WP/00/167

IMF Working Paper

ARCHIVES
ROOM IS11-400 0441

The Plutocratic Bias in the CPI: Evidence from Spain

Javier Ruiz-Castillo, Eduardo Ley, and Mario Izquierdo

INTERNATIONAL MONETARY FUND

© 2000 International Monetary Fund WP/00/167

IMF Working Paper

Fiscal Affairs Department

The Plutocratic Bias in the CPI: Evidence from Spain

Prepared by Javier Ruiz-Castillo, Eduardo Ley, and Mario Izquierdo*

Authorized for distribution by Liam P Ebrill

October 2000

Abstract

The views expressed in this Working Paper are those of the author(s) and do not necessarily represent those of the IMF or IMF policy. Working Papers describe research in progress by the author(s) and are published to elicit comments and to further debate.

We define the plutocratic bias as the difference between inflation measured according to the current official CPI and a democratic index in which all households receive the same weight. We estimate that during the 1990s the plutocratic bias in Spain amounts to 0.055 percent per year. However, positive and negative biases cancel off when averaging over the whole period. The mean absolute bias is significantly larger, 0.090. We can explain most of the oscillations experimented by the plutocratic bias by the price behavior of three goods: a luxury good and two necessities.

JEL Classification Numbers: C43, D31, D63

Keywords: Consumer price index, cost-of-living index, aggregation

Authors' E-Mail Addresses: jrc@eco.uc3m.es, eley@imf.org, izquierdo@fedea.es

* Javier Ruiz-Castillo is Professor of Economics, Universidad Carlos III de Madrid, Spain; Eduardo Ley is an Economist, Tax Policy Division, Fiscal Affairs Department, IMF; and Mario Izquierdo is an Economist at Fundación de Estudios de Economía Aplicada (FEDEA), Madrid, Spain. We thank Liam Ebrill and Marshall Reinsdorf for useful comments. This research has been partially funded by "la Caixa".

Contents Page

I. INTRODUCTION

In all countries, the official consumer price index (CPI), 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 United States Senate Commission headed by Michael Boskin (Boskin and others, 1996), this is not all that is wrong with the United States CPI elaborated by the Bureau of Labor Statistics (BLS).

The Boskin Commission focused on five sources of bias in the CPI, all of which are presumed to contribute to an overstatement of the true inflation in the cost of living in the United States. The Boskin Commission estimated that, on average, during the last few decades the United States CPI has been overstating inflation by 1.1 percent per year. This bias 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 dramatic.[1] The report has become 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 and others, 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 nonelderly, 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.

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 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

[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 and others (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."

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.[3]

In this paper, we argue 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. One reason in favor of such a concept is that 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).

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 percent per year—or about one-third of the classical substitution bias estimated by the Boskin Commission for the United States. 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 percent per year. The mean absolute bias is much larger, 0.090. 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 three dimensions, consisting of a luxury good and two necessities. The price behavior of these three goods provides a convincing explanation of the oscillations experimented by the plutocratic bias.

The paper studies the robustness of these results in two dimensions. First, 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 period and large in the second: 0.091 and 0.264 percent per year, respectively. Second, 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, 0.064, and 0.254 percent per year, respectively.

[3] As pointed out by Pollak (1998), the first two issues are given a cursory treatment in footnote 2 and page 71 of Boskin and others (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.

II. INDIVIDUAL AND GROUP INDEXES

A. 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 $\mathbf{p}_0$,[4]

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

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.

---

[4] 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 $\ell(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_0^h) = \sum_i w_{i0}^h (p_{it}/p_{i0})$.

B. 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:

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

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 Appendix II, the CPI based on period 0 is an aggregate SPI defined by[5]

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

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:

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

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 Appendix II,

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

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

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,

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

then the aggregate CPI provides an upper bound to a plutocratic-weighted mean of modified Laspeyres COLIs:[6] $CPI(\mathbf{p}_t, \mathbf{p}_0; \mathbf{Q}_\tau) \geq \sum_h \phi^h \kappa(\mathbf{p}_t, \mathbf{p}_0; \mathbf{q}_\tau^h)$. 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.

III. THE PLUTOCRATIC BIAS

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. 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. (See Appendix I for a description of these data sources and some issues regarding the definition of household expenditures.)

A. A Definition of the Plutocratic Bias

We will divide the period Winter 1991–January 1998 in the seven 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 $\boldsymbol{\Pi} = (\Pi^1, \ldots, \Pi^H)$, where $\Pi^h = (\ell_T^h - 1)$, and T = Jan 98. The aggregate inflation for the population as a whole according to the plutocratic scheme is

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

[6] 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.

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 PLUT $= \sum_h \phi^h (\ell_T^h - 1)$, and 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 $B = \text{PLUT} - \text{DEM}$.[7] 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).

B. The Main Findings

The first two columns of Table 1 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 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. Column 3 shows 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.

The main findings are the following: (1) For the period as a whole, B is positive and equal to 0.055 percent per year. This is, approximately, one-third of the substitution bias estimated by the Boskin Commission for the United States economy, which is equal to 0.15 percent per year. However, positive and negative biases cancel off when averaging over the whole period. The mean absolute bias is much larger, 0.090.[8] (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.

[7] 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.

[8] These figures might be also compared with the overall upward bias for the Spanish economy which, following the Boskin Commission procedures, we have estimated as 0.61 percent per year. For a discussion of the economic consequences of such a bias, see Ruiz-Castillo and others (1999b).

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
	Jan 93 to Jan 98, average absolute bias			0.090

Source: Authors' calculations.

C. 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 and others (1988) and Coulter and others (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 household 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.

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 less than 1 (necessities I, dominated by Food expenditures), or slightly less 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

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 1/	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 2/	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

Source: Authors' calculations.

1/ Except "Other Food Products" (beef, processed fish, fruit preserves and other unclassified foods).

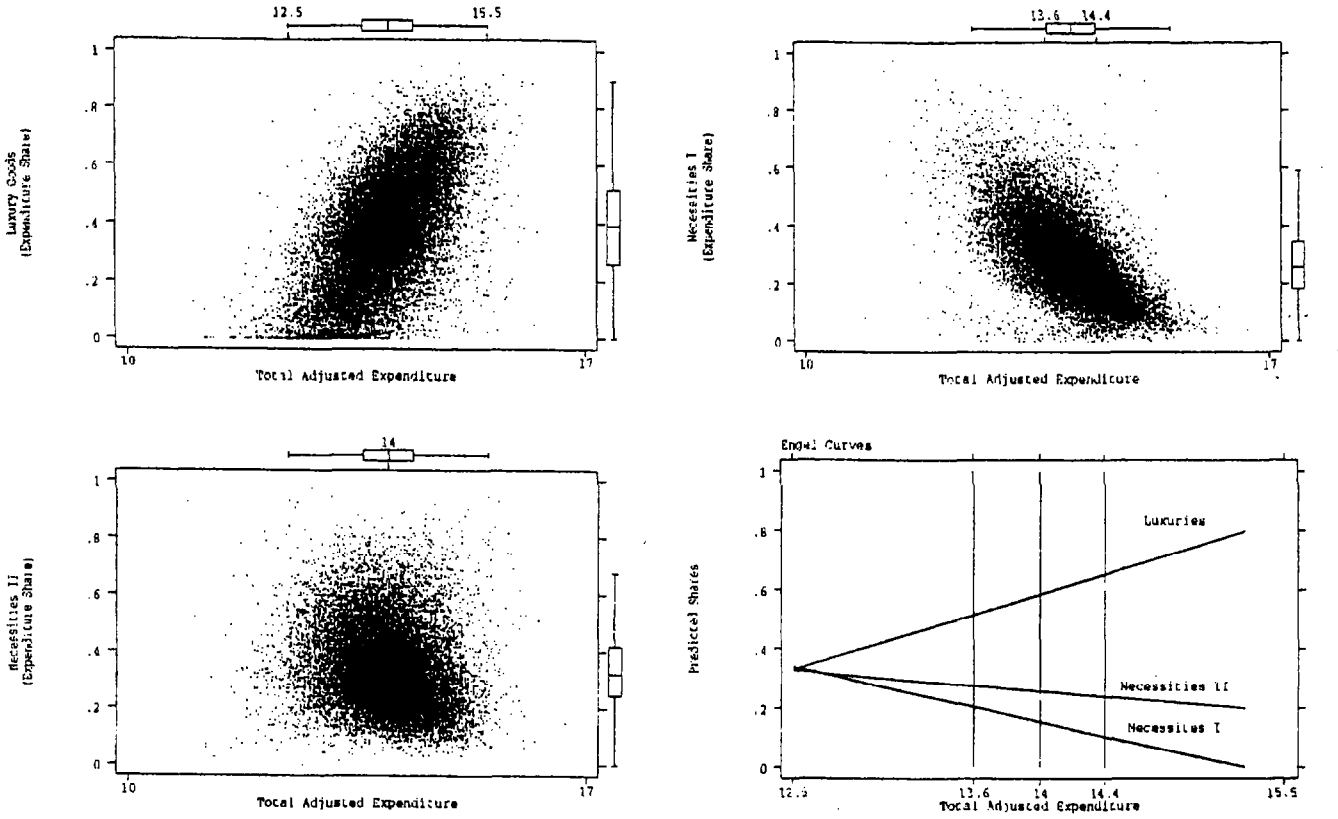
2/ "Non-alcoholic Drinks" and "Other Food Products."

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,[9] the last panel shows the estimated Engel curves (trimming the 1 percent tails off the support of the adjusted household expenditures).

Intuitively, the evolution of prices would tend to damage to a relatively greater extent the richer households over the poorer ones depending on 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

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

Figure 1. Individual Budget Shares and Adjusted Household Expenditures



Source: Author's calculations.

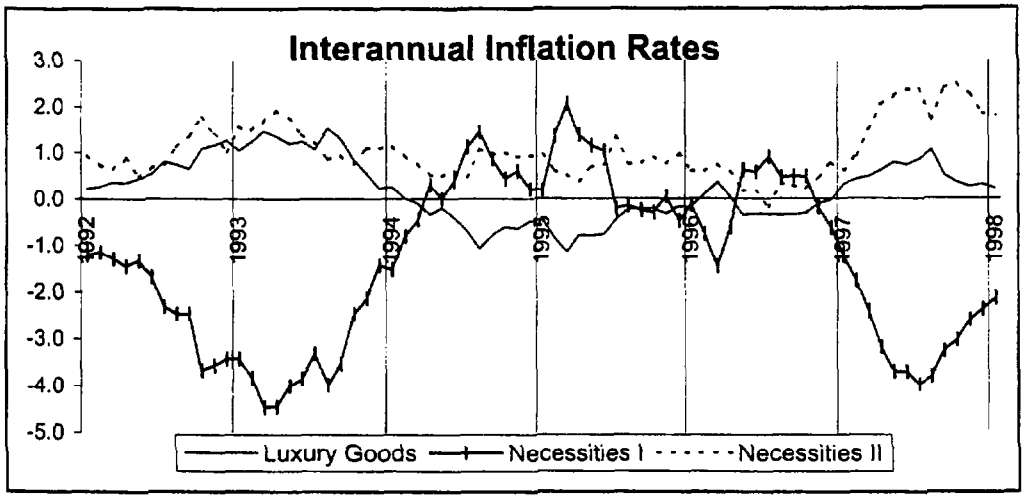
of the three goods in relation to the general inflation as well as the inter-annual B_t, $t =$ January 1992,...,January 1998.

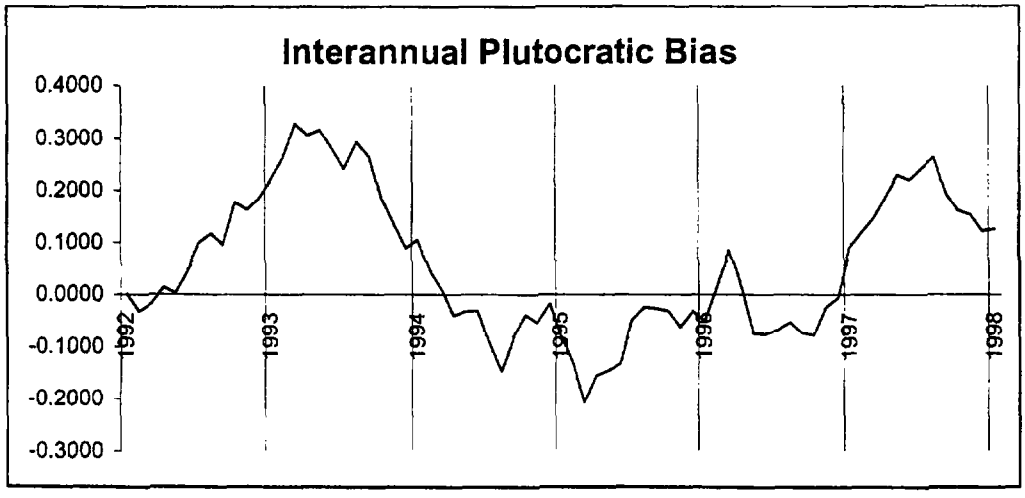
In spite of the fact that the second necessity shows the stronger price growth, the behavior of the luxury good and the first necessity is the main explanatory force behind the positive sign of the plutocratic bias. To test this, we run a regression of the inter-annual (January-to-January) plutocratic bias B_t, from $t =$ January 1992,..., January 1998, on the corresponding monthly inter-annual price subindexes for the three 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:

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

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).

Figure 2. Inflation Rates of Different Goods: January 1992–January 1998





Source: Author's calculations.

IV. ROBUSTNESS

A. 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. (See Appendix I for details.)

The main findings are the following:[10] (1) From Winter 1981 to Winter 1991 we estimate that $B = 0.091$ percent per year, a positive bias larger than witnessed for the 1990s. During different subperiods the bias is negative, and oscillates from a maximum of 0.380 to a minimum, in absolute value, of -0.025 percentage point. (2) From 1973–74 to Winter 1981 the plutocratic bias is always positive and reaches high annual maxima from 1976 to 1979, in 1979 it equals 0.833 percentage point. For the period as a whole, $B = 0.264$ percent per year, a bias equal in size to the sum of the classical substitution bias and the outlet bias reported by the Boskin Commission. (3) We must point out that the Spanish inflation during the second part of the 1970s and 1980s is considerably greater than during the 1990's: the mean annual inflation from the midpoint of 1973 and 1974 to Winter 1981 is 17.9 percent, and from Winter 1981 to Winter 1991 is 8.5 percent. However, as before, there is no relationship between the size of the aggregate inflation in a given subperiod and the sign or the magnitude of the plutocratic bias. Regressing the bias in absolute values against inflation yields a nonsignificant coefficient (0.002 with a standard error of 0.006, using generalized least squares correcting for autocorrelation).

Finally, 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, ..., January 1998, as well as the inter-annual inflation rate.

B. The Aggregation Scheme

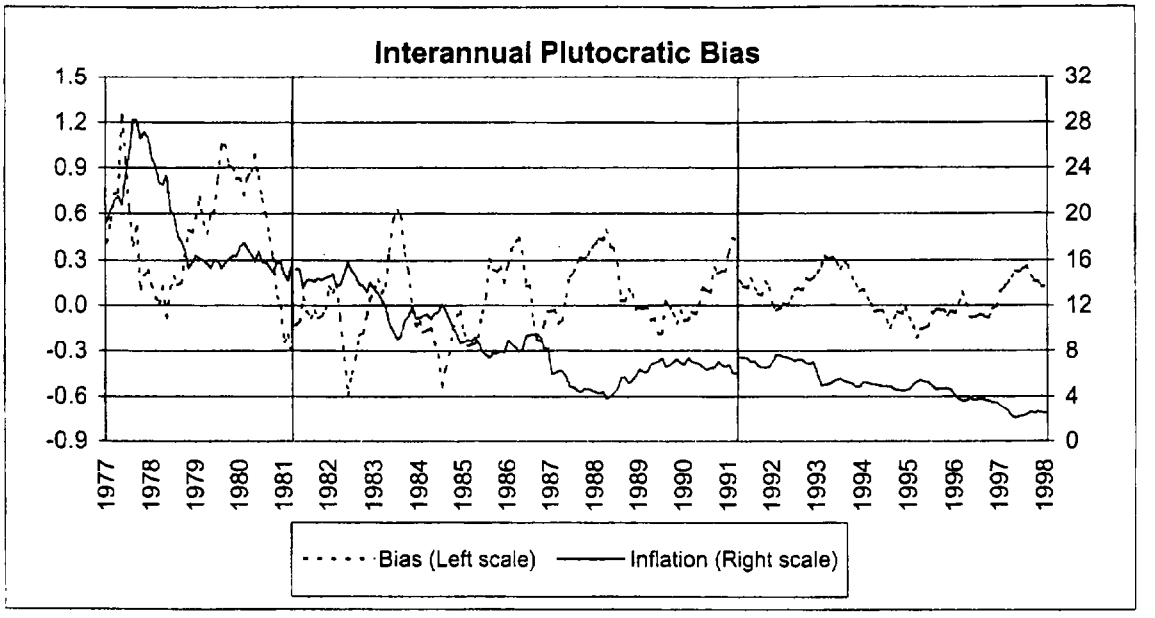
In the second place, it is interesting to experiment with other aggregation schemes to map a distribution of individual inflations to an aggregate index. Given that the discipline of welfare economics is more interested in personal rather than household welfare, it is natural to ask for the consequences of estimating the inflation for the population as a whole as the weighted mean of individual inflations with weights proportional to household size.[11]

In Table 3 we present mean total household expenditures at Winter 1991 prices by house-

[10] For more detailed results, see Ruiz-Castillo and others (1999c).

[11] Alternatively, one could define an aggregate price index which gives greater weight to poorer households. Given the results of this paper, the corresponding bias in the measurement of inflation should be greater than the plutocratic one.

Figure 3. Plutocratic Bias and Interannual Inflation: 1976–98



Source: Author's calculations.

Table 3. Average Household Expenditures at Winter 1991 Prices and Average Annual Inflation in the Partition by Household Size

Household size size	Frequency Distribution (percentages)	Average Expenditures (pesetas)	Average Annual Inflation (percentages)
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

Source: Authors' calculations.

hold size in the 1990–91 EPF, as well as the mean annual inflation from the Winter 1991 to January of 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 two, four, 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 percent per year. That this figure is greater than the previously estimated 0.055 percent 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 0.064 and 0.254 for the new bias *versus* a plutocratic bias of 0.091 and 0.264. 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.

V. IMPLICATIONS

In 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 the distributional consequences of such a complex multidimensional process be grasped? In this paper we propose a procedure which combines two elements. First, 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 inflation is measured using the current plutocratic CPI instead of using an alternative group index in which all households receive equal weight. Second, the estimation of an Engel curve system in a 16 goods commodity space, would permit reduction in 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.[12]

[12] As we show in Ruiz-Castillo and others (2000), the gap between the change in money income inequality and the socially relevant change in real income inequality is given by a term which captures the distributional role of price changes. The sign of this term is largely determined by the sign of the plutocratic bias. In consonance with the results of the present paper, using the mean logarithmic deviation and a value of 0.5 for the parameter which reflects the economies of scale within the household, in Ruiz-Castillo and others (2000) we find that this term is positive. Thus, we conclude that the decrease in real household expenditures inequality in Spain during the second part of the 1970s, the 1980s, and the 1990s has been 9.07, 4.82, and 2.97 percent larger than the decrease in money household expenditures inequality due to the distributive role of price changes during these periods.

Beyond all the measurement issues, what are the practical policy consequences of our research? The first question to consider is how to adjust income taxes and public transfers annually. At this point, we have little to add to the arguments offered by others[13] but note that, in most countries, 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 person in this matter? Perhaps because both people and experts believe that the CPI represents an "average consumer." However, when in an important paper Muellbauer (1976a) asked for the consumer whose budget shares are equal to the official CPI aggregate weights, he answered that in the United Kingdom this consumer occupied the 71 percentile in the household expenditures distribution.[14] At any rate, indexing by the current CPI has the following unintended effects which may not have 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 perspective, the current plutocratic formula is open to some critique. Admittedly, these critiques 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 magnitude during certain periods of time. There is relatively little information on this issue in other countries,[15] in particular in underdeveloped countries where the relative price of a few staples may loom large in the standard of living of the majority of

---

[13] See Triplett (1983), Fry and Pashardes (1985), Griliches (1995), and Pollak (1998) and, in connection to the poverty line, see the National Research Council (1995).

[14] See Muellbauer (1975, 1976b) for the theoretical basis of this work. For the United States 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 percent 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.

[15] For the United Kingdom, Carruthers and others (1980) indicate that from January 1975 to January 1979 the democratic index has increased by around 0.1 percent 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 point higher than the average for all households; Newberry (1994) found that the distributional effects were negligible and not significantly different from zero in Hungary and the United Kingdom during the 1980s. For the United States, 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; Slesnick (1991) finds that cost of living indexes are surprisingly insensitive to the choice of the form of the index; Garner and others (1999) find evidence that the plutocratic bias during the 1980s is slightly anti-rich.

the population. It is likely 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 approach would have the following advantages:

1. The farther down one goes toward the elementary price level, the greater will be the dispersion of the distribution of household-specific price indexes. However, one 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 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 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 correct? In Ruiz-Castillo and others (1999d), we have put this idea to a test by combining the structure of the bias for the United States 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), (1973–74, Winter 1981) is 0.035. 0.073, and 0.249 percent per year, respectively. Since, as we have seen, the plutocratic bias before the correction is 0.055, 0.091, and 0.264 percent per year, we can conclude that there is some evidence indicating that the point made by those critics is well taken.

4. Muellbauer (1976a) 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.

Finally, it could be argued 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. As noted by Muellbauer (1976a): "aggregate index numbers are **not** neutral political indicators." However, as shown here, anyone can come up with a reasonable version of these indexes using already publicly available information. After all, in an open society, the dissemination of relevant—albeit controversial—information should always be encouraged for the sake of transparency.

THE DATA

The Data for the 1990s

The *Encuesta de Presupuestos Familiares* (EPF) collected by the Spanish statistical agency, *Instituto Nacional de Estadística* (INE), 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 Appendix II) 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 and others (1999a). 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 and others (1999a,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 nonrental 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 and others (1999b), the official CPI understates the true Spanish inflation from 1992 to January of 1998 in 0.241 percent 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 Appendix II. 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.

If we normalize this series at prices of period 0 = 1992, we can obtain the conceptually correct CPI, that is,

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

For the details of this construction, see Ruiz-Castillo and others (1999a). The series of modified Laspeyres price indexes are available at www.eco.uc3m.es/investigacion/epf.html.

The Data for the 1970s and the 1980s

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, respectively. In this case, we do not depart from the official definition of household total 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 106 and 88 *subclases* at the national level in the 1983 and 1976 bases, respectively; and *(iii)* a series of adjustment factors for 52 goods which constitute the minimum common denominator between the 58 official *rúbricas* and the 60 goods in Catasús and others (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. For the details of these constructions, see Ruiz-Castillo and others (1999a). Both series of modified Laspeyres price indexes are available at http://www.eco.uc3m.es/investigacion/epf.html.

THE MODIFIED LASPEYRES 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. (This is where Pollak places the "beer vs champagne" issue.) These price quotes are aggregated in elementary price indexes E_{ijt}. (This is where the Boskin Commission places the so called "lower substitution level" problem. Neither this nor "beer vs champagne" issue should concern us in this paper.) 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, $x_{i\tau}^h$, total household expenditures, $x_\tau^h = \sum_i x_{i\tau}^h$, and budget shares $w_{i\tau}^h = x_{i\tau}^h / x_\tau^h$. In each area j, we can observe the aggregate expenditures on each good, $X_{ij\tau} = \sum_{h \in j} x_{i\tau}^h$, and aggregate budget shares $W_{ij\tau} = X_{ij\tau}/X_\tau$, where $X_\tau = \sum_h x_\tau^h$ 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 $q_{i\tau}^h$, *i.e.*, $x_{i\tau}^h = p_{ij\tau} q_{i\tau}^h$. 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 $Q_{ij\tau} = \sum_{h \in j} q_{i\tau}^h$, 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 $\phi_\tau^h = x_\tau^h / X_\tau$, then

$$W_{ij\tau} = \sum_{h \in j} \frac{x_\tau^h}{X_\tau} \frac{x_{i\tau}^h}{x_\tau^h} = \sum_{h \in j} \phi_\tau^h w_{i\tau}^h.$$

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 $p_{it} = \sum_j (Q_{ij\tau}/Q_{i\tau}) p_{ijt}$. Simlarly, 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:

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

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

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

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 $\phi^h = \mathbf{p}_{j0} \cdot \mathbf{q}^h_\tau / \mathbf{p}_0 \cdot \mathbf{Q}_\tau$, and budget shares $w^h_i = p_{ij0} q^h_{i\tau} / \mathbf{p}_{j0} \cdot \mathbf{q}^h_\tau$. Then, as before, aggregate expenditure shares can be expressed as a plutocratic-weighted mean of individual expenditure shares:

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

and

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

REFERENCES

Abraham, K., J. Greenlees and B. Moulton, 1998, "Working to Improve the Consumer Price Index," *Journal of Economic Perspectives*, 12, pp. 27–36.

Boskin, M., E. Dulberger, R. Gordon, Z. Griliches, and D. Jorgenson, 1996, *Toward a More Accurate Measure of the Cost of Living*, Final Report (Washington D.C.: Senate Finance Committee).

Buhmann, B., L. Rainwater, G. Schmaus and T. Smeeding, 1988, "Equivalence Scales, Well-Being, Inequality and Poverty: Sensitivity Estimates Across Ten Countries Using the Luxembourg Income Study Database," *Review of Income and Wealth*, 34, pp. 115–42.

Carruthers, A., D. Sellwood and P. Ward, 1980, "Recent Developments in the Retail Price Index," *The Statistician*, 29, pp. 1–32.

Catasús, V., J.L. Malo de Molina, M. Martínez and E. Ortega, 1986, "Cambio de base del Indice de Precios de Consumo," (unpublished, Madrid: Banco de España).

Coulter, F., F. Cowell and S. Jenkins, 1992a, "Differences in Needs and Assessment of Income Distributions," *Bulletin of Economic Research*, 44, pp. 77–124.

_____, 1992b, "Equivalence Scale Relativities and the Extent of Inequality and Poverty," *Economic Journal*, 102, pp. 1067-1082.

Crawford, V., 1996, "United Kingdom Household Cost-of-Living Indices, 1979-92," in *New Inequalities: the Changing Distribution of Income and Wealth in the United Kingdom*, ed. by J. Hills (Cambridge: Cambridge University Press).

Deaton, A., 1998, "Getting Prices Right:What Should Be Done?," *Journal of Economic Perspectives*, 12, pp. 37-46.

Deaton, A. and J. Muellbauer, 1990, *Economics and Consumer Behavior*, (New York: Cambridge University Press).

Diewert, E., 1995, "Axiomatic and Economic Approaches to Elementary Price Indexes," NBER Working Paper No. 5104 (Cambridge, Massachusetts: National Bureau of Economic Research).

_____, "Index Number Issues in the Consumer Price Index," *Journal of Economic Perspectives*, 12, pp. 47-58.

Fry, V. And P. Pashardes, 1985, *The RPI and the Cost of Living*, Report Series No. 22 (London: Institute for Fiscal Studies).

García España, E. and J.M. Serrano, 1980, *Indices de Precios de Consumo, Instituto Nacional de Estadística* (Madrid: Ministerio de Economía y Comercio).

Garner, T, J. Ruiz-Castillo and M. Sastre, 1999, "The Influence of Demographics and Household Specific Price Indices on Expenditure Based inequality and Welfare: A Comparison of Spain and the United States," Working Paper 9963 Economic Series, (Madrid: Universidad Carlos III de Madrid).

Griliches, Z., 1996, "Prepared Statement," in *Consumer Price Index: Hearings Before the Committee on Finance*, United States Senate, Senate Hearing 104-69 (Washington D.C.: United States Printing Office).

Kokowski, J., 1987, "Consumer Price Indices by Demographic Group," BLS Working Paper 167 (Washington D.C.: Bureau of Labor Statistics).

Konüs, A.A., 1924, "The Problem of the True Index of the Cost of Living," English version in *Econometrica*, 7, pp. 10-29.

Lorenzo, F., 1998, "Modelización de la inflación con fines de predicción y diagnóstico" (Ph.D. Dissertation, Madrid: Universidad Carlos III de Madrid).

Madrick, J., 1997, "The Cost of Living: A New Mith" and "Cost of Living: An Exchange," *New York Review of Books*, March 6 and June 26, respectively.

Moulton, B., 1996, "Bias in the Consumer Price Index: What is the Evidence?," *Journal of Economic Perspectives*, 12, pp. 133–50.

Muellbauer, J., 1975, "Aggregation, Income Distribution and Consumer Demand," *Review of Economic Studies*, 62, pp. 525–43.

_____, J., 1976a, "The Political Economy of Price Indices," Birbeck Discussion Paper 22 (London: University College).

_____, 1976b, "Community Preferences and the Representative Consumer," *Econometrica*, 44, pp. 979–99.

National Research Council, 1995, *Measuring Poverty: A New Approach*, (Washington D.C.: National Academic Press).

Newberry, David M., 1994, "The Distributional Impact of Price Changes in Hungary and the United Kingdom," University of Cambridge Discussion Papers on Economic Transition: DPET 9402 (Cambridge: University of Cambridge).

Pollak, R., 1995, "Prepared Statement and Testimony," in *Consumer Price Index: Hearings Before the Committee on Finance*, United States Senate, Senate Hearing 104-69 (Washington D.C.: United States Printing Office).

_____, 1998, "The Consumer Price Index: A Research Agenda and Three Proposals," *Journal of Economic Perspectives*, 12, pp. 69–78.

Prais, S., 1958, "Whose Cost of Living?" *Review of Economic Studies*, 26, pp. 126–34.

Ruiz-Castillo, J., C. Higuera, M. Izquierdo and M. Sastre 1999a, "Series de precios individuales para las EPF de 1973–74, 1980–81 y 1990–91 con base en 1976, 1983 y 1992," (unpublished, Madrid: Universidad Carlos III de Madrid).

Ruiz-Castillo, J., E. Ley and M. Izquierdo 1999b, "The Boskin Commission and Other Sources of Bias in the Spanish Consumer Price Index," (unpublished, Madrid: Universidad Carlos III de Madrid).

_____, 1999c, "The Plutocratic Bias in the CPI: Evidence from Spain," FEDEA Working Paper 99-15 (Madrid: FEDEA).

_____, 1999d, "Distributional Aspects of the Quality Change Bias in the CPI. Evidence from Spain", FEDEA Working Paper 2000-08 (Madrid: FEDEA).

_____, 2000, "The Distributive Effects of Prices Changes in the Measurement of Income Inequality. The Spanish Case, 1973-74 to 2000"(Madrid: FEDEA).

Triplett, J. , 1983, "Escalation Measures: What Is the Answer, What Is the Question?" in W. E. Diewert and C. Montmarquette (eds.), *Price Level Measurement* (Ottawa: Statistics Canada).