# THE IMPACT OF FEMALE LABOR SUPPLY ON THE BRAZILIAN INCOME DISTRIBUTION

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## 1- Introduction

The strong growth of female participation in Brazilian labor market in the last twenty-five years occurred over a period of outstanding structural shifts in Brazilian economy. In the first half of the 80's there were external adjusting process and increasing inflation rates. In the second half of the 80's, several attempts to stabilize prices failed, so that the monthly inflation rate went to 80% on March 1990. In the 90's, however, there was a process of economic opening and privatization of state-owned firms. After 1994, *Plano Real* marked the beginning of economic stabilization. Despite the abundance of economic phenomena, Brazilian income distribution remained high and stable in that period (Barros, Henrique and Mendonça, 2000).

In spite of the stability, we may question whether the strong women's entry into the labor market in that period - pointed out, among others, by Scorzafave and Menezes-Filho (2001) - has influenced the evolution of Brazilian income distribution. In another way, we would like to know what the women's contribution was towards the high level of persistent inequality. Have women contributed to a more equanimous distribution, equalizing eventual trend towards the growth of existing inequality among men? Has difference between sexes contributed to the growth of inequality? Thus, this paper has the aim to estimate how the increase of female participation has affected the growth of inequality of Brazilian income distribution from 1982 to 1997.

Despite the last twenty-five years were marked with two trends - increase of female participation and constant inequality - surprisingly, there are few studies on the relationship between women in the labor market and the Brazilian income distribution. One exception is Ferreira and Litchfield's (2001) paper which develops the dynamic decomposition of *Theil-L* index to investigate which characteristics of households would explain the shifts in income inequality in that period. Particularly, the authors divide the sample into two subgroups - households headed by women and by men - concluding that all the growth of income

inequality can be neither explained by shift in the composition of the two groups (allocation effect), nor by shift in groups' relative mean income (income effect). Almost all the growth of inequality is explained by shifts in inequality within the groups (pure inequality effect).

## 2- Theoretical Background and Methodology

### 2.1 - Theoretical Background

During the 70's, the discussion about the decomposability of the inequality measures according to subgroups, mainly Gini index, has grown in literature[1]. Nevertheless, only in late 70's, Bourguignon (1979) and Shorrocks (1980, 1982)'s studies brought into the field the discussion on what measures of income distribution would be decomposable. Particularly, Shorrocks (1982) states some basic axioms, which should be respected by any probable index, to be decomposable in the inequality between groups and in the inequality within groups, at a certain period of time. The author shows that, if those axioms were followed, the parcel of inequality of the subgroup's responsibility would not depend on the chosen index of inequality.

However, these methods are limited because they try to decompose the inequality indexes at a certain period of time, or better, the decomposition is just static. Nevertheless, Mookherjee and Shorrocks (1982) surpass such problem, proposing a dynamic decomposition of inequality indexes, in order to estimate the contribution of different factors to the growth of inequality. Ferreira and Litchfield (2001) applied such methodology to the Brazilian context, but did not find significant results on the number of female-headed households in the growth of income inequality. However, the authors do not specify married women's participation, for they dissolve households, separating them into types of families, householder's sex, etc.

Cancian and Reed (1998) also attempt to estimate the effect of women's income on the growth of inequality. Nevertheless, the authors highlight that this type of practice needs to undergo *a priori* a clear definition of what we understand by 'effect' or 'impact'. In other words, we could only estimate if a particular income source contributes to the growth of inequality if there were any comparison measures. Such reference would be supplied by the construction of counterfactual income distributions, allowing the effect estimation on each income source of the growth of inequality. By applying the decomposition of the coefficient of variation (CV) only to a sample of married couples, they show that the growth of female income distribution contributed to decrease the inequality of American income.

Reed and Cancian (2001) try to go beyond the bounds of working on CV and propose a new approach. For each period, they classify the women into 1000 groups, according to their rank in female income distribution and estimate the mean income in each group. Thus, they can attribute the one year female income distribution to another year, which would allow the identification of the effect on the household income distribution, mainly the one resulting from the shift in women's income distribution in that period. They confirm previous results of women's contribution to decrease family income inequality. They conclude, for example, that if female income distribution from 1969 to 1999 did not shift, Gini coefficient would have increased 33% instead of the 21% observed. On the other hand, following similar approach, Burtless (1999) concluded that the growth of female inequality did not influence the shift of inequality in American income between 1979 and 1999.

Nevertheless, the fact that women's participation increased considerably between 1982 and 1997 leads us to investigate, at individual level, women's impact on income distribution. Besides, our analysis considers both: the role of both married and unmarried women, which constitute one advantage compared to Ferreira and Litchfield's (2001) work. Similarly, Cancian and Reed (1998) also use only the CV as inequality measure, besides following only a sample of couples, which does not reflect all the possible family arrangements that may influence the country's inequality income distribution, as a whole. This means that only working women are taken into account (Burtless, 1999; Cancian, 2001) to estimate the impact of women on the income distribution.

## 2.2 Methodology

In order to estimate the effect of women's participation in Brazilian income distribution, we followed the methodology based on Ferreira and Barros (1999). First, we assessed the conditional probability of work for 1982 and 1997, through a Probit model:

$$Y = Z\gamma + u \tag{1}$$

where Y is a dummy variable which is 1, for working women; Z is a explanatory vector which includes age, age$^2$, educational level, number of kids; a dummy indicating white individuals and regional dummies; $\gamma$ is a vector of parameters to be estimated and $u$ is a random error

[1] See Fei, Ranis and Kuo (1978).

term. By the estimated results, each woman's probability of being employed (at work) was estimated in both periods:

$$P_{82} = \Phi\ (Z_{82}\,\hat{\gamma}_{82})$$
$$P_{97} = \Phi\ (Z_{97}\,\hat{\gamma}_{97})$$

After that, one woman's probability of being employed (at work) in 1997 was assessed, according to the observable characteristics, but coefficients referred to 1982. Or rather, if women's characteristics of 1997 were paid under 1982 coefficients, what would be their probabilities of being employed (at work) ? Thus, a simulated probability was obtained for all 1997 sample.

$$PS_{97} = \Phi\ (Z_{97}\,\hat{\gamma}_{82})$$

Therefore, we could obtain the probability of women's participation if the returns to observable characteristics remained steady in that period. It is obvious that the values $P_{97}$ and $PS_{97}$ differ. Besides, we already know which women worked in 1997 and which women did not work. Thus, we built a set of decision rules, which allow 'selecting' the women who would work in 1997 at prices of 1982, in order to estimate the effect of female participation in income distribution.

For each woman who effectively worked in 1997:

a) If $PS_{97} > P_{97}$, we assume the woman works.
b) If $PS_{97} < P_{97}$, we select a random number (RN) between $[0, P_{97}]$:
   b1) If $PS_{97} < RN$, we assume the woman does not work;
   b2) If $PS_{97} > RN$, we assume the woman works.

For each woman who did not work in 1997:

a) If $PS_{97} < P_{97}$, we assume the woman does not work.
b) If $PS_{97} > P_{97}$, we select a random number (RN) between $[P_{97},1]$:
   b1) If $PS_{97} < RN$, we assume the woman does not work ;
   b2) If $PS_{97} > RN$, we assume the woman works.

Based on that, we obtained a sample of working women in 1997, taking into account the coefficients estimated in 1982. The last step of the approach attributes income to women who did not work, but the rules of decision labeled them as working women. For that, we estimated a regression of the monthly labor income using a set of explanatory variables for women in 1997. By using the estimated coefficients and the estimated variance of the error, we attributed a income value to women who did not have income in 1997. Therefore, the application of that methodology allowed building a 'simulated' sample of women with positive income for 1997. Based on that sample and on the results effectively verified in 1997, we estimated the impact of the increase of women's participation in the evolution of Brazilian income distribution.

The estimation of women's impact on income distribution can be assessed, using some measures of income inequality with different properties. In this study, we used the following generalized entropy measures:

a) Theil – L or mean log deviation $$GE(0) = \frac{1}{n}\sum_{i=1}^{n} \log \frac{y_i}{\bar{y}}$$

b) Theil – T $$GE(1) = \frac{1}{n}\sum_{i=1}^{n} \left(\frac{y_i}{\bar{y}}\right) \log \frac{y_i}{\bar{y}}$$

c) Coefficient of Variation $$CV = \frac{1}{\bar{y}}\left[\frac{1}{n}\sum_{i=1}^{n}\left(y_i - \bar{y}\right)^2\right]^{\frac{1}{2}}$$

d) GE (-1) $$GE(-1) = \frac{1}{2}\left[\frac{1}{n}\sum_{i=1}^{n}\frac{\bar{y}}{y_i} - 1\right]$$

In addition, we used Gini coefficient and standard deviation of logs:

e) Gini coefficient $$Gini = \frac{1}{2n^2\bar{y}}\sum_{i=1}^{n}\sum_{j=1}^{n}\left|y_i - y_j\right|$$

f) Standard deviation of logs $$SDLOG = \frac{1}{\sqrt{n}}\left[\sum_{i=1}^{n}\left(\log y_i - \frac{1}{n}\left(\sum_{i=1}^{n}\log y_i\right)\right)^2\right]^{1/2}$$

### 2.3 Income Measures

From the simulated results related to the impact of the shift in female participation in all the women's labor income, we will build other measures of income, in order to have a thorough view of the phenomenon. This way, we will adopt two different income measures: (i) *per capita* familiar income, defined as all the income produced by the family members, divided by number of family members; and (ii) individual labor income, which comprises aspects related to labor market without worrying about intra-household distribution aspects, as well as other income sources.

## 3- Data and Results

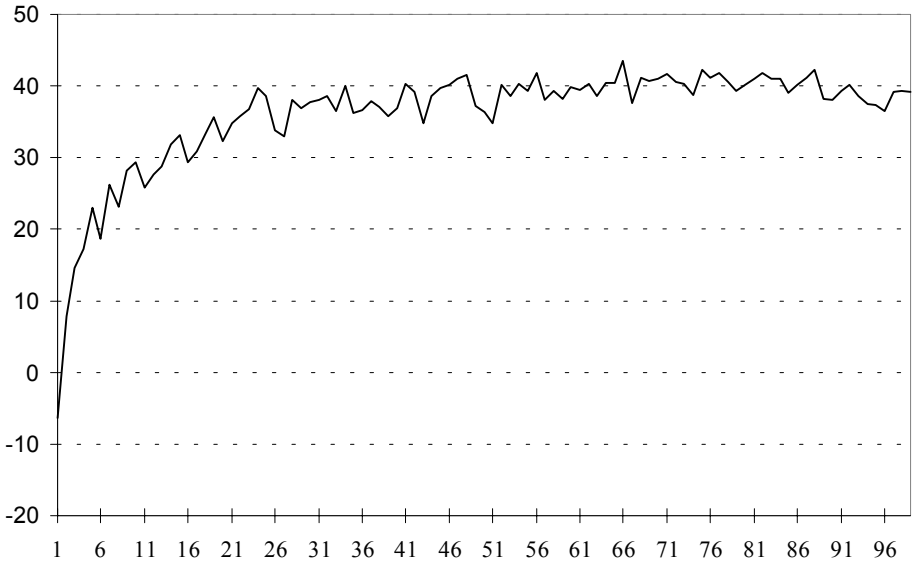
For the estimation of per capita familiar income, data include, firstly, information about all the families whose head's age (or partner's age) was between 25 and 64 years old, in 1982 and 1997. As mentioned before, in this case, the income measure used includes not only the one coming from labor but the ones from other sources such as rent, pensions, etc. However, for the analysis referring to individual income, only the women's income at ages between 25 and 64 were considered.

### 3.1 - Descriptive Analysis of the Data

First of all, we will analyze the data on the income distribution considering per capita familiar income measure for the individuals on positive income[2]. Figure 1 shows the growth of the *per capita* familiar income for all the centiles of distribution between 1982 and 1997. Data show that individuals of centile 1 were submitted to a real income decrease of 6% in that period. The median *per capita* familiar income, by its turn, increased 36.4%. There is, still, a positive relationship between the centiles and the income variation up to centile 20. Therefore, the poorest 20% were submitted to a minor increase of per capita familiar income, leading them to a relative inferior condition.

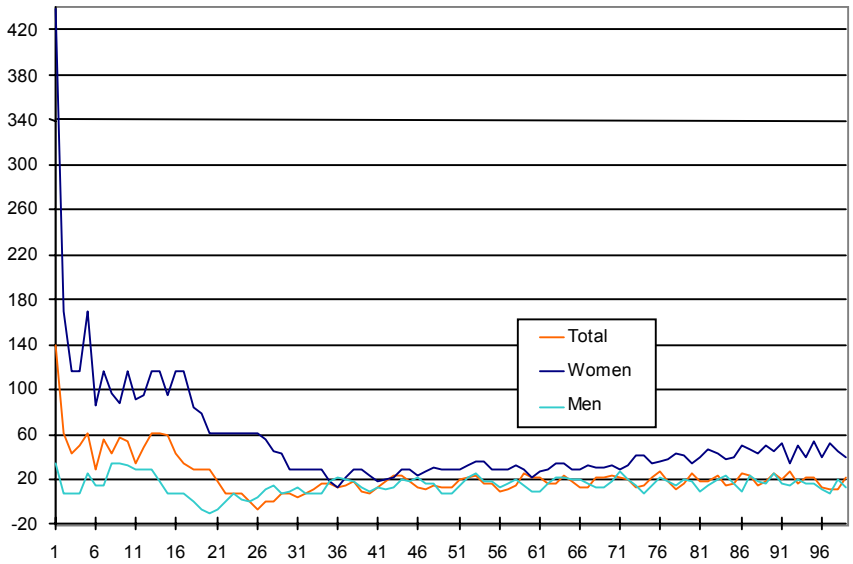
[2] The figures are in *Reais* (the Brazilian currency) of 1996 and were deflated by IPCA/IBGE.

**Figure 1 - Growth of Per Capita Familiar Income by Centiles**




The familiar income is estimated by the aggregation of all family members' individual income. This way, it is important to point out the shifts occurred in women's individual income, in order to have a clear view of the relationship between shifts in individual income and familiar income. Therefore, we will analyze now the behavior of the individual income of people who had some type of income.

**Figure 2 - Growth of Individual Income - 1982/1997**




The chart analysis of the income growth according to centiles can provide a clear view of the individuals' income behavior for each centiles. The remarkable aspect of the figure is the high growth of poorest women's income (up to centile 30). In comparison, it also shows the growth of men's income and also men and women together (total).

Initially, the results would point out to a fall in inequality, mainly between women. However, this fact would not be reflecting on Gini coefficient, which increased 1.4% in that period, ranging from 0.578 to 0.586. To solve the problem, we estimated the ratio of total income accumulated by women at specific deciles of female income distribution.

**Table 1 - Accumulated Income by Deciles - 1982/97**

| | 1982 | 1997 |
|---|---|---|
| **1 to 3** | 5.85% | 8.16% |
| **4 to 7** | 24.01% | 19.27% |
| **8 to 10** | 70.14% | 72.57% |

Data show that, despite the outstanding increase of women's mean income in the first deciles, the parcel of total income, appropriated by those women, increased from 5.85% to 8.16% of the total women's income. However, the women of intermediate deciles (4 to 7) decreased the parcel of their appropriated income. Finally, the women of 8 to 10 deciles increased their income parcel. Thus, the slight growth of women's Gini index occurred, despite the strong percentage increase of the mean income of lower deciles, the sum of the income appropriated by women of these deciles increased much less, which is not enough for Gini index decrease.

### 3.2 - Simulation Results

After the descriptive data analysis, we will present the results of simulation discussed in section 2.2. First, we will show the estimation results of female participation model for 1982 and 1997, according to equation (1).

It is important to highlight the negative result for the white dummy coefficient, explaining that white women have smaller probabilities of being working or employed compared to others. Concerning the regional dummies, southeast women have more chances of working, except for the ones from Center-West in 1982. The other coefficients show the expected results.

**Table 2 – Probability of Being Employed - Women – 1982/97**

| | 1982 | | | 1997 | | |
|---|---|---|---|---|---|---|
| | **Marginal Effect** | **Coefficient** | **Standard Deviation** | **Marginal Effect** | **Coefficient** | **Standard Deviation** |
| **Education** | 0.024 | 0.063 | 7.6E-05 | 0.025 | 0.063 | 5.5E-05 |
| **Age** | 0.049 | 0.128 | 2.3E-04 | 0.057 | 0.142 | 1.8E-04 |
| **Age$^2$** | -0.001 | -0.002 | 2.7E-06 | -0.001 | -0.002 | 2.1E-06 |
| **Kids** | -0.026 | -0.068 | 1.4E-04 | -0.019 | -0.048 | 1.6E-04 |
| **North** | 0.032 | 0.084 | 1.8E-03 | 0.059 | 0.149 | 1.2E-03 |
| **Northeast** | 0.053 | 0.138 | 7.1E-04 | 0.062 | 0.156 | 5.7E-04 |
| **Center-West** | -0.029 | -0.078 | 1.2E-03 | 0.012 | 0.029 | 8.8E-04 |
| **South** | 0.088 | 0.227 | 7.8E-04 | 0.088 | 0.222 | 6.4E-04 |
| **White** | -0.089 | -0.235 | 6.3E-04 | -0.039 | -0.097 | 5.0E-04 |
| **Constant** | | -2.638 | 4.6E-03 | | -2.925 | 3.6E-03 |
| **Pseudo-$R^2$** | | 0.0614 | | | 0.0699 | |

OBS $_1$: All coefficients are significant at 1% level.

Next, we carried out the process described by the decision rule in order to verify which women would be working in 1997 if the coefficients related to 1982 were taken into account. Besides, we attributed wages to the women without income based on the coefficients and on the estimated variance of the random error, based on the regression estimates whose results are in Table 3.

**Table 3 – Labor Income Regression – 1997**

| | **Coefficient** | **Standard Deviation** |
|---|---|---|
| **Education** | 0.128 | 2.1E-04 |
| **Age** | 0.064 | 2.1E-04 |
| **Age$^2$** | -0.001 | 2.5E-06 |
| **Age*Education** | 1.2E-04 | 5.2E-06 |
| **Number of Kids** | -0.024 | 1.6E-04 |
| **North** | -0.075 | 1.1E-03 |
| **Northeast** | -0.475 | 5.5E-04 |
| **Center-West** | -0.055 | 8.4E-04 |
| **South** | -0.013 | 6.1E-04 |
| **Urban** | 0.293 | 7.3E-04 |
| **Metropolitan** | 0.275 | 4.7E-04 |
| **Constant** | 2.860 | 4.6E-03 |
| **$R^2$** | 0.507 | |

OBS: All coefficients are significant at 1% level

Therefore, we built one sample of women aging from 25 to 64 in 1997, all of them on positive income. From that on, we simulated some exercises to find the effect of growth of women's participation in the distribution:

a) of individual income;
b) of per capita familiar income.

The referent simulations were prepared in three different samples:

I) women from 25 to 64 years old on positive corresponding income (of labor or per capita familiar);
II) men and women from 25 to 64 years old on positive corresponding income;
III) men and women on positive corresponding income, independent of age.

Table 4 shows the results of simulations related to all labor individual income. The table is divided into three parts according to samples (I), (II) and (III), mentioned before. In each part of the table, lines 1 and 2 show the values effectively verified of several inequality measures in 1982 and 1997, respectively, while line 3 shows the results of simulation of the effect of changing women's participation in the income distribution in that period. More specifically, how the income distribution would be in 1997 if the observable female characteristics were estimated using the1982 coefficients.

Thus, if we compare, for example, the simulated income distribution to the real one observed in 1997, the only difference between them are the returns to the observable characteristics. Therefore, if the simulated inequality is superior to the one estimated, it means that the shift of the return to the observable characteristics of women, between 1982 and 1997, contributed to decrease the income inequality and *vice versa.* In Table 4 this is expressed on the line "contribution" with an "↓" when the contribution points to decrease of inequality and "↑" the opposite.

First, we observed that there was difference in the growth of inequality depending on the measure used. On the one hand, CV, Gini and Theil-T show the increase of income inequality between women from 25 to 64. On the other hand, SDLOG, Theil-L and GE (-1) signal decrease. When men from 25 to 64 were incorporated to the analysis, only GE (-1) and SDLOG remained signaling the inequality decrease, result that is kept when all individuals with positive labor income were considered. In the case of GE (-1), the salient decrease can be explained by the substantial increment of poorest women's labor income, pointed out in Figure 2.

**Table 4 - Labor Income Inequality**

| | CV | SDLOG | Gini | Theil-T | Theil-L | GE(-1) |
|---|---|---|---|---|---|---|
| **Sample I** | | | | | | |
| **1982** | 1,483 | 1,140 | 0,563 | 0,594 | 0,618 | 1,404 |
| **1997** | 1,551 | 1,065 | 0,565 | 0,620 | 0,594 | 1,118 |
| **Simulated 1997** | 1,587 | 1,083 | 0,572 | 0,637 | 0,612 | 1,170 |
| **Contribution** | ↓ | ↓ | ↓ | ↓ | ↓ | ↓ |
| **Sample II** | | | | | | |
| **1982** | 1,608 | 1,077 | 0,564 | 0,619 | 0,595 | 1,206 |
| **1997** | 1,770 | 1,071 | 0,575 | 0,661 | 0,614 | 1,151 |
| **Simulated 1997** | 1,787 | 1,086 | 0,579 | 0,670 | 0,626 | 1,200 |
| **Contribution** | ↓ | ↓ | ↓ | ↓ | ↓ | ↓ |
| **Sample III** | | | | | | |
| **1982** | 1,712 | 1,090 | 0,573 | 0,656 | 0,619 | 1,257 |
| **1997** | 1,872 | 1,071 | 0,581 | 0,689 | 0,626 | 1,163 |
| **Simulated 1997** | 1,883 | 1,080 | 0,583 | 0,695 | 0,634 | 1,190 |
| **Contribution** | ↓ | ↓ | ↓ | ↓ | ↓ | ↓ |

Besides, results show that the shift of coefficients, associated with observable women's characteristics related to female activity, contributed for a decrease of inequality income, independent of the sample and of the inequality measure adopted. In another way, if vector $\hat{\gamma}$ had not changed from 1982 to 1997, the inequality would be much greater.

The same exercise was repeated in the case of per capita familiar income. There was an increase of inequality according to all the measures, except GE (-1) and Gini coefficient, when Sample I was considered. However, Gini indicates increase while CV points to inequality fall, when men from 24 to 64 on positive per capita familiar income are incorporated. At the presence of all individuals (Sample III), all the indexes indicate growth of inequality, including GE (-1). That means that when the poorest women's labor income is just one more element constitutive of the income measure (together with other types and other members' income, as well as the number of family individuals), the high increase of labor income experimented by the poorest women is not enough to guarantee the fall of per capita familiar income inequality between 1982 and 1997.

**Table 5 - Per Capita Familiar Income Inequality**

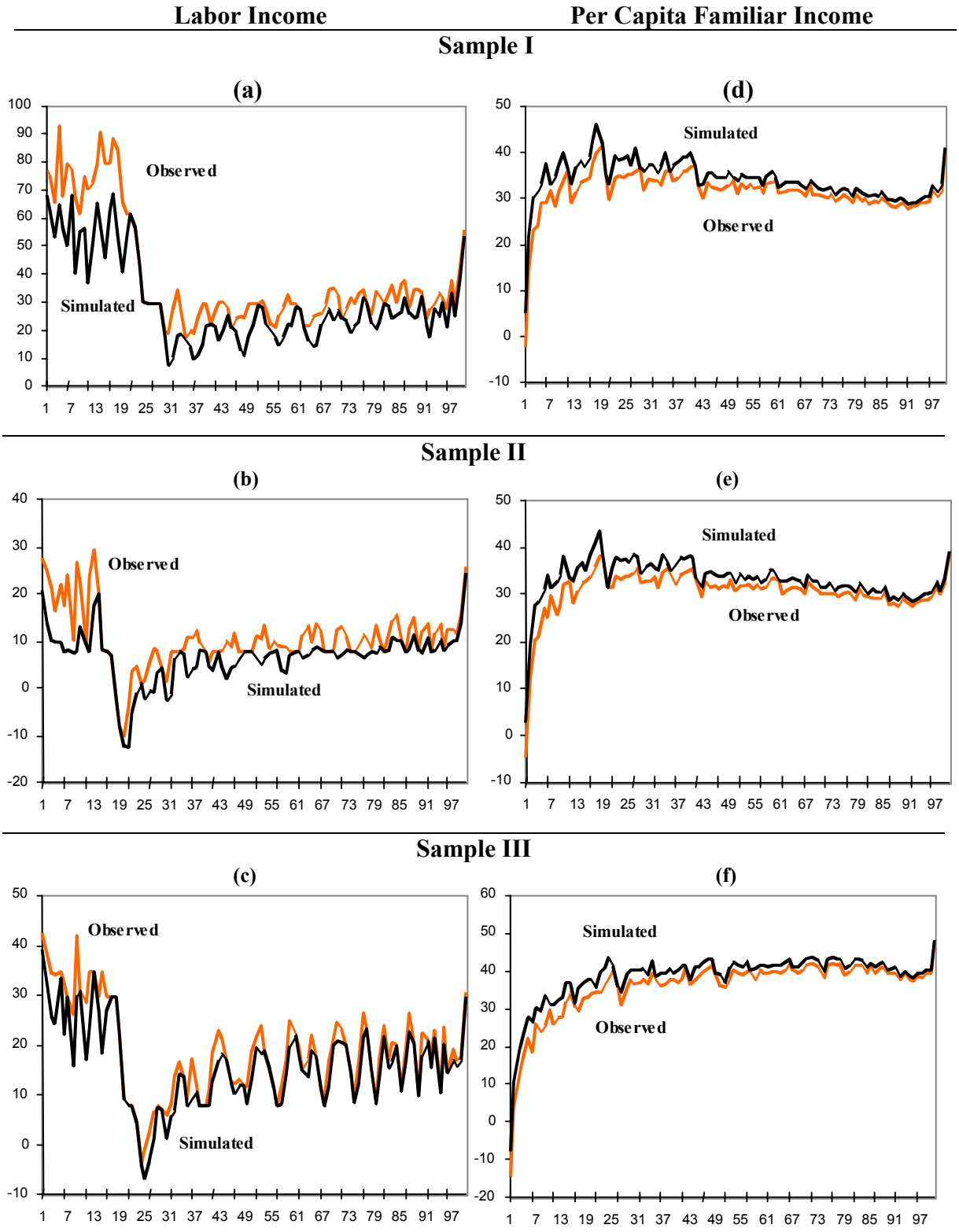
| | CV | SDLOG | Gini | Theil-T | Theil-L | GE(-1) |
|---|---|---|---|---|---|---|
| **Sample I** | | | | | | |
| **1982** | 1.719 | 1.128 | 0.590 | 0.680 | 0.662 | 1.479 |
| **1997** | 1.749 | 1.133 | 0.590 | 0.690 | 0.664 | 1.389 |
| **Simulated 1997** | 1.741 | 1.123 | 0.587 | 0.685 | 0.656 | 1.342 |
| **Contribution** | ↑ | ↑ | ↑ | ↑ | ↑ | ↑ |
| **Sample II** | | | | | | |
| **1982** | 1.840 | 1.128 | 0.592 | 0.697 | 0.667 | 1.525 |
| **1997** | 1.834 | 1.139 | 0.593 | 0.707 | 0.674 | 1.425 |
| **Simulated 1997** | 1.822 | 1.127 | 0.591 | 0.701 | 0.665 | 1.373 |
| **Contribution** | ↑ | ↑ | ↑ | ↑ | ↑ | ↑ |
| **Sample III** | | | | | | |
| **1982** | 1.845 | 1.109 | 0.592 | 0.701 | 0.660 | 1.356 |
| **1997** | 1.890 | 1.156 | 0.599 | 0.723 | 0.692 | 1.486 |
| **Simulated 1997** | 1.878 | 1.144 | 0.596 | 0.717 | 0.682 | 1.430 |
| **Contribution** | ↑ | ↑ | ↑ | ↑ | ↑ | ↑ |

Concerning the simulations, results are opposite to the ones found for labor income. In all the indexes and in different samples, the change of equation coefficient of activity between 1982 and 1997 contribute for the inequality growth. For example, Gini coefficient would increase from 0.596 to 0.599.

An alternative mode of estimating the effect of simulations, as well as the differences between the weight that various entropy measures attribute to different parts of distribution, is by comparing the verified growth of the mean income of each centile of income distribution to the simulated growth, or better, the growth which would occur in case the coefficients of the participation equation had not altered from 1982 to 1997. We carried these experiments for the labor income and per capita familiar income for samples I, II and III, and the results are presented in Figure 3.

Charts (a), (b) and (c) show the individual labor income mean of each centile according to samples I, II and III, respectively. Chart (a) points the strong income growth to women up to centile 20, effect which decreases considerably when men between 25 and 64 (Figure b) and individuals over 64 and under 25 (Figure c) are incorporated to the analysis.

**Figure 3 - Percentage Growth in Individual Labor Income and in Observed and Simulated Per Capita Familiar Income - 1982/1997.**




Comparing the results and the simulations, we confirmed the results of Table 4. Particularly, in Figure (a), minor income growth in simulated lower centiles than in the

observed results allow us to say that the shift of the parameters of participation equation between 1982 to 1997 contributed for increasing the poorest women's income better than the others, contributing for the decrease of inequality. The same fact is also observed in figures (b) and (c), although in smaller scale.

Figures (d), (e) and (f), on their turn, built for the centiles of per capita familiar income are similar in the three samples. There is a small growth of income in the lower centiles, seeing that, the individuals in the following centiles present growth, until reaching a peak in centile 19 for Samples I and II and in centile 25, for Sample III. From that on, there is a relative stability in the income growth up to centile 98, when a new growth occurs. Comparing the simulated result to the observed one, we can say that the change of coefficients of participation equation would lead to a smaller growth of income in all centiles. However, the process would be more accentuated for the individuals of lower centiles of distribution. This way, results of Figures (d), (e) and (f) point out to the same direction as to the results of Table 5, that is, the shift of women's participation contributed for the inequality growth of the Brazilian per capita familiar income.

## 4- Final Remarks

The 80's and 90's were known by the development of female activity in Brazil. Besides that shift, intense modification of Brazilian women's profile occurred, such as: decrease of married women ratio and increase of female headed households, increase of women residing in urban areas and decrease of mean average of kids. Besides, female educational level also improved.

Several shifts women went through that period - in particular, the ratio increase of female participation - may have potentially influenced the impact on the growth of income inequality. The growth of participation affected greatly the inequality measures of individual income, seeing that, the growth of participation contributed to decrease, or at least not to increase even more, the income inequality. On the other hand, this result changed when we analyzed the effect of female activity shift on per capita familiar income, leading us to conclude that women would have contributed to increase the inequality of per capita income distribution. This is very important considering that there is a need for further studies which may allow a better understanding of the family members inter relation, concerning their decisions of participation in the labor market and the consequences of their decisions in terms of familiar income production.

In short, we can say that the shift in women's characteristics in the last twenty five years provides a challenge of how to translate female participation into a better income distribution. Female rising educational level marches to that direction. However, this element is not sufficient to alter the present picture of low income appropriation by the poorest families, at least in a short period of time. The truth is that no matter how much poorer women have experimented income growth, this is not enough to change the perverse picture of Brazilian per capita familiar income.

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