The plutocratic bias in the CPI: Evidence from Spain

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Abstract. We define the plutocratic bias as the difference between the inflation measured according to the current official CPI and a democratic index in which all households receive the same weight. *(i)* We estimate that during the 1990s the plutocratic bias in Spain amounts to 0.055 per cent per year, or about one third of the classical substitution bias estimated by the Boskin Commission for the U.S. *(ii)* We find that a 16-dimensional commodity space can be conveniently reduced to 3 dimensions, consisting of a luxury good and two necessities. The price behavior of these 3 goods provides a convincing explanation of the oscillations experimented by the plutocratic bias. *(iii)* Finally, the fact that the plutocratic bias is positive during this period, implies that the change in money income inequality is between 2 and 5 per cent greater than the change in real income inequality. We study the robustness of these results to the time period considered and to the definition of the group index which serves as an alternative to the CPI. We estimate that during the 1980s and the second part of the 1970s in Spain, the plutocratic bias is 0.033 and 0.239 per cent per year, respectively.

Keywords. Consumer price index, cost-of-living index, aggregation, inequality.

JEL Classification System. C43, D31, D63

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1. INTRODUCTION

In all countries, the official CPI (Consumer Price Index), which is meant to be representative for a certain reference population, is a fixed-weight price index. At least since Konüs (1924), economists have known that a fixed-weight CPI suffers from a "substitution bias" relative to a true cost-of-living index which, instead of maintaining constant the budget shares of the households represented in the index, maintains constant their living standards or welfare levels. But according to the review of the literature carried out by a U.S. Senate Commission headed by Michael Boskin (Boskin *et al.*, 1996), this is not all that is wrong with the U.S. CPI elaborated by the BLS (Bureau of Labor Statistics).

The Boskin Commission focused on five sources of bias in the CPI, all of which are supposed to contribute to an overstatement of the true inflation in the cost of living in the U.S.: (1) The substitution bias among commodities, or the "upper level substitution" problem, causes an estimated upward bias of 0.15 per cent per year. (2) The way elementary price quotations are aggregated within each geographic zone, or the "lower level substitution" problem, is responsible for a bias of 0.25 per cent per year. (3) Consumers adjust their behavior on where to buy in response to price differences between the outlets which happen to be sampled by the BLS and other outlets which are competing with them by lowering prices. The outlet substitution bias due to the failure of the CPI to reflect this aspect of consumer behavior, is estimated at 0.10 per cent per year. The last two sources of bias have to do with the alleged failure of the BLS (4) to take fully into account the quality changes experienced in many sectors of the economy, and (5) to introduce in a timely fashion the new products constantly appearing on the market. These two sources together are supposed to cause a bias of 0.60 per cent per year.

Thus, the Boskin Commission estimates that, on average, during the last few decades the U.S. CPI has been overstating the inflation by 1.1 per cent per year. This bias

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might seem small. However, when compounded over time, the implications for *(i)* the public deficit created through an indexed budget, *(ii)* the wage bargaining process and the determination of the nominal interest rates in the private sector, and *(iii)* the measurement of the economic performance in real terms, are little short of catastrophic.[1] Be this as it may, the fact is that the report has become already very influential.[2] This does not mean, of course, that it has escaped criticism. Some critics question the Commission's analysis of each and every one of the five sources of bias (Moulton *et al.*, 1998). Others point out toward neglected issues and, in particular, the scant attention paid to distributional issues to which we now turn our attention —see, *e.g.*, Pollak (1998), Deaton (1998), and Madrick (1997).

In the CPI context, the issues raised by the heterogeneity of the population are usually identified by asking "Whose cost-of-living index?," a question which is seen to contain three issues in Pollak (1998). "How many cost-of-living indexes?," "Beer or champagne?," and "What type of group indexes?" The first issue refers to whether we should have different indexes for different groups —rich and poor, elderly and non-elderly, urban and rural, etc. The second issue refers to the selection of the appropriate set of items, qualities and outlets that are to be reflected in the index.[3]

Assume that the population of households (individuals or consumers) for whom a representative index must be constructed has been decided, and that a solution has been found for the beer-vs-champagne issue. The third issue, which is the topic of this paper, originates with the nature of the CPI as a group index. Given the commodity space and a household budget survey representative of the reference population, we can use

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[1] For an evaluation of these sources of bias in the measurement of inflation through the official Spanish price index and its implications for the Spanish economy, see Ruiz-Castillo *et al.* (1999b).

[2] As Diewert (1998) puts it: ". . . with a total budget of $25,000, Boskin, Dulberger, Griliches, Gordon and Jorgenson have probably written the most important measurement paper of the century in terms of its impact: Every statistical agency in the world is revaluating its price measurement techniques as a direct result of their report and the widespread publicity it has received."

[3] Against the view of the Boskin Commission and Diewert (1995) that the "lower level substitution" problem is primarily a problem of choosing an appropriate formula for combining the prices of items, Pollak (1995, 1998) argues that it is primarily a problem of selecting the "items" to be priced, and reflects fundamental ambiguities in the meaning of "goods," "commodities," and "items" at theoretical and empirical levels.

each household's budget shares as the fixed weights for the construction of household-specific price indexes. Since Prais (1958), we know that the CPI is the weighted average of such individual price indexes with weights proportional to each household's total expenditures. Because richer households weigh more than poor ones, Prais baptized the CPI as a *plutocratic* price index. The question is whether we can think of a better alternative to this particular construction.[4]

In this paper, we defend that the so-called democratic index, in which all households receive the same weight, is an option worth pursuing. Thus, we define the plutocratic bias as the difference between the inflation measured according to the current official CPI and a democratic index. We offer two reasons for being interested in such a concept. In the first place, it is always interesting to know who suffers the greatest inflation: those households with the largest total expenditures, or those at the bottom of the distribution, in which case we would say that prices have behaved in an anti-rich or an anti-poor manner, respectively. In the first (second) case we should expect that the mean inflation weighted by the total household expenditures would be greater (smaller) than the simple mean. Thus, the plutocratic bias would be positive or negative according to whether prices have behaved in an anti-rich or in an anti-poor manner, respectively —this idea can be traced back to Fry and Pashardes (1985).

In the second place, when two distributions of household expenditure, or income, are expressed at constant prices using household-specific price indexes, the change in nominal income inequality —which is the magnitude usually estimated in the empirical literature— is seen to be equal to the change in real income inequality plus a price term which captures the distributional impact of price changes.[5] Knowledge of the sign of the plutocratic bias takes us a long way in the direction of knowing the sign of the price term. Thus, whether a given change in money income inequality is smaller (larger)

[4] As pointed out by Pollak (1998), the first two issues are given a cursory treatment in footnote 2 and page 71 of Boskin *et al.* (1996). The Boskin Commission never addresses the third issue directly, although Pollak selects some passages of its report which appear to reflect an implicit judgement that the CPI ought to be a plutocratic price index.

[5] The idea that price movements should be included in intertemporal income inequality comparisons was originally suggested by Iyengar and Battacharya (1965). Subsequently, in a social welfare context Muellbauer (1974) showed that under general assumptions on individual preferences real income inequality comparisons are not price independent.

than the socially relevant change in real income inequality depends to a large extent on whether the plutocratic bias is positive (negative).

Nevertheless, the importance of this concept depends crucially on its empirical magnitude. Our main result is that the plutocratic bias in Spain during the 1990s is equal to 0.055 per cent per year —or about one third of the classical substitution bias estimated by the Boskin Commission for the U.S. Nonetheless, averaging magnitudes of different signs underestimates the real importance of this bias. The bias in specific years oscillates from a maximum of 0.150 to a minimum of −0.080 per cent per year. Interestingly, neither the sign nor the magnitude of the bias in a given subperiod depends on the magnitude of the inflation in that subperiod. Using the total expenditures elasticities estimated in an Engel curve system, we find that a 16-dimensional commodity space can be conveniently reduced to 3 dimensions, consisting of a luxury good and two necessities. The price behavior of these 3 goods provides a convincing explanation of the oscillations experimented by the plutocratic bias. Finally, the fact that the plutocratic bias is positive, implies that the gap between the changes in nominal and real household expenditures inequality during the 1990s is between 2 and 5 per cent, depending on the inequality measure and the importance we give to the scale economies in consumption within the household.

The paper studies the robustness of these results in two dimensions. In the first place, we estimate the plutocratic bias for the 1980s and the second part of the 1970s in Spain. We find that, on average, the bias is small in the first case and large in the second: 0.033 and 0.239 per cent per year, respectively.[6] In the second place, we ask what would have been the bias in the measurement of inflation if instead of using the plutocratic CPI we were to use a group index equal to the weighted mean of the household-specific indexes with weights proportional to the household size. We find that such a bias for the 1990s, the 1980s and the second part of the 1970s would be equal to 0.088, 0.015 and 0.223 per cent per year, respectively.

[6] The sign and the magnitude of the bias for these two periods are consistent with previous findings about the fact that the decrease in the real household's total expenditures inequality is greater than the decrease in the money inequality —see Del Río and Ruiz-Castillo (1996), and Ruiz-Castillo and Sastre (1999).

The rest of the paper is organized as follows. Section 2 presents the essentials on individual and aggregate price indexes. Section 3 presents the empirical results on the plutocratic bias in Spain during the 1990s, while Section 4 studies the implications for total household expenditures inequality measurement. Section 5 is devoted to the robustness of those results during previous periods, and the use of weighted group indexes with weights proportional to household size. Section 6 summarizes and discusses the political implications of our results in a heavily indexed economy.

2. INDIVIDUAL AND GROUP INDEXES

2.1. *Individual Price Indexes*

Let there be I goods and H households indexed by $i = 1, \ldots, I$ and $h = 1, \ldots, H$, respectively, and let $\mathbf{q} = (q_1, \ldots, q_I)$ be a commodity vector. Each household h is characterized by her total expenditures, x^h, and her preferences represented by a utility function, $u = U^h(\mathbf{q})$. Assume that all households have the same preferences, so that $u = U^h(\mathbf{q}) = U(\mathbf{q})$ for all h, and let $c(u, \mathbf{p})$ be the cost function, which gives the minimum cost of achieving the utility level u at prices $\mathbf{p}$. Under general conditions, we know that $x^h = c(U(\mathbf{q}^h), \mathbf{p})$, where $\mathbf{q}^h$ is the utility-maximizing commodity vector at prices $\mathbf{p}$ when the household expenditures are x^h.

Consider two price vectors $\mathbf{p}_0$ and $\mathbf{p}_t$ in periods 0 and t. A true or a Konüs cost-of-living index (COLI for short) which takes as its reference the utility level u^h, is defined as the ratio of the minimum cost of achieving that utility level at prices $\mathbf{p}_t$ and $\mathbf{p}_0$, *i.e.*,

$$\kappa(\mathbf{p}_t, \mathbf{p}_0; u^h) = \frac{c(\mathbf{p}_t, u^h)}{c(\mathbf{p}_0, u^h)}.$$

When the reference utility is the utility-maximizing level at prices $\mathbf{p}_0$, denoted by u_0^h, we say that the COLI $\kappa(\mathbf{p}_t, \mathbf{p}_0; u_0^h) = c(\mathbf{p}_t, u_0^h)/c(\mathbf{p}_0, u_0^h)$ is a Laspeyres type index.

Given a reference commodity vector, $\mathbf{q}^h$, we can define a statistical price index (SPI) as the ratio of the cost of acquiring $\mathbf{q}^h$ at prices $\mathbf{p}_t$ and $\mathbf{p}_0$,[7]

$$\ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}^h) = \frac{\mathbf{p}_t \cdot \mathbf{q}^h}{\mathbf{p}_0 \cdot \mathbf{q}^h}.$$

[7] An SPI can also be written as a weighted average of individual-commodity indexes. Let w_{i0}^h be the good-i household budget share at prices $\mathbf{p}_0$, *i.e.*, $w_{i0}^h = p_{i0}q_{i0}^h/\mathbf{p}_0\mathbf{q}_0^h$. Then we have that $\ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_0^h) = \sum_i w_{i0}^h(p_{it}/p_{i0})$.

When $\mathbf{q}^h = \mathbf{q}_0^h$, the utility-maximizing consumption bundle at prices $\mathbf{p}_0$, we say that the SPI $\ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_0^h) = \mathbf{p}_t \cdot \mathbf{q}_0^h / \mathbf{p}_0 \cdot \mathbf{q}_0^h$ is a Laspeyres type index.

A fundamental theorem in Konüs (1924) establishes that, under general assumptions, the Laspeyres SPI provides an upper bound to the Laspeyres COLI,

$$\kappa(\mathbf{p}_t, \mathbf{p}_0; u_0^h) \leq \ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_0^h).$$

Equality is obtained when preferences are of the Leontief type, *i.e.*, when there is no substitution between goods.

2.2. The CPI

Define the vector of aggregate quantities bought in situation 0 by $\mathbf{Q}_0 = (Q_{10}, \ldots, Q_{I0})$, where $Q_{i0} = \sum_h q_{i0}^h$, and let $W_{i0} = p_{i0} Q_{i0} / \mathbf{p}_0 \cdot \mathbf{Q}_0$. The aggregate Laspeyres SPI —for period t based on period 0— is then defined as follows:

$$\mathcal{L}(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_0) = \sum_i W_{i0} \frac{p_{it}}{p_{i0}} = \frac{\mathbf{p}_t \cdot \mathbf{Q}_0}{\mathbf{p}_0 \cdot \mathbf{Q}_0}. \tag{1}$$

However, the CPI actually computed by statistical agencies is not exactly an aggregate price index of the type defined in equation (1). The reason is that individual behavior is typically investigated by means of a household budget survey conducted in a period τ prior to the index base period 0. As it is shown in the Appendix, the CPI based on period 0 is an aggregate SPI defined by[8]

$$CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) = \frac{\mathcal{L}(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{Q}_\tau)}{\mathcal{L}(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{Q}_\tau)} = \frac{\mathbf{p}_t \cdot \mathbf{Q}_\tau}{\mathbf{p}_0 \cdot \mathbf{Q}_\tau}. \tag{2}$$

This is what the BLS calls a *modified Laspeyres* aggregate price index (Moulton, 1996). What are the normative bases for such a construction? To answer this question we need to define a set of household-specific modified Laspeyres price indexes:

$$cpi(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h) = \frac{\ell(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{q}_\tau^h)}{\ell(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{q}_\tau^h)} = \frac{\mathbf{p}_t \cdot \mathbf{q}_\tau^h}{\mathbf{p}_0 \cdot \mathbf{q}_\tau^h}.$$

[8] Note that we could instead use *average* quantities, $\bar{\mathbf{Q}}_\tau$, with elements $\bar{Q}_{i\tau} = \frac{1}{H} Q_{i\tau}$, since the $\frac{1}{H}$ terms in the numerator and denominator would cancel off. Hence the notion of the CPI being referred to an 'average consumer.'

For each h, let $u_\tau^h = U(\mathbf{q}_\tau^h)$. It is easy to see that the ratio of the corresponding Laspeyres COLIs leads to what we can call a modified Laspeyres COLI:

$$\frac{\kappa(\mathbf{p}_t, \mathbf{p}_\tau; u_\tau^h)}{\kappa(\mathbf{p}_0, \mathbf{p}_\tau; u_\tau^h)} = \frac{c(\mathbf{p}_t, u_\tau^h)}{c(\mathbf{p}_0, u_\tau^h)} = \kappa(\mathbf{p}_t, \mathbf{p}_0; u_\tau^h).$$

Konüs theorem assures that, for each h, $\ell(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{q}_\tau^h) - \kappa(\mathbf{p}_0, \mathbf{p}_\tau; u_\tau^h) \geq 0$ and $\ell(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{q}_\tau^h) - \kappa(\mathbf{p}_t, \mathbf{p}_\tau; u_\tau^h) \geq 0$, but it says nothing about the ratio of the Laspeyres indexes which give rise to an individual CPI. However, the household budget survey collection period τ is typically not far apart from the base year 0 of the CPI system. Thus, under the assumption that the substitution bias $\ell(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{q}_\tau^h) - \kappa(\mathbf{p}_0, \mathbf{p}_\tau; u_\tau^h)$ is smaller than $\ell(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{q}_\tau^h) - \kappa(\mathbf{p}_t, \mathbf{p}_\tau; u_\tau^h)$, we have that a household-specific CPI provides an upper bound to a modified Laspeyres COLI. As shown in the Appendix,

$$CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) = \sum_h \phi^h \, cpi(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$$

where $\phi^h = \mathbf{p}_0 \cdot \mathbf{q}_\tau^h / \mathbf{p}_0 \cdot \mathbf{Q}_\tau$. Thus, only under the assumption that, for a sufficiently large number of households,

$$cpi(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h) = \frac{\ell(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{q}_\tau^h)}{\ell(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{q}_\tau^h)} \geq \frac{\kappa(\mathbf{p}_t, \mathbf{p}_\tau; u_\tau^h)}{\kappa(\mathbf{p}_0, \mathbf{p}_\tau; u_\tau^h)} = \kappa(\mathbf{p}_t, \mathbf{p}_0; \mathbf{u}_\tau^h),$$

then the aggregate CPI provides an upper bound to a plutocratic-weighted mean of modified Laspeyres COLIs:[9] $CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) \geq \sum_h \phi^h \, \kappa(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$. Otherwise it would instead provide a lower bound. Nonetheless, the proximity of the theoretical construct —*i.e.*, a COLI— and the empirical counterpart —*i.e.*, the CPI— constitutes a rather remarkable situation.

[9] In the democratic case, we have that $\frac{1}{H}\sum_h \ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h) \geq \frac{1}{H}\sum_h \kappa(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$. Under the same assumption, the simple mean of modified Laspeyres SPIs constitutes an upper bound to the simple mean of modified Laspeyres COLIs.

3. THE PLUTOCRATIC BIAS

3.1. The Data

In order to estimate the plutocratic bias defined below, we need to construct a series of household-specific Laspeyres price indexes. For that purpose, we use the following two pieces of publicly available information in Spain: the 1990–91 household-budget survey (EPF) used to estimate the weights of the official CPI, and a set of price subindexes at a certain level of spatial and commodity disaggregation.

The EPF *(Encuesta de Presupuestos Familiares)* collected by the Spanish statistical agency, INE *(Instituto Nacional de Estadística),* from April 1990 to March 1991, is a household budget survey of 21,155 household sample points, representative of a population of approximately 11 milllion households and 38 million persons occupying residential housing in all of Spain, including the North African cities of Ceuta and Melilla.

The INE collects elementary price indexes (denoted by E_{ijt} in the Appendix) for a commodity basket consisting of 471 items in each of the 52 provinces under the CPI present system, based in 1992. For confidentiality reasons, the INE does not publish this information at the maximum disaggregation level. Instead, it publishes on a monthly basis price subindexes for the period January 1993 to January 1998 for a commodity breakdown of 110 *subclases*, 57 *rúbricas*, 33 *subgrupos* and 8 *grupos* at the national level, the *rúbricas*, *subgrupos* and *grupos* at the 18 Autonomous Community level, and the *subgrupos* and *grupos* at the 52 province level.

For any commodity breakdown, it is possible to reconstruct the official CPI series using an appropriately defined aggregate budget shares vector. Similarly, defining a budget share vector for every household in the 1990–91 sample, we can obtain a series of household-specific CPIs for any commodity breakdown. In principle, the only difference between alternative specifications of the commodity space, is that the dispersion of the set of individual CPIs should be greater the greater the disaggregation level of the price information used in their construction. Unfortunately, in spite of using the same informational basis as the INE —namely, the 1990–91 EPF— we find several small discrepancies between our estimates of the aggregate budget share vectors and those published by the INE —for the details, see Ruiz-Castillo *et al.* (1999b). Thus, the CPI series which we can reconstruct vary slightly depending on the different commodity breakdowns characterizing the price information we use. In Ruiz-Castillo *et al.* (1999b) we find that the specification consisting of the 21 food *rúbricas* at the Autonomous Community level, and the 32 non-food *subgrupos* at the provincial level outperforms the rest of the alternatives according to various statistical and economic criteria.

It should be emphasized that our series of household-specific price indexes defined over this 53 commodity space differ from the series underlying the official CPI in two ways. In the first place, there are a number of aspects in the official definition of total household expenditures for which we believe there are superior alternatives. We refer to: *(i)* the definition of housing expenditures for households occupying non-rental housing; *(ii)* the inclusion of imputations for home production, wages in kind and subsidized meals, and *(iii)* the estimation of annual food and drink expenditures using all the available information on bulk purchases in the 1990–91 EPF. The joint impact of these modifications is important: according to Ruiz-Castillo *et al.* (1999b), the official CPI understates the true Spanish inflation from 1992 to January of 1998 in 0.241 per cent per year.

In the second place, it should be noticed that the Spanish CPI is not the modified Laspeyres price index defined in equation (2), which takes as a reference the mean quantity vector actually acquired by the EPF households at the time they were interviewed in the 1990–91 survey period. The reason is that the INE does not use the adjustment factors $A_{ij\tau}$ defined in the Appendix. Fortunately, Lorenzo (1998) provides such factors for the 110 subclases at the national level. Using this information, for each household h interviewed in a quarter τ during the 1990–91 period (τ = Spring, Summer, Autumn of 1990, and Winter of 1991), we construct a series of modified Laspeyres SPIs, $\ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$, based on period 0 = Winter of 1991, which takes as a reference the commodity vector $\mathbf{q}_\tau^h$ actually acquired during the interview quarter τ:[10]

[10] If we normalize this series at prices of period 0 = 1992, we can obtain the conceptually correct CPI, that is, $\ell(\mathbf{p}_t, \mathbf{p}_\tau; q_\tau^h)/\ell(\mathbf{p}_0, \mathbf{p}_\tau; q_\tau^h) = \mathbf{p}_t \cdot \mathbf{q}_\tau^h / \mathbf{p}_0 \cdot \mathbf{q}_\tau^h = CPI^h(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$. For the details of this construction, see Ruiz-Castillo *et al.* (1999a). This series of modified Laspeyres price indexes is available at http://www.eco.uc3m.es/investigacion/epf.html.

3.2. A Definition of the Plutocratic Bias

We will divide the period Winter 1991–January 1998 in the 7 subperiods shown on table 1 below. For each h we define the inflation (or deflation) caused by the evolution of prices in a given subperiod by:

$$\pi_t^h = \frac{\ell_t^h - \ell_{t-1}^h}{\ell_{t-1}^h}$$

The distribution of individual inflations in each subperiod is denoted by $\boldsymbol{\pi}_t = (\pi_t^1, \ldots, \pi_t^H)$. For the entire period, we have $\mathbf{\Pi} = (\Pi^1, \ldots, \Pi^H)$, where $\Pi^h = (\ell_T^h - 1)$, where $T = \text{Jan98}$. The aggregate inflation for the population as a whole according to the plutocratic scheme is

$$\text{PLUT}_t = \frac{\sum_h \phi^h (\ell_t^h - \ell_{t-1}^h)}{\sum_h \phi^h \ell_{t-1}^h} = \frac{\sum_h (\phi^h \ell_{t-1}^h)(\ell_t^h - \ell_{t-1}^h)/\ell_{t-1}^h}{\sum_h \phi^h \ell_{t-1}^h} = \sum_h \psi_t^h \, \pi_t^h,$$

where $\psi_t^h = \phi^h \ell_{t-1}^h / \sum_h \phi^h \ell_{t-1}^h$. For the democratic scheme,

$$\text{DEM}_t = \frac{\sum_h \ell_t^h - \ell_{t-1}^h}{\sum_h \ell_{t-1}^h} = \sum_h \xi_t^h \, \pi_t^h,$$

where $\xi_t^h = \ell_{t-1}^h / \sum_h \ell_{t-1}^h$ —note that ψ_t^h is proportional to $\phi^h \xi_t^h$. Since $\ell_0^h = 1$, for the overall period from 0 to T the weights simplify to ϕ^h and $\frac{1}{H}$ and we have $\text{PLUT} = \sum_h \phi^h (\ell_T^h - 1)$, and $\text{DEM} = \frac{1}{H} \sum_h (\ell_T^h - 1)$. We define the plutocratic bias in the measurement of inflation in subperiod t by $B_t = \text{PLUT}_t - \text{DEM}_t$, and for the overall period by $B = \text{PLUT} - \text{DEM}$.[11] Notice that, as pointed out in the Introduction, if price changes in subperiod t (or for the entire period) are relatively more detrimental to the rich, *i.e.*, if π_t^h (or Π^h) are greater for the rich than for the poor households, then we expect the plutocratic mean of individual inflations in the plutocratic case to be greater than the democratic mean. That is, B_t or B are positive (negative) according to whether the price change in the corresponding time interval is anti-rich (anti-poor).

3.3. The Main Findings

In the first two columns of Table 1 we show the plutocratic and the democratic means of both $\mathbf{\Pi}$ and $\boldsymbol{\pi}_t$. For comparative purposes with the measurement units used by the

[11] Note that the inflation rate does not display temporal separability —*i.e.*, the inflation for a given period does not equal the sum of inflations for a partition of that period. If the inflation rate were defined instead as the log price change, then temporal separabilty would hold but group separability would be lost.

Boskin Commission, all figures are expressed in annual terms. Notice that the aggregate inflation keeps decreasing over time, from a high 6.9 percentage points during the first subperiod to a low 2.4 percentage points during 1997. In column 3 we measure the plutocratic bias as the difference between the plutocratic and the democratic means of distributions $\mathbf{\Pi}$ and π_t. Note, however, that this summary for the whole period understates the true importance of the plutocratic bias since the positive and negative biases in various subperiods offset each other.

Table 1. The Plutocratic Bias During the 1990s
(In Percent Per Year)

		Inflation		
t	Subperiods	Plutocratic	Democratic	Plutocratic bias
1	Winter 91 to 1992	6.989	6.911	0.078
2	1992 to Jan 1993	5.394	5.244	0.150
3	Jan 93 to Jan 94	5.271	5.165	0.105
4	Jan 94 to Jan 95	4.621	4.701	-0.080
5	Jan 95 to Jan 96	4.079	4.130	-0.050
6	Jan 96 to Jan 97	3.180	3.090	0.090
7	Jan 97 to Jan 98	2.494	2.369	0.125
	Winter 91 to Jan 98	4.632	4.577	0.055

The main findings are the followings three: (1) For the period as a whole, B is positive and equal to 0.055 per cent per year. This is, approximately, one third of the substitution bias estimated by the Boskin Commission for the U.S. economy, which is equal to 0.15 per cent per year. (2) Price behavior is not uniform over the entire period: B_t is negative during 1994 and 1995, indicating that during these two years prices have caused relatively more damage to the poor than to the rich households. (3) Neither the sign nor the magnitude of B_t in a given period depends on whether inflation is large or small during that period.

3.4. An Economic Interpretation

Which goods are primarily consumed by the poor or the rich households? To answer this question, we must begin by recognizing the fact that, in a heterogeneous world, total expenditures of households with different characteristics are not directly comparable. Following Buhmann *et al.* (1988) and Coulter *et al.* (1992a, 1992b), we adopt an equivalence scale model in which scale economies in consumption depend only on household size, s^h, and adjusted total household expenditures are defined by

$$y^h = \frac{x^h}{(s^h)^\theta}, \quad \theta \in [0,1] \tag{3}$$

When $\theta = 0$, adjusted expenditures coincide with unadjusted household expenditures, while if $\theta = 1$, it becomes *per capita* household expenditures. Taking a single adult as the reference type, the expression s^θ can be interpreted as the number of equivalent adults in a household of size s. Thus, the greater the equivalence elasticity θ, the smaller the scale economies in consumption or, in other words, the larger the number of equivalent adults.

Table 2. Budget Shares in the Distribution of Adjusted Household Expenditures ($\theta = 0.5$), and Total Expenditure Elasticities for 16 Goods

	Quintiles						
GOODS	Q1	Q2	Q3	Q4	Q5	All	Elasticities
1. Personal Transportation	5.00	7.65	9.51	11.51	14.87	11.54	1.655
2. Clothing	4.98	6.24	7.52	8.36	8.58	7.79	1.593
3. Furniture	0.55	0.85	1.05	1.30	1.64	1.28	1.734
4. Domestic Services	0.14	0.19	0.25	0.57	1.46	0.78	2.242
5. Leisure, Education, Cultural	3.27	4.74	5.80	6.68	7.34	6.30	1.189
6. Other Personal Services	7.88	10.55	11.94	13.43	14.71	12.92	1.340
7. Other Household Goods	1.65	1.86	1.94	1.94	2.09	1.97	1.239
8. Medicine	2.07	2.37	2.65	2.65	2.76	2.62	1.253
LUXURY GOODS (1 + ... + 8)	25.54	34.45	40.66	46.44	53.45	45.20	1.451
9. Food[a]	33.75	27.41	23.37	19.39	13.23	19.69	0.566
10. Housing Utilities	4.63	3.61	3.11	2.61	2.04	2.74	0.482
NECESSITIES I (9 + 10)	38.38	31.02	26.48	22.0	15.27	22.43	0.555
11. Alcoholic Drinks and Tobacco	3.16	3.02	2.82	2.57	1.97	2.48	0.847
12. Remainder of Group I[b]	4.48	4.53	4.58	4.14	3.25	3.94	0.811
13. Shoes	1.79	2.03	1.98	1.95	1.61	1.82	1.097
14. Housing	21.72	20.31	19.24	18.81	21.16	20.20	0.874
15. Other Transport and Comm.	2.46	2.40	2.37	2.44	2.15	2.31	0.775
16. Household Maintenance	2.46	2.24	1.88	1.64	1.14	1.62	0.795
NECESSITIES II (11 + ... + 16)	36.07	34.53	32.87	31.55	31.28	32.37	0.866

[a] Except "Other Food Products" (beef, prepared fish, fruit preserves and other unclassified foods)

[b] "Non-alcoholic Drinks" and "Other Food Products."

In Table 2, we present the budget shares for the quintiles of the distribution of adjusted total expenditures for an intermediate value of $\theta = \frac{1}{2}$. The commodity space consists of 16 goods, classified in three groups according to whether their total expenditures elasticity is greater than 1 (luxuries), considerably smaller than 1 (necessities I, dominated by Food expenditures), or weakly smaller than 1 (necessities II, dominated by Housing expenditures). The total expenditure elasticities are estimated at the mean of the variables in the following system of Engel-curve regressions:

$$w_i^h = \alpha_i + \beta_i \ln(y^h) + \gamma_i \mathbf{z}^h + \varepsilon_i^h, \quad i = 1, \ldots, 16,$$

where: ε_i^h is an error term; $y^h = x^h/\sqrt{s^h}$ is total household expenditures adjusted for household size with parameter $\theta = \frac{1}{2}$; and $\mathbf{z}^h$ is a vector of household characteristics including *(i)* demographic variables (household size and composition, the household head's age and age squared), *(ii)* socioeconomic variables (number of income earners, educational level and socioeconomic category of the household head, educational level and labor status of the spouse, number of dwellings and characteristics of the residential unit), as well as *(iii)* seasonal and geographic variables (municipality size and Autonomous Community of residence). In Figure 1 we display the joint distribution of the individual budget shares for these three goods and the logarithm of the adjusted total household expenditures,[12] the last panel shows the estimated Engel curves (trimming the 1 percent tails off the support of the adjusted household expenditures).

[Insert Figure 1]

Intuitively, the evolution of prices would tend to damage to relatively greater extent the richer households over the poorer ones according to whether the luxury good or the necessities experience the greatest relative increase. For the entire period, the inflation experienced by the luxury good and the two necessities are 31.59, 21.08, and 38.46 index points, respectively. In Figure 2 we represent the evolution of the inter-annual inflation of the three goods in relation to the general inflation as well as the inter-annual B_t, $t =$ January 1992,…,January 1998.

[Insert Figure 2]

In spite of the fact that the second necessity shows the stronger price growth, the behavior of the luxury good and the first necessity is the main explanatory force behind

[12] The boxplots on the top margins show the 1, 25, 50, 75 and 99 percentiles.

the positive sign of the plutocratic bias. To test this idea, we run a regression of the inter-annual (January to January) plutocratic bias B_t, from $t =$ January 1992,..., January 1998, on the corresponding monthly inter-annual price subindexes for the 3 goods and a constant. The results, with robust t-ratios in parentheses (generalized least-squares and Cochrane-Orcutt regressions yield identical results), are the following:

$$\hat{B}_t = \underset{(1.80)}{0.025} + \underset{(9.95)}{0.050\, L_t} - \underset{(-42.14)}{0.056\, NI_t} - \underset{(-0.75)}{0.0043\, NII_t} \qquad R^2 = 0.96$$

All the coefficients have the expected sign —although the one corresponding to Necessities II (NII) is not statistically significant— and the results corroborate the explanatory power of the Luxury good (L) and Necessities I (NI).

4. THE IMPLICATIONS FOR INEQUALITY MEASUREMENT

4.1. The Change in Money and Real Inequality

Let us assume that we want to compare the household income or expenditures distributions in two different time periods, $\mathbf{x}_0 = (x_0^1, \ldots, x_0^H)$ and $\mathbf{x}_t = (x_t^1, \ldots, x_t^{H'})$, with H' not necessarily equal to H. Let $\mathbf{p}_0$ and $\mathbf{p}_t$ be the price vectors in the two situations. For each h, we can express the household's total expenditures in situation 0 at prices $\mathbf{p}_t$, $x_{0,t}^h$, by multiplying her original money income in period 0, x_0^h, by an SPI of the Laspeyres type:

$$x_{0,t}^h = x_0^h\, \ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_0^h) = \mathbf{p}_0 \cdot \mathbf{q}_0^h\, \frac{\mathbf{p}_t \cdot \mathbf{q}_0^h}{\mathbf{p}_0 \cdot \mathbf{q}_0^h} = \mathbf{p}_t \cdot \mathbf{q}_0^h.$$

For any $H \geq 2$, let $\mathcal{I} : \Re^H \mapsto \Re$ be any convenient inequality index satisfying continuity, S-concavity, scale independence and replication population invariance. The change in money income inequality, ΔM, can be expressed as the sum of two terms:

$$\Delta M = \mathcal{I}(\mathbf{x}_t) - \mathcal{I}(\mathbf{x}_0) = [\mathcal{I}(\mathbf{x}_t) - \mathcal{I}(\mathbf{x}_{0,t})] + [\mathcal{I}(\mathbf{x}_{0,t}) - \mathcal{I}(\mathbf{x}_t)] = \Delta R + \Delta P, \qquad (4)$$

where $\Delta R = \mathcal{I}(\mathbf{x}_t) - \mathcal{I}(\mathbf{x}_{0,t})$ is the change in real income inequality and $\Delta P = \mathcal{I}(\mathbf{x}_{0,t}) - \mathcal{I}(\mathbf{x}_t)$ captures the distributional impact of price changes on inequality measurement according to the households' preferences at period 0.[13] ¿From a social point of view, we

[13] We could have expressed the distribution in situation t at prices $\mathbf{p}_0$ by using an appropriate Paassche type price index. In this case, $\Delta P = \mathcal{I}(\mathbf{x}_t) - \mathcal{I}(\mathbf{x}_{t,0})$ would have measured the distributional impact of price changes according to the households' preferences at t.

are primarily interested in the sign of ΔR. However, in the absence of household-specific price indexes —which is the dominant situation in the empirical literature— researchers usually measure only ΔM. Therefore, it is important to know the relationship between ΔR and ΔM when the change from $\mathbf{p}_0$ to $\mathbf{p}_t$ is not neutral, that is, when ΔP is different from 0.

The compensating variation introduced by Hicks (1940), CV^h, is the amount of money that household h must receive in compensation for the price change from $\mathbf{p}_0$ to $\mathbf{p}_t$, *i.e.*,

$$CV^h = x^h_{0,t} - x^h_0 = (\ell(\mathbf{p}_t, \mathbf{p}_0; q^h_0) - 1) \times x^h_0 = \pi^h_t \, x^h_0. \tag{5}$$

As before, the plutocratic bias is defined as $B_t = \sum_h (\phi^h - \frac{1}{H})\pi^h_t$. Intuitively, when $B_t > 0$, for example, household inflation π^h_t tends to be greater for the rich than for the poor households. Taking equation (5) into account, in this case we expect the compensating variation to be also greater for the rich than for the poor households. Given that $\mathbf{x}_{0,t} = \mathbf{x}_0 + \mathbf{CV}$, where $\mathbf{CV} = (CV^1, \ldots, CV^H)$, as long as $\mathcal{I}(\mathbf{CV}) > \mathcal{I}(\mathbf{x}_0)$ we have that $\mathcal{I}(\mathbf{x}_{0,t}) > \mathcal{I}(\mathbf{x}_0)$ or, in other terms, $\Delta P > 0$.[14] In view of (4), we expect that $\Delta M \gtrless \Delta R$ as long as $B_t \gtrless 0$. This means that when, for instance, $B_t > 0$ —so that the price change from $\mathbf{p}_0$ to $\mathbf{p}_t$ is anti-rich— then the term ΔP is positive and $\Delta M > \Delta R$. In this case, *(i)* if the normatively significant term ΔR is positive, so that there has been an increase in real income inequality, then the change in money income inequality would overstate the socially relevant magnitude. *(ii)* Conversely, if $\Delta R < 0$, then ΔM would understate the reduction in real income inequality.

4.2. Results

Let us identify situation with the Winter 1991 and t with T, *i.e.*, January 1998 —so that, as before, we drop the subscript from B. Since, as we have seen, $B = 0.055$ per cent per year during this period, we expect the term ΔP to be positive. To verify this circumstance, we must select an inequality index.

[14] Note that the reverse need not be the case. It might be that $\mathcal{I}(\mathbf{x}_{0,t}) > \mathcal{I}(\mathbf{x}_0)$ because the vector $\mathbf{CV}$ is very unequal and there are a lot of re-orderings between $\mathbf{x}_t$ and $\mathbf{CV}$. A re-ordering between $\mathbf{x}_0$ and $\mathbf{CV}$ means that, for a pair of households h and h', $x^h_0 > x^{h'}_0$ but $CV^h = \pi^h_t x^h_0 < CV^{h'} = \pi^{h'}_t x^{h'}_0$. Therefore, $\pi^h_t > \pi^{h'}_t$, *i.e.*, the inflation would be greater for the rich than for the poor household and the plutocratic mean of $\boldsymbol{\pi}$ could be smaller than the democratic mean. Thus, in spite of $\mathcal{I}(\mathbf{x}_{0,t}) > \mathcal{I}(\mathbf{x}_0)$, we can have that $B_t < 0$.

It is well known that the Generalized Entropy family of inequality indexes are the only measures of relative inequality that satisfy the usual normative properties required from any inequality index and, in addition, are decomposable by population subgroups —see, *e.g.*, Shorrocks (1984). The family can be described by means of the following convenient cardinalization:

$$\begin{aligned}
\mathcal{I}_c(\mathbf{y}(\theta)) &= \frac{1}{H}\frac{1}{c^2-c}\sum_h\left(\left(\frac{y^h(\theta)}{\mu(\mathbf{y}(\theta))}\right)^c-1\right), \quad c \in (0,1) \\
\mathcal{I}_0(\mathbf{y}(\theta)) &= \frac{1}{H}\sum_h -\ln\left(\frac{y^h(\theta)}{\mu(\mathbf{y}(\theta))}\right); \\
\mathcal{I}_1(\mathbf{y}(\theta)) &= \frac{1}{H}\sum_h\left(\frac{y^h(\theta)}{\mu(\mathbf{y}(\theta))}\right)\ln\left(\frac{y^h(\theta)}{\mu(\mathbf{y}(\theta))}\right),
\end{aligned} \tag{6}$$

where $\mathbf{y}(\theta)$ is the distribution of adjusted total household expenditures defined in equation (3), and $\mu(\cdot)$ stands for the mean of the distribution. The parameter c summarizes the sensitivity of $\mathcal{I}_c$ in different parts of the total household expenditures distribution: the more positive (negative) c is, the more sensitive $\mathcal{I}_c$ is to differences at the top (bottom) of the distribution —see Cowell and Kuga (1981). $\mathcal{I}_1$ is the original Theil index, while $\mathcal{I}_0$ is the mean logarithmic deviation.

In Table 3 we present the estimates for $\Delta P_c(\theta)$ for the parameter values $c = -1, 0, 1, 2$ and $\theta = 0.0, 0.5, 1.0$. These values have been estimated using 1,000 Bootstrap samples.[15] In each Bootstrap sample we draw 21,155 households using stratified resampling according to the 260 strata in the original survey. For each (c, θ) pair, we obtain a Bootstrap distribution of 1,000 inequality measures.

Table 3. The Distributional Effect of Price Changes on Adjusted Total Household Inequality During the 1990s
(Bootstrap Standard Errors in parentheses)

	$c=-1$	$c=0$	$c=1$	$c=2$
$\theta=0.0$	2.43	2.06	2.36	3.15
	(0.1619)	(0.1264)	(0.2102)	(0.6065)
$\theta=0.5$	3.19	2.97	3.27	4.09
	(0.1692)	(0.1506)	(0.2505)	(0.6727)
$\theta=1.0$	3.29	3.24	3.67	5.02
	(0.1669)	(0.1542)	(0.2506)	(0.6242)

[15] See Hall (1992) for a description of Bootstrap methods and Mills and Zandvakili (1997) for an application to income inequality measurement.

For every θ, the increase in income inequality is slightly greater for $c = -1$ and, above all, for $c = 2$. This means that prices have affected primarily households at both ends of the distribution and, above all, to those at the upper tail. On the other hand, for every c, $\Delta P_c(\theta)$ increases slightly as θ increases, that is, as the importance of scale economies diminishes. At any rate, we find that, as expected, $\Delta P_c(\theta) > 0$ for all c and θ and, in view of the standard errors, this increase in inequality is statistically significant. Thus, because prices have behaved in an anti-rich manner during 1991–1998, we conclude that the inequality of the 1990–91 adjusted household expenditures inequality is between 2 and 5 per cent greater in January 1998 than in the Winter of 1991. Assume that we have data on household expenditures in 1998 and that the change in money household expenditures inequality is to be equal to 10 per cent, for example. According to equation (4), our results on $\Delta P_c(\theta)$ imply that the increase in real household expenditures inequality would be only between 5 and 8 per cent depending on the values of θ and c. If the change in money household expenditures inequality were to be equal to -10 per cent, then the decrease in real household expenditures inequality would be 12 or 15 per cent.

5. ROBUSTNESS

5.1. The Time Period

In this subsection we study the robustness of our results on the B trend in two different directions. In the first place, we consider the period covered by the two previous Spanish CPI systems, which run from August 1985 to December 1992 (base year = 1983) , and from January 1977 to July 1985 (base year = 1976), respectively. The EPFs which serve to estimate the official weights were conducted from April 1980 to March 1981, and from July 1973 to June 1974, respectively. These are household budget surveys strictly comparable to the 1990–91 EPF, containing 23,972 and 24,151 household sample units, representative of, approximately, a population of 10 or 9 million households and 37 or 34 million persons in 1980–81 or 1973–74. In this case, we do not have reasons to depart from the official definition of total household expenditures, but we must take into account that, as before, the Spanish CPI is not a modified Laspeyres price index. We construct two series of appropriate household-specific price indexes with the information provided by: *(i)* the 1980-81 and 1973-74 EPFs; *(ii)* the official monthly price information for 58 and 57 *rúbricas* at the national level in the 1983 and 1976 bases, respectively, and *(iii)* a series of adjustment factors for 60 goods at the national level provided by Catasús *et al.* (1986) for the first period, and for only 5 goods at the national level provided by García España and Serrano (1980) for the second period.[16]

For the 1980s, we base the individual modified Laspeyres price indexes at Winter 1981, and distinguish between the following nine subperiods: 1) From Winter 1981 to 1983, the base year of the CPI system; 2) from 1983 to December 1984; and 3) - 9) from r to $r+1$, where $r =$ December 1984,...,December 1990. For the second part of the 1970s, we base the individual modified Laspeyres price indexes at the midpoint of 1973 and 1974, and distinguish between the following six subperiods: 1) From the 1973-1974 to 1976, the base year of the CPI system; 2) from 1976 to December 1977; and 3) - 6) from r to $r+1$, where $r =$ December 1977,..., December 1980. The results on the plutocratic bias in the measurement of inflation for these two periods are presented on Tables 4 and 5.

Table 4. The Plutocratic Bias During the 1980s
(In Percent Per Year)

		Inflation		
t	Subperiods	Plutocratic	Democratic	Plutocratic bias
1	Winter 81 to 1983	13.002	12.886	0.116
2	1983 to December 1984	9.484	9.655	-0.171
3	1985	7.782	7.857	-0.075
4	1986	8.230	8.420	-0.189
5	1987	4.587	4.292	0.295
6	1988	5.866	5.956	-0.089
7	1989	6.922	6.962	-0.040
8	1990	6.641	6.474	0.167
9	1991	5.604	5.350	0.254
	Winter 81 to Winter 91	8.557	8.524	0.033

The first thing to notice is that the Spanish inflation during the 1970s and 1980s is considerably greater than during the 1990s. However, as before, there is no relationship

[16] For the 1980s, we had to work with a set of 52 goods which constitute the minimum common denominator between the 58 official *rúbricas* and the 60 goods in Catasús *et al.* (1986). For the 1970s, the minimum common denominator is given by the 5 goods in García España and Serrano (1980). For the details of these constructions, see Ruiz-Castillo *et al.* (1999a). Both series of modified Laspeyres price indexes are available at http://www.eco.uc3m.es/investigacion/epf.html.

Table 5. The Plutocratic Bias From 1973–74 to Winter 1981
(In Percent Per Year)

t	Subperiods	Inflation: Plutocratic	Inflation: Democratic	Plutocratic bias
1	1973–74 to 1976	16.869	16.816	0.053
2	1976 to December 1977	23.024	22.497	0.527
3	1978	16.529	16.246	0.283
4	1979	15.523	14.806	0.717
5	1980	15.311	15.323	-0.012
6	1981	14.541	14.578	-0.037
	1973–74 to Winter 81	17.746	17.506	0.239

between the size of the aggregate inflation in a given subperiod and the sign or the magnitude of the plutocratic bias. As far as the 1980s is concerned, from 1983 to 1986 and from 1988 to 1989, $B_t < 0$. The sum of these magnitudes is more than offset by the anti-rich price behavior during the remaining 3 subperiods. Thus, for the 1980s we observe that $B = 0.033$ per cent per year, a positive bias smaller than what we saw for the 1990s. However, from 1973-74 to the Winter of 1991, the plutocratic bias is always positive and reaches very high annual maxima from 1976 to 1979. For the period as a whole, $B = 0.239$ per cent per year, a bias whose size is equal to the sum of the classical substitution bias and the outlet bias according to the Boskin Commission.

To appreciate the variability of the plutocratic bias during the entire period considered in this paper, Figure 3 shows the evolution of the inter-annual (month to month) B_t, $t =$ January $1977, \ldots,$ January 1998, as well as the inter-annual inflation rate. Regressing the bias in absolute value against inflation yields a nonsignificant coefficient (0.005 with a standard error of 0.006, using generalized least squares correcting for autocorrelation).

[Insert Figure 3]

These results can be related to those obtained in previous papers about the distributional impact of price changes on household expenditures inequality measurement, where household expenditures are defined as total household expenditures net of the acquisition of some durables but inclusive of a number of expenditures which, although not included in the CPI definition, we believe are part of the best possible approximation to current consumption of private goods and services using the information available in the EPF. Using the inequality index $\mathcal{I}_1(\cdot)$ defined in equation (6), Ruiz-Castillo and

Sastre (1999) point out that the term ΔP in equation (4) is almost 5 times greater for the second part of the 1970s than for the 1980s. As a matter of fact, Del Río and Ruiz-Castillo (1996) reports that although the Lorenz curve of the 1980-81 adjusted household expenditures distribution at Winter 1981 prices dominates in a numerical sense the Lorenz curve of that same distribution at Winter 1991 prices, from a statistical point of view both distributions are indistinguishable, so that the distributional role of prices during the 1980s is essentially neutral. Nevertheless, for the 1973-74 to Winter 1981 period as a whole, Ruiz-Castillo and Sastre (1999) find that the decrease in real household expenditures inequality —which is equal to 27 per cent— is 37.7 per cent greater than the decrease in nominal household expenditures inequality; that is to say, during this period the distributional impact of price changes on inequality measurement is very important.

5.2. The Aggregation Scheme

In the second place, it is interesting to experiment with other aggregation schemes to map a distribution of individual inflations to an aggregate index. Given that the discipline of welfare economics is more interested in personal rather than household welfare, it is natural to ask for the consequences of estimating the inflation for the population as a whole as the weighted mean of individual inflations with weights proportional to household size.

Table 6. Average Total Household Expenditures at Winter 1991 Prices and Average Annual Inflation in the Partition by Household Size

Household size	Frequency Distribution	Average Expenditures	Average Annual Inflation
1 member	9.99%	1,147,338	4.842%
2 members	22.30%	1,795,808	4.625%
3 members	20.77%	2,559,993	4.634%
4 members	24.97%	3,091,959	4.611%
5 members	13.22%	3,277,244	4.623%
6 members	5.44%	3,516,374	4.627%
$\geq$ 7 members	3.31%	3,629,602	4.619%
ALL	100.00%	2,563,502	4.632%

On Table 6 we present mean total household expenditures at Winter 1991 prices by household size in the 1990–91 EPF, as well as the mean annual inflation from Winter

1991 to January 1998 for that same partition. As in the majority of other countries, we observe a positive association between total expenditures and household size. Therefore weighting household inflation by household size should have a similar effect, although of a lesser magnitude, than weighting directly by total household expenditures as in the plutocratic scheme. On the other hand, the fact that 2, 4 and more member households have a mean annual inflation below the population as a whole works in the opposite direction. The end result is that the new bias —defined as the difference between the plutocratic and the household size weighted mean— is equal to 0.088 per cent per year. That this figure is greater than the previously estimated of 0.055 per cent per year for the plutocratic bias, indicates that during this period the second factor has had a greater impact than the first one.

The same computations for the 1980s and 1970s lead to an estimate of 0.015 and 0.223 for the new bias *versus* a plutocratic bias of 0.033 and 0.239, respectively. The fact that the new bias is smaller than the plutocratic bias indicates that the positive association between total household expenditures and household size dominates the size of the new bias during these two periods.

6. CONCLUSIONS

In a country like Spain, a commodity basket of 471 goods is priced in each of the 52 provinces in order to construct the set of elementary price indexes which form the core of the current 1992 CPI system. We have been able to work in a 53 dimensional commodity space, consisting of the 21 food *rúbricas* at the 18 Autonomous Community level, and the 32 non-food *subgrupos* at the 52 province level. For such a commodity breakdown, we construct 21,155 household-specific Laspeyres price indexes representative of a 1990–91 population of about 11 million households. Because of the fixed-weight nature of our construction, the individual inflation variation we observe during the Winter 1991-January 1998 period is the consequence of the price variation publicly disseminated by the INE in this 53 commodity space.

How can we grasp the distributional consequences of such a complex multidimensional process? In this paper we propose a procedure which combines three elements. In the first place, whether price behavior in a given period hurts relatively more the rich or the poor households can be expressed in terms of a single scalar: the so-called plutocratic bias, incurred when we measure inflation using the current plutocratic CPI instead of using an alternative group index in which all households weight equally. In the second place, the estimation of an Engel curve system in a 16 goods commodity space, allows us to reduce the size of the price universe to only three dimensions: a luxury good and two necessities with considerably different total expenditures elasticities. Price behavior at this level provides an intelligible explanation of the sign and magnitude of the plutocratic bias. Finally, the change in money income inequality —which is the only magnitude usually estimated in the empirical literature on income inequality— is seen to be equal to the change in real income inequality —which is the socially relevant magnitude— and a price term which captures the distributional impact of price changes according to consumers' preferences as manifested via their budget share vectors in a given moment in time. The price term, and hence the gap between the change in money and real income inequality, depends on the size and magnitude of the plutocratic bias.

Knowing the plutocratic bias is also important for two other topics in positive economics. On one hand, in Spain —like in many other countries— the CPI is not really a modified Laspeyres price index. Thus, one can define a "Laspeyres bias" as the difference between measuring inflation using an appropriate Laspeyres type index or using the CPI actually constructed in these countries. As shown in Ruiz-Castillo *et al.* (1999d), the sign of the plutocratic bias determines the sign of the Laspeyres bias. On the other hand, recent theoretical results have opened up the way to the possibility of estimating the classical upper level substitution bias in the CPI as the difference between the inflation measured according to two readily computable statistical price indexes: a Laspeyres price index, and a Törnquist one for which we only need to observe consumer choices during the two periods under comparison. In Ruiz-Castillo *et al.* (1999e) we show how the knowledge of the sign of the plutocratic bias helps us to overcome some difficulties which arise when we attempt to apply this idea in practice.

But beyond all of the above measurement issues, what are the political consequences of our research? The first question we need to rise is how should we adjust our income taxes and public transfers annually. At this moment, we do not have anything to add to the arguments offered by others.[17] But we should recognize that, in most countries,

[17] See Triplett (1983), Fry and Pashardes (1985), Griliches (1995), and Pollak (1998) and, in connection

income taxes, public pensions, other public transfers, and minimum wages are revised in terms of a plutocratic CPI. Why should we follow a dollar rather than a household or a personal logic in this matter? Perhaps because both people and experts believe that the CPI represents an "average consumer." However, when in an important but unpublished paper Muellbauer (1976a) asked for the consumer whose budget shares are equal to the official CPI aggregate weights, he answered that in the UK this consumer occupied the 71 percentile in the household expenditures distribution.[18] At any rate, indexing by the current CPI has the following perverse effects which, in our opinion, have not been sufficiently emphasized before: when prices behave in a anti-poor (anti-rich) way, *i.e.*, when the plutocratic bias is negative (positive), then we revise public programs, which primarily benefit the poor, below (above) what would be the case with a democratic group index. Similarly, if the plutocratic bias is negative (positive), then direct tax revenues would be larger (smaller) than what would be the case under the democratic alternative.

From this point of view, the current plutocratic formula can be conceptually critized. Admittedly, this issue would be more important the greater the size of the plutocratic bias (and perhaps, depending on the sign of the bias). In the Spanish case, we have shown that this bias: *(i)* has had a positive sign over an extended period of time, *(ii)* presents a rather unstable pattern over the short run, and *(iii)* has had a considerable size during certain periods of time. There is relatively little information on this issue in other countries,[19] in particular in underdeveloped countries where the relative price of a

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to the poverty line, see the National Research Council (1995).

[18] See Muellbauer (1975, 1976b) for the theoretical basis of this work. For the U.S. in 1990, Deaton (1998) estimates that this consumer occupies the 75 percentile. In our case, we have simply computed the location of Spanish consumers who have an inflation in a 5 per cent interval of the official one during the 1990s; the answer is that their mean adjusted household expenditures is in the 61 percentile of such distribution.

[19] For the U.K., Carruthers *et al.* (1980) indicate that from January 1975 to January 1979 the democratic index has increased by around 0.1 per cent per year faster than the official CPI; Fry and Pashardes (1985) obtain also that from 1974 to 1982 the plutocratic bias was negative; for 1975-76, Deaton and Muellbauer (1980) report that the inflation rate for the poor was around two points higher than for the rich; however, Crawford (1996) finds that, between 1979 and the end of 1992, inflation for richer households was 0.16 percentage points higher than the average for all households. For the U.S., Kokowski (1987) finds that from 1972 to 1980 the democratic and the plutocratic Laspeyres indexes are rather close in value for most demographic groups but, in general, the first measure exceeds its counterpart by 1 to 3 index points; Garner *et al.* (1999) find evidence that the plutocratic bias during the 1980s is

few staples may cause havoc in the standard of living of the majority of the population. In our opinion, it is advisable to estimate the plutocratic bias on a regular basis. For this and other purposes, we recommend that statistical agencies in charge of the CPI compute and make available, at least annually, a set of household-specific price indexes. This idea has the following four advantages:

1. We expect that the farther down we go towards the elementary price level, the greater will be the dispersion of the distribution of household-specific price indexes. However, we would also expect a larger number of zero expenditures in most households. Therefore, there are advantages and disadvantages in enlarging the commodity space. Given that, for confidentiality reasons, the price information at the elementary level is not publicly available, the statistical agencies are the only institutions in a position to determine the optimal disaggregation level for the construction of individual price indexes.

2. Given the set of (official) individual price indexes, anyone —except the statistical agencies themselves, which should not become involved in political issues— can study the differential inflation suffered by the subgroups of interesting population partitions, an issue to be considered prior to the political solution to the issue of "How many cost of living indexes". Similarly, anyone would be in a position to estimate the bias in the measurement of inflation created by the use of the current plutocratic CPI, instead of other politically interesting definitions of what a group index should be.

3. Perhaps more importantly, statistical offices (and others) can evaluate the distributional consequences of their methodological decisions. Take, for example, the Boskin Commission's analysis of the quality issue and the introduction of new products, surely the most debated and critized part of their report. Different social critics —Madrick (1997) and Deaton (1998), for instance— conjecture that new goods and goods affected by quality effects are disproportionately consumed by the rich. In our own terms, this implies that the set of household-specific price indexes after the correction of this bias should exhibit a smaller plutocratic bias. Are these critics

slightly anti-rich.

correct? In Ruiz-Castillo *et al.* (1999f), we have put this idea to a test by combining the structure of the bias for the U.S. economy with the consumer behavior of Spanish households as given in the 1990–91, 1980-81 and 1973-74 EPFs. The plutocratic bias after the correction of the quality bias in the intervals (Winter 1991, January 1998), (Winter 1981, Winter 1991), (1973-1974, Winter 1981) is 0.035. 0.024, and 0.227 per cent per year, respectively. Since, as we have seen, the plutocratic bias before the correction is 0.055, 0.033 and 0.240 per cent per year, we can conclude that there is some evidence indicating that the point made by those social critics is well taken.

4. Muellbauer (1976a) indicates that he does not regard the historical bias of inflation as the most important issue. Given that keeping down inflation is such an important policy goal, it is natural that any government should be very sensitive to the effects of policy change on the official CPI. Thus, the aggregate weights are the forces which push government policy affecting relative prices into particular directions. Within this context, armed with a set of publicly available household-specific price indexes, both the government (and others) would be in a position to evaluate, both *ex ante* and *ex post*, the distributional consequences on the CPI of certain policy actions.

Some would argue that, given the public opinion's potential sensitivity to the distributional issues embedded in the construction of a single CPI, officially publishing a set of household-specific price indexes would ultimately affect the credibility of the CPI itself. But we have shown that anyone can come up with a reasonable version of these indexes using already publicly available information. In an open society we should not fear the dissemination of relevant, albeit controversial, information. Or is it the case that, precisely because "aggregate index numbers are **not** neutral political indicators" (Muellbauer 1976a), we should resist making public the statistical basis of this issue?

APPENDIX: The Modified Laspeyres Price Index

To understand the relation between a CPI and an aggregate Laspeyres SPI, we have to start by recognizing that statistical agencies partition the physical space into a set of J geographical areas, which we index by $j = 1, \ldots, J$. For every item $i = 1, \ldots, I$ in every area $j = 1, \ldots, J$, during each period t (typically a month), statistical agencies collect price quotes for a number of previously determined item specifications in a certain predetermined sample of outlets.[20] These price quotes are aggregated in elementary price indexes E_{ijt}.[21] Conceptually, we can view an elementary price index as the relative price of item i in area j in period t with respect to the base period 0, *i.e.*,

$$E_{ijt} = \frac{p_{ijt}}{p_{ij0}}.$$

On the other hand, household budget surveys provide information, not on individual prices and quantities which are often hard to define, but on individual expenditures in each good, $x^h_{i\tau}$, total household expenditures, $x^h_\tau = \sum_i x^h_{i\tau}$, and budget shares $w^h_{i\tau} = x^h_{i\tau}/x^h_\tau$. In each area j, we can observe the aggregate expenditures on each good, $X_{ij\tau} = \sum_{h\in j} x^h_{i\tau}$, and aggregate budget shares $W_{ij\tau} = X_{ij\tau}/X_\tau$, where $X_\tau = \sum_h x^h_\tau$ is the aggregate total expenditure for the entire population. Under the assumption that all households living in the same area face the same prices, we can view observable household expenditures on item i by a household h living in area j and interviewed in period τ, as the product of a price $p_{ij\tau}$ and a quantity $q^h_{i\tau}$, *i.e.*, $x^h_{i\tau} = p_{ij\tau} q^h_{i\tau}$. Denote the vector of aggregate quantities actually purchased during the survey period τ by $\mathbf{Q}_\tau = (Q_{1\tau}, \ldots, Q_{I\tau})$ where $Q_{i\tau} = \sum_j Q_{ij\tau}$ and $Q_{ij\tau} = \sum_{h\in j} q^h_{i\tau}$, then we have

$$W_{ij\tau} = \frac{X_{ij\tau}}{X_\tau} = \frac{p_{ij\tau} Q_{ij\tau}}{\mathbf{p}_\tau \cdot \mathbf{Q}_\tau}.$$

If we define the plutocratic weights $\phi^h_\tau = x^h_\tau / X_\tau$, then

$$W_{ij\tau} = \sum_{h\in j} \frac{x^h_\tau}{X_\tau} \frac{x^h_{i\tau}}{x^h_\tau} = \sum_{h\in j} \phi^h_\tau w^h_{i\tau}.$$

[20] This is where Pollak places the "beer vs champagne" issue.

[21] This is where the Boskin Commission places the so called "lower substitution level" problem. Neither this nor the issue in the previous footnote should concern us in this paper.

If we have information on what we will call the adjustment factors for each i, $A_{ij\tau} = (p_{ij\tau}/p_{ij0})$, then one can define the elementary price index based in period τ,

$$E_{ijt}(\tau) = \frac{E_{ijt}}{A_{ij\tau}} = \frac{p_{ijt}}{p_{ij\tau}}.$$

For each household h living in area j, the Laspeyres SPI which takes as a reference the quantity vector $\mathbf{q}_\tau^h$, is defined by

$$\ell(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{q}_\tau^h) = \sum_i w_{i\tau}^h E_{ijt}(\tau) = \frac{\mathbf{p}_{jt} \cdot \mathbf{q}_\tau^h}{\mathbf{p}_{j\tau} \cdot \mathbf{q}_\tau^h},$$

where $\mathbf{p}_{jt} = (p_{1jt}, \ldots, p_{Ijt})$.

At the aggregate level, let $\mathbf{p}_t = (p_{1t}, \ldots, p_{It})$, where $p_{it} = \sum_j (Q_{ij\tau}/Q_{i\tau}) p_{ijt}$. Similarly, let $\mathbf{p}_\tau = (p_{1\tau}, \ldots, p_{I\tau})$, where $p_{i\tau} = \sum_j (Q_{ij\tau}/Q_{i\tau}) p_{ij\tau}$. Then the aggregate Laspeyres SPI which takes as a reference the vector $\mathbf{Q}_\tau$ is seen to be:

$$\begin{aligned}
\mathcal{L}(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{Q}_\tau) &= \sum_i \sum_j W_{ij\tau} E_{ijt}(\tau) = \frac{\sum_i \sum_j p_{ijt} Q_{ij\tau}}{\sum_i \sum_j p_{ij\tau} Q_{ij\tau}} = \frac{\mathbf{p}_t \cdot \mathbf{Q}_\tau}{\mathbf{p}_\tau \cdot \mathbf{Q}_\tau} \\
&= \sum_i \left(\sum_j \sum_{h \in j} \phi_\tau^h w_{i\tau}^h \right) E_{ijt}(\tau) = \sum_j \sum_{h \in j} \phi_\tau^h \sum_i w_{i\tau}^h E_{ijt}(\tau) \\
&= \sum_h \phi_\tau^h \frac{\mathbf{p}_{jt} \cdot \mathbf{q}_\tau^h}{\mathbf{p}_{j\tau} \cdot \mathbf{q}_\tau^h} = \sum_h \phi_\tau^h \ell(\mathbf{p}_{jt}, \mathbf{p}_{j\tau}; \mathbf{q}_\tau^h).
\end{aligned}$$

For each good i in an area j, let $W_{ij} = p_{ij0} Q_{ij\tau} / \mathbf{p}_0 \cdot \mathbf{Q}_\tau$. The CPI based on period 0 is an aggregate SPI defined by

$$CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) = \sum_i \sum_j W_{ij} E_{ijt} = \frac{\mathcal{L}(\mathbf{p}_t, \mathbf{p}_\tau; \mathbf{Q}_\tau)}{\mathcal{L}(\mathbf{p}_0, \mathbf{p}_\tau; \mathbf{Q}_\tau)} = \frac{\mathbf{p}_t \cdot \mathbf{Q}_\tau}{\mathbf{p}_0 \cdot \mathbf{Q}_\tau},$$

which is what the BLS calls a *modified Laspeyres* aggregate price index (Moulton, 1996), with base year 0 and reference consumption patterns surveyed at τ.

Finally, for household h in area j we now redefine the plutocratic weights by $\phi^h = \mathbf{p}_{j0} \cdot \mathbf{q}_\tau^h / \mathbf{p}_0 \cdot \mathbf{Q}_\tau$, and budget shares $w_i^h = p_{ij0} q_{i\tau}^h / \mathbf{p}_{j0} \cdot \mathbf{q}_\tau^h$. Then, as before, aggregate expenditure shares can be expressed as a plutocratic-weighted mean of individual expenditure shares:

$$\sum_{h \in j} \phi^h w_i^h = \sum_{h \in j} \frac{\mathbf{p}_{j0} \cdot \mathbf{q}_\tau^h}{\mathbf{p}_0 \cdot \mathbf{Q}_\tau} \frac{p_{ij0} q_{i\tau}^h}{\mathbf{p}_{j0} \cdot \mathbf{q}_\tau^h} = \frac{p_{ij0} Q_{ij\tau}}{\mathbf{p}_0 \cdot \mathbf{Q}_\tau} = W_{ij}$$

and

$$\begin{aligned}
CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) &= \sum_i \sum_j W_{ij} E_{ijt} = \sum_i \sum_j \sum_{h \in j} \phi^h w_i^h E_{ijt} \\
&= \sum_j \sum_{h \in j} \phi^h \sum_i w_i^h E_{ijt} = \sum_h \phi^h cpi(\mathbf{p}_{jt}, \mathbf{p}_{j0}; \mathbf{q}_\tau^h).
\end{aligned}$$

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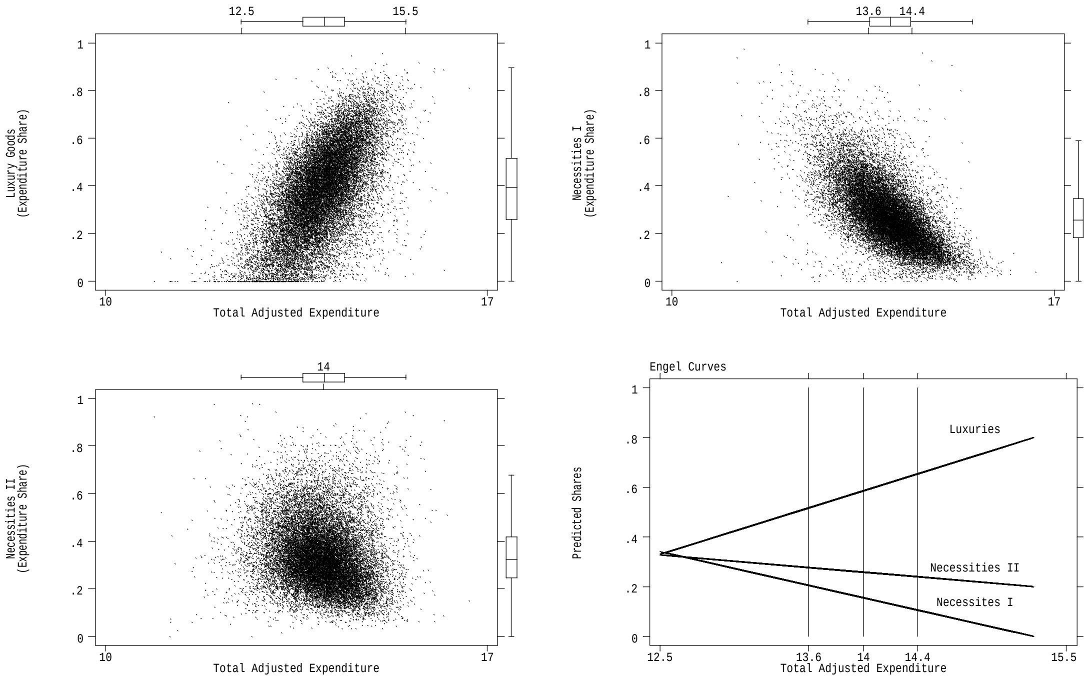
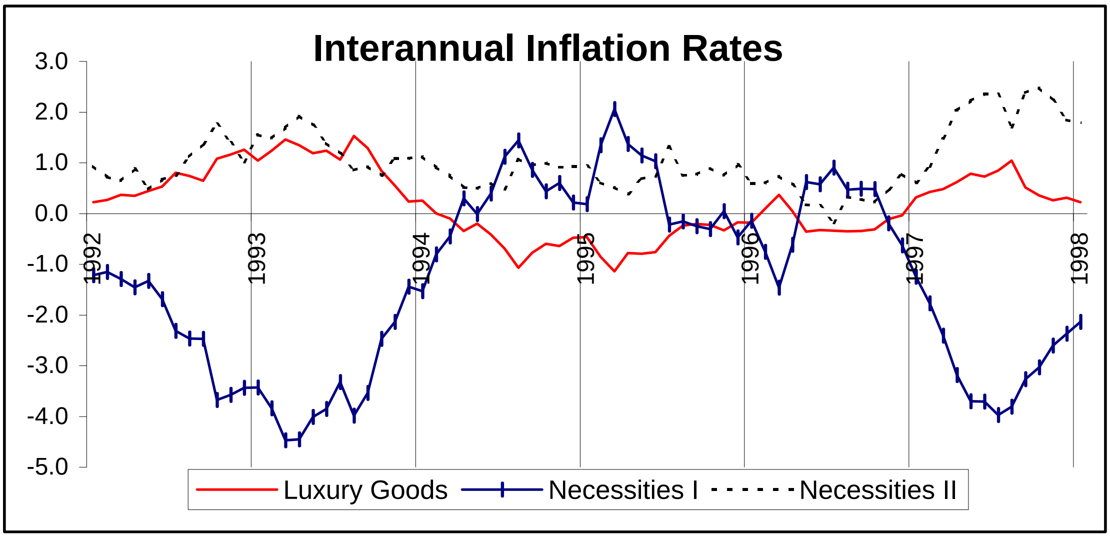


Fig1. Individual budget shares and adjusted household expenditures



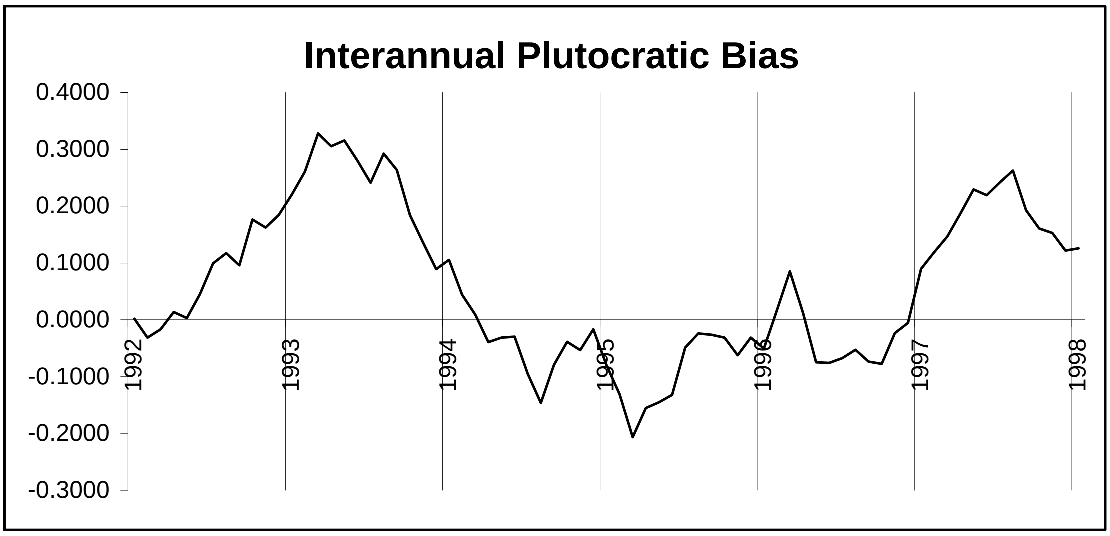


Fig. 2. Inflation rates of different goods: January 92- January 98

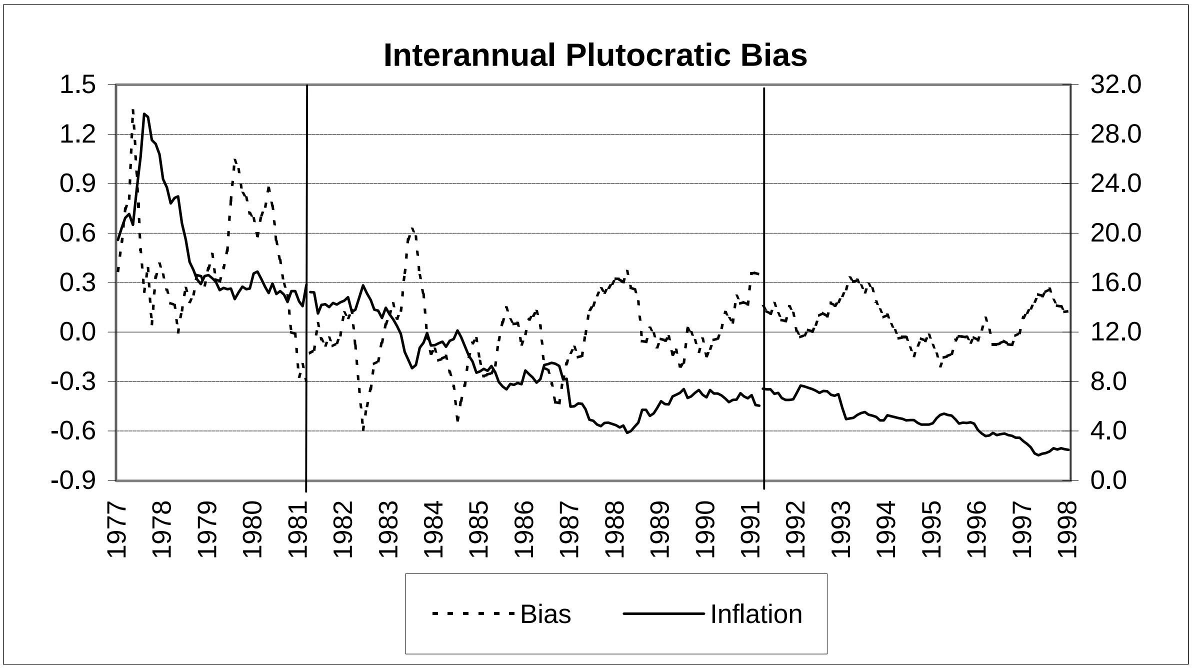


Fig.3. Plutocratic Bias and Interannual Inflation 1976-98

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