The myth of post-reform income stagnation: Evidence from Brazil and Mexico

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The Myth of Post-Reform Income Stagnation: Evidence from Brazil and Mexico

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Abstract:

Economic policies are often judged by a handful of statistics, some of which may be biased during periods of change. We estimate the income growth implied by the evolution of food demand and durable good ownership in post-reform Brazil and Mexico, and find that changes in consumption patterns are inconsistent with official estimates of near stagnant incomes. That is attributed to biases in the price deflator. The estimated unmeasured income gains are higher for poorer households, implying marked reductions in "real" inequality. These findings challenge the conventional wisdom that post-reform income growth was low and did not benefit the poor.

JEL Classification Numbers: D12; E01; I32; O10.

Keywords: Household consumption; Measurement Error; CPI Bias; Economic reform; Trade liberalization; Inflation stabilization.

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1. Introduction

The merits of economic policies are often judged by their impact on a handful of statistics, such as unemployment, GDP growth, inflation, and poverty incidence. While these statistics are often treated as if they were very precise numbers, their construction involves a number of assumptions, competing methodological choices, and resource constraints. This inherent imprecision can be exacerbated during periods of significant changes in economic policies. More worryingly, some changes in policy may systematically bias the errors in one direction. For example, market reforms which liberalize trade and introduce new and higher quality goods to the domestic market can cause the Consumer Price Index (CPI) to overstate inflation if the CPI basket is outdated, fails to account for new goods, or if there are limited or no provisions for quality adjustments. In this paper, we explore two different methods to gauge the improvement in real household income and consumption after reform episodes in Brazil and Mexico: the evolution of Engel curves for food consumption and changes in patterns of durable goods ownership.

During the late 1980s/1990s much of the developing world embarked in a market-oriented reform process. Latin America provides an excellent setting to study these reforms.[2] Throughout the region, countries opened, privatized and liberalized much of their economies (these changes are documented and quantified in Lora 1997). However, beginning as early as 1995, several observers have questioned whether Latin America's post-reform growth had been disappointing (e.g., Krugman 1995 and Easterly, Loayza, and Montiel 1997). This disappointment can be illustrated by the experience of Brazil and Mexico, which account for over half of Latin America's population and for over 60 percent of its output. After decades of high growth, these economies contracted in the 1980s after the balance of payments crises in the beginning of that decade (Figures 1-2).[3] The conventional wisdom at the time

[2] The post-reform experience in other regions is influenced by other factors likely conducive to low growth. Output collapse in Eastern Europe may result from the particular challenges of moving from a centrally planned to a market economy (Blanchard and Kremer 1997). Africa's growth was hindered by conflict (Easterly and Levine 1997).

[3] Although the focus of this paper is on real household income growth, GDP growth was used in Figures 1-2 owing to its broader historical coverage. The evolution of real household income is similar to the latter for the years in which it is available.

attributed their stagnation to the failure of their import-substitution strategy to cope with adverse external shocks. When Brazil and Mexico began to lower trade barriers and liberalize their economies (late 1980s/early 1990s in the case of Brazil and mid-1980s in the case of Mexico), there was an expectation that strong economic growth would resume. After all, while there is debate about the optimal sequencing and implementation of reforms, most economists would agree that removing trade barriers that choke most manufacturing trade including intermediate and capital goods (e.g. Brazil's outright ban on imports of personal computers), dismantling of inefficient state monopolies, reducing black market premia, and ending high- or hyper-inflation should noticeably improve economic performance over the medium-term. However, it appears that was not the case (Figures 1-2). We propose an alternative explanation to these puzzling developments: official statistics underestimated the growth in real income after the market-oriented reforms.

We compute real per capita household income growth in Brazil and Mexico implied by Engel curves for food consumption, following the method in Nakamura (1997), Hamilton (1998, 2001a), and Costa (2001) for the United States. One of the strongest empirical regularities in economics is that the share of food in total household expenditures declines as (real) income grows (Engel's law). We estimate a model for the household-level budget share of food as a function of real expenditure, relative prices, and household characteristics, using different cross-sectional surveys of household expenditure. Assuming nominal expenditure is measured accurately and preferences are stable, we attribute the difference between the real expenditure growth based on our Engel curve estimates and the "headline" real expenditure growth obtained by deflating nominal income by the CPI to measurement error in the latter. Using this method, Hamilton (2001a) and Costa (2001) estimate real household income growth in the U.S. since the 1980s to be roughly 1 percent per year higher than implied by nominal income deflated by the CPI. Their estimates are similar to those of the Boskin Commission, which estimated CPI bias at 1.1 percentage points per year in 1995–96 (Boskin and others, 1996). Our estimates for Brazil during 1987/88–2002/03 point to an underestimation of real household income growth of at least 3 percent per year. Our estimates for Mexico during 1984–2006 point to an underestimation of 2½–3½ percent per year.[4]

It is difficult to identify the source of the estimated biases without replicating this paper for a large sample of countries with different experiences and characteristics. Nevertheless, the transformation brought upon by trade liberalization seems to be a plausible explanation. There is an extensive literature arguing that changes in the CPI overestimate the increase in cost of living in the United States.[5] The main sources of CPI bias include consumer substitution, improvements in the distribution of goods, the late introduction of new goods into the CPI basket and failure to (fully) account for improvements in quality. When Brazil and Mexico opened their economies to international trade, these sources of bias were amplified. Many goods that were previously expensive (relative to world prices) suddenly became more affordable as domestic prices of tradables converged towards the lower international levels. This large shock in relative prices led to changes in consumption patterns contributing to substitution bias in the CPI.[6] This could help explain the large CPI bias we estimate for Brazil and Mexico. It is worth noting that trade liberalization also led to a substantial improvement in the quality of goods available (not only through imported goods, but also through competitive pressures on import-competing domestic goods), which is not captured by the empirical strategy we use (suggesting an even larger bias).

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[4] While the empirical strategy used only allows us to identify the bias in the CPI, it is unlikely that other price deflators (e.g. GDP deflator) are accurately measured if the bias in the CPI is large (which could explain the apparent stagnation in Figures 1-2).

[5] For an overview of this literature, see National Research Council (2002), Lebow and Rudd (2003), and Hausman (2003).

[6] Some of this decline in price was never captured by the CPI because of outdated consumption baskets. Although the time intervals between updates in Brazil's CPI basket are not unusually long, their timing was such that the first update reflecting post-liberalization consumption only took place in 1999 (based on 1995/96 consumption patterns), a decade after the liberalizing reforms began. In Mexico, the first post-liberalization update took place in 1994 (based on 1989 consumption patterns). For instance, personal computers were available in Brazil, albeit at a very high price, in the 1980s, yet their prices were only included in the CPI in 1999.

The hypothesis that reforms led to large one-time improvements in real household income is supported by our results for early reformer Mexico, where the estimated bias is concentrated in the 1984-1996 period, and declines to about 1 percent per year from 1996 onwards, a level comparable to bias estimates for the U.S. The hypothesis of one-time gains is also supported by the evolution of food consumption in Brazil between pre-reform expenditure surveys in 1974/75 and 1987/88, which does not suggest a bias in the CPI.

Our paper also focuses on distributional issues as it uses a semi-parametric methodology that allows for variation in the estimated bias at different points in the expenditure distribution. Our methodological contribution to this literature is to weight these expenditure-specific bias estimates by household expenditure, which is the relevant aggregation for comparison with the CPI (which is itself based on an aggregate consumption bundle where individual household expenditure shares are weighted by their total expenditure). This weighting is particularly important when income inequality is high (as in Brazil and Mexico). We find larger unmeasured improvements in real income for poorer households, implying a substantial reduction in income inequality when measured in real terms, particularly for Brazil. The poor have benefited from a decline in the relative price of food in Brazil (since food has a large weight in their consumption basket, e.g. de Carvalho Filho and Chamon, 2008b). Perhaps more importantly, they may have disproportionately benefited from improvements in the distribution of goods.[7] Finally, the end of hyper-inflation in Brazil in 1994 also disproportionately benefited the poor. Richer households had access to interest-earning bank accounts and government bonds with daily compounding. Poorer households on the other hand, kept more of their monthly wages in cash, bearing a higher inflation-tax, which made them relatively worse-off even after adjusting their income by CPI inflation (Neri 1995)[8].

[7] This is somewhat similar to the finding in Hamilton (2001b) that evidence from Engel curves suggest that black standard of living in the U.S. improved by 15 percent in 1974-1991 relative to whites even after accounting for gains in nominal income, perhaps because blacks initially faced higher prices than whites.

[8] Sturzenegger (1992) presents a model where endogenous financial adaptation protects the rich from inflation while leaving the poor vulnerable, and affects the socially optimal level of inflation.

Other papers have used food Engel curves to estimate CPI bias. Beatty and Larsen (2005) use semi-parametric Engel curves to estimate a bias of 1.5 percent per year in Canada during 1978-2000. Larsen (2007) estimates a negative CPI bias in Norway (i.e. food Engel curves implying the CPI understates the increase in the true cost of living). Barrett and Brzozowski (2010) estimated CPI bias in Australia of about 1 percent per year over the 1975/76 – 2003/04 period. For Korea, Chung, Gibson and Kim (2010) find a bias of slightly lower than 1 percent per year in 2000-2005. Among developing countries, Gibson, Stillman and Le (2008) estimated CPI bias in Russia in 1994-2001 as about 1 percent per month, and argue for a substantial understatement of growth in Russia during the transition period.[9] In a separate paper, we estimate the CPI bias in urban China during 1993-2005 to be less than 1 percent per year (de Carvalho Filho and Chamon 2008a).[10] Langebaek Rueda and Edgar Caicedo Garcia (2007) found CPI bias in Colombia, for the years 1984/85 to 1994/95, a period of market oriented reforms, of about 1.6-1.7 percentage points per year. Their findings match the pattern we find of relatively large biases in reforming Latin American countries.

In addition to the contributions on the distributional dimension of the CPI bias and its implications for CPI aggregation, this paper also innovates by using estimates of the durable goods demand sensitivity to income, combined with actual changes in durable good ownership, as an additional source of evidence on CPI bias. The estimates based on the demand for durable goods closely match the ones based on the Engel curve for food. Thus the evidence amassed in this paper points to a substantial improvement in the material well-being of households, and stands in sharp contrast to the conventional wisdom of post-reform economic stagnation.

The remainder of this paper is organized as follows. Section 2 describes the methodology. Sections 3 and 4 provide some background information on policy changes,

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[9] It is possible that challenges related to deflating past expenditures in a high-inflation environment have caused an overestimation of the food budget share during years of high inflation (since typically the recall window to measure non-food expenditures is longer), thereby leading to implausibly large CPI bias estimates in Russia. It is also possible that estimating an expenditure level-specific bias and weighting those estimates by household expenditure as we do would lower their bias estimates for Russia.

[10] Gong and Meng (2008) use Engel curves to estimate regional price differences in China.

describe the data and present the results for Brazil and Mexico respectively. Finally Section 5 concludes.

2. Empirical Methodology

Our parametric estimates of real income growth follow the method developed in Hamilton (1998, 2001a), building on an insight by Nakamura (1997). We start with the demand function for food that emerges from Deaton and Muellbauer's (1980) Almost Ideal Demand System:

$$\begin{aligned} w_{i,j,t} &= \phi + \gamma(\ln P_{F,j,t} - \ln P_{N,j,t}) \\ &+ \beta(\ln Y_{i,j,t} - \ln P_{G,j,t}) \\ &+ \sum_x \theta_x \mathbf{X}_{i,j,t} + \mu_{i,j,t}\,, \end{aligned} \tag{1}$$

where the subscripts refer to household i, region j, and period t; w is the share of food in total household expenditures; P_F, P_N and P_G are the true but unobservable price indices of food, nonfood and the general index for all goods; Y is the household's nominal expenditure; $\mathbf{X}$ is a vector of household characteristics; and μ is the residual. A negative β (downward sloping Engel curve) characterizes a necessity good while a positive β (upward sloping Engel curve) characterizes a luxury good. The true price index P_G is measured with CPI error. Let $\Pi_{G,j,t}$ denote the percent cumulative increase in the CPI measured price and $E_{G,j,t}$ denote the percent cumulative measurement error from period *0* to period t, for food, nonfood or all goods, as indicated by the subscript, where $P = (1+\Pi)(1+E)$. Equation (1) can be rewritten as:

$$\begin{aligned} w_{i,j,t} &= \phi + \gamma(\ln(1+\Pi_{F,j,t}) - \ln(1+\Pi_{N,j,t})) \\ &+ \beta(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t})) \\ &+ \gamma(\ln P_{F,j,0} - \ln P_{N,j,0}) - \beta \ln P_{G,j,0} \\ &+ \gamma(\ln(1+E_{F,j,t}) - \ln(1+E_{N,j,t})) - \beta \ln(1+E_{G,j,t}) \\ &+ \sum_x \theta_x \mathbf{X}_{i,j,t} + \mu_{i,j,t}\,. \end{aligned} \tag{2}$$

We assume that the CPI measurement error does not vary geographically, and rewrite (2) as:

$$
\begin{aligned}
w_{i,j,t} = {} & \phi + \gamma(\ln(1+\Pi_{F,j,t}) - \ln(1+\Pi_{N,j,t})) \\
& + \beta(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t})) \\
& + \sum_j \delta_j D_j + \sum_t \delta_t D_t \\
& + \sum_x \theta_x \mathbf{X}_{i,j,t} + \mu_{i,j,t},
\end{aligned}
\tag{3}
$$

where D_j and D_t are regional and time dummies and:

$$\delta_j = \gamma(\ln P_{F,j,0} - \ln P_{N,j,0}) - \beta \ln P_{G,j,0} \tag{4}$$

$$\delta_t = \gamma(\ln(1+E_{F,t}) - \ln(1+E_{N,t})) - \beta \ln(1+E_{G,t}) \tag{5}$$

All the terms in Equation (3) are observable and once the equation above has been estimated, we are ready to compute the cumulative CPI bias. If food and nonfood are equally biased (i.e. $E_{F,t} = E_{N,t}$), then:

$$\ln(1+E_{G,t}) = -\delta_t / \beta \tag{6}$$

It seems likely that mismeasurement is less of a problem for food than for nonfood goods. As a result, to the extent that food is a necessity ($\beta < 0$) and food shares increase with the relative price of food ($\gamma > 0$), one can show that equation (6) understates the bias for small positive values of γ as in our estimates.[11]

The parametric specification discussed above assumes that all households at a given date face the same bias. In the context of a high inequality country, it is particularly relevant to inquire about differential effects across the income distribution. The estimation of (3) through minimization of squared errors yields an estimate of the bias for the average household. However, the actual CPI index is based on an aggregate consumption bundle that by design disproportionately represents richer households as they account for a disproportionate share of aggregate consumption. Thus, to the extent that the discrepancy

[11] Note that this formula yields a multiplicative bias. That is, if the change in the CPI in a given period was 10 percent and its estimated bias is 3 percent, the estimated change in the true cost of living would be 6.7 percent, since (1-0.03)·1.1=1.067. If the change in the CPI was 100 percent, the estimated change in the true cost of living would be 94 percent (since 1-0.03) ·2=1.94.

between the true cost of living index and the headline CPI varies across the income distribution, there might be substantial differences between the bias facing the average household and the bias for the aggregate consumption bundle, which is the one that maps to the CPI.

The model in equation (3) can be extended to address this concern, by assuming the bias is a linear function of the log of real expenditure:

$$\ln(1+E_{G,i,t}) = a_t + b_t(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}))$$

Maintaining the same assumption that food and non-food are equally biased and that the bias does not vary by region, one can estimate:

$$\begin{aligned} w_{i,j,t} &= \phi + \gamma(\ln(1+\Pi_{F,j,t}) - \ln(1+\Pi_{N,j,t})) \\ &+(\beta + \sum_t \lambda_t D_t)(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t})) \\ &+\sum_j \delta_j D_j + \sum_t \delta_t D_t \\ &+\sum_x \theta_x \mathbf{X}_{i,j,t} + \mu_{i,j,t}, \end{aligned}$$

and obtain the following expression for CPI bias at different points in the expenditure distribution.

$$\ln(1+E_{G,i,t}) = \frac{-\delta_t}{\beta} - \frac{\lambda_t}{\beta}(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t})) \tag{7}$$

Finally, we can use a flexible non-parametric approach for the bias by estimating a semi-parametric version of the demand function (1), allowing for estimation of the bias at different levels of expenditure. Still under the assumption that food and non-food are equally biased and that the bias does not vary by region, we can rewrite (3) as:

$$\begin{aligned} w_{i,j,t} &= \phi + \gamma(\ln(1+\Pi_{F,j,t}) - \ln(1+\Pi_{N,j,t})) \\ &+\sum_t D_t f_t(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}) - \ln(1+E_{G,i,t})) \\ &+\sum_x \theta_x \mathbf{X}_{i,j,t} + \mu_{i,j,t} \end{aligned}$$

We estimate $f_t(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}) - \ln(1+E_{G,i,t}))$ non-parametrically using the differencing method proposed in Yatchew (1997). In a nutshell, that method consists of ordering households by their (CPI-measured) real income, and (higher-order) differencing the equation above so as, to an approximation, eliminate the terms involving *f()*. We are then able to estimate the parametric terms $\hat{\phi}, \hat{\gamma}, \hat{\delta}_j$ and $\hat{\theta}_X$, and thus obtain the non-parametric term:

$$f_t(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}) - \ln(1+E_{G,t}) = w_{i,j,t} - \hat{\phi} - \hat{\gamma}(\ln(1+\Pi_{F,j,t}) - \ln(1+\Pi_{N,j,t})) - \sum_x \hat{\theta}_x \mathbf{X}_{i,j,t},$$

and estimate $f_t()$ using a locally-weighted linear regression with quartic kernel weights. The bias at a given CPI-measured real income level at time *t* is then estimated as the increase in CPI-measured real income that would have prevented the Engel curves from shifting. That is, we solve at each expenditure level for the value of $E_{G,i,t}$ that satisfies:

$$\hat{f}_t(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}) - \ln(1+E_{G,i,t})) = \hat{f}_0(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t}))$$

3. Brazil

3.1 Background on Economic Policy Changes

For decades, Brazil was one of the world's most closed economies, with high tariff and non-tariff barriers, including outright bans on the importation of several goods for which a domestic variety was available. In the late 1980s/early 1990s, Brazil began to cut tariffs substantially and to dismantle most non-tariff barriers; and by the mid-1990s it was already a relatively open economy (particularly in comparison with its former self).[12] Although import penetration remained relatively small as a percentage of GDP, competitive pressures following trade liberalization led to dramatic improvements in the productivity of Brazilian manufacturing firms.[13] Hyper- and high-inflation was finally halted in 1994, with the *Plano*

[12] The (unweighted) average tariff for 53 final goods collected by Kume, Piani, and de Souza (2000) declines from 55 percent in 1987 to 31 percent in 1990, and to only 11 percent in 1995. Using different sources, Lora (2001) documents a similar pattern of average tariff reductions in Brazil.

[13] Ferreira and Rossi (2003) attribute an increase in annual Total Factor Productivity (TFP) growth in Brazilian industry of more than 5 percentage points per year during 1991–97 to the effects of trade liberalization. Muendler (2001) finds that increased foreign competition pressured Brazilian

(continued…)

Real, and since then, annual inflation has remained mostly in the single digits. Privatization efforts were concentrated in the late nineties (Lora, 2001).

3.2 Data

The *Pesquisa de Orçamentos Familiares* (henceforth, POF) is the household income and expenditure survey carried out by the *Instituto Brasileiro de Geografia e Estatística* (IBGE), based on probabilistic sample and stratified design. The POF surveys were conducted in 1987/88, 1995/96, and 2002/03 (trade liberalization was concentrated between the first and second surveys; while privatization of services such as telecommunication was concentrated between the second and third surveys). The 1987/88 and 1995/96 POF surveys are representative of the household population in 9 metropolitan regions, encompassing more than 30 percent of the Brazilian population in 2000 and matching the geographical coverage of the IPCA consumer price index.[14] The 2002/03 POF was the first to cover a nationally representative sample. In order to make the samples comparable over time, we limit the 2002/03 sample to the geographical sub-sample also covered by the previous surveys. The POF contains household expenditure information on individual goods at a disaggregated level, households' inventory of durable goods, and also demographic, schooling and income characteristics of each household member.

IBGE produces and disseminates the *Índice de Preços ao Consumidor Ampliado* (IPCA), which is considered Brazil's official consumer price index, and is based on the basket consumed by the POF families (which is therefore, by construction, the relevant sample in which to base our exercise). The IPCA is also the broadest measure of consumer prices available, covering all families earning between 1 and 40 times the minimum wage in the geographical areas covered by the POF. A few years after each POF is collected, the IPCA index is reweighed, hence catching-up with changing consumption patterns of

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manufacturers to raise productivity markedly during the same period. Hay (2001) documents dramatic efficiency improvements in large Brazilian manufacturing firms.

[14] Belém in the Northern region; Fortaleza, Recife and Salvador in the Northeast; Rio de Janeiro, Belo Horizonte and São Paulo in the Southeast; Curitiba and Porto Alegre in the South, plus the Distrito Federal and the municipality of Goiânia, in the Center-West region. Goiânia is excluded from our sample because it did not have its own region-specific price index prior to 1991.

Brazilian households. Throughout the paper, we will be referring to the IPCA whenever the CPI in Brazil is mentioned.

The POFs are conducted over a 12-month period. The field work took place in: March 1987–February 1988, October 1995–September 1996, and July 2002–June 2003. Households are asked to provide information on their expenditures, which are collected based on different reference periods depending on the type of expenditure and its frequency. Four reference periods are used: 7 days, 30 days, 90 days, and a longer recall period of either 12 months (2002/03 POF) or 6 months (1987/88 and 1995/96 POFs).[15]

The POF uses six collection instruments for its data: Housing conditions questionnaire, collective expenditures questionnaire, collective expenditures notebook, individual expenditures questionnaire, individual income questionnaire and living conditions questionnaire. Expenditures on frequently purchased items by the household were recorded on the collective expenditures notebook. Those items include mainly food, beverages, cleaning and personal hygiene products. That notebook was kept during a 7-day period by the person who manages those expenditures in the household budget. During that period, the expenditures on those frequently purchased products were recorded, as well as the amount and location of purchase. This notebook is the source of information for expenditures on food and beverages to be consumed at home. Expenditures on food and beverage consumption outside the home were recorded on the individual expenditures questionnaire, which was based on recall over a seven-day reference period. Our measure of food consumption covers food and beverages. Our measure of expenditure covers all expenditures except financial savings, but does not include the rental value of owner-occupied housing (which is not available for all the years of the POF).[16] Data limitations also prevent us from considering the consumption of self-produced food. Given the sample only covers metropolitan areas, it is safe to assume that the latter is negligible.

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[15] For the potential effects of differences in recall windows on the measurement of expenditure, see Deaton and Kozel (2005) on Indian data.

[16] The finding of a large positive CPI bias for Brazil in this paper is robust to restricting to a sample of tenants (de Carvalho Filho and Chamon, 2006).

Given the length of the collection, reference and recall periods, expenditure data can span a 24-month period. In order to make the data comparable across households surveyed over different periods, IBGE deflates or inflates all expenses using item-specific deflators to a same reference date.[17] The reference dates for the POFs are: October 15, 1987; September 15, 1996; and January 15, 2003.

Figure 3 shows the relative prices for different groups of goods (vis-à-vis the aggregate price index). Note the persistent increase in the relative price of most "nontradable" goods (e.g., housing and health care), and the large decline in the relative price of most "tradable" goods (e.g., clothing and household supplies) over time. Similar trends are experienced in other countries, and fit the prior of greater technological progress in the tradable sectors. This trend seems to sharpen around the early 1990s, suggesting that trade barriers were previously dampening these secular changes in the relative price of tradable and nontradable goods. It is also remarkable that the changes in relative prices in the aftermath of the January 1999 exchange rate devaluation (when the real lost about 35 percent of its value in real effective terms) are dwarfed by previous movements in relative prices such as those observed right after stabilization plans.[18] The divergence in relative prices shown in Figure 3 highlights the potential for CPI mismeasurement from outdated weights in the consumption basket and substitution bias.

3.3 Results

Our first estimates are based on the specification on equation (3). Following Costa (2001), we use total expenditure instead of income because expenditure better reflects permanent

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[17] For example, if a household bought rice and beans on March 4, 2003, those two expenditures are deflated to January 15, 2003 values using two different deflators. If prices are not collected for a given item, the price for its sub-group is used.

[18] The pass-through from the exchange rate to inflation has been relatively small in Brazil, although it tends to have a larger effect on tradable goods (e.g. food). For estimates of pass-through to different goods for both Brazil and Mexico, as well as a discussion of its distributional implications, please refer to de Carvalho and Chamon (2008).

income.[19] Our specification allows for regional variation in relative prices. The controls for household characteristics include dummies for home ownership and rental; gender of the head of the household; presence of a spouse; whether the head of the household, the spouse, or both have labor income; and the number of household members in age groups: 0 to 4; 5 to 9; 10 to 14; 15 to 19; and 20+ years old.

Hamilton (1998, 2001a) uses the education of the head and of the spouse as controls, finding a significant negative effect of education on the food budget share. If we add dummies for the education attainment of the head and spouse, the coefficient on log expenditure becomes less precisely estimated since education is highly correlated with the latter.[20]

Our discussion will consider two samples, one including all households (henceforth, "full" sample), the other including all but the 3 percent poorest (measured by expenditure levels) in each survey year (henceforth, "97 percent" sample). The rationale for the latter is to provide a robustness check with regards to the finding of non-monotonicity of Engel curve at the left tail of the expenditure distribution (we will elaborate on this later on the paper). Our overall results are not affected by excluding the 3 percent poorest because those households are not included at the target population of the inflation index, and their inclusion or exclusion in the sample has a negligible effect on the aggregate (expenditure-weighted) CPI bias estimates.

Table 1 provides summary statistics for the "full" and "97 percent" samples. The average budget share of food was 30¼ percent, 25½ percent, and 22½ percent in the full sample of the 1987/88, 1995/96, and 2002/03 surveys, respectively. The mean food shares for the "97 percent" sample are close to the full sample. The relative price of food declined

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[19] Expenditures better reflect permanent income particularly in places and for people that have access to market or non-market mechanisms to smooth consumption. To the extent that the poor have less access to consumption smoothing and food is a necessity, then their food shares would increase relatively more after a bad shock.

[20] When head and spouse education are added as dependent variables, bias estimates in general become larger and less precisely estimated. Estimating the coefficient on the education variables with the differencing method used in the semi-parametric regressions would also be challenging since households at a similar level of expenditure tend to have similar educational attainment.

between the first two surveys but remained stable afterwards. Per capita expenditure on food deflated by the food CPI has increased, even though the food share in the budget has declined. Average household size declined considerably, as indicated by the stagnant average total household expenditures despite a sizable increase in per capita expenditure. Family composition changed over time, with a reduction in the households where a spouse was present and an increase in the households not headed by a male. The surveys also indicate an increase in the likelihood of the head of the household not having labor income, perhaps due to a combination of weaker labor market conditions, shifting demographics, and expansion of entitlement programs.[21] Spousal labor remained more stable.

Since expenditures are potentially measured with error, we also present estimates where current income is used as an instrument. This is particularly important because attenuation bias would tend to drive down the absolute value of the coefficient on the log of expenditure (slope of the Engel curve), hence increasing the estimate of the CPI bias in equation (6).

Parametric Model

Table 2 reports the linear regression results for the full sample, pooling the three surveys. The first two columns estimate the bias assuming it constant across households in a same year (as in Hamilton, 1998, 2001a; Costa, 2001). The third and fourth column present bias estimates that vary linearly with the log of real expenditure, as in equation (7). The coefficients on the time dummies are negative and statistically significant, suggesting that food shares declined by more than would be predicted based on relative price and household characteristics. The estimated coefficients on log expenditure range from -0.061 to -0.097, which implies income elasticities in line with previous estimates for Brazil (in the 0.6-0.7 range, e.g., Asano and Fiuza, 2001).[22] The coefficients on the log of the relative price of food

[21] For evidence on a sizeable effect of cash benefits on labor participation in Brazil, see de Carvalho Filho, 2008.

[22] The formulas for income and price elasticities are: $\eta_y = 1 + \beta/w$ and $\eta_p = -1 + (\gamma - \alpha\beta)/w$ where α is the share of food in the price index and w is the share of food in total expenditure.

are not statistically significantly different from zero, which may reflect competing income and substitution effects, or limited regional variation in relative prices of food.[23] Table 2 also reports the resulting estimate for the cumulative CPI bias, $E_{G,t}$, which is negative as expected (the values reported correspond to its absolute value). Given that negative bias, the implied gross change in the true cost of living is $(1-|E_{G,t}|)$ times the gross change in the CPI, and the resulting gross true real income growth is $1/(1-|E_{G,t}|)$ times the gross real income growth obtained by deflating nominal income by the CPI.

When the CPI bias is assumed constant across households, the OLS based estimates imply an annualized CPI bias of 9½ percent per year in 1987/88-1995/96 and about 5 percent per year in 1995/96-2002/03 (column 1). Those numbers represent the upper bound of all our estimates across different methods and samples. When instrumental variables are used to reduce attenuation bias and endogeneity (expenditure levels are in the denominator of the food share as well as in the right-hand side), these figures drop to about 6½ and 3 percent per year respectively (column 2).

When the CPI bias is assumed to vary linearly with the log of (CPI-measured) real expenditures, the bias for the average household (population weighted average) is about 7½ percent per year in 1987/88-1995/96 and 3¾ percent per year in 1995/96-2002/03 (column 3). Those estimates drop to 6 and 2¾ percent respectively when income is used as an instrument (column 4). Finally, the expenditure weighted averages are 2.9 percent per year in 1987/88-1995/96 and 3½ percent per year in 1995/96-2002/03 (column 3). These estimates increase to 3.1 and 4¼ percent per year respectively when income is used as an instrument (column 4). For the whole period 1987/88-2002/03, the expenditure-weighted CPI bias is 3.6 in the preferred parametric specification (IV estimation with bias varying linearly with log of expenditure).[24]

[23] This is not likely to affect the thrust of our results since the relative price of food remained roughly constant between the second and third surveys, yet we find substantial CPI bias during that period.

[24] The bias estimates become larger if we include dummies for whether the head and the spouse have secondary or higher education as additional controls in the OLS based specifications, but are virtually unchanged in the IV ones.

Semi-Parametric Model

We now turn to the semi-parametric estimation of the model. Figure 4 shows the non-parametric estimates of the relationship between the food budget share and the log of real expenditure, for the full and the 97 percent samples. As expected, the food budget share declines with real expenditure, and the curves have shifted downward with each survey. However, using the full sample, for low levels of expenditure, this relationship is non-monotonic (Figure 4, left chart), but the proportion of households in that expenditure range is relatively small and their contribution to the estimated expenditure-weighted aggregate CPI bias is negligible. To illustrate this point, Figure 4, right chart shows that the non-monotonic portion of the Engel curve disappears when the observations with expenditure below the 3rd percentile of each survey are excluded (97 percent sample).

As discussed in Section 2, the food share-real expenditure profiles in Figure 4 map into a CPI bias-real expenditure profile by computing the necessary change in real expenditure, at each real expenditure level that would maintain the Engel curves in the same position. For example, for each level of expenditure in the 1995/96 survey, CPI bias is determined by its difference to the real expenditure level in 1987/88 that was associated to the same food budget share. For CPI measured real expenditure R we can solve for the corresponding bias in $E_{1995/1996}(R)$ as the solution to:

$$\hat{f}_{1995/96}(\ln R - \ln(1 + E_{1995/96}(R)) = \hat{f}_{1987/88}(\ln R), \quad (8)$$

where $\hat{f}_{1995/96}$ and $\hat{f}_{1987/88}$ are the estimated Engel curves for 1995/96 and 1987/88, controlling for all the right-hand side variables used in the regressions reported in Table 2.

Since we rely on Engel's Law, such mapping is only meaningful in the range where the food share is declining on income. The data, however, sometimes shows non-monotonicity of the Engel curve for low levels of expenditure. To address this issue, at the left tail of the expenditure distribution, we set the bias to equal that of the first expenditure level R^L in the declining range of the curve for which the food share is below the one for the lowest level of expenditure $R_{\min}$. While we acknowledge that is an imperfect fix, this adjustment has a negligible effect on the aggregate (expenditure-weighted) bias estimates. In more precise terms, we impose the constraint:

$$E_{1995/96}(R) = E_{1995/96}(R^L) \text{ if } R < R^L,$$
$$\text{where } R^L = \hat{f}^{-1}_{1995/96}(\hat{f}_{1995/96}(R_{\min})), \hat{f}'_{1995/96}(R^L) < 0$$

At the right-tail of the expenditure distribution, there are levels of the food share in 1995/96 for which there are no counterparts in the 1987/88 Engel curve. In those cases we set the bias to equal that of the highest expenditure level R^U in the 1995/96 curve for which the mapping into the 1987/88 curve is still possible (to the highest value of expenditure $R_{\max}$). That is:

$$E_{1995/96}(R) = E_{1995/96}(R^U) \text{ if } R > R^U,$$
$$\text{where } R^U = R_{\max}(1 + E_{1995/96}(R^U))$$

Similarly, the bias from 1995/96 to 2002/03 is estimated by computing the real expenditure growth that would make the 2002/03 Engel curve match the one for 1995/1996, subject to the same adjustments at the tails of the expenditure distribution.

Figure 5A shows the estimated annual bias from 1987/88 to 1995/96 as a function of headline real expenditure, as well as the estimated density function of the log of CPI-deflated real expenditure in 1995/96 for the full and 97 percent samples. The bias is higher for the poorest households, and declines gradually as real expenditure increases. The annualized bias for the average household is above 6 percent per year, whereas the expenditure weighted aggregate bias is slightly below 3 percent per year, for both samples. Figure 5B shows the estimated annual bias from 1995/96 to 2002/03. The profile is flatter than the one in Figure 5A, suggesting that the differences in CPI bias across the income distribution have narrowed. The bias for the average household is about 4¾ percent while the expenditure weighted aggregate bias is 4.3 percent for both samples.

Robustness

Table 3 reports the estimated biases under different methods and samples. The first two groups of estimates (Panels A and B) are based on the parametric models from Table 2. The third set of estimates (Panel C) is based on the semi-parametric approach discussed above. It presents results for four different samples: the full sample, the 97 percent sample which excludes the bottom 3 percent of the expenditure distribution, a “compliant” sample (e.g. comprising only those households that turned their collective expenditure notebook with at least one expenditure recorded), and a sample constrained to households whose income is between 1 and 40 times the minimum wage (the target population for the IPCA).[25]

The bias estimates under the assumption of constant bias across the income distribution are reported in Table 3, Panel A. Bias estimates are always larger for the full sample than elsewhere; and for all sub-samples, OLS based bias estimates were larger than the IV based ones. For the first period, from 1987/88 to 1995/96, the population sample weighted bias estimates are always above 4½ percent per year; whereas for the second period, they are always larger than 2½ percent per year. For each of the samples and estimation methods, the estimated bias declined during the second period.

At Panel B, we report the estimates for the bias varying linearly with log of expenditure as in equation (7) above. During the first period (1987/88–1995/96), the weight of the evidence suggests that CPI during this period was more overstated for poorer than for richer households. Expenditure-weighted bias estimates range from about 2 to 3¼ percent. During the second period (1995/96 to 2002/03), the results do not clearly indicate whether the bias was more pro-poor or not. OLS estimates show slightly smaller point estimates for the expenditure-weighted bias, but IV estimates show otherwise. The population-weighted bias estimate ranged from 2½ to 3¾ percent for all samples and methods, and the estimates of expenditure-weighted bias, which maps to the CPI aggregate basket, ranged from 2¾ to 4¼ percent. Thus the expenditure-weighted bias estimates show a pattern of increasing CPI bias over time.

At Panel C, we show that for all different samples, the semi-parametric estimates imply declining population-weighted bias over time, from a range of 5¼-6½ percent in the first period to 3½-5 percent in the second period. However, the opposite trend was found for expenditure-weighted bias, which increased from a range of 1½-3 percent to 4¼-5 percent.

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[25] Such constraint excludes from the regression sample 2.7 percent, 2.7 percent and 4.7 percent of the weighted households respectively in the 1987/88, 1995/96, and 2002/03 original samples due to total income lower than 1 minimum wage, and 6.6 percent, 5.3 percent and 5.5 percent, respectively, because household income is greater than 40 minimum wages.

The bias estimates for all but one sample and period were statistically significant. The only exception was the estimate for the expenditure-weighted bias during the first period, using the sample with income between 1 and 40 minimum wages. Finally, in the 97 percent sample, which has the desirable feature of focusing only on the monotonic portion of the Engel curves, the bias mapping to the CPI was 2¾ percent in 1987/88 and 4¼ percent between 1995/96–2002/03 (which bracket 3.6 percent, the expenditure-weighted bias for the 1987/88-2002/03 period in the full sample, IV estimation and bias varying linearly in log of real expenditure).[26]

Distributional Implications

The bias estimates shown in Figures 5A and 5B imply a substantial decline in "real" inequality. Figure 6A plots the distribution of CPI-deflated real expenditures, while Figure 6B plots the distribution of our estimates of the true real expenditure implied by the Engel curves for each of the surveys. The curve for 2002/03 is the same as the one in Figure 6A, while the curves for 1987/88 and 1995/06 are adjusted based on our estimates (with the bias being measured relative to 2002/03).[27] While the distributions in Figure 6A are virtually overlapping,[28] Figure 6B indicates a marked improvement, with the distribution moving to

[26] It would have been interesting to check the robustness of our estimates when the rental value of owner-occupied housing is considered among the household expenditures (and when computing the food budget share). Unfortunately, that information was not available. Our results are very similar if we exclude expenditures on rent when computing total expenditures and the food budget share (which would be a rough substitute for including the rental value of owner-occupied housing as far as improving comparability of expenditures of home-owner and tenant households). Our results are also robust to inclusion of financial savings to the definition of expenditure; to inclusion of interactions of time dummies with demographic variables as suggested by Logan (2008); and to the non-parametric modeling of household demographic characteristics.

[27] The choice of holding the final distribution fixed (as opposed to the earlier ones) is arbitrary, but is consistent with the result of more unmeasured inequality in the past. Note that while we can estimate the decline in this unmeasured inequality, we cannot quantify how much unmeasured inequality remains in the final distribution.

[28] Note that survey measurement error was likely higher in the 1987/88 survey due to hyper-inflation. Glewwe (2007) shows that a decline in measurement error over time tends to increase the measured income growth among the poor (even if the "true" growth is zero). Since we cannot quantify how

(continued…)

the right and becoming more equal. It indicates that the bias-corrected "real" inequality (based on 2002/03 prices) was much worse in the past once we account for different CPI bias between the rich and the poor.[29]

The decline in the relative prices of tradable goods (and in food prices in particular) shown in Figure 3 contributed to the narrowing in real expenditure inequality since tradable goods take a larger share of the consumption basket of the poor (de Carvalho Filho and Chamon, 2008b). Improvements in the distribution of goods may have disproportionately benefited the poor, and that would be an interesting topic for future research. Finally, some of the improvements stemmed from the inflation stabilization that took place in 1994. The cumulative inflation in the 12 months covered in the 1987/88 POF is 400 percent, whereas it is only 13 percent in the 1995/96 one. Since the poor did not have in general access to financial instruments, they were disproportionately burdened by the inflationary tax (unlike richer households whose deposits were compounded daily by an overnight rate).[30]

Table 4 reports the CPI-deflated "headline" and the bias adjusted real expenditure per capita for the average, median households, as well as for the bottom and top quintiles of the expenditure distribution. The latter illustrates the growth experienced by a household that has remained at that particular quintile during our sample. When expenditure is deflated by the CPI, the largest gains over the sample period are experienced by the top quintile and the gains for the average household are larger than for the median and the bottom quintile. After correcting for CPI bias, the largest gains are now experienced by the bottom quintile,

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survey measurement error has evolved over time, we cannot incorporate these considerations into our estimation.

[29] While the assumptions made to deal with the non-monotonicity in the Engel curves at the left tail of the distribution have a negligible impact on the expenditure-weighted bias estimates, its distributional impacts are not negligible. However, even if we exclude the households in that left-tail, the distributional effects remain strong and relevant (since the remaining households in the bottom of the distribution are still fairly poor in absolute terms).

[30] The poor could protect their income from inflation to some extent by stocking up on goods. But this was limited by the fact that some goods are perishable and cannot be easily stocked, and the fact that certain types of expenditures cannot be easily concentrated around pay day.

followed by the median household, the average and the top quintile (but growth is higher for all groups).

3.4 Evidence From Durable Goods Ownership

Table 5 presents the percentage of households that own different durable goods using data from the POF. It suggests a substantial improvement in the living conditions of Brazilian households. For example, while only 29 percent of households in the POF sample owned a washing machine in 1987/88, 53 percent did by 2002/03. During this same period, the ownership of color TVs increased from 57.4 percent to 93.4 percent. The number of households that owns either a car or a motorcycle increased from 34.2 percent to 38.1 percent. At first, one may be inclined to dismiss this type of evidence as being driven mainly by declines in the price for these goods and not necessarily by income growth. However, Figure 7 shows that the increase in the ownership of durable goods was skewed toward those goods that are more of a luxury. Figure 7A shows the change in average holdings of a given good between 1987/88 and 2002/03 and the sensitivity to income of holdings of that good in the 1987/88 cross-section. That sensitivity is obtained by regressing the household's holdings of that good on the log of total expenditure (a pseudo-Engel curve):

$$G_{i,t} = \phi + \beta_G(\ln Y_{i,j,t} - \ln(1+\Pi_{G,j,t})) + \sum_x \theta_x \mathbf{X}_{i,t} + \mu_{i,t},$$

where $G_{i,t}$ is household i's holding of good G at time t, β_G is the sensitivity with respect to income of that good and $\mathbf{X}_{i,t}$ are the same household controls used in the previous regressions. We are then ready to estimate:

$$\Delta_{G,t,T} = \alpha + \psi \bullet \hat{\beta}_G + \varepsilon_G,$$

where $\Delta_{G,t,T}$ is the change in average ownership of good G between times t and T. The slope of that regression gives the expected change in the true real income:

$$\hat{\psi} = \overline{(\ln Y_{2002/03}/(1+P_{2002/03}))} - \overline{(\ln Y_{1987/88}/(1+\Pi_{1987/88}))},$$

from which we can obtain the estimated CPI bias during that period:

$$\ln(1+E_{G,2002/03}) = \overline{(\ln Y_{2002/03} /(1+\Pi_{2002/03}))} - \overline{(\ln Y_{1987/88} /(1+\Pi_{1987/88}))} - \hat{\psi}$$

This method assumes that households face the same bias (in a given year). Since richer households have a disproportionate influence in the average log real expenditure, this method gives them a larger weight in the bias estimate, although the result is not directly comparable to our expenditure-weighted estimates. There are a number of dynamic aspects related to the consumption of durables (including access to consumer credit and interest rate levels) that are not captured in the setting above. Therefore, this result must be interpreted with caution, and is used mainly as rough cross-check for our bias estimates based on Engel curves for food consumption.

We estimate the equation above using the sensitivity of holdings to income estimated on the 1987/88 and 2002/03 cross-sections. When the 1987/88 income sensitivities are used, the implied bias estimate is 2.2 percent per year during the period in question (Figure 7A). On the other hand, when the 2002/03 income sensitivities are used, the implied bias is 3.7 percent per year (Figure 7B). Thus, the observed pattern of changes in durable good ownership suggests a bias in the CPI similar to the one estimated based on the evolution of food demand. The fact that we are able to obtain similar estimates through substantially different approaches is reassuring.

3.5 Food Consumption in the 1970s

Our hypothesis that the large unmeasured gains in real income stem from one-off effects from market reforms could be tested if the coverage of our sample included at least one pair of pre-reform surveys. In 1974/75 IBGE carried out the *Estudo Nacional da Despesa Familiar* (ENDEF) survey, a precursor of the recent POF expenditure surveys. We obtained household-level data of the ENDEF, which allows some comparison with the recent surveys.

The comparability of Engel curves from the 1974/75 ENDEF with the 1987/88 POF is affected not only by changes in survey design, but also by the unavailability of a price index covering different metropolitan areas prior to the IPCA's introduction in 1980.[31] With those caveats in mind, we find a negative pre-reform CPI bias point estimate (suggesting that inflation indices underestimated the increase in the true cost of living in 1974/75-1987/88).[32] That is the opposite of our findings for the post-reform period, supporting the view that the large bias stems from the reforms. But as cautioned above, this bias estimate in the pre-reform period is likely not comparable to the other estimates for the reasons above.

Figure 8 shows the evolution of the average food budget share over time, which does not require deflating expenditures from one survey to the other. The average food budget shares in the 1974/75 ENDEF and the 1987/88 POF are nearly identical, but there is a marked decline in the two post-reform POFs, notwithstanding a faster measured real income growth between 1974/75 and 1987/88. This pattern supports our hypothesis that large CPI bias is a post-reform phenomenon.

4. Mexico

4.1 Background on Economic Policy Changes

Mexico began a gradual liberalization process in 1983, with moderate reductions in non-tariff barriers and a simplification of the tariff schedule. The liberalization process accelerated in July 1985, with a reduction in the coverage of import permit requirements from 75 to 39 percent of total import value, accompanied by the announcement of a 30-month schedule of

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[31] Prior to 1980, there were indices collected for a few cities by different institutions, notably FGV in Rio and FIPE in Sao Paulo. There is substantial variation on how these alternative indices compare with the IPCA for the years when both were available. For example, between January 1980 to October 1987 (the reference date for the first POF), the cumulative inflation according to the IPCA is 1.33 times that of the IPC-FIPE index covering São Paulo. On the other hand, in this same period, the cumulative inflation according to the IPCA is only 0.83 times that of the FGV-IPC index covering Rio de Janeiro.

[32] Due to the limitations on the availability of a price deflator we considered three alternative approaches: (i) using only the metropolitan areas where an alternative price index spanning 1974-1987 is available and using that index as the deflator; (ii) using the same sample as (i) but splicing the alternative price indices with the IPCA once the latter becomes available; (iii) using all metropolitan areas covered in the ENDEF, deflating expenditures by an index splicing the IPCA with an imputed pre-1980 inflation based on alternative indices from other metropolitan areas. In all cases the bias estimates remained negative.

tariff reductions. By the end of 1987, only 23 percent of imports were subject to prior licensing requirements. The average tariff rate declined from 25 percent in 1985 to 13 percent in 1990.[33] This effort culminated in 1994 when the North American Free Trade Agreement (NAFTA) came into effect. Most of the privatization efforts were concentrated in the first half of the 1990s (Lora 2001). Similarly to Brazil, Mexico experienced a growth disappointment: GDP per capita in 1994 was lower than its pre-1982 crisis level, and Mexico's experience during that period is summarized in the title of Dornbusch and Werner (1994): "Stabilization, Reform and no Growth." Matters became worse in 1995, when Mexico experienced a severe currency crisis. Since then, its economy has recovered, but its performance still has not made the desired transition to sustained high growth rates (e.g. Tornell, Westermann and Martinez, 2004; Hanson 2010).

4.2 Data

The *Encuesta Nacional de Ingresos y Gastos de los Hogares* (henceforth, ENIGH) is the household income and expenditure survey carried out by the *Instituto de Instituto Nacional de Estadística Geografía e Informática* (INEGI). ENIGH surveys are based on probabilistic sample and stratified design, and were conducted in 1984, 1989, 1992, 1994, 1996, 1998, 2000, 2002, 2004, 2005 and 2006.

Mexico's Central Bank produces and disseminates the *Índice Nacional de Precios al Consumidor* (INPC), which is Mexico's consumer price index. It is based on the consumption pattern of the families in the ENIGH (which is therefore, by construction, the relevant sample in which to base our exercise).[34] Its target population are the families in the municipalities where at least one locality has 20,000 or more habitants. The index is available for 46 cities, although the coverage in 11 of those cities begins in 1995. Those 11 cities are

[33] For a description of Mexico's trade liberalization process, please refer to Symczak (1992). The index of trade reforms in Lora (2001) concentrates most of the trade reform effort in Mexico in the middle of the 1980's.

[34] From 1980 to 1994, the INPC was based on the consumption in the 1977 ENIGH; From 1994 to 2002 it was based on the 1989 ENIGH; And from 2002 onwards it was based on the 2000 ENIGH.

excluded from our sample (reducing the number of household-survey observations by about 15 percent, from 49,701 to 41,805).[35]

The ENIGHs contain household expenditure information on individual goods at a disaggregated level, the households' inventory of durable goods, and also demographic, schooling and income characteristics of each household member. They were conducted over 100 days (10 collection periods covering 10 days each), typically in the third quarter of each survey year (August-November for all the years except 1994 when it ran from September-December). Households were asked to provide information on their expenditures, which are collected based on different reference periods depending on the type of expenditure and its frequency. Four reference periods are used: 7 days, 30 days, 90 days, and a longer recall period of 6 months.

The surveys use two collection instruments. A basic questionnaire which records living conditions, characteristics of the residents, income and recalled expenditures. Expenditures on food, beverage, tobacco (inside and outside the house), and public transportation were recorded in a daily expenditure notebook. Unlike in the case of Brazil, the survey also covers the consumption of self-produced food (which is useful since the Mexican sample is not limited to the major metropolitan areas).[36] It also covers consumption through in-kind transfers and gifts and includes cash transfers (e.g. remittances from absent household members) in our measures of income. Only 0.3 percent of households in our sample did not report any expenditure on food, and the results are nearly identical when those households are dropped. The ENIGH provides the imputed rental value of owner-occupied housing in all surveys. For the sake of comparison with the results for Brazil, we do not consider that imputed rent as part of expenditures in our main estimates. But we do include them as part of our robustness checks, and their consideration tends to increase the bias

[35] Our sample includes the metropolitan area of Mexico City, Acapulco, Aguascalientes, Cd. Juárez, Colima, Córdoba, Culiacán, Chetumal, Chihuahua, Guadalajara, Hermosillo, La Paz, Léon, Matamoros, Mérida, Mexicali, Monclova, Monterrey, Morelia, Puebla, San Luis Potosí, Tampico, Tapachula, Tijuana, Toluca, Torreón, Tulancingo, Veracruz and Villahermosa.

[36] The rural poor, for whom self-produced food accounts for a substantial share of consumption, have a small weight in our expenditure-weighted estimates, and our results are robust to their exclusion from the sample.

estimates slightly. As in the case of Brazil, our measure of food consumption covers food and beverages, and our measure of expenditure covers all expenditures except financial savings.

Unlike in the Brazilian POF, nominal expenditures are not deflated to a common reference date. Since Mexico experienced double-digit inflation during most of our sample period, it is of paramount importance that the timing of expenditures be taken into account when comparing or aggregating expenditure groups subject to different recall periods. In particular, ENIGH food expenditures measure purchases that occurred later than the recall period for most non-food purchases. If that difference in timing is not taken into account when deflating the expenditures, we would systematically underestimate the real value of expenditures on items with a longer recall window, thereby artificially increasing the food budget share, and biasing our estimates when comparing that share across years with different inflation rates.[37] The ENIGH indicates the date when the households were surveyed. Based on that date and on in the inflation in the preceding months, we distribute the nominal expenditures so as to smooth them in real terms through the recall window.[38] For example, an expenditure in the 6 month recall window would be smoothed into 6 monthly expenditures. All expenditures are then deflated to a same date, so they can be compared.

4.3 Results

Figure 9 shows the relative prices for different groups of goods (vis-à-vis the aggregate price index). As in the case of Brazil, there were substantial relative price movements during the 1980s and early 1990s. But the relative prices for most groups have been more stable at least since the large Mexican peso devaluation occurred in late 1994, which is consistent with the view that trade liberalization was an important factor behind those movements. As in the case of Brazil, the change in relative prices is not as large as one might have expected following

[37] For example, the inflation rate in the third quarter of 1994 was 7 percent per year while the corresponding one for 1996 was 19 percent. If the food budget share was computed based on nominal expenditures (or if all expenditures were deflated using the third quarter price level), there would be a substantial increase in the average food budget share between those two surveys. But once the deflation takes into account the difference in the timing of purchases, the increase is much smaller.

[38] The surveys do not indicate the date in which items in the recall window were purchased.

the sharp depreciation of the peso in late 1994 (when it lost about 35 percent of its value in real effective terms).[39]

Table 6 provides summary statistics for our regression sample. The average budget share of food declines substantially over time, with most of the decline occurring between 1984 and 1992. Household income and expenditure increase until the 1994 survey, experience a sharp drop in the 1996 survey as a result of the 1994/1995 crisis, only recovering at the end of our sample. The effect of the crisis on household income and expenditure is much stronger than that indicated in the national accounts, with per capita expenditure declining by about 30 percent from 1994 and 1996.[40] In fact, the crisis is so severe that real per capita expenditure on food declines in its aftermath and has not recovered to its previous level as of 2006 (of course, if the food CPI overstates the increase in food prices, perhaps due to strong substitution bias in that group, the true decline is not as severe).[41] While the average food budget increases from 1994 to 1996 by 1.4 percentage points, that increase is relatively small and is quickly reversed. The aggregate food budget share, however, experiences a more substantial increase (from 26.1 to 31.7 percent), and only returns to pre-crises levels by the 2002 survey. Still, the behavior of real food consumption suggests caution before interpreting changes in the food budget shares as the result of an unmeasured income effect. We initially present results pooling all years together, but as a robustness check, we also present estimates based on two separate samples: a pre-crisis (1984-1994) and a post-crisis one (1996-2006). The results are qualitatively similar to those of our baseline sample (except for the 1994-1996 period for which there is no estimate when the sample is broken in two). Finally, it is also worth noting the substantial decline in household size over our sample, from 4.8 in 1984 to 3.8 in 2006.

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[39] For exchange rate pass-through estimates for different groups of goods, please refer to de Carvalho and Chamon (2008).

[40] Similar declines are documented in other papers that use the ENIGH. The decline in real private expenditures from 1994 to 1996 in the national accounts is “only” 7.5 percent.

[41] McKenzie (2006) documents changes in Mexican household consumption patterns in the aftermath of the Crisis.

Parametric Model

Table 7 reports our parametric estimates for Mexico, as well as the implied cumulative bias since 1984 for every survey year. As in the case of Brazil, we control for household demographic characteristics, labor market participation, and for whether the household owns or rents its dwelling. The first two columns estimate the bias assuming it is constant across households in a same year (as in Hamilton 1998, 2001a; Costa 2001). The resulting estimates indicate a substantial bias, particularly in the first years covered by the survey. From 1984 to 1996, the estimated cumulative bias is 51 percent. That estimate drops to 45 percent when expenditure is instrumented by current income. Note the substantial increase in the cumulative bias from 1994 to 1996. The CPI inflation in that period (1994Q3-1996Q3) was 85 percent, in large part owing to the currency crisis in January 1995. A large bias estimate during that period could be the result of substitution bias, whereby the CPI exaggerates the impact of the exchange rate depreciation on the cost of living by failing to take into account the full extent of consumer substitution. However, the magnitude of that bias remains large despite those considerations. It is worth noting that this empirical strategy can be too noisy for short-term comparisons. Since the cumulative bias slows down substantially from 1996 onwards (2.3 and 1.2 percent per year in the first and second columns respectively), the cumulative bias from 1994 to 2006 yields a more moderate estimate. If we drop 1996 from the sample, we would obtain similar results: a large change in the cumulative bias from 1994 to 1998, with the bias slowing down afterwards.

The last two columns of Table 7 present the estimates when the bias is allowed to vary linearly with the (CPI deflated) log real expenditure. The estimated cumulative bias from 1984 to 1996 is 48 percent both under population and expenditure weights. Once expenditure is instrumented by current income, the estimated bias drops to 43 (under both weights). From 1996 to 2004, the changes in the cumulative population weighted bias are comparable to those in columns 1 and 2, while the change in the expenditure-weighted bias is not statistically significant (with point estimates implying less than 1 percent per year).[42]

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[42] The bias estimates are similar if we include dummies for whether the head and the spouse have primary, secondary or higher education as additional controls. The resulting cumulative bias in 1984-

(continued…)

As in the case of Brazil, these results imply a substantial underestimation of the true real income growth (a 40 percent bias in the CPI implies an unmeasured 67 percent growth in real income). However, the difference between the average bias using population and expenditure weights is substantially smaller than in the case of Brazil. Finally, the fact that the expenditure-weighted bias (the relevant measure for the aggregate economy) declines to U.S. levels from the 1996 survey onwards supports our hypothesis that the large bias stems from transitory effects of the reforms.[43]

Semi-Parametric Model

Figure 10 shows the non-parametric estimates of the relationship between the food budget share and the log of real expenditure. For clarity purposes the figure only plots the Engel curves for 1984, 1989, 1994, 1996, 1998, 2004 and 2006. As expected, the food budget share declines with real expenditure and on average the curves have shifted downward over time (the latter curves seem to slightly pivot, implying a downward shift for lower levels of income and an upward shift for higher levels of income). It is worth noting that while the Engel curve shifts significantly from 1994 to 1996, it does not shift much from 1996 onwards (as in our findings from the parametric model).

Figure 11A shows the estimated annual bias from 1984 to 1998 as a function of headline real expenditure, as well as the estimated density function of the log of CPI-deflated real expenditure in 1998. The bias for the average household is 5.0 percent per year whereas the expenditure-weighted aggregate bias is 4.6 percent per year. Figure 11B shows the estimated annual bias from 1998 to 2006. The bias for the average household is 1.6 percent per year whereas the expenditure-weighted bias is 0.9 percent per year (standard errors are 0.3 and 0.4 percent per year respectively). The results are similar if we use 1996 as a basis for comparison instead of 1998, but we chose the latter out of concern that 1996 was an

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2006 would change by less than 3 (1½) percentage points for the population-weighted (expenditure-weighted) estimates, with the difference being less than ½ percentage points in the IV specifications.

[43] While the population-weighted bias does not vanish after 1996, it is not statistically significantly different than the 1 percent per year bias associated with the U.S.

unusual year due to the 1995 crisis.[44] These results are also fairly robust when the sample is broken down by proximity to the U.S. border and education.[45]

Robustness

Figure 12 reports the estimated biases under different methods and samples. We present estimates using: (i) Our baseline sample; (ii) A Winsorized sample; and (iii) A sample where the imputed rental value of owner-occupied housing is included among the expenditures. The Winsorization sets the value of food, total expenditures and income for observations below the 5th and above the 95th percentile to the level of that percentile. The thick line corresponds to our preferred estimates, based on the semi-parametric estimation. Different samples and methods illustrate the robustness of the finding of a large bias, concentrated in the 1984-1996 period. Among the parametric specifications, the ones using IV yield lower bias estimates. The estimated expenditure-weighted cumulative bias from 1984 to 2006 based on the semi-parametric estimation is 51 percent in the baseline sample, 50 percent in the Winsorized sample and 49 percent in the sample that includes the rental value of owner-occupied housing among expenditures. The lowest estimate for the 1984-2006 cumulative bias across all methods and samples is 38 percent, which still implies an underestimation of the "true" real income growth during those 22 years by 2.2 percent per year.[46]

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[44] The estimated bias for 1984-1996 is 5.4 percent per year for the average household and 5.1 percent per year for the expenditure weighted bias. For 1996-2006 those estimates are 1.4 and 0.8 percent per year respectively

[45] For example, the estimated expenditure weighted bias is 4.5 percent per year in 1984-1998 and 0.9 percent per year in 1998-2006 for a subsample of 10 cities close to the U.S. border. The estimates are also fairly comparable between a subsample where the head of household has more/less education (e.g. the expenditure weighted bias is 4.7 percent per year in 1984-1998 and 0.4 percent per year in 1998-2006 for a subsample where the head of household had 7 years or more of education).

[46]Our results are also robust to inclusion of financial savings to the definition of expenditure; to inclusion of interactions of time dummies with demographic variables as suggested by Logan (2008); and to the non-parametric modeling of household demographic characteristics.

In order to address potential concerns that the aftermath of the 1995 crisis may have biased our estimates, we perform the same exercise in two separate, non-overlapping sub-samples. The first covers 1984-1994 (the pre-crisis period) and the second 1996-2006 (the post-crisis period). Tests based on the pooled sample for the interaction of each variable on a post-crisis dummy indicate no statistically significant change except for the coefficient on the log of the relative price of food, which all else equal would only decrease the post-crisis bias by a small amount. Our semi-parametric estimates for the pre- and post-crisis baseline sample indicate an expenditure-weighted cumulative bias of 40 percent in 1984-1994, and 9 percent in 1996-2006 (less than 1 percent per year). Thus, aside from the 1994-1996 bias (for which there is no estimate based on these two separate sub-samples), these estimates confirm the results from our pooled estimates. If we assume that there is no bias after 1994, the estimates from the 1984-1994 sub-sample (4.9 percent per year) would still imply an underestimation of real income growth by 2.3 percent per year in 1984-2006. Thus, even this conservative estimate still implies a respectable post-reform real income growth.

Distributional Implications

Figure 13A plots the distribution of CPI-deflated real expenditures in 1984, 1994 and 2006. These curves imply an increase in inequality over time. Figure 13B plots the distribution of our estimates of the true real expenditure implied by the changes in the food budget shares. The curve for 2006 is the same as the one in Figure 13A, while the curves for 1984 and 1994 are adjusted based on our estimates (with the bias being measured relative to 2006). As in the case of Brazil, "real" inequality was much worse in the past once it is measured in constant prices correcting for the CPI bias. The curves indicate widespread gains over time (as shown by the movement of mass from the left to the right-end of the distribution). They also indicate a reduction in inequality.

Table 8 reports the CPI-deflated "headline" and the bias adjusted real expenditure per capita for the average and median households, as well as for the bottom and top quintiles of the expenditure distribution. As in the case of Brazil, when expenditure is deflated by the CPI, the largest gains over the sample period are experienced by the top quintile and the gains for the average household are larger than for the median household and the bottom quintile. After correcting for CPI bias (based on semi-parametric estimates for 1984-2006), the largest gains are now experienced by the bottom quintile, followed by the median household, the average and the top quintile (but growth is higher for all groups).

4.3 Evidence From Durable Goods Ownership

Table 9 presents the average durable good ownership of Mexican households. For example, while in 1984 there were only 55 washing machines per 100 households, by 2004 there were 76. The average number of automobiles increased from 35 per 100 households to 50 during that period, and there were 29 PCs per 100 households in 2004. We perform a similar exercise to the one described in Section 3.4 for durable ownership in Brazil. Figure 14A shows the change in the average durable good ownership from 1984 to 2004 and the good's sensitivity to income based on the 1984 cross-section. It implies a bias of 2.3 percent per year. Figure 14B is similar to Figure 14A but estimates the sensitivity to income based on the 2004 cross-section. It yields a bias of 2.7 percent per year. These estimates match those based on food demand fairly well.

This alternative method can also serve as a check on the sudden disappearance of the bias in the second half of the 1990s, as implied by our food demand estimates. The estimated bias based on changes in durable good ownership from 1984 to 1998 is 4.6 (4.7) percent per year when income sensitivity is based on the 1984 (1998) cross-section. The estimated bias for 1998-2004 is 0.5 (2.0) percent per year when income sensitivity is based on the 1998 (2004) cross-section. Thus, this alternative method corroborates our previous finding that the bias has declined in Mexico since the late nineties (consistent with our prior of a one-off effects from the reforms).

5. Discussion and Conclusion

This paper uses household-level data from Brazil and Mexico to present evidence that post-reform growth in real household income has been substantially underestimated by standard measures, and that these "unmeasured" income gains have been stronger for poorer households. Our estimates indicate that average household per capita income grew by 4½ percent per year in Brazil during 1987/88-2002/03. This figure is substantially higher than the 1½ percent annual "headline" growth obtained by deflating nominal per capita household income by the CPI. In the case of Mexico, our estimates imply a growth rate of 4½-5½ percent per year in 1984-2006, which is again substantially higher than the 2 percent implied by standard methods. The estimated bias for Mexico is concentrated in the 1984-1996 period, and declines to U.S. levels from 1996 onwards. The estimated biases prove fairly robust across a number of specifications. While it is difficult to identify their source, the pattern for early reformer Mexico supports our hypothesis that large biases stem from one-off level effects of the market reforms. This hypothesis is also supported by the evolution of food budget shares during pre-reform years in Brazil.

We also find that evidence based on ownership of durable goods corroborates our findings based on the food Engel curve. Our estimates of real income growth based on the relationship between changes in individual durable goods ownership and their demand sensitivity to income yields similar estimates of real income growth, and a working paper version of this paper also reported anthropometric tabulations suggestive of stronger income growth. [47]

An important question for future research is whether similar biases exist in other developing countries. For example, if household income in Brazil and Mexico grew faster than previously thought during this period, would the same apply to Asia? Our prior is no, since our preferred explanation for the large CPI bias in Brazil and Mexico are one-off effect of trade liberalization and inflation stabilization. If that is indeed the case, we should expect a significant bias to be present in many other Latin American countries, but less so in Asia. However a precise answer to this question would require future research to replicate this analysis to other countries.

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[47] de Carvalho Filho and Chamon (2008) reports height-for-age tabulations for children available through the World Health Organization (WHO) Global Database on Child Growth and Malnutrition. There is a significant improvement in Brazil, with the percentage of urban children 0-60 months below 2 standard deviations from the reference median declining from 12.3 percent in 1989 to 7.8 percent in 1996. In Mexico, that share declines from 19 percent in 1988 to 11.6 percent in 1998/99. To the extent that malnutrition is associated with deprivation among the poorest households, these sizable improvements are inconsistent with the official finding of income stagnation and consistent with the conclusions of this paper about reduction in real expenditure inequality.

Some rough comparisons using aggregate consumption data for Korea and Taiwan also suggest a more limited scope for CPI bias in East Asia.[48] These comparisons should be taken with caution since they are based on aggregate data and do not control for a number of differences across these countries (for example, differences in the relative price of food, the slope of Engel curves, the starting level of income and the definition of expenditures based on which food shares are constructed). Between 1984 and 2006, the aggregate food share of consumption declined from 38 percent to 26 percent in Korea and from 39 percent to 24 percent in Taiwan. These declines are comparable to the decline in the food budget share in our sample for Mexico during this period (from 41 to 25 percent of expenditures). But real household income roughly tripled in Korea and doubled in Taiwan during that period. Between 1987 and 2003, the aggregate food share of consumption in Korea declined from 35 to 27 percent, which is comparable to the decline experienced in the food budget share in our sample for Brazil (from 20 to 14 percent of expenditures). However, during that same period, real household income doubled in Korea. It is reasonable to assume that it is unlikely that changes in food demand in Korea and Taiwan would imply a bias in the magnitude of 3 percent per year as we find for Brazil and Mexico.[49]

This paper focuses on the mismeasurement of households' real income, but similar sources of bias apply for the measurement of production price deflators (e.g., new and better goods). It is unlikely that the GDP deflator was accurately measured if the CPI bias was large, particularly because household consumption, which is deflated by the CPI, is a large component of GDP.[50] Quantity index estimates for GDP also suffer from severe measurement

[48] The data source for Korea is the Annual Report on the Household Income and Expenditure Survey covering urban households, and for Taiwan is the Report on the Survey of Family Income and Expenditure. The ratios reported correspond to the food share in household consumption expenditures.

[49] Chung, Gibson and Kim (2010) find CPI bias in Korea for the period 2000-2005 of slightly less than 1 percent per year, a magnitude more similar to the estimates for advanced economies than to our estimates for Brazil and Mexico.

[50] The national accounts systems in both Brazil and Mexico compute GDP from the production and the expenditure sides, reconciling the two measures. Household consumption accounts for about 55 percent of the expenditure GDP in Brazil, and for about 70 percent in Mexico.

problems, which could also help explain the stagnation of per capita GDP despite the strong household consumption growth we estimate.[51, 52]

If trade liberalization is indeed the source of these large biases, then a similar effect might be at play in other reforming countries. Since Brazil and Mexico account for over half of Latin America's population and for over 60 percent of its output, our findings already suggest a significant correction to population- or output- weighted regional household income growth averages. More generally, this paper suggests caution should be used when reading aggregate economic statistics during periods of large changes in economic policies. This can be difficult since practitioners and policy makers are often pressed to make important decisions using whatever limited data is available. This paper suggests that, to the extent possible, aggregate figures should be checked with patterns observed in micro-level data and other alternative measures.

[51] The quantity indices are either obtained by deflating a sector's nominal output by a relevant price deflator, or through a measure of volume of its output. In both cases, there is a potential for measurement problems to bias the estimate (e.g. improvements in quality or new products).

[52] A revised GDP series for Brazil was released on March 26, 2007. The revision increased the level of GDP by about 11 percent, and was the result of a number of methodological improvements (including a broader coverage of economic activity by the National Accounts). In their paper on the implementation of the 1993 System of National Accounts (SNA 93) in Latin America, Olinto Ramos et al (2008) argue that the Brazilian GDP revision is roughly consistent with our bias estimates. Similarly, a revised GDP series for Mexico was released on April 29, 2008. The revision increased the level of GDP by 12.5 percent, and was also the result of a number of methodological improvements (with the number of activities covered increasing from 362 to 750).

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Table 1. Descriptive Statistics for Brazil

	Full Sample			97 Percent Sample		
	1987-88	1995-96	2002-03	1987-88	1995-96	2002-03
Share of food	0.303 [0.163]	0.254 [0.176]	0.224 [0.16]	0.301 [0.159]	0.253 [0.173]	0.223 [0.157]
Share of food at home	0.248 [0.158]	0.208 [0.169]	0.175 [0.152]	0.245 [0.154]	0.206 [0.165]	0.173 [0.149]
Share of food at restaurant	0.055 [0.069]	0.046 [0.072]	0.049 [0.072]	0.056 [0.069]	0.047 [0.071]	0.049 [0.071]
Relative price of food	100 [0]	74.38 [5.09]	72.66 [5.69]	100 [0]	74.4 [5.1]	72.7 [5.69]
Ln (Relative price of food)	4.6 [0]	4.3 [0.06]	4.28 [0.07]	4.6 [0]	4.3 [0.06]	4.28 [0.07]
Expenditure on food, in 1996 R$	2875 [2444]	2800 [3303]	2638 [2819]	2936 [2438]	2868 [3319]	2714 [2831]
Per capita expenditure on food, in 1996 R$	813 [851]	887 [1180]	915 [1125]	829 [855]	907 [1189]	939 [1134]
Expenditure, in 1996 R$	17328 [25306]	15786 [23978]	16672 [24414]	17719 [25473]	16185 [24164]	17172 [24637]
Per capita expenditure, in 1996 R$	5025 [7707]	5282 [10217]	6001 [10700]	5133 [7765]	5411 [10319]	6174 [10824]
Ln (Expenditure/CPI)	9.226 [1.032]	9.085 [1.078]	9.168 [1.03]	9.292 [0.944]	9.157 [0.989]	9.243 [0.953]
Ln (After-tax income/CPI)	9.284 [1.049]	9.059 [1.172]	9.094 [1.232]	9.334 [0.999]	9.107 [1.129]	9.15 [1.187]
Ln (Household size)	1.270	1.183	1.106	1.287	1.195	1.120
# ages 0 to 4	0.435	0.334	0.281	0.440	0.335	0.281
# ages 5 to 9	0.462	0.352	0.302	0.468	0.352	0.302
# ages 10 to 14	0.392	0.380	0.314	0.399	0.384	0.317
# ages 15 to 19	0.388	0.370	0.341	0.395	0.375	0.347
# ages 20 and up	2.387	2.260	2.212	2.407	2.277	2.235
Male head	0.790	0.748	0.678	0.797	0.753	0.683
Spouse present	0.733	0.689	0.647	0.743	0.697	0.657
Head has some income from work	0.813	0.763	0.754	0.822	0.771	0.762
Spouse has some income from work	0.282	0.282	0.286	0.288	0.288	0.293
Head and spouse have income from work	0.260	0.256	0.251	0.265	0.261	0.258
Rental unit	0.294	0.187	0.173	0.297	0.191	0.177
Owner occupied	0.588	0.700	0.713	0.590	0.700	0.713
Other living arrangement (ceded housing)	0.118	0.113	0.114	0.114	0.109	0.110
Sample size	12417	14528	6757	12045	14093	6555

Notes: The 97 percent sample excludes the households whose expenditures are below the 3rd percentile of the expenditure distribution. Standard deviations in brackets.

Table 2. Regression Results for Brazil, Full Sample
(Dependent variable = expenditure share of food in 1987, 1996, and 2003)

	Bias invariant on expenditure		Bias linear on Ln(Expenditure/CPI)	
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
Dummy for 1996	-0.053 [0.013]	-0.048 [0.014]	-0.335 [0.052]	-0.244 [0.047]
Dummy for 2003	-0.074 [0.015]	-0.066 [0.014]	-0.363 [0.049]	-0.181 [0.066]
Ln (Expenditure/CPI)	-0.061 [0.004]	-0.085 [0.004]	-0.084 [0.004]	-0.097 [0.003]
Ln (Expenditure/CPI) x Dummy for 1996			0.031 [0.005]	0.021 [0.005]
Ln (Expenditure/CPI) x Dummy for 2003			0.031 [0.005]	0.012 [0.007]
Ln(Relative price of food)	0.008 [0.042]	0.031 [0.041]	0.003 [0.041]	0.025 [0.041]
Ln(Household size)	0.027 [0.008]	0.038 [0.009]	0.027 [0.009]	0.038 [0.009]
Number ages 0 to 4	0.011 [0.003]	0.004 [0.003]	0.011 [0.003]	0.005 [0.003]
Number ages 5 to 9	0.009 [0.003]	0.005 [0.003]	0.009 [0.003]	0.005 [0.003]
Number ages 10 to 14	0.009 [0.003]	0.005 [0.003]	0.009 [0.003]	0.005 [0.003]
Number ages 15 to 19	0.006 [0.003]	0.005 [0.003]	0.006 [0.003]	0.005 [0.003]
Number ages 20 and up	0.002 [0.002]	0.006 [0.003]	0.003 [0.002]	0.006 [0.003]
Male head	0.033 [0.005]	0.034 [0.006]	0.034 [0.005]	0.035 [0.006]
Spouse present	-0.017 [0.005]	-0.017 [0.005]	-0.017 [0.005]	-0.017 [0.005]
Head of household has some income from work	-0.006 [0.004]	0.001 [0.004]	-0.007 [0.004]	0.001 [0.004]
Spouse has some income from work	-0.012 [0.006]	-0.008 [0.006]	-0.014 [0.006]	-0.009 [0.006]
Head and spouse have income from work	0.006 [0.007]	0.01 [0.007]	0.006 [0.007]	0.01 [0.007]
Housing unit ceded by family, employer	0.054 [0.006]	0.042 [0.006]	0.054 [0.006]	0.043 [0.006]
Owner occupied unit	0.026 [0.004]	0.024 [0.004]	0.026 [0.004]	0.025 [0.004]
Constant	0.807 [0.183]	0.899 [0.178]	1.038 [0.192]	1.033 [0.185]
Observations	33702	33702	33702	33702
R-squared	0.250	0.232	0.257	0.237
Anderson canon. corr. LR statistic (p-value)	N/A	0.0000	N/A	0.0000
Population Weighted Bias				
Cumulative bias 87-96 (%)	57.72 [5.99]	43.29 [5.71]	45.64 [5.51]	39.86 [5.30]
Cumulative bias 87-03 (%)	69.85 [4.52]	53.90 [4.89]	56.64 [4.63]	50.37 [4.57]
Cumulative bias 96-03 (%)	28.69 [3.73]	18.71 [3.10]	22.47 [2.84]	17.05 [2.77]
Annual equivalent 87-96 (%)	9.54 [1.47]	6.40 [1.09]	7.58 [1.08]	6.02 [0.96]
Annual equivalent 87-03 (%)	7.52 [0.89]	4.93 [0.65]	5.73 [0.64]	4.52 [0.57]
Annual equivalent 96-03 (%)	4.89 [0.74]	3.02 [0.55]	3.70 [0.52]	2.79 [0.46]
Expenditure Weighted Bias				
Cumulative bias 87-96 (%)			16.38 [9.32]	22.19 [7.57]
Cumulative bias 87-03 (%)			33.71 [7.94]	42.73 [6.05]
Cumulative bias 96-03 (%)			21.74 [3.71]	25.13 [3.50]
Annual equivalent 87-96 (%)			2.87 [1.20]	3.17 [1.05]
Annual equivalent 87-03 (%)			3.12 [0.71]	3.63 [0.64]
Annual equivalent 96-03 (%)			3.57 [0.67]	4.26 [0.70]

Notes: Robust standard errors for the regression coefficients and bootstrapped standard errors for bias estimates in brackets. Controls also include regional dummies. Regressions (3) and (4) also include the interactions of time dummies with the log of real expenditure. Current real income is used as an instrument to total real expenditure in the IV regressions. Cumulative bias reported corresponds to $|E_{G,t}|$. The implied gross change in the true cost of living is (1-$|E_{G,t}|$) times the gross change in the CPI, and the resulting gross true real income growth is 1/(1-$|E_{G,t}|$) times the gross real income growth obtained by deflating nominal income by the CPI.

Table 3. Annual Bias Estimates for Brazil Across Different Methods and Samples

	Population Weighted		**Expenditure Weighted**	
	1987/88–95/96	1995/96–2002/03	1987/88–95/96	1995/96–2002/03
Parametric Estimates				
Panel A. Bias Constant Across Households				
OLS, Full Sample	9.54	4.89		
	[6.82 12.25]	[3.52 6.55]		
OLS, Compliant	6.61	4.23		
	[4.29 8.96]	[3.07 5.35]		
OLS, 97 Pct Sample	8.32	4.05		
	[5.92 10.63]	[2.84 5.33]		
IV, Full Sample	6.40	3.02		
	[4.38 8.35]	[2.00 4.24]		
IV, Compliant	4.73	2.91		
	[2.80 6.61]	[1.97 3.84]		
IV, 97 Pct Sample	6.04	2.66		
	[4.17 7.91]	[1.71 3.66]		
Panel B. Bias Linear Function of Real Expenditure				
OLS, Full Sample	7.58	3.70	2.87	3.57
	[5.49 9.67]	[2.67 4.71]	[0.60 5.26]	[2.18 4.88]
OLS, Compliant	5.58	3.39	1.96	2.83
	[3.73 7.37]	[2.45 4.30]	[-0.04 3.88]	[1.72 4.00]
OLS, 97 Pct Sample	6.67	3.10	2.37	2.81
	[4.94 8.43]	[2.22 3.96]	[0.49 4.26]	[1.73 3.86]
IV, Full Sample	6.02	2.79	3.17	4.26
	[4.14 7.90]	[1.86 3.68]	[1.15 5.25]	[2.86 5.69]
IV, Compliant	4.52	2.70	2.43	3.31
	[2.83 6.19]	[1.84 3.52]	[0.59 4.26]	[1.97 4.59]
IV, 97 Pct Sample	5.64	2.45	2.97	3.55
	[4.04 7.25]	[1.64 3.24]	[1.20 4.72]	[2.27 4.82]
Panel C. Semi-Parametric Estimates				
Full Sample	6.15	4.85	2.87	4.34
	[4.39 8.31]	[3.01 6.22]	[0.89 5.27]	[2.73 6.25]
Compliant	5.61	4.56	2.02	4.29
	[4.07 7.55]	[3.28 5.77]	[0.08 4.45]	[2.50 6.17]
Top 97 Percent Sample	6.43	4.63	2.70	4.29
	[3.79 7.25]	[3.16 5.76]	[0.41 4.66]	[2.54 6.12]
1-40 Minimum Wages	5.23	3.52	1.43	4.92
	[2.85 7.73]	[1.35 7.08]	[-0.89 4.58]	[1.89 10.21]

Notes: 95 percent confidence interval in square brackets. Full sample stands for all the metropolitan areas for which CPI is available for the three surveys. Compliant sample corresponds to those observations that turned their collective expenditure notebook with at least one expenditure recorded. The 1-40 Minimum Wages sample corresponds to households whose income lies in that range (the target population for the IPCA). For all the specifications estimated by IV methods, the p-value of Anderson canonical correlation LR statistic for testing the relevance of the instruments was 0.0000.

Table 4. Household Per Capita Expenditure and Net Income: Headline and Corrected
(in Reais of January 2003)

		1987-88	1996-95	2002-03	Annual percent change
Using official CPI as the deflator					
Household per capita expenditure	Mean	7,826	8,203	9,330	1.2
	Median	4,253	4,086	4,579	0.5
	Bottom 20%	1,210	1,049	1,284	0.4
	Top 20%	23,466	26,091	29,592	1.5
Correcting for estimated CPI bias					
Household per capita expenditure	Mean	4,785	6,087	9,330	4.5
	Median	1,872	2,928	4,579	6.0
	Bottom 20%	359	696	1,284	8.7
	Top 20%	16,482	19,386	29,592	3.9

Notes: Based on estimates of the semi-parametric specification in the “full” sample. The bottom and top 20 percent refer to quintiles of expenditure per survey year in the “full” sample.

Table 5. Brazil: Ownership of Durable Goods, and Sensitivity to Income

	Owns at least one (percent)			Average numbers of units/household			Regression coefficient income sensitivity	
	1987-88	1995-96	2002-03	1987-88	1995-96	2002-03	1987-88	2002-03
Air Conditioner	6.5	8.7	11.2	0.10	0.13	0.16	0.15 [0.005]	0.20 [0.009]
Bicycle	30.1	40.7	39.5	0.43	0.60	0.58	0.22 [0.008]	0.07 [0.015]
Black and White TV	57.5	25.7	5.9	0.70	0.27	0.06	0.51 [0.007]	-0.02 [0.004]
Blender	83.7	84.8	85.6	0.89	0.88	0.91	0.19 [0.005]	0.09 [0.021]
Cake Mixer	35.1	39.2	42.3	0.36	0.40	0.43	0.25 [0.005]	0.22 [0.009]
Car	33.0	35.5	36.0	0.42	0.45	0.45	0.43 [0.006]	0.42 [0.010]
Car or Motorcycle	34.1	36.8	37.9	0.46	0.47	0.48	0.47 [0.009]	0.44 [0.011]
CD Player	n.a.	7.8	7.3	n.a.	0.09	0.08	n.a.	0.06 [0.006]
Color TV	57.5	83.2	93.4	0.70	1.21	1.48	0.51 [0.007]	0.53 [0.027]
Desk Radio	39.8	36.3	21.0	0.46	0.44	0.25	0.18 [0.007]	0.10 [0.010]
Dish Antenna	n.a.	1.8	4.6	n.a.	0.02	0.05	n.a.	0.02 [0.004]
Dishwasher	2.0	6.7	5.9	0.02	0.07	0.06	0.04 [0.002]	0.09 [0.004]
Dryer	4.1	9.8	7.0	0.04	0.10	0.07	0.05 [0.002]	0.07 [0.005]
DVD Player	n.a.	n.a.	6.6	n.a.	n.a.	0.07	n.a.	0.11 [0.005]
Fan	52.7	64.9	68.3	0.73	1.00	1.15	0.25 [0.010]	0.26 [0.028]
Floor Waxer	41.5	23.0	10.0	0.42	0.24	0.10	0.17 [0.006]	0.06 [0.006]
Freezer	6.8	18.7	18.7	0.07	0.19	0.19	0.11 [0.003]	0.14 [0.007]
Hair Dryer	39.7	36.2	31.1	0.48	0.43	0.35	0.36 [0.007]	0.29 [0.010]
Iron	90.2	92.1	91.3	1.07	1.08	1.04	0.26 [0.006]	0.17 [0.029]
LP Player	12.8	5.2	0.0	0.13	0.05	n.a.	0.04 [0.004]	n.a.
Microwave Oven	n.a.	16.2	30.4	n.a.	0.16	0.31	n.a.	0.27 [0.008]
Motorcycle	3.2	2.7	3.6	0.04	0.03	0.04	0.03 [0.007]	0.02 [0.004]
Ozonizer Filter	n.a.	6.3	9.3	n.a.	0.06	0.09	n.a.	0.09 [0.005]
Personal Computer	n.a.	7.0	22.3	n.a.	0.07	0.25	n.a.	0.30 [0.008]
Portable Radio	38.1	28.6	20.8	0.47	0.37	0.25	0.13 [0.008]	0.11 [0.010]
Refrigerator	87.7	91.2	94.4	0.92	0.95	0.98	0.18 [0.004]	0.09 [0.006]
Sewing Machine	47.0	35.4	24.5	0.50	0.39	0.28	0.11 [0.006]	0.06 [0.021]
Sound System	46.0	64.1	66.9	0.49	0.72	0.76	0.23 [0.006]	0.23 [0.011]
Stove	98.7	99.3	99.3	1.04	1.04	1.03	0.02 [0.003]	0.01 [0.004]
Tape Recorder	19.4	17.8	10.3	0.22	0.20	0.11	0.12 [0.003]	0.14 [0.006]
Toaster	9.7	9.3	12.5	0.10	0.09	0.13	0.16 [0.005]	0.01 [0.007]
TV	89.0	94.6	95.2	1.22	1.47	1.54	0.44 [0.008]	0.51 [0.027]
Vacuum Cleaner	22.7	22.1	19.3	0.24	0.23	0.20	0.26 [0.005]	0.23 [0.007]
VCR Player	n.a.	38.2	48.2	n.a.	0.41	0.52	n.a.	0.33 [0.010]
Washing Machine	29.4	46.5	53.2	0.30	0.47	0.54	0.27 [0.005]	0.29 [0.008]
Water Filter	n.a.	n.a.	35.8	n.a.	n.a.	0.37	n.a.	-0.01 [0.009]

Notes: Based on the full sample. Regression coefficients are the sensitivity to log of total expenditure for the number of units of each durable. Coefficient was estimated using log of income as an instrument for log of total expenditure, and using the same controls as the regressions in Table 2. Standard errors in brackets. The abbreviation "n.a." indicates the relevant data were not available for that durable good.

Table 6. Descriptive Statistics for Mexico

	1984	1989	1994	1996	2000	2006
Share of food	0.485[0.18]	0.424[0.18]	0.395[0.178]	0.409[0.166]	0.37[0.158]	0.333[0.158]
Relative price of food	112.427[8.268]	116.975[8.178]	97.947[5.765]	106.567[6.469]	100.729[2.782]	105.263[3.145]
Real Expenditure on food	38199[25517]	34313[42438]	38392[28209]	27239[18825]	29914[18297]	30825[23414]
Real per capita expenditure on food	9768[8445]	9145[12433]	10836[10701]	7725[6336]	8993[7274]	10012[10046]
Real Expenditure	107688[93426]	120265[171545]	142223[214456]	98779[142991]	121650[161620]	143075[181613]
Real Per capita expenditure	28022[33073]	32513[55656]	39937[64405]	28888[52697]	37087[58654]	47780[78219]
Ln (Expenditure/CPI)	11.263[0.746]	11.272[0.816]	11.365[0.83]	11.081[0.729]	11.243[0.76]	11.406[0.809]
Ln (Current Monetary Income/CPI)	11.181[1.015]	11.135[1.327]	11.266[1.054]	10.957[0.999]	11.206[0.995]	11.306[0.971]
Ln (Household size)	1.426[0.574]	1.394[0.534]	1.339[0.53]	1.31[0.536]	1.253[0.523]	1.195[0.561]
# ages 0 to 4	0.521[0.786]	0.468[0.709]	0.444[0.714]	0.448[0.708]	0.366[0.647]	0.34[0.634]
# ages 5 to 9	0.622[0.872]	0.495[0.751]	0.46[0.721]	0.425[0.701]	0.369[0.634]	0.33[0.6]
# ages 10 to 14	0.613[0.901]	0.508[0.806]	0.454[0.746]	0.43[0.736]	0.394[0.682]	0.357[0.629]
# ages 15 to 19	0.594[0.95]	0.566[0.882]	0.463[0.769]	0.462[0.781]	0.395[0.684]	0.375[0.66]
# ages 20 and up	2.457[1.15]	2.538[1.26]	2.498[1.204]	2.452[1.185]	2.427[1.155]	2.393[1.148]
Male head	0.781[0.414]	0.808[0.394]	0.801[0.399]	0.796[0.403]	0.782[0.413]	0.726[0.446]
Spouse present	0.722[0.448]	0.753[0.431]	0.741[0.438]	0.734[0.442]	0.715[0.452]	0.647[0.478]
Head has labor income	0.794[0.405]	0.799[0.401]	0.797[0.402]	0.817[0.387]	0.806[0.396]	0.799[0.401]
Spouse has labor income	0.196[0.397]	0.196[0.397]	0.204[0.403]	0.246[0.431]	0.265[0.442]	0.308[0.462]
Head and spouse have labor income	0.18[0.384]	0.173[0.378]	0.184[0.387]	0.225[0.417]	0.245[0.43]	0.282[0.45]
Owner occupied house	0.528[0.499]	0.625[0.484]	0.691[0.462]	0.686[0.464]	0.692[0.462]	0.652[0.476]
Rental Unit	0.334[0.472]	0.251[0.434]	0.183[0.387]	0.19[0.393]	0.188[0.391]	0.195[0.396]
Other living arrangement	0.138[0.345]	0.124[0.33]	0.126[0.332]	0.124[0.329]	0.119[0.324]	0.153[0.36]
Sample Size	1492	3191	3309	3074	2403	5094

Notes: Expenditure and income data deflated to 2002 Pesos using the INPC. Data for 1992, 1998, 2000, 2002, 2004 and 2005 omitted for presentation purposes

Table 7. Regression Results for Mexico, Pooled Sample
(Dependent Variable: Expenditure Share on Food)

	Bias invariant on income		**Bias linear on income**	
	(1)	(2)	(3)	(4)
	OLS	IV	OLS	IV
Dummy for 1989	-0.065 [0.007]	-0.065 [0.007]	-0.357 [0.107]	-0.397 [0.184]
Dummy for 1992	-0.055 [0.007]	-0.051 [0.007]	-0.154 [0.108]	-0.026 [0.186]
Dummy for 1994	-0.058 [0.008]	-0.054 [0.008]	-0.124 [0.103]	-0.132 [0.170]
Dummy for 1996	-0.084 [0.007]	-0.089 [0.007]	-0.084 [0.103]	-0.097 [0.177]
Dummy for 1998	-0.092 [0.007]	-0.095 [0.007]	-0.144 [0.111]	-0.123 [0.180]
Dummy for 2000	-0.092 [0.007]	-0.092 [0.007]	-0.074 [0.101]	-0.223 [0.167]
Dummy for 2002	-0.105 [0.007]	-0.104 [0.007]	-0.256 [0.100]	-0.361 [0.167]
Dummy for 2004	-0.098 [0.007]	-0.094 [0.007]	-0.269 [0.098]	-0.108 [0.166]
Dummy for 2005	-0.12 [0.007]	-0.118 [0.007]	-0.335 [0.097]	-0.401 [0.167]
Dummy for 2006	-0.112 [0.006]	-0.107 [0.006]	-0.309 [0.098]	-0.336 [0.167]
Ln(Relative price of food)	0.127 [0.028]	0.13 [0.029]	0.123 [0.028]	0.124 [0.029]
Ln(real expenditure)	-0.119 [0.001]	-0.149 [0.002]	-0.13 [0.008]	-0.161 [0.014]
Ln(Household size)	0.007 [0.006]	0.021 [0.006]	0.007 [0.006]	0.021 [0.006]
Number ages 0 to 4	0.009 [0.002]	0.005 [0.002]	0.009 [0.002]	0.005 [0.002]
Number ages 5 to 9	0.01 [0.002]	0.007 [0.002]	0.01 [0.002]	0.007 [0.002]
Number ages 10 to 14	0.01 [0.002]	0.007 [0.002]	0.01 [0.002]	0.007 [0.002]
Number ages 15 to 19	0.007 [0.002]	0.006 [0.002]	0.007 [0.002]	0.006 [0.002]
Number ages 20 and up	0.007 [0.002]	0.009 [0.002]	0.008 [0.002]	0.009 [0.002]
Male head	0.018 [0.004]	0.021 [0.004]	0.019 [0.004]	0.02 [0.004]
Spouse present	-0.011 [0.004]	-0.015 [0.004]	-0.011 [0.004]	-0.015 [0.004]
Head of household has labor income	0.006 [0.003]	0.011 [0.003]	0.005 [0.003]	0.011 [0.003]
Spouse has labour income	-0.003 [0.006]	0.001 [0.006]	-0.003 [0.006]	0.001 [0.006]
Head and spouse have labor income	0 [0.007]	0.004 [0.007]	0 [0.007]	0.004 [0.007]
Owner occupied unit	-0.019 [0.003]	-0.01 [0.003]	-0.019 [0.003]	-0.01 [0.003]
Rental unit	-0.042 [0.003]	-0.03 [0.004]	-0.042 [0.003]	-0.03 [0.004]
Observations	41805	41805	41805	41805
R-squared	0.39	0.37	0.39	0.37
Population Weighted Bias				
Cumulative bias 1984-89 (%)	42.08 [3.25]	35.49 [2.87]	38.43 [3.66]	32.51 [3.67]
Cumulative bias 1984-92 (%)	36.9 [3.93]	28.73 [3.57]	33.92 [4.11]	26.67 [3.92]
Cumulative bias 1984-94 (%)	38.4 [4.15]	30.15 [3.78]	35.51 [4.33]	28.15 [4.35]
Cumulative bias 1984-96 (%)	50.58 [2.8]	44.88 [2.49]	48.45 [2.91]	43.34 [2.83]
Cumulative bias 1984-98 (%)	53.77 [2.71]	47.24 [2.46]	51.28 [2.96]	45.39 [2.96]
Cumulative bias 1984-2000 (%)	54.02 [2.86]	46.08 [2.7]	51.17 [3.24]	43.88 [3.62]
Cumulative bias 1984-2002 (%)	58.56 [2.58]	50.24 [2.5]	55.49 [3.24]	47.73 [3.76]
Cumulative bias 1984-2004 (%)	55.97 [2.47]	46.9 [2.4]	52.5 [3.3]	44.37 [3.61]
Cumulative bias 1984-2005 (%)	63.6 [2.07]	54.73 [2.07]	60.1 [3.12]	51.76 [3.83]
Cumulative bias 1984-2006 (%)	60.91 [2.18]	51.07 [2.21]	57.05 [3.39]	47.85 [4.21]
Expenditure Weighted Bias				
Cumulative bias 1984-89 (%)			26.74 [7.25]	20.92 [9.44]
Cumulative bias 1984-92 (%)			29.56 [7.6]	27.54 [8.97]
Cumulative bias 1984-94 (%)			32.98 [6.86]	25.49 [8.84]
Cumulative bias 1984-96 (%)			48.49 [4.31]	43.22 [5.38]
Cumulative bias 1984-98 (%)			50.16 [4.44]	44.87 [5.61]
Cumulative bias 1984-2000 (%)			51.59 [4.75]	41.03 [6.56]
Cumulative bias 1984-2002 (%)			52.12 [4.94]	42.22 [6.75]
Cumulative bias 1984-2004 (%)			48.03 [5.5]	44.05 [6.87]
Cumulative bias 1984-2005 (%)			53.82 [6.1]	43.6 [8.99]
Cumulative bias 1984-2006 (%)			52.2 [5.61]	42.35 [7.94]

Notes: Robust standard errors for the regression coefficients and bootstrapped standard errors for bias estimates in brackets. Controls also include regional dummies. Regressions (3) and (4) also include the interactions of time dummies with the log of real expenditure. Current real income is used as an instrument to total real expenditure in the IV regressions. For all the specifications estimated by IV methods, the p-value of Anderson canonical correlation LR statistic for testing the relevance of the instruments was 0.0000.

Table 8. Household Per Capita Expenditure and Net Income in Mexico: Headline and Corrected (in 2002 Pesos)

		1984	1994	2006	Annual percent change 1984-2006
Using official CPI as the deflator					
Household per capita expenditure	Mean	26,371	35,389	43,132	2.3
	Median	17,395	20,868	24,115	1.5
	Bottom 20%	6,868	7,618	9,631	1.5
	Top 20%	67,066	100,041	125,101	2.9
Correcting for estimated CPI bias					
Household per capita expenditure	Mean	12,646	25,743	43,132	5.7
	Median	7,830	14,591	24,115	5.2
	Bottom 20%	2,532	4,986	9,631	6.3
	Top 20%	34,811	74,862	125,101	6.0

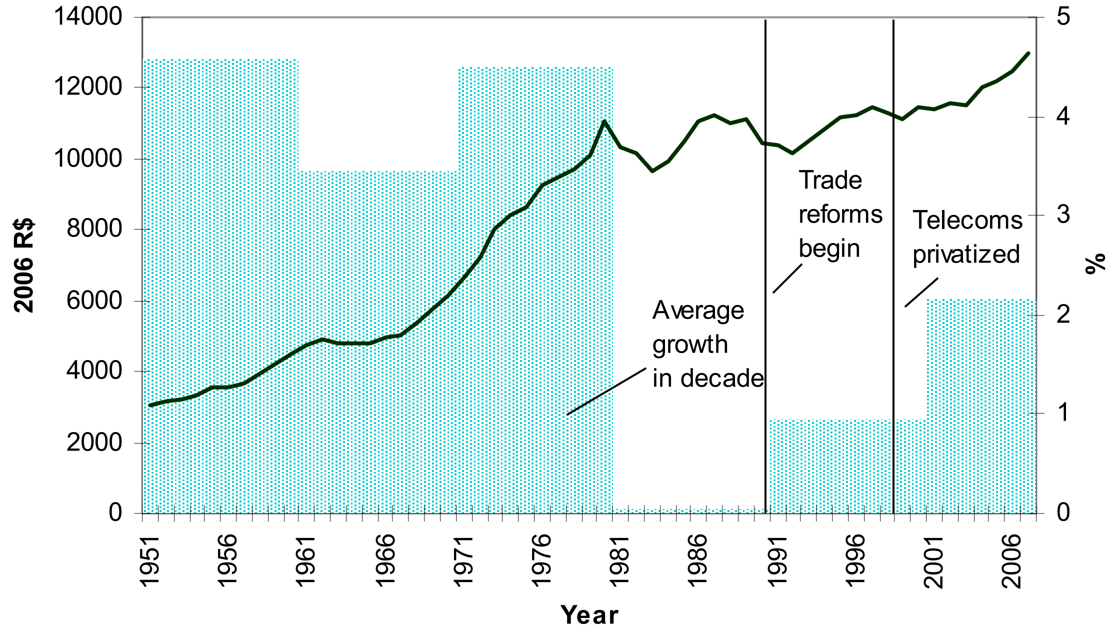
Notes: Based on estimate of CPI bias from semi-parametric model in 1984-2006. The bottom and top 20 percent refer to quintiles of expenditure per survey year

Table 9. Mexico: Ownership of Durable Goods, and Sensitivity to Income

	Owns at least one (percent)			Average numbers of units/household			Regression coefficient income sensitivity	
	1984	1998	2004	1984	1998	2004	1984	2004
Air conditioner	4.9	13.7	14.6	0.05	0.17	0.20	0.07 [0.012]	0.18 [0.011]
Animal traction vehicle	0.3	0.1	0.5	0.00	0.00	0.01	0.00 [0.004]	-0.01 [0.003]
Bicycle	8.6	11.0	7.6	0.11	0.15	0.10	0.00 [0.019]	-0.07 [0.008]
Black and white TV	72.6	n.a.	10.7	0.78	n.a.	0.12	0.05 [0.027]	-0.03 [0.007]
Blender	85.6	89.4	90.4	0.88	0.92	0.93	0.21 [0.020]	0.10 [0.006]
Boat	n.a.	0.5	0.4	n.a.	0.00	0.00	n.a.	0.00 [0.002]
Car	27.8	37.0	39.3	0.35	0.46	0.50	0.53 [0.026]	0.55 [0.012]
CD player	n.a.	14.0	20.1	n.a.	0.15	0.22	n.a.	0.20 [0.009]
Color TV	28.1	n.a.	93.8	0.31	n.a.	1.56	0.42 [0.023]	0.68 [0.015]
Fan	33.7	59.2	58.0	0.42	0.98	1.01	0.11 [0.034]	0.36 [0.024]
Gas heater	46.4	59.7	n.a.	0.47	0.61	n.a.	0.37 [0.022]	n.a.
Gas oven	91.3	96.6	n.a.	0.93	0.97	n.a.	0.15 [0.015]	n.a.
Hand mill	5.7	4.9	5.9	0.06	0.05	0.06	-0.02 [0.012]	-0.01 [0.005]
Heater	n.a.	4.8	6.0	n.a.	0.06	0.07	n.a.	0.11 [0.006]
Heater, other	10.1	1.7	n.a.	0.10	0.02	n.a.	0.01 [0.015]	n.a.
Iron board	91.3	93.4	92.7	0.97	0.98	0.97	0.17 [0.018]	0.12 [0.007]
LP player	n.a.	52.4	67.1	n.a.	0.55	0.73	n.a.	0.29 [0.011]
Microwave oven	n.a.	26.3	53.4	n.a.	0.26	0.54	n.a.	0.34 [0.009]
Motorcycle	1.0	0.8	1.4	0.01	0.01	0.02	0.00 [0.006]	0.00 [0.003]
Other oven	5.9	1.6	n.a.	0.06	0.02	n.a.	-0.10 [0.012]	n.a.
Other vehicle	0.3	0.3	0.3	0.00	0.00	0.00	0.00 [0.003]	0.00 [0.002]
Personal computer	n.a.	10.9	26.8	n.a.	0.12	0.29	n.a.	0.39 [0.009]
Radio	51.2	43.2	29.5	0.57	0.51	0.33	0.09 [0.034]	0.09 [0.011]
Recorder	33.5	55.0	41.1	0.36	0.63	0.47	0.12 [0.027]	0.16 [0.012]
Refrigerator	74.1	87.4	90.2	0.75	0.90	0.92	0.30 [0.020]	0.12 [0.006]
Sewing machine	46.9	33.2	28.1	0.49	0.34	0.30	0.12 [0.026]	0.07 [0.010]
Telephone	n.a.	50.2	63.4	n.a.	0.50	0.63	n.a.	0.28 [0.008]
Truck	5.3	9.8	9.6	0.06	0.11	0.11	0.05 [0.012]	0.13 [0.007]
TV	88.4	95.8	97.1	1.09	1.36	1.68	0.47 [0.028]	0.65 [0.016]
Vaccum cleaner	15.2	15.1	14.6	0.15	0.15	0.15	0.28 [0.016]	0.25 [0.007]
VCR	n.a.	50.2	47.1	n.a.	0.56	0.52	n.a.	0.36 [0.011]
Videogame	n.a.	13.6	15.6	n.a.	0.14	0.17	n.a.	0.18 [0.008]
Washing machine	53.7	64.8	74.7	0.55	0.66	0.76	0.34 [0.024]	0.22 [0.008]
Water pump	7.9	19.2	n.a.	0.08	0.19	n.a.	0.10 [0.013]	n.a.

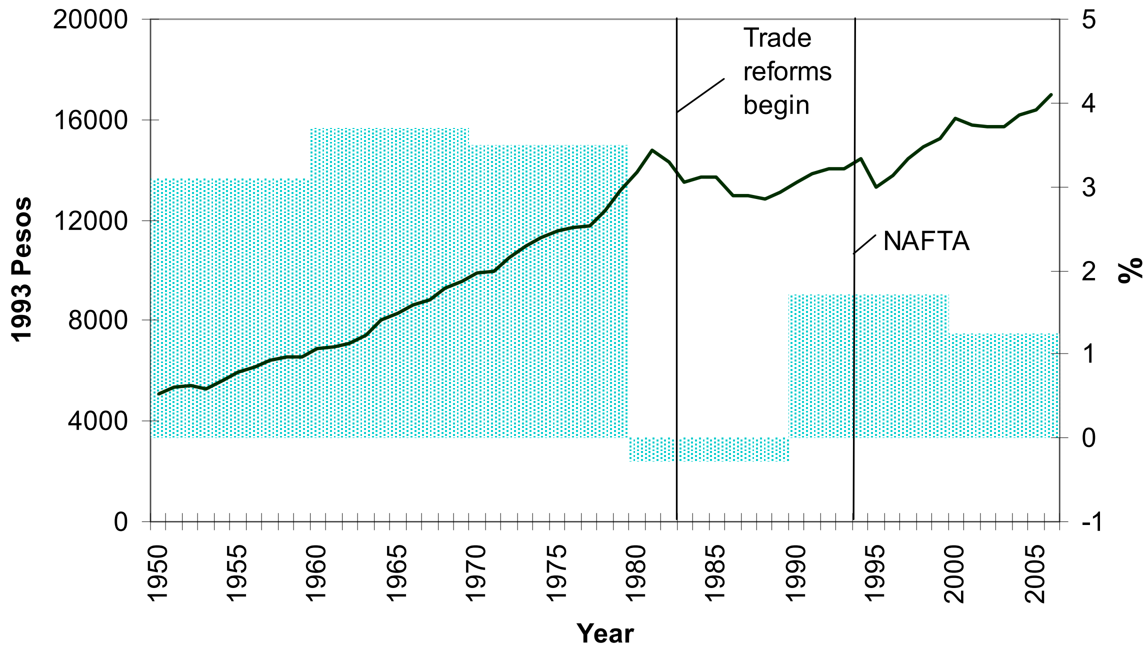
Notes: Based on same sample as in Table 7. Regression coefficients are the sensitivity to log of total expenditure for the number of units of each durable. Coefficient was estimated using log of income as an instrument for log of total expenditure, and same controls as the regressions in Table 7. Standard errors in brackets. The abbreviation “n.a.” indicates the relevant data were not available for that durable good.

Figure 1. Brazil: GDP Per Capita and Average Growth In Decade (Constant 2006 *reais*)



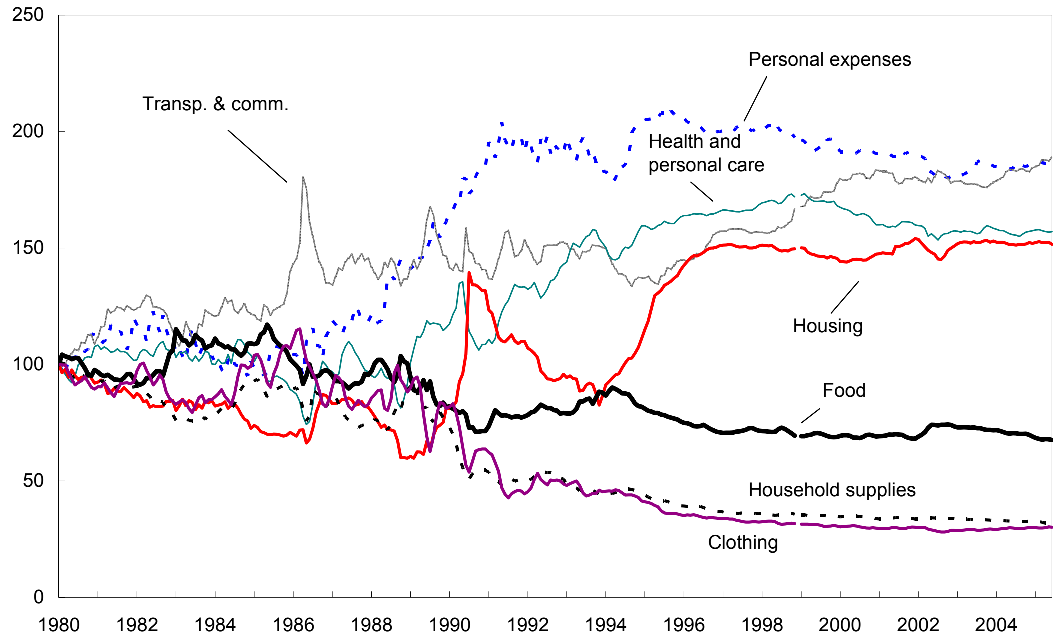
Source: *Instituto de Pesquisa Econômica Aplicada* (IPEA).
Note: Solid line indicates level while columns indicate average growth rate in decade.

Figure 2. Mexico: GDP Per Capita and Average Growth In Decade (Constant 1993 pesos)



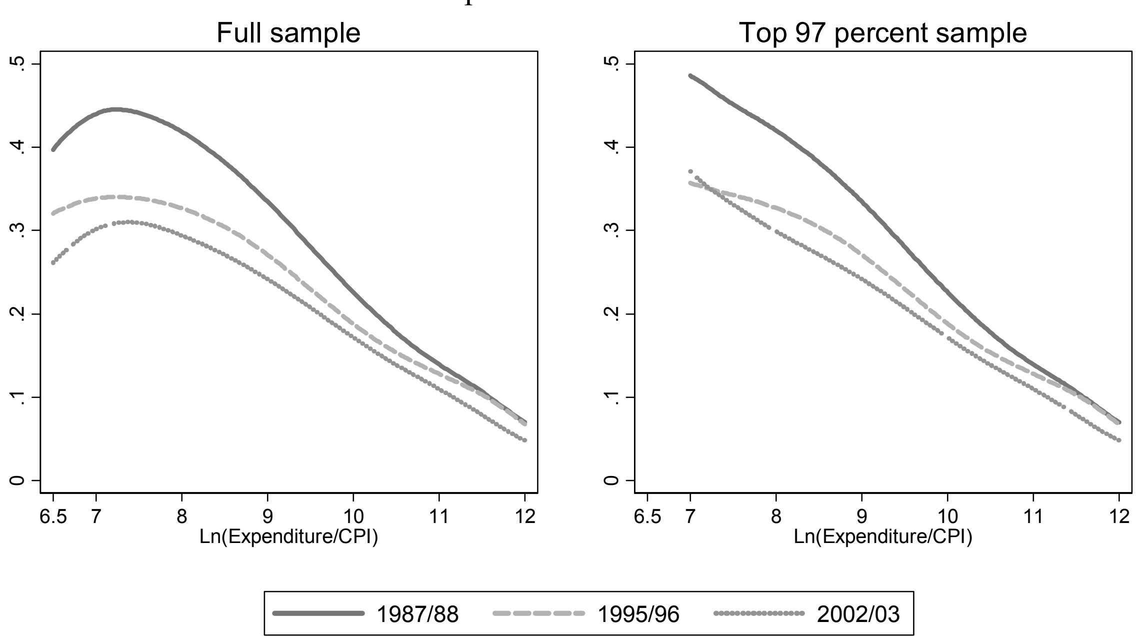
Source: International Financial Statistics (IFS). Note: Solid line indicates level while columns indicate average growth rate in decade.

Figure 3. Changes in Relative Prices in Brazil



Sources: IBGE; and authors' calculations.
Note: Relative price defined as the ratio of the price level for the group to the level of the overall index.

Figure 4. Non-Parametric Estimates of Relationship between Food Shares and Household Expenditure in Brazil



Note: Curves obtained from locally weighted linear regressions using quartic kernel weights.

Figure 5A. Estimated Bias in Brazil in 1987/88-1995/96 as a Function of CPI-Measured Real Expenditure in 1995/96

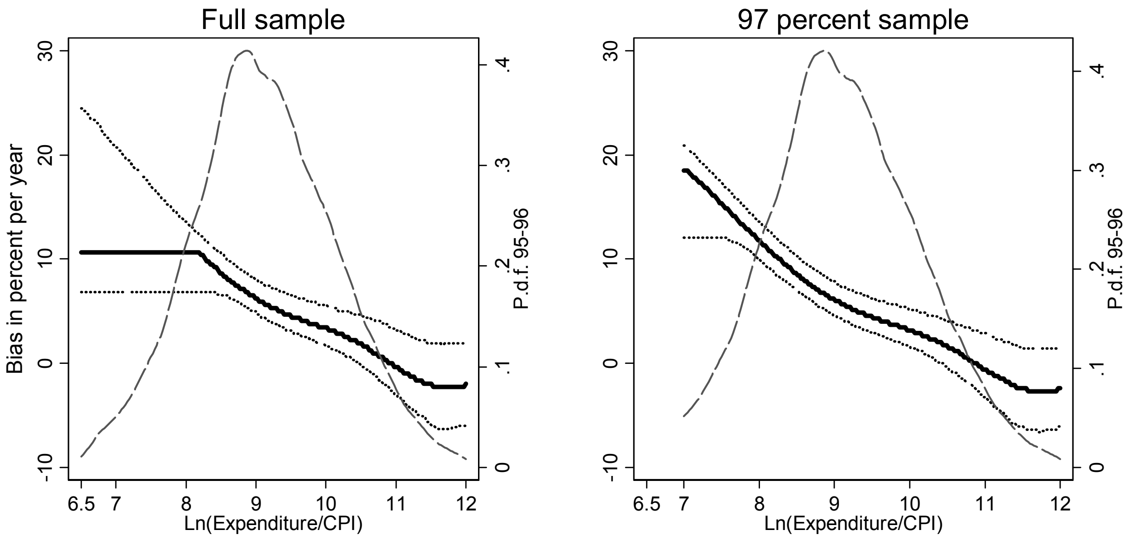
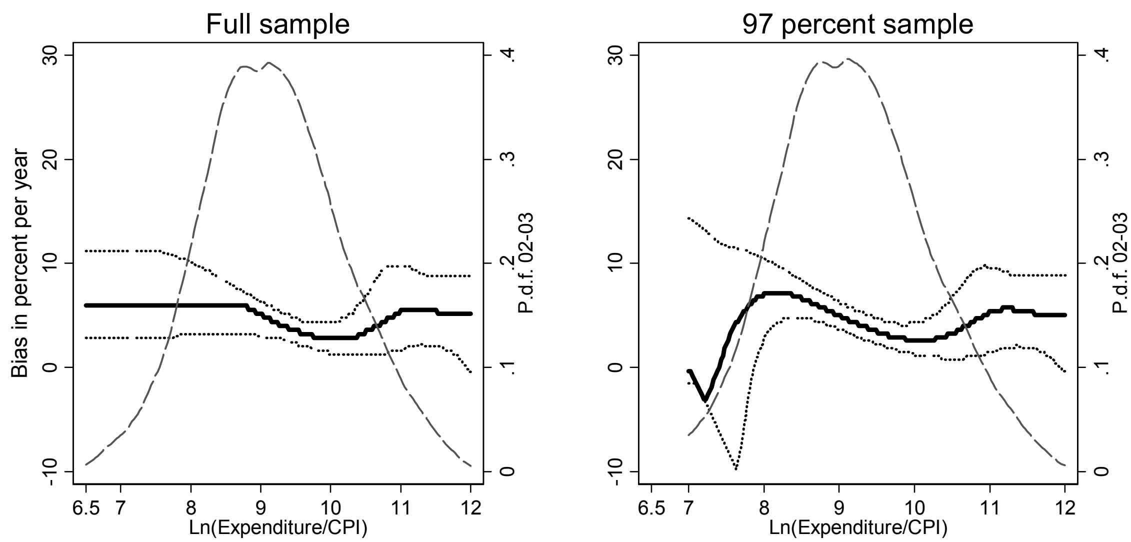


Figure 5B. Estimated Bias in Brazil 1995/96–2002/03 as a Function of CPI-Measured Real Expenditure in 2002/03



Note: Based on semi-parametric bias estimates from the full and the 97 percent samples. Solid line corresponds to bias estimates and dotted lines to 95 percent confidence interval. Dashed line corresponds to the distribution of real expenditures. Bias estimates based on shift of semi-parametrically estimated Engel curves.

Figure 6A. Distribution of Expenditure in Brazil Deflated by the CPI: 87/88, 95/96 and 02/03

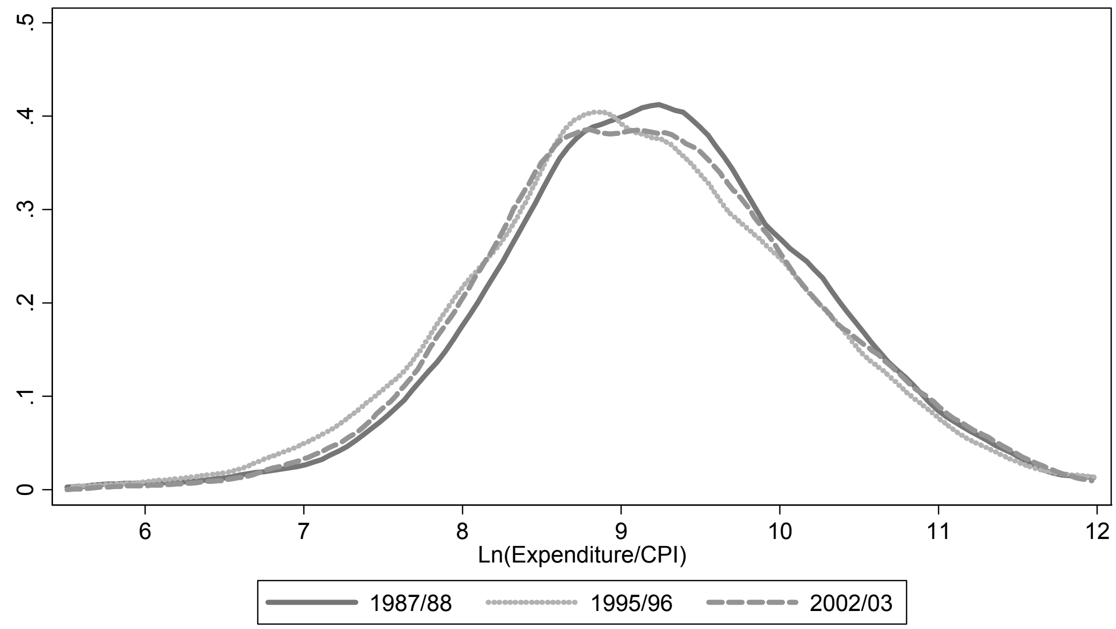
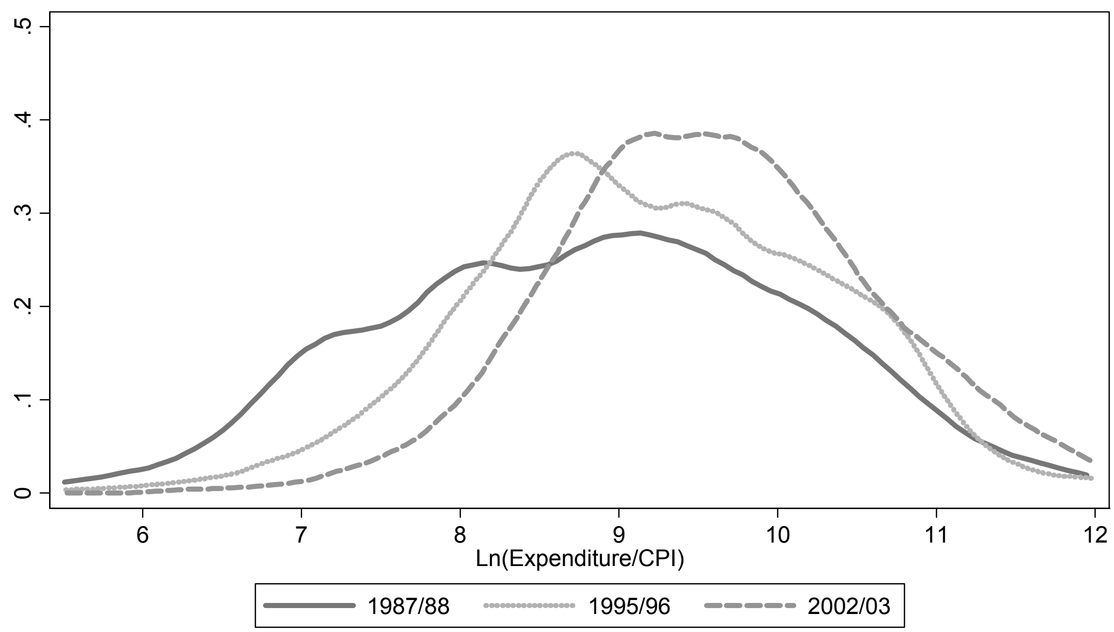


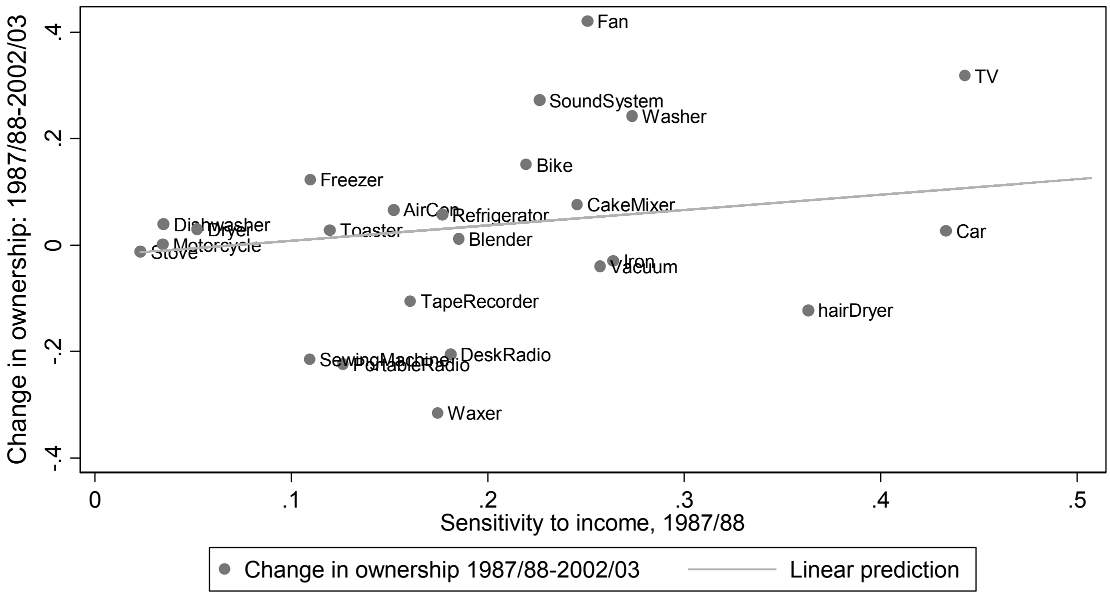
Figure 6B. Distribution of Expenditure in Brazil Deflated by the Estimated True Cost of Living Index: 87/88, 95/96 and 02/03



Note: Based on the semi-parametric model and the full sample. Estimated True Cost of Living Index adjusts CPI-Measured Expenditures in 1995/96 and 2002/03 by the estimated bias since 1987/88.

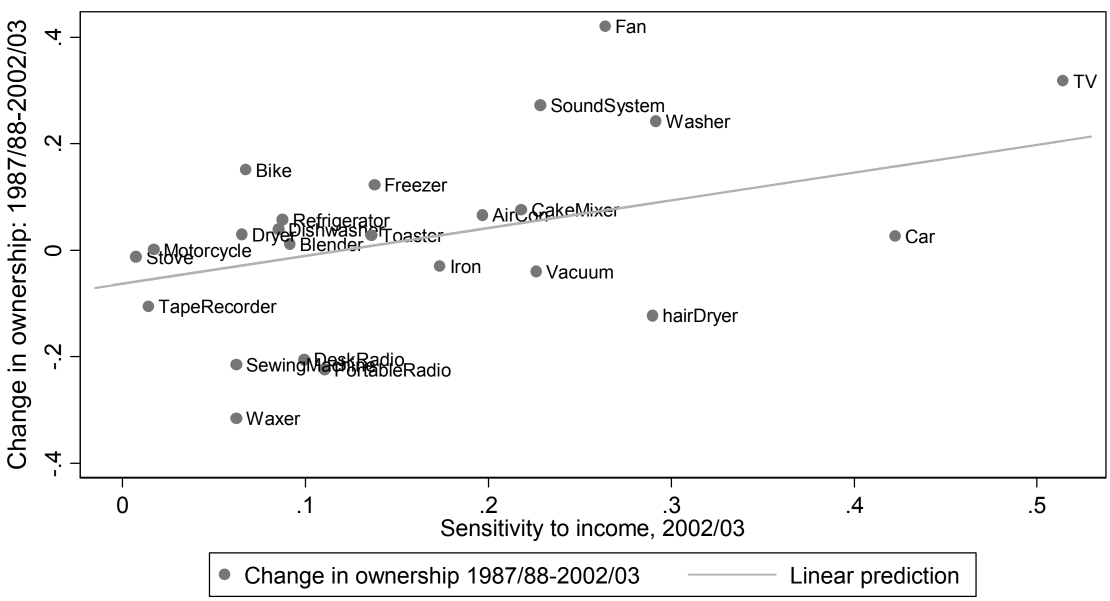
Figure 7. Changes in Durable Goods Holdings in Brazil and Sensitivity to Income

Figure 7A



Slope: .28 [.27] Implied annual bias: 2.17

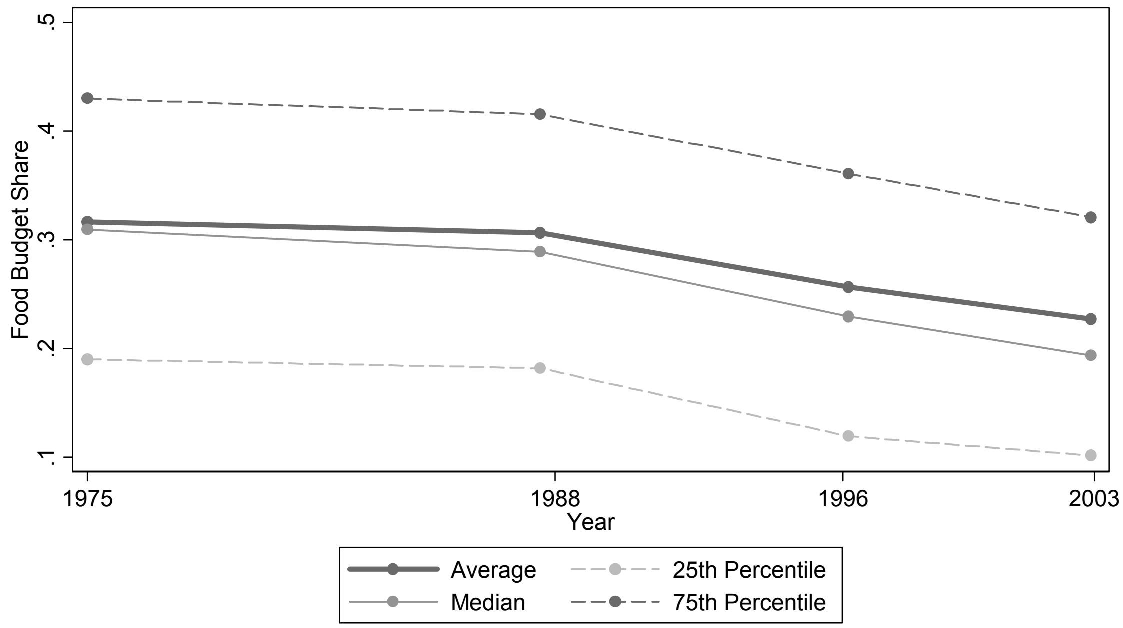
Figure 7B



Slope: .52 [.25] Implied annual bias: 3.69

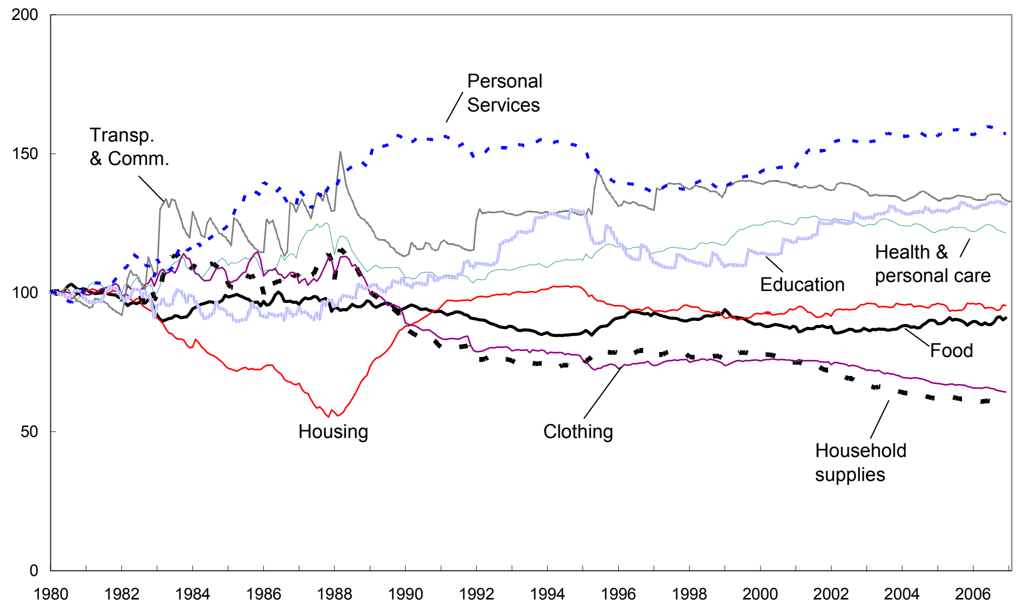
Note: Sensitivity to income defined as the coefficient obtained when average ownership or ownership dummy is regressed on log of total expenditure instrumented by current income and the same controls as in the regressions in Table 2.

Figure 8. Evolution of the Food Budget Share in Brazil Since in 1974/75–2002/03



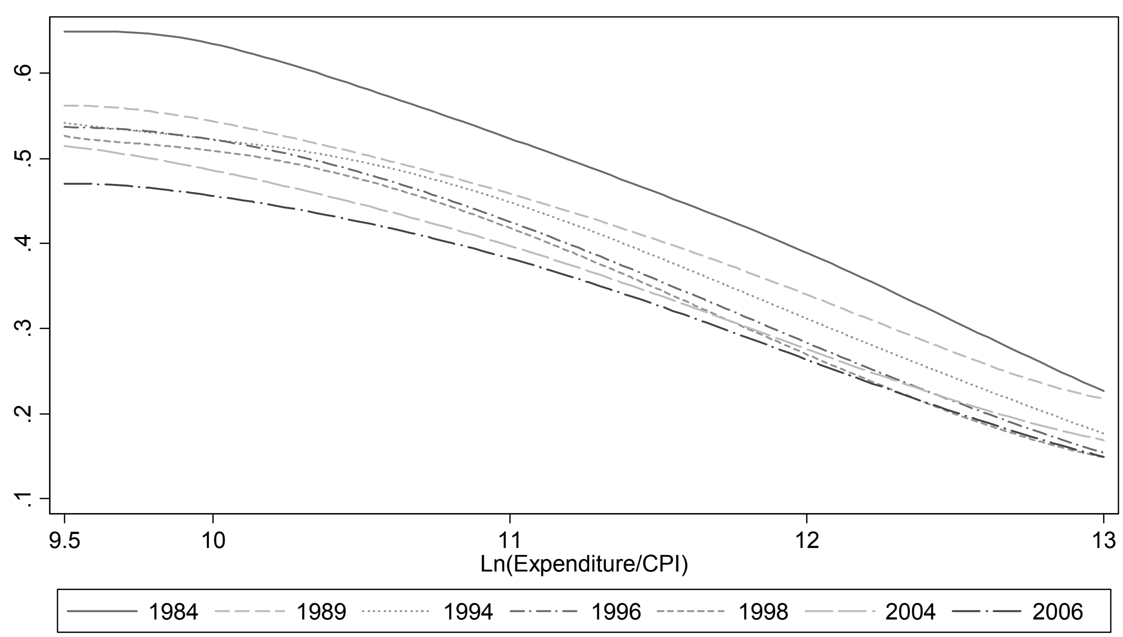
Notes: Estimates based on the 1974/75 ENDEF, and 1987/88, 1995/96 and 2002/03 POF surveys. The comparability of the ENDEF and POF surveys is affected by changes in survey design.

Figure 9. Changes in Relative Prices in Mexico



Sources: Banco de Mexico; and authors' calculations.
Note: Relative price defined as the ratio of the price level for the group to the level of the overall index.

Figure 10. Non-Parametric Estimates of Relationship Between Food Budget Shares and Household Expenditure in Mexico.



Notes: Curves obtained from locally weighted linear regressions using quartic kernel weights.

Figure 11A. Estimated Bias in Mexico in 1984-1998 as a Function of CPI-Measured Real Expenditure in 1998 and Distribution of the Latter.

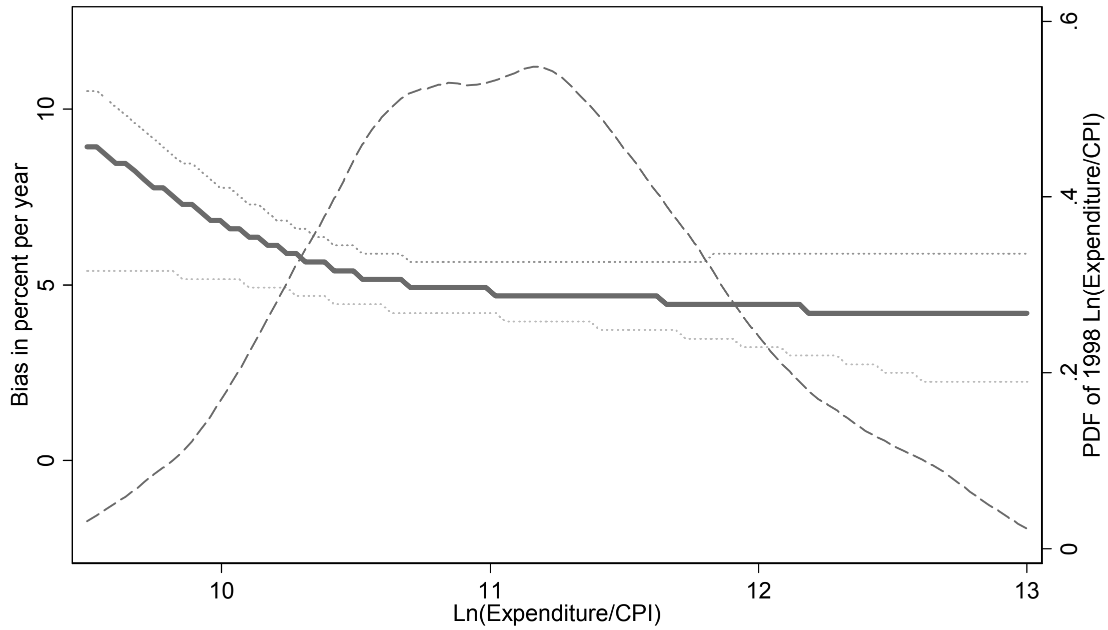
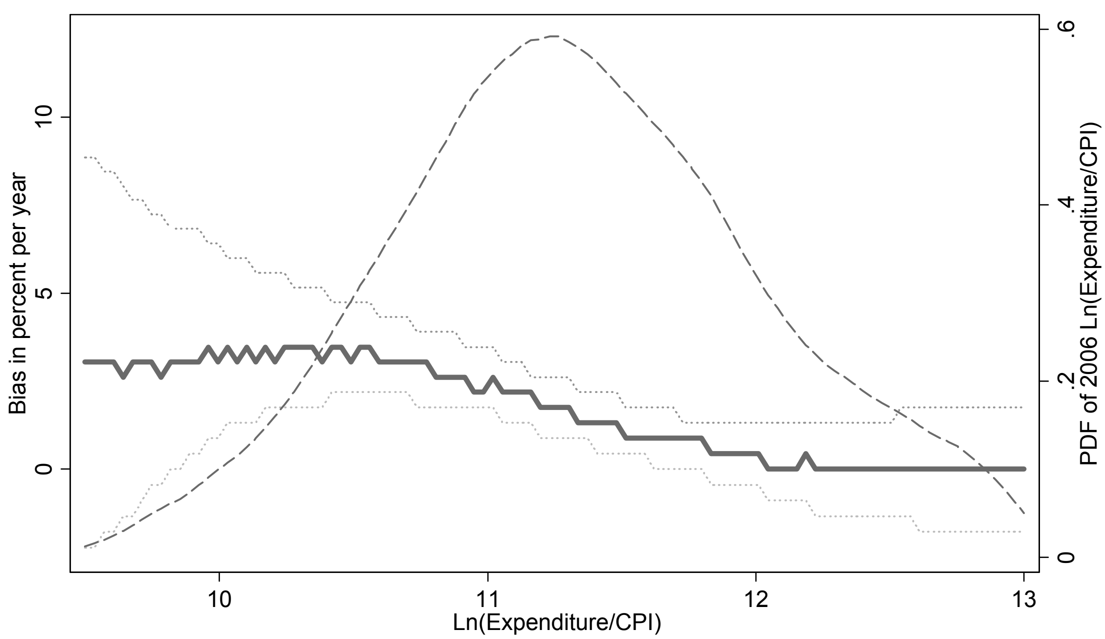
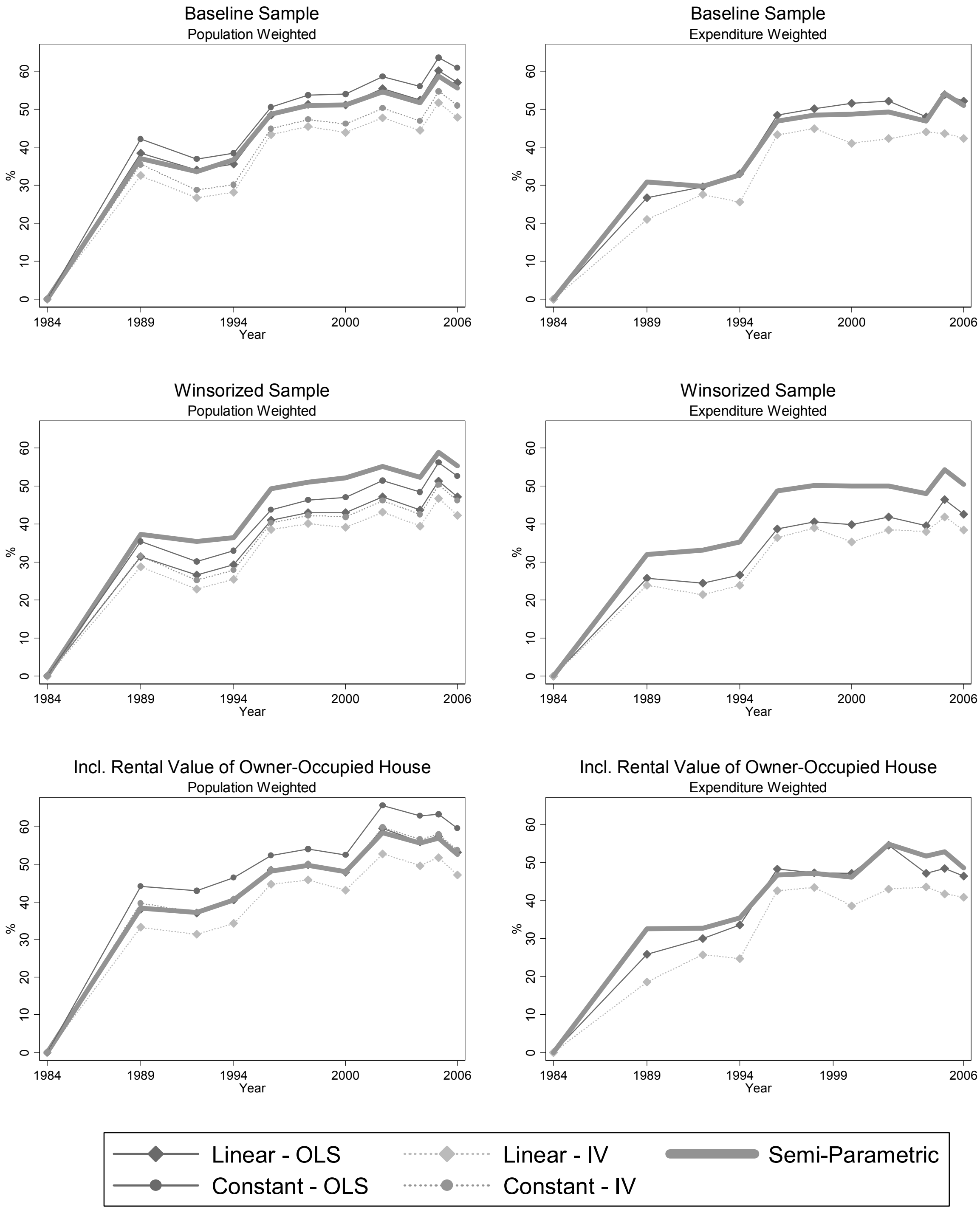


Figure 11B. Estimated Bias in Mexico 1998-2006 as a Function of CPI-Measured Real Expenditure in 2006 and Distribution of the Latter.



Notes: Solid lines corresponds to estimated bias, dotted lines to its 95 percent confidence interval and dashed line to the distribution of CPI-Measured Real Expenditure. Bias estimates based on shift of semi-parametrically estimated Engel curves.

Figure 12. Estimated Cumulative Bias in Mexico Since 1984 Across Different Methods and Samples.



Notes: "Constant" Refers to specification where the bias is constant across all households in a given year; "Linear" to the one where it is linear on (CPI-measured) real expenditure; and "Semi-Parametric" where it is a non-parametric function of (CPI-measured) real expenditure.

Figure 13A. Distribution of CPI-Measured Real Expenditure in Mexico

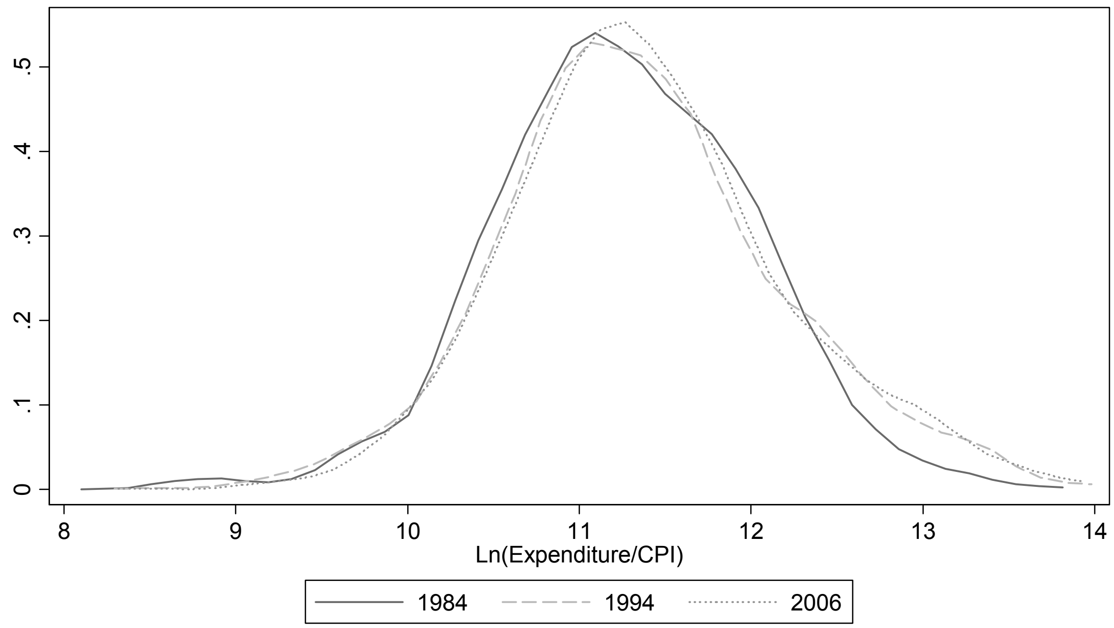
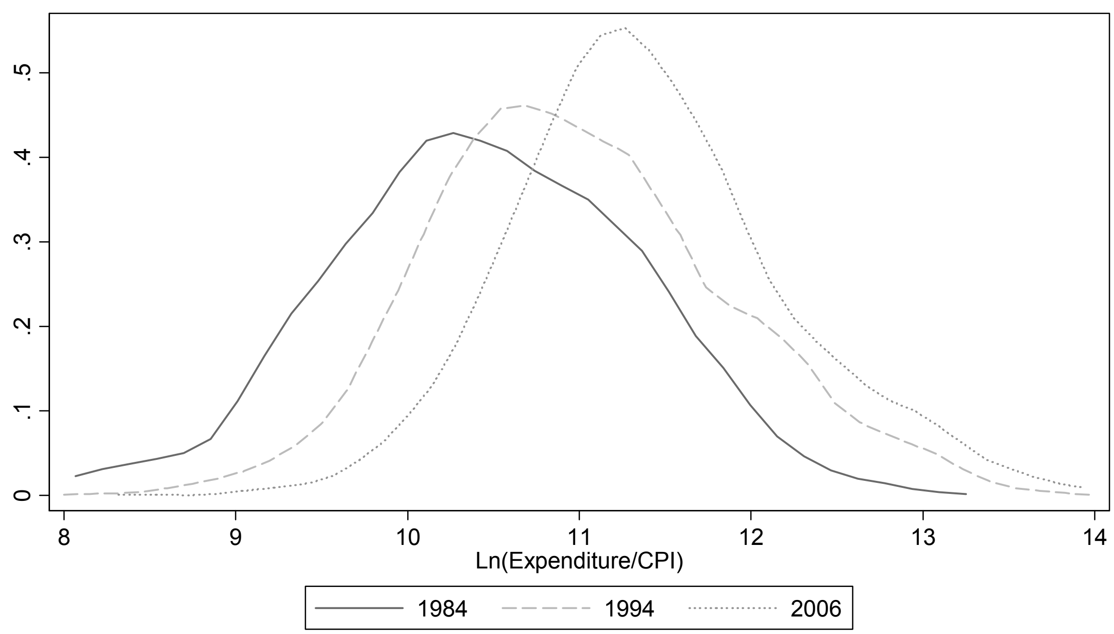
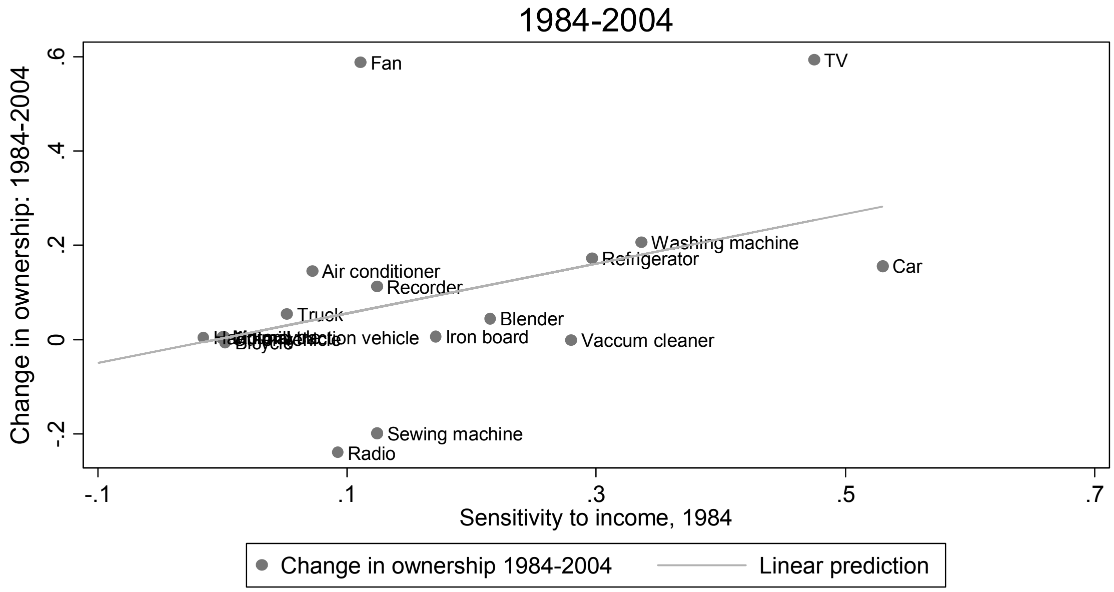


Figure 13B. Distribution of Expenditure in Mexico Deflated by the Estimated True Cost of Living Index



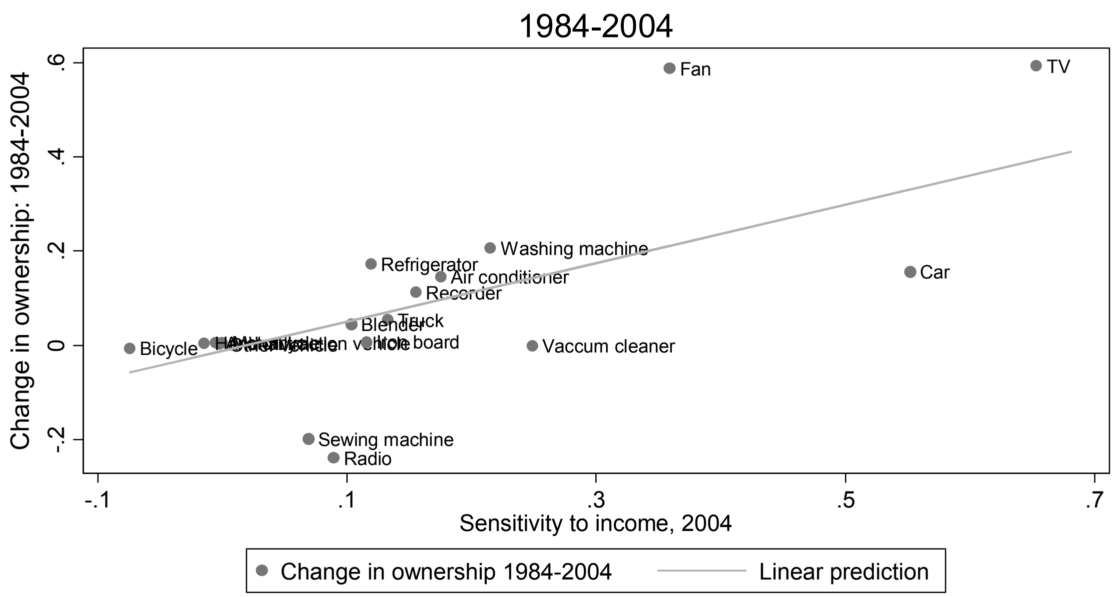
Note: Based on the semi-parametric model in 1984-2006. Estimated True Cost of Living Index adjusts CPI-Measured Expenditures in 1984 and 1994 by the estimated bias from those years to 2006 (the baseline year in this comparison).

Figure 14. Changes in Durable Goods Holdings in Mexico and Sensitivity to Income
Figure 14A



Slope: .52 [.21] Implied annual bias: 2.25

Figure 14B



Slope: .62 [.17] Implied annual bias: 2.74

Note: Sensitivity to income defined as the coefficient obtained when average ownership or ownership dummy is regressed on log of total expenditure instrumented by current income and the same controls used in the regressions in Table 8.