowe (2006) and Albanesi (2007) highlight the political-economy channel of high rates of inflation and inequality, i.e. that macroeconomic stabilisation took so long to be implemented in non-democratic societies like Brazil at the time because the rich had always benefited from high inflation, and therefore they would not lobby for stabilisation. Coincidentally enough, stabilisation came in 1994-1995 only after full democratisation took place in 1989.

[4] Beck, T., A. Demirguc-Kunt, et al. (2007) document that the ratio private credit over GDP in the US and Brazil during the period 1960-1999 was of .944 and .272 respectively.

[5] See Agénor and Montiel (1999) for more on wage-contract indexation in Brazil and also other developing countries during their high-inflation periods. Also, Cardoso (1992) documents the imperfect wage-indexation mechanism in Latin America between 1977 and 1989 and its regressive consequences on poverty.

That said, this paper distinguishes itself from the previous studies for some important reasons. First, it fills in a gap in this literature on Brazil—concentrating its attention on the high-inflation period and hence avoiding the contamination from a different economic regime—which can be mirrored not only to other developing countries that presented similar poor macroeconomic conditions at the time, but also to emerging developing countries that still do not present credible anti-inflationary institutions[6]. Second, it extends the specifications previously estimated not only with an important and much debated policy variable not included before, the minimum wage, but also with a different concept of inflation, past inflation or anticipated inflation. Third, it makes use of both the time-series and panel variations present in the data. More fundamentally, it takes advantage of the relatively novel panel time-series analysis that deals with new empirical issues—which is a significant step forward compared to previous studies in terms of estimation—and therefore it is believed that better and more insightful estimates are reported[7].

The remainder of the paper has the following structure: Section 2 deals with the data set used. Firstly it explains how the variables are obtained and provides some descriptive statistics of the data, and secondly it shows some stylised facts in the data. In Section 3 we discuss the empirical strategy and also present and discuss the results. Finally, Section 4 concludes. It summarises the evidence, highlights the differences between developed and developing countries on the subject, suggests extensions and raises policy implications that arise from the empirical results.

2 The Data

2.1 Data Description

The data set comes from the Brazilian Institute of Geography and Statistics (IBGE), which is the Brazilian Census Bureau, and also from the Institute of Applied Economic Research (IPEA) files. The IBGE is the most impor-

[6]For instance, the Reserve Bank of Zimbabwe has been recently plagued by political interference, which has resulted in high rates of inflation and adverse consequences on economic welfare. See Muñoz (2007) or Coorey, S., J. Clausen, et al. (2007).

[7]For instance, Barros et al. (2000) use pooled data and analysis. However, they do not deal with non-stationarity, nor with a possible heterogeneity bias present in their dynamic models, nor with the possibility of cross-region dependence in panels.

tant institution for data collection and dissemination, and is the body that covers the Brazilian territory most thoroughly. The IPEA is an agency of the Brazilian government that, among other things, compiles primary and provides secondary data coming from the IBGE itself and also other national sources.

The data on earnings come from the Monthly Employment Survey (PME) files produced by the IBGE—which is a monthly rotative survey that follows ILO recommendations for international comparability—and that covers six regions over time and approximately 38,500 households. The six regions are, from north to south: Pernambuco (PE), Bahia (BA), Minas Gerais (MG), Rio de Janeiro (RJ), São Paulo (SP) and Rio Grande do Sul (RS). The concept of before-tax earnings adopted by the PME includes wages, monetary bonuses and fringe benefits earned by those at work, profits made by those who are self employed and employers, and the monetary value of goods for those earning in kind. Hence, this concept of earnings is broader and less restrictive than what is usually understood by more conservative definitions of earnings.

In a country which presented high rates of inflation for such a long period of time the way the data are deflated is rather important. The earnings data are deflated by the IBGE's National Index of Consumer Prices (INPC). One important prior adjustment is the use of a converter to express all data in Real (R$) mainly because Brazil had many monetary reforms attempting to tackle high inflation, especially between 1986 and 1994. Some adjustments in the INPC itself are also implemented. These include a correction of 22.25 percent for the inflation incurred in June 1994, a month before the full implementation of the R$. The reason is that the INPC calculated inflation using the price variations of a virtual, but not fully implemented R$, which was lower than the price variation incurred by the still widely used Cruzeiro (CR$).

Another important adjustment is the need to centre the INPC as if it was measuring inflation starting on the first day of each month, which is the date that most people get their paycheques. Hence, taking into consideration that the information on earnings reported in the questionnaires of the PME is related to the first day of a particular reference month t, earnings are in fact corrected by the deflator of month $t + 1$ to allow the inflation incurred in t to be accounted for. All in all, these corrections are particularly

important because otherwise, in a country with such high rates of inflation, the information on earnings would be severely distorted by inflation and the computed estimates not as reliable[8].

Given that, we use the information of individual earnings from people between fifteen and sixty five years of age to obtain the Coefficient of Variation (CV), the Gini coefficient ($GINI$) and the respective shares of each quintile of the earnings distribution (Q_i). These measures of inequality are used for their attractive properties. The Coefficient of Variation and the Gini coefficient are simultaneously consistent with the Anonymity, Population, Relative Income and Dalton principles, and are therefore Lorenz consistent. Furthermore, according to the Relative Income principle, the earnings shares are sufficient to measure inequality[9].

Regarding the information on the rates of inflation ($INFL$), we use the variation in the IBGE's regional Consumer Price Indexes (IPCs). A second concept of inflation used is the past or anticipated inflation ($PASTINFL$), which consists of a four-month average of the rates of inflation measured by the regional IPCs.

The unemployment rates ($UNEMP$) used as a control variable also come from the PME files. Unemployment is calculated by the IBGE following the method of the number of people unemployed and who are currently looking for employment over the labour force, who are at least fifteen years old.

The regional minimum-wage Kaitz index ($MINWAGE$) is the national minimum wage divided by the average earnings of each region covered by the PME. The minimum wage data are from the IPEA files and deflated by the INPC.

Table 1 provides the descriptive statistics using the national time-series variation in the data, and also the correlations between the inequality measures and the rates of inflation in Brazil. It is worth mentioning the high means of the Coefficient of Variation and Gini coefficient (1.642 and .548 respectively)—with both measures reaching their maximum values in August 1990 and January 1989 respectively—and of the inflation rates, on average 18.46 percent *per month*, during the period in the first half of the Table.

[8] See Corseuil and Foguel (2002) for more details on how best to deflate earnings and income data from Brazil.

[9] For more on inequality measures and their properties, see Sen and Foster (1997).

No less important is the fact that the richest twenty percent ($Q5$) of those in the sample appropriate, on average, an astounding 43 percent of the total earnings—reaching its maximum in November 1989, and the poorest forty percent ($Q12$) appropriate a mere 18 percent of the total earnings—reaching its minimum in December 1989. Considering that the rates of inflation reached the maximum of 82 percent per month in March 1990, it can be initially said that inequality deteriorated considerably during the first burst of hyperinflation.

Additionally, in the second half of the Table we can see the positive correlations between the Coefficient of Variation and the Gini coefficient with inflation. Also important to mention are the negative correlations between the shares of the first four quintiles ($Q12$ and $Q34$) of the earnings distribution with inflation and, in contrast, the positive correlation between the shares of the fifth quintile of the distribution with the very same rates of inflation. Most correlations are statistically significant at either the 5 or 10 percent level.

Table 1: Descriptive Statistics and the Correlation Matrix, Brazil 1983-1994.

Variables	Observations	Mean	Std. Dev.	Min	Max
CV	144	1.642	.211	1.277	2.984
GINI	144	.548	.016	.510	.609
Q12	144	.181	.010	.157	.211
Q34	144	.392	.011	.325	.409
Q5	144	.428	.019	.396	.521
INFL	144	18.466	14.065	.430	82.180
UNEMP	144	5.220	1.420	2.540	9.770
MINWAGE	144	206.700	42.820	115.030	321.500

Correlations	CV	GINI	Q12	Q34	Q5	INFL
CV	1					
GINI	.657*	1				
Q12	-.157**	-.698*	1			
Q34	-.298*	-.341*	.235*	1		
Q5	.289*	.618*	-.754*	-.080*	1	
INFL	.270*	.276*	-.091	-.304*	.271*	1

Source: PME, IPC, IBGE, IPEA and author's own calculations. * significant at the 5 percent level. ** significant at the 10 percent level.

2.2 Behaviour of the Variables

The behaviour of the rates of inflation in Brazil was notoriously erratic in the 1980s and first half of the 1990s. The rates of inflation cover a range that goes from a rate of virtually zero per cent, .43 percent in April 1986, up to around 80 percent, 82.18 percent in March 1990 *per month*. For example, the accumulated inflation rate during the period between January 1983 and December 1994 is a staggering 2,659 percent, with an average of 18.46 percent per month. To illustrate it further, the annual rate of inflation in 1989 alone was 1,863 percent.

Figure 1 illustrates, using the national time-series variation in the data, some important inflationary events that took place during the period. It shows the period of relatively low inflation after the implementation of the Cruzado Plan in February 1986[10]—nine months before regional elections

[10] See Agénor and Montiel (1999) for more on this plan.

took place—and the hyperinflationary period that happened by the years of 1989-1990 when inflation reached its peak of around 80 percent per month, and then the sudden, but not durable, drop due to the Collor Plan[11]. Another particular feature is the rising inflation, especially from 1991 onwards, which culminated with the implementation of the Real Plan in 1994[12]. The duration of the price stabilisation after those stabilisation plans is also significant. The drop due to the Real Plan has been not only much deeper, but also more durable than any other before, and the behaviour of inflation has actually been relatively low and stable in Brazil since then.

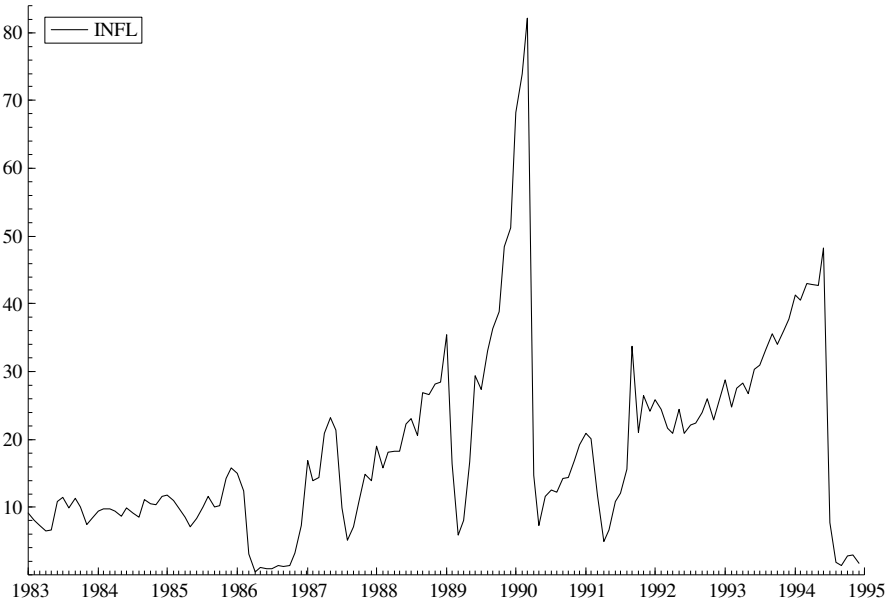

Figure 1: Monthly Inflation Rates in Brazil, 1983-1994. Source: IPC, IBGE. $INFL$ is the inflation rates.

[11] The ill-fated stabilisation plan implemented by the then newly elected President Fernando Collor, and which literally confiscated most financial assets held by the public. See Kiguel and Liviatan (1992) for more on this plan.

[12] The Real Plan was gradually implemented during the first half of 1994 and the Real (R$) itself implemented in July 1994. See Agénor and Montiel (1999) for a treatment of the Real Plan.

Regarding the behaviour of the Coefficient of Variation and Gini coefficient of the earnings distribution combined with inflation, the main feature in the data is that both inequality measures markedly increased during the hyperinflationary bursts. For instance, both measures of inequality presented increases of 43.71 and 9.19 percent between January 1988 and August 1990, and June 1988 and January 1989 respectively. The effects are symmetric though, i.e. when the hyperinflationary periods come to an end inequality also returns to its previous figures. Figure 2 illustrates the above.

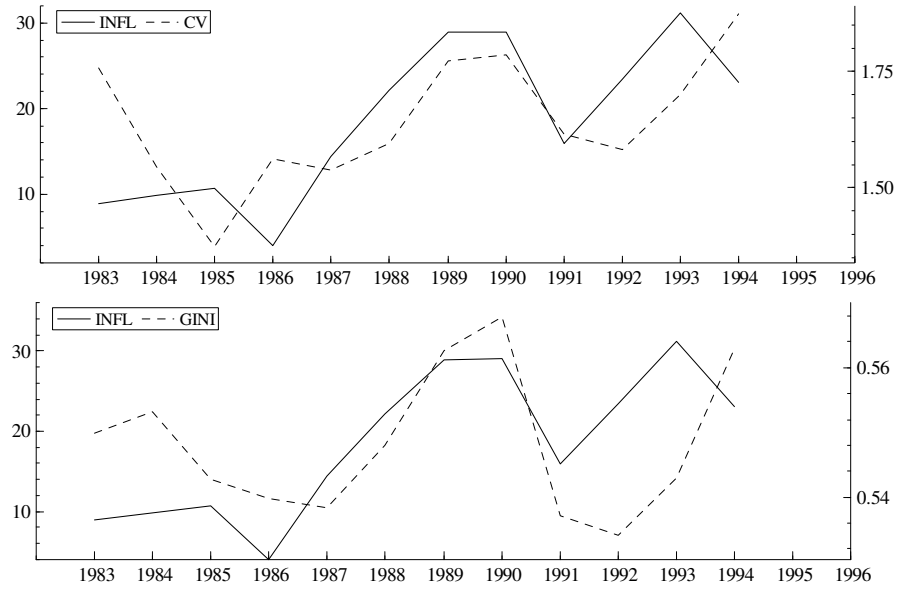

Figure 2: Annual Averages of Monthly Inflation and Inequality in Brazil, 1983-1994. Source: PME, IPC, IBGE and author's own calculations. The measures of inequality are the Coefficient of Variation (CV) and the Gini coefficient ($GINI$), and $INFL$ is the inflation rates.

Moreover, when we plot the shares of the earnings of the low-middle ($Q23$) and top fifth ($Q5$) quintiles against inflation, the data show that during the hyperinflationary peak of 1989-1990 the shares of the earnings of the poor and middle classes fell markedly. For example, the decrease between July 1988 and November 1989 was 24.28 percent. However, after this hyperinflationary peak there was a considerable recovery, to their previous figures at least, in the earnings shares of the second and third quintiles. With respect to the earnings of the top fifth quintile, its share increased significantly during the hyperinflation of 1989-1990 and then dropped when inflation fell. In this case, the increase between April 1988 and November 1989 was 26.61 percent. Figure 3 illustrates the above.

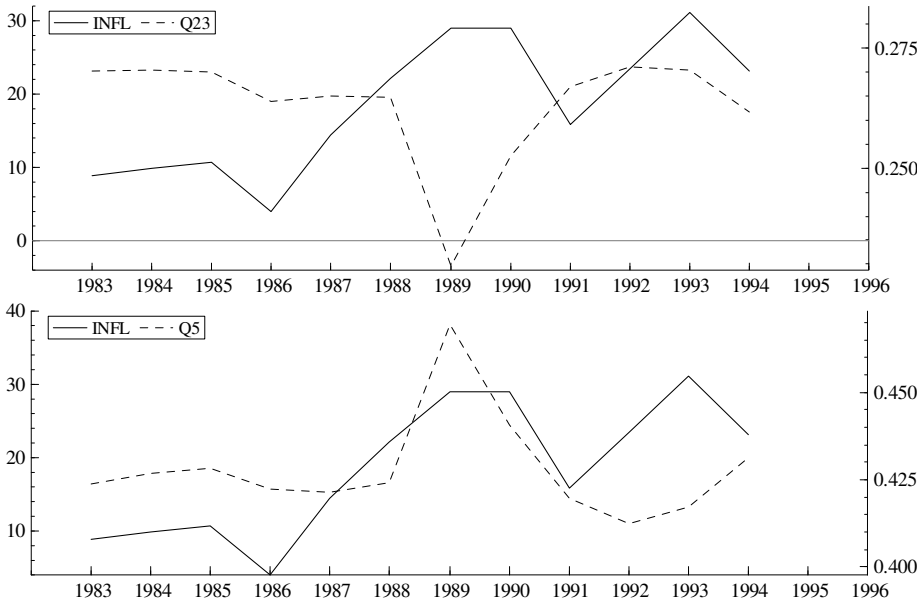

Figure 3: Annual Averages of Monthly Inflation and Inequality in Brazil, 1983-1994. Source: PME, IPC, IBGE and author's own calculations. The measures of inequality are the shares of the second and third quintiles ($Q23$) and the fifth quintile ($Q5$) of the earnings distribution, and $INFL$ is the inflation rates.

To enhance this initial inspection of the data, we plot the OLS regression lines between the measures of inequality and inflation. The Coefficient of Variation, the Gini and the shares of earnings of the fifth quintile of the distribution, as expected by now, display positive relationships with inflation. On the other hand, the shares of the second and third quintiles present a negative relationship with inflation. Figure 4 illustrates the results.

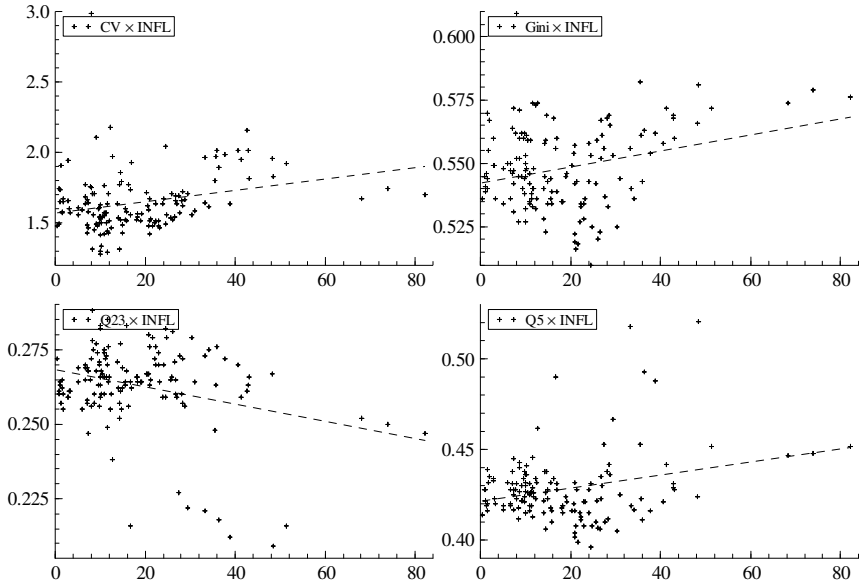

Figure 4: OLS Regression Lines, Inequality and Inflation, Brazil 1983-1994. Source: PME, IPC, IBGE and author's own calculations. The estimated equation is $INEQUALITY_t = \alpha + INFL_t + u_t$. The measures of inequality are the Coefficient of Variation (CV), the Gini coefficient ($GINI$), and the shares of the second and third quintiles ($Q23$) and the fifth quintile ($Q5$) of the earnings distribution, and $INFL$ is the inflation rates. All estimates are statistically significant at the 1% level.

Thus, what can be drawn from the above preliminary descriptive and visual evidence from the time-series variation in the national data is that high inflation widened the earnings distribution during the period. Moreover, the inequality measures clearly presented the ability to decrease to their previous figures when inflation fell, inequality and inflation have literally moved together at the time, which suggests that low and stable rates of inflation at least do not have a regressive effect on inequality[13].

3 Empirical Strategy and Results

In this Section we use the sub-national $T \succ N$ data to estimate the impact of macroeconomic performance on inequality, and also discuss some important issues present in panel time-series analysis. We then report and discuss the results.

Firstly, the centred twelve-point moving averages are used to deal with any possible seasonality and to smooth the irregular component in the series. These transformed data have information from January 1983 to December 1994 ($T = 132$) covering six regions of Brazil ($N = 6$). Secondly, for non-stationarity in the regional time series we use the Im, Pesaran and Shin [IPS (2003)] test, which allows for heterogeneous parameters and serial correlation[14]. Thirdly, the issue of heterogeneity bias in dynamic $T > N$ panels—caused for under wrongly assumed homogeneity of the slopes the composite disturbance term is serially correlated and the explanatory variables x_s are not independent of the lagged variable y_{t-1}—is dealt with Swamy's (1970) Random Coefficients (RC-GLS) estimator, which gives consistent estimates of the expected values. Moreover, the one-way Fixed Effects (FE) estimator also provides consistent estimates in dynamic models when $T \to \infty$, but only when the slopes are homogeneous[15]. Fourthly, since the data present

[13] Bulíř (2001) suggests that there is a "free lunch", i.e. that there are no disinflation costs on inequality, but only benefits. Furthermore, Easterly and Fischer (2001) document that the poor from thirty-eight countries in 1995 considered inflation as a more pressing macroeconomic problem than their richer counterparts.

[14] An alternative to IPS (2003) is the test by Levin, Lin and Chu (2002). However, this test assumes parameter *homogeneity*, and therefore does not consider a possible heterogeneity bias present in the data.

[15] When heterogeneous slopes are present, the Mean Group (MG) estimator, proposed by Pesaran and Smith (1995) is also an alternative, however it is sensitive to outliers, a problem not faced by the RC-GLS estimator. A second alternative would be the Instrumental Variable estimator, however an instrument uncorrelated with the residuals will be

$T > N$ variation, between-region dependence is believed to be through the disturbances, i.e. $E(u_{it}u_{jt}) \neq 0$. This is accounted for with Zellner's (1962) Seemingly Unrelated Regressions (SUR-FGLS) estimator[16] [17].

Given the above, the IPS test for unit roots is based on an Augmented Dickey-Fuller (ADF) regression for each region of each variable, which are then averaged. The moments of the mean E and variance var of the average $\bar{t}$ to be plugged into the IPS test are taken from IPS (2003) and in this case are -1.504 and .683 respectively. Equations 1 and 2 illustrate the regional ADF equations of a particular variable y and the IPS test, respectively.

$$\Delta y_{it} = \alpha_i + \beta_i y_{it-1} + \sum_{j=1}^{k} \gamma_{ij} \Delta y_{i,t-j} + u_{it}, \tag{1}$$

$$IPS = \frac{\sqrt{N(\bar{t} - E(\bar{t}))}}{\sqrt{var(\bar{t})}}, \tag{2}$$

where N accounts for the number of regions. The IPS statistics suggest that we can reject the null hypothesis of unit roots in all variables and accept in favour of the alternative that at least *one* region of each variable is stationary at the 5 percent level. Table 2 reports the results.

uncorrelated with the explanatory variable, and hence not a valid instrument. Finally, GMM-type estimators are not an alternative under $T \succ N$ for the overfitting problem. For more, see Pesaran and Smith (1995) or Bond (2002).

[16] An alternative to SUR-FGLS is the Common Effects Estimator proposed by Pesaran (2006). However, N is assumed to be large and in our data set *N=6*. Furthemore, Kapoor, M., H. H. Kelejian, et al. (2007) propose a FGLS estimator that also works under the $N \to \infty$ assumption.

[17] For a more thorough discussion about panel time-series analysis in general, see Smith and Fuertes (2004), or the various papers in Mills and Patterson (2006).

Table 2: Panel Unit-Root Tests

Variables	IPS Statistics
CV	-2.02
GINI	-4.46
Q12	-3.27
Q34	-3.51
Q5	-2.95
INFL	-3.98
UNEMP	-5.78
MINWAGE	-2.81

The moments of the mean E and variance var of the average $\bar{t}$ are respectively: -1.504 and .683. Source: Im, Pesaran and Shin (2003) and author's own calculations.

Given that all variables are stationary, and no cointegration analysis needs to be pursued nor other data transformations needed, we proceed to the issue of heterogeneity bias in dynamic models and also to static models[18].

We first estimate dynamic equations using the FE estimator, which assumes heterogeneous intercepts and homogeneous slopes. Equation 3 illustrates the main estimated dynamic equation.

$$CV_{it} = \alpha_i + \beta CV_{it-1} + \gamma INFL_{it-1} + \delta UNEMP_{it-1} + \epsilon MINWAGE_{it} + u_{it}, \tag{3}$$

where the explained CV_{it} is the Coefficient of Variation of the earnings distribution. The explanatory variables include lagged inflation ($INFL_{it-1}$), the lagged unemployment rates ($UNEMP_{it-1}$), the minimum-wage index ($MINWAGE_{it}$), and the lagged values of the Coefficient of Variation (CV_{it-1}) and lagged past inflation ($PASTINFL$), which is estimated against the next period CV_{it}. We then move to the RC-GLS estimator, which assumes the existence of heterogeneous intercepts and slopes. The RC-GLS estimator consists of a weighted average of $\hat{\alpha}_i$ and $\hat{\beta}_i$ and the weight is a modified variance-covariance matrix of the heterogeneous α_i and β_i.

The results in Table 3 show that in most equations and estimators the dynamic estimates of inflation and past inflation are positive and statistically

[18] Zellner (1969) states that for static models all panel estimators give unbiased estimates of the expected values.

significant. For instance, using the dynamic RC-GLS estimates from the first specification, a point increase in inflation increases inequality in .062 points per year. Furthermore, the estimates of lagged past inflation at the bottom of the Table confirm the fact that inequality is regressively affected not only by unanticipated, but also anticipated high rates of inflation.

Regarding the estimates of the lagged measure of inequality, they are positive and significant, confirming the fact that inequality is persistent[19]. The static estimates of the unemployment rates are positive, somehow confirming the theoretical prediction that the poor are the ones to be displaced first when a recession occurs. However, the dynamic estimates of unemployment are in fact negative, which suggests that the underground economy and migration might be playing a crucial role in buffering the regressive impact of short-run unemployment on inequality in the medium run[20].

The estimates of the minimum-wage index are all negative and significant, which suggests that this policy does not increase inequality[21]. The Likelihood Ratio (LR) tests for homogeneity of intercepts and slopes indicate that we can accept the alternative hypothesis that the parameters are in fact *heterogeneous*, which makes the RC-GLS estimator the most appropriate for these dynamic models.

In static specifications we first estimate equations using the Pooled Ordinary Least Squares (POLS) estimator—which assumes homogeneous intercepts and slopes—and then move to the FE estimator. In all specifications and estimators the rates of inflation remain regressive and statistically significant. For instance, using the FE estimates of the more general second specification, inflation increases inequality in .060 points per year.

[19] Corroborating the fact that according to the IPS test all variables are stationary, it is important to mention that under $T > N$ a spurious regression is less of a problem anyway. Phillips and Moon (1999), argue that since these pooled estimators are averaging over the regions, the noise is attenuated and the estimates are consistent. Moreover, Smith and Fuertes (2004) suggest that this result holds even under between-region dependence.

[20] Ferreira and Litchfield (2001) highlight the importance of the underground economy in, to some extent, buffering the prospective regressive effect of higher unemployment on inequality in Brazil. Moreover, Satchi and Temple (2006) argue that a sizeable underground economy can be an equilibrium outcome. Also, Fiess, Fugazza, et al. (2006) report evidence using Brazilian data covering the period of 1983-1989, which suggests that the underground economy played an important role in absorbing displaced workers from the formal sector.

[21] Incidentally, Lemos (2004) argues in a study on the effects of the Brazilian minimum wage on employment that the minimum did not significantly create job losses between 1982 and 2000.

The unemployment rates estimates are significant and, as expected in the short run, regressive. The minimum wage is progressive and significant in the FE estimator, which confirms that this particular policy does not increase inequality. The LR tests reject the null of homogeneous intercepts, suggesting the presence of regional fixed effects. Table 3 reports the results.

Table 3: Estimates of Macroeconomic Performance on Inequality, 1983-1994.

	Dynamic Models		Static Models	
CV	FE	RC	POLS	FE
INFL			.0998 (16.64)	.0924 (21.24)
INFL (1)	.0554 (12.37)	.0522 (2.37)		
UNEMP	.0826 (2.93)	.0218 (.20)	.3204 (8.28)	.0561 (1.67)
MINWAGE	-.0274 (-15.86)	-.0270 (-6.78)		
Constant		17.263 (30.77)	.1259 (49.73)	
LR test	771.98	2320.78		525.30
F test	274.34	NA	153.06	189.75
R^2	.73	.72	.27	.62
CV (1)	.8872 (80.21)	.8393 (29.00)		
INFL			.1027 (15.32)	.0504 (11.33)
INFL (1)	.0085 (5.34)	.0093 (1.53)		
UNEMP			.3079 (7.57)	.0945 (3.29)
UNEMP(1)	-.0170 (-1.82)	-.0303 (-1.47)		
MINWAGE	-.0040 (-6.28)	-.0059 (-2.28)	.0022 (.98)	-.0291 (-17.11)
Constant		2.9981 (5.09)	.12396 (38.53)	
LR test	84.89	231.48		776.08
F test	2954.02	NA	102.36	264.5
R^2	.97	.94	.28	.72
CV (1)	.9129 (67.47)	.8969 (49.33)		
PASTINFL (1)	.0096 (4.64)	.0121 (2.79)		
Constant		.1444 (5.45)		
LR test	26.78	56.33		
F test	2301.51	NA		
R^2	.95	.95		

T-ratios in parentheses, number of observations: $NT = 792$. The basic estimated equation is $CV_{it} = \alpha_i + \beta INFL_{it} + \gamma UNEMP_{it} + \delta MINWAGE_{it} + u_{it}$, where CV is the Coefficient of Variation of the earnings distribution, $INFL$ the inflation rates, $UNEMP$ the unemployment rates and $MINWAGE$ the minimum-wage index. Source: author's own calculations.

Additionally, we look at the rather important issue of between-region

dependence, which is dealt with the two-step SUR-FGLS estimator. This estimator presents greater efficiency, the greater the correlation among the disturbances and it estimates different regional time series, which are then weighted by the covariance matrix of the disturbances[22]. Furthermore, this estimator delivers more insightful estimates since it disaggregates the analysis. Equation 4 illustrates the general dynamic equation estimated for each region.

$$CV_t = \alpha_t + \beta INFL_{t-1} + \gamma UNEMP_{t-1} + \delta MINWAGE_t + u_t, \quad (4)$$

where CV_t is the Coefficient of Variation of the earnings distribution, $INFL_{t-1}$ accounts for lagged inflation, $UNEMP_{t-1}$ for the lagged unemployment rates and $MINWAGE_t$ for the minimum-wage index.

The dynamic and static effects of inflation are positive and significant in most regions. An interesting feature seen in those effects is that the poorer regions of the Northeast, i.e. Pernambuco (PE) and Bahia (BA), and to a lesser extent Rio de Janeiro (RJ), present the largest estimates of all, which indicates that the poorer the region, the more regressive inflation is. For instance, the SUR-FGLS estimates from the first specification indicate that a point increase in inflation increases inequality in .150 points per year in Pernambuco, which is the poorest region in the sample[23].

Unemployment presents regressive effects in those poor regions of the Northeast, i.e. Pernambuco and Bahia, and also Rio de Janeiro. In the more affluent regions of the South the effect of this variable is not clear cut, which possibly indicates the existence of more organised dual labour markets attenuating the regressiveness of short-run unemployment.

Regarding the minimum-wage index, the results show that the minimum wage does not have any regressive effect on inequality. The Lagrange Multiplier (LM) tests reject the null hypothesis that the variance-covariance matrices are diagonal, which suggests that these regions are, in fact, related

[22] Phillips and Sul (2003) argue that when between-region dependence is present there is very little gain in using pooled analysis.

[23] Related to that, Guitián (1998), and Romer and Romer (1999) show in cross-sections of countries that inflation presents regressive effects on inequality, the poorer the countries. Moreover, Bulír (2001) reports that in countries that present hyperinflationary periods, inflation presents stronger regressive effects on inequality.

to each other through the disturbances[24]. Table 4 reports the results.

[24]The IPS test reported in Table 2 above assumes the existence of between-region *independence*. An alternative that considers the existence of between-region dependence is proposed by Pesaran (2006), the cross-section IPS (CIPS) test. However, CIPS assumes that $N > 10$ and we have $N = 6$ in our data set. It is therefore thought that the IPS test in this case is slightly biased but still informative and the best alternative available. See Baltagi, Bresson, et al. (2005) for more on panel unit-root tests and between-region dependence.

Table 4: SUR-FGLS Estimates of Macroeconomic Performance on Inequality, 1983-1994.

	SUR-FGLS		
CV	PE	BA	MG
INFL(1)	.1250 (15.29)	.0781 (8.77)	.0350 (3.27)
UNEMP(1)	.3606 (10.20)	.4411 (7.58)	-.2137 (-4.01)
MINWAGE	-.0151 (-6.58)	-.0371 (-11.71)	-.0174 (-4.45)
Constant	1.4540 (33.52)	1.7088 (32.88)	1.8715 (33.04)
LM test	455. 42		
INFL	.1144 (13.71)	.0680 (7.31)	.0265 (2.45)
UNEMP	.0382 (10.61)	.0452 (7.49)	-.0222 (-4.21)
MINWAGE	-.0018 (-7.81)	-.0039 (-11.94)	-.0019 (-4.88)
Constant	1.4967 (34.69)	1.7417 (32.53)	1.9079 (34.11)
LM test	513.88		
	RJ	SP	RS
INFL(1)	.0886 (11.76)	.0023 (0.30)	-.0168 (-1.38)
UNEMP(1)	.1295 (2.58)	-.1716 (-3.94)	-.1436 (-1.82)
MINWAGE	-.0202 (-7.01)	-.0379 (-9.94)	-.0297 (-5.96)
Constant	1.5523 (36.31)	1.7625 (42.43)	1.8043 (25.43)
LM test	455.42		
INFL	.0867 (11.67)	.0037 (.49)	-.0097 (-.79)
UNEMP	.01711 (3.40)	-.01711 (-3.91)	-.0092 (-1.13)
MINWAGE	-.0021 (-7.77)	-.0037 (-10.17)	-.0027 (-5.59)
Constant	1.5547 (37.36)	1.7579 (43.04)	1.7543 (24.64)
LM test	513.88		

T-ratios in parentheses, number of observations: $NT = 792$. The basic estimated equation is: $CV_t = \alpha_t + \beta INFL_t + \gamma UNEMP_t + \delta MINWAGE_t + u$, where CV is the Coefficient of Variation of the earnings distribution, $INFL$ the inflation rates, $UNEMP$ the unemployment rates and $MINWAGE$ the minimumwage index. Source: author's own calculations.

The economic intuition behind the above evidence is: firstly, chronic high rates of inflation are detrimental to those who are not at the very top of the distribution[25]; secondly, although the regional inflation rates follow

[25] When we estimate inflation and past inflation against the earnings shares of the quin-

a similar national trend, they affect different regions differentially. Loosely speaking, the poorer the region, the more regressive inflation tends to be, which highlights the fact that poor regions are more vulnerable to extreme inflation for not presenting the right mechanisms—indexation coverage in all its forms—against it.

Thirdly, in terms of unemployment effects, the static pooled evidence somehow confirms the standard prediction that those at the bottom of the distribution present lower turnover costs, and the more disaggregated SUR-FGLS evidence highlights that the poorer the region, the more regressive unemployment is. Furthermore, the dynamic pooled evidence also suggests the existence of dual labour markets and migration, and their role on buffering the impact of poor macroeconomic performance on inequality in the medium run.

Fourthly, regarding the minimum-wage index, the estimates suggest that this policy does not have a regressive impact on inequality. It is worth mentioning though, that the minimum wage had been kept reasonably low in Brazil until the stabilisation in 1995. Moreover, the minimum wage suffered severe restrictions from the government in terms of indexation, i.e. it would only be readjusted if inflation had reached a particular threshold.

All in all, the evidence—based on panel time-series data and analysis—confirms the initial inspection of the data based on the national time-series variation in the data.

4 Concluding Remarks

We investigated the impact that macroeconomic performance had on earnings inequality in Brazil in the 1980s and first half of the 1990s. The evidence, based firstly on national time-series and then on the relatively novel sub-national panel time-series $T > N$ data and analysis, suggests that extreme rates of unanticipated and anticipated inflation had significant effects in raising inequality during the period. The evidence shows that the poorer did not have access to indexed financial assets, and all that it entails, to protect themselves against accelerating inflation, and nor fully monthly-indexed wages. Moreover, the panel time-series analysis permits us to disentangle

tiles, the results suggest that the only group that manages to increase its share is the twenty percent richest in the distribution. Also, when the Gini coefficient is the measure of inequality used, the same sort of qualitative results arise. Available on request.

the effects of inflation on regions with different levels of development, and the regions suffering most with high inflation are the poorest ones in the Northeast.

The other two explanatory variables regressed against inequality alongside inflation, i.e. the unemployment rates and the minimum-wage index, respectively presented mixed and progressive effects on inequality. Initially, the results confirm the fact that the poor present lower turnover costs in the short run, and hence lose their formal jobs and earnings first when a recession occurs, and that a minimum-wage policy does not increase inequality when the minimum is low enough. Still, with regards to unemployment—it can be said that the poorest the region, the more regressive unemployment tends to be, which suggests the importance of organised dual labour markets in buffering the regressive effects of unemployment in better-off areas, and somehow the dynamic pooled estimates confirm that.

Another important issue raised is the need to differentiate the impact of inflation on inequality in countries that present different economic conditions. The review presented in Section 1 above from previous studies on the US, and to a lesser extent the Netherlands and UK, suggests that moderate rates of inflation would be beneficial for the poor, since they would benefit from the decreasing amounts of their debts via the debtor and creditor channel. Moreover, slightly higher rates of inflation are also associated with an expansive monetary policy, normally used to boost employment in developed countries, and which would decrease inequality through lower rates of unemployment in the very short run.

On the contrary, in a country with rampant rates of inflation such as Brazil, and other developing countries in the 1980s and early 1990s, any possible gain coming from the debtor and creditor channel was offset by the poor macroeconomic performance. The evidence presented in Sections 2 and 3 from a range of inequality measures, specifications and estimators, shows the regressive effects of high inflation on inequality, and therefore the importance of having sound monetary and fiscal policies—not to mention independent monetary authorities and fiscal rules—that actually keep

inflation consistently low and under control in the very long run[26] [27].

Moreover, the quality of the results are to a certain extent boosted not only by the inclusion of the minimum-wage index in the equations, but also by the novel analytical approach used. The evidence—particularly the one based on sub-national panel time-series $T > N$ data and analysis—deals with issues such as non-stationarity in panels, heterogeneity bias in dynamic panels and between-region dependence. None of these issues has been considered before in any other study of the impact of macroeconomic performance on inequality. Therefore, it is believed that—given the usual caveat that panel time-series is an area developing rapidly—the sort of analysis carried out here can be regarded as a significant step forward in terms of achieving better and more insightful estimates.

Regarding future work, the use of Brazilian data from 1995 onwards to check whether the low rates of inflation generated by the Real Plan have actually had a progressive impact on inequality, as in developed countries, would naturally complement this study. Another extension is an investigation of the importance of financial development on inequality in Brazil, i.e. whether access to finance would really present the poor and the middle classes not only with credit that could be used to invest in all sorts of short and long-run productive activities, but also with some sort of indexed protection of their earnings against high inflation. Further, a study on the importance of high inequality in keeping inflation high in non-democratic societies is also worth pursuing. Finally, panel-var analysis is a possible extension given the sort of $T \succ N$ data available.

To conclude, first we understand that in a country that presents high inequality like Brazil, the unstable macroeconomic performance of the 1980s and 1990s, although important, is not the whole story behind inequality. Second, however, when we take into consideration the high rates of inflation per month seen at the time and the size of the estimates presented in

[26] Singh (2006), Singh and Cerisola (2006), and Santiso (2006) highlight the role of macroeconomic stability on a range of positive economic outcomes in Latin America in general and in Brazil in particular recently. Furthermore, Carvalho and Chamon (2006) argue that real income growth in Brazil after the reforms of the 1990s has been, for methodological reasons, severely underestimated, which further highlights the importance of macroeconomic stability on improved welfare.

[27] Although the Brazilian Central Bank has implemented, e.g. inflation targeting in the late 1990s, Carstens and Jácome (2005) argue that Brazil still possess one of the *least* independent central banks in the whole of Latin America.

Section 3, the impact of bad macroeconomic performance on inequality is considerable. Therefore, the moral to be drawn from the evidence presented is that a stable macroeconomic environment—which is only to be achieved by the implementation of sound monetary and fiscal policies—is certainly a *necessary* condition to achieve lower inequality in Brazil.

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