Working Paper 08
Economics Series 02
April 1993

Departamento de Economía
Universidad Carlos III de Madrid
Calle Madrid 126
28903 Getafe (Spain)
Fax (341) 624-9875

THE DISTRIBUTION OF EXPENDITURE IN SPAIN, 1973-74 TO 1980-81

Javier Ruiz-Castillo*

Abstract

This paper examines how far we can go in establishing normative conclusions with regard to the evolution of inequality over time, making use of statistical methods which do not need either too specific assumptions about individual preferences, or their recovery by means of complex and expensive econometric methods. The decomposition of the change in money inequality into a real and a price effect occupies the center of the analysis. We have used statistical Laspeyres type price indices which are household specific, and a parametrization which captures the weight one is prepared to give to household size in the definition of equivalent expenditure per person. The central finding is that the improvement in real inequality in Spain from 1973-74 to 1980-81 is always greater than the improvement in money inequality. Changes in relative prices have been less damaging to the poor than to the rich, and have had a uniform impact across groups from different partitions. The explanatory power of overall inequality provided by different characteristics is studied by means of statistical constructs independent of the equivalence scale used.

Key Words
Real and monetary inequality measurement; Household specific price indices; Equivalence scales

*Universidad Carlos III de Madrid. This is the final draft of a paper presented at the international conference on "Microeconometric Models and Tax Policy", sponsored by the *Instituto de Estudios Fiscales* in Segovia on the 18-20th of December of 1991. This work is part of a larger project carried on under the auspicies of the chair *"Gumersindo de Azcárate"*, presently occupied by the author and funded by the *Ministerio de Asuntos Sociales* of Spain. The excelent research assistanship provided by Carmen Higueras and Carlos Ballano, whose work has been partly financed by the Project PB-890103 of the Spanish DGICYT, is gratefully acnowledged, as well as numerous conversations with Juan Romo and Eduardo Ley. However, only the author is responsable for all remaining shortcomings.

INTRODUCTION

This paper compares two large household budget samples, of about 24.000 observations each, for a population of approximately 10 million households occupying private housing: the *Encuestas de Presupuestos Familiares* (EPF from here on) for 1973-74 and 1980-81, collected by the Spanish *Instituto Nacional de Estadística*.

Ruiz-Castillo (1987) measured relative money inequality and poverty in Spain using the 1980-81 EPF. Subsequently, Bosch, Escribano, and Sánchez (1989) measured the same phenomena for the 1973-74 distribution following identical procedures, and established the pertinent comparisons. At the national level, their main finding was that "the Spanish society has experienced a slight increase in mean per capita total expenditure, ..., and such an increase has been distributed so that it has led to a reduction in the inequality between the Autonomous Communities, as well as a reduction of the inequality for the country as a whole".

However, the two money distributions being compared belong to a period of high inflation: according to official estimates, the Spanish *Indice de Precios de Consumo* increased by a factor of 5,6 from 1973 to July of 1985, the last month of the index system based in 1976. What about the distributional impact of the changes in relative prices between these two dates?

Abadía (1986a) has studied the evolution of aggregate price indices from 1976 to 1981 for different household types. Using true cost-of-living indices, based on the estimation of a linear expenditure demand system, Abadía (1986b, 1987) found that inflation during the period 1976-1985 had been biased against small households and those with greater total expenditure.

Given these results, the purpose of this paper is twofold: 1) to determine the importance of changes in relative prices, measured through household specific price indices, on the evolution of measured inequality from 1973-74 to 1984; and 2) to study the implications of such changes when we consider relevant partitions of the population.

The main finding is that the improvement in real inequality is considerably larger than the improvement in money inequality. However, this qualitative result has to be understood and evaluated in the light of the following research decisions in many fronts:

- we have chosen household total expenditures, net of the acquisition of certain durables, as the best variable to represent the household's standard of living;

- we have taken the person as the unit of analysis, and we have considered the distributions which assign to each person the 'equivalent' total expenditure of the household to which she belongs, obtained with the help of a variety of equivalence scales;

- in recognition of the index number problem inherent in comparisons over time, we have estimated the change in real inequality at the time both surveys were taken, in 1973-74 and 1980-81, as well as in the middle and outside this period, in the years 1978 and 1984, respectively;

- to express the two original distributions in money units of the time periods just mentioned, we have used household specific statistical price indices rather than true cost-of-living indices;

- we have measured only relative inequality by means of the General Entropy family of additively decomposable inequality indices.

As we will see, at the national level the central result is reasonably robust to variations in the value judgements used. However, to complete this approach which pools all household types into a single distribution, we have explored also the inequality experienced by households of the same size which are presumed to have comparable needs.

Moreover, given the heterogeneity in a population of a country as large and complex as Spain, we have performed other disaggregated analysis to study which groups of the population experienced gains or losses in real inequality, as well as the role of different characteristics in explaining total inequality. We present results for the partitions by Autonomous Community, municipality size, and educational attainment of the household head. Special attention has been paid to measurement procedures independent of the equivalence scale used.

Methodologically, this is a paper in the area of applied descriptive statistics. As a matter of fact, our main interest is in carefully examining how far we can go in establishing normative conclusions with regard to the evolution of inequality over time within an explicit microeconomic framework, making use of statistical methods which do not need either too specific assumptions about individual preferences, or their recovery by means of complex and expensive econometric methods.

The rest of the paper is organized in four sections. The first section summarizes the conceptual framework and the difficulties involved in the rigorous comparison of income distributions which pertain to different groups of people confronting different price structures. The second section provides a justification for the procedures we favor, as well as the details of their empirical implementation. The third section contains the empirical results, while the final section offers some concluding comments.

I. THE CONCEPTUAL FRAMEWORK

I. 1. Money *versus* real inequality

Although the realization that different groups are not affected in the same manner by the evolution of relative prices had long been recognized in the empirical literature(1), it appears that the idea that price movements should be included in intertemporal inequality comparisons was originally suggested by Iyengar and Bhattacharya (1965), some time before the path-breaking work of Atkinson (1970), Kolm (1976a,b) and Sen (1973) on the axiomatic foundations of inequality measurement .

Subsequently, Muellbauer (1974a) proved that an inequality measure derived from a strictly quasi-concave social welfare function, defined over the distribution of individual indirect utility levels, is independent of price changes if and only if preferences are identical and homothetic for all consuming units. The conclusion is inescapable: if we do not accept such strong restrictions on individual preferences, real inequality comparisons will depend on the reference price vector.

To be precise, let us start with the comparison of "income" distributions in two moments of time for a population with constant tastes consisting of H consuming units, referred to as "individuals". Let x_τ^h be the income of the h-th individual in situation τ, with h = 1,...,H and τ = 1, 2. Under general conditions on individual preferences

$$x_\tau^h = c^h(u_\tau^h, p_\tau) \text{ for all h and } \tau,$$

where $c^h(\cdot)$ is the h-th individual cost function, p_τ an n-dimensional vector of prices in period τ, and u_τ^h the maximum utility level achievable by the h-th individual in situation τ. Finally, let us denote by $\{x_\tau^h\}$ an income distribution, consisting of H real numbers, and let $I(\cdot)$ be a real valued index of inequality defined in the space of such distributions.

We are interested in two empirical matters: 1) the measurement of changes in real inequality according to $I(\cdot)$, in other words, in the expression $\Delta R = I(\{u_2^h\}) - I(\{u_1^h\})$ where the two utility distributions are evaluated at common prices; and 2) the distributional impact of the change in relative prices from p_1 to p_2.

Of course, utility levels are not observable, but if individual preferences have been estimated we can reprice the second period utility distribution as follows:

$$x_{21}^{h} = c^{h}(u_{2}^{h}, p_{1}), \quad h = 1, \ldots, H,$$

where x_{21}^{h} is the minimum total expenditure necessary for individual h to reach the utility level u_{2}^{h} at prices p_{1}. We shall denote the change in real inequality at prices p_1 by

$$\Delta R_{1} = I(\{x_{21}^{h}\}) - I(\{x_{1}^{h}\}).$$

To evaluate the quantitative significance of the relative price changes, taking the utility levels achieved in period 2 as the reference standards of living, one may use the expression

$$\Delta P_{21} = I(\{x_{2}^{h}\}) - I(\{x_{21}^{h}\}).$$

Alternatively, we can reprice the period 1 utility distribution at prices p_2 and define the change in real inequality by

$$\Delta R_{2} = I(\{x_{2}^{h}\}) - I(\{x_{12}^{h}\}).$$

Then, the change in inequality attributable to the change of prices from p_1 to p_2, taking as reference the utility levels achieved in period 1, will be given by

$$\Delta P_{12} = I(\{x_{12}^{h}\}) - I(\{x_{1}^{h}\}).$$

As a matter of fact, if we know the prices at other time periods, we can estimate the change in real inequality at prices p_t by means of

$$\Delta R_{t} = I(\{x_{2t}^{h}\}) - I(\{x_{1t}^{h}\}),$$

where $$x_{\tau t}^{h} = c^{h}(u_{\tau}^{h}, p_{t}), \quad \tau = 1, 2^{(2)}.$$

Quite apart from the interest of verifying whether our empirical conclusions are robust to the time period used, that is, whether we obtain similar values for ΔR_1, ΔR_2 and ΔR_t, we can estimate the distributional impact of price changes at time periods within or outside the original interval [1, 2] by means of

$$\Delta P_{1t} = I(\{x_{1t}^{h}\}) - I(\{x_{1}^{h}\})$$

and $$\Delta P_{2t} = I(\{x_{2}^{h}\}) - I(\{x_{2t}^{h}\})$$

which measure the change in inequality attributable to the change of prices from p_1 (or p_2) to p_t, taking as reference the utility distributions of periods 1 (or 2), respectively.

If we define the change in money inequality by

$$\Delta M = I(\{x_2^h\}) - I(\{x_1^h\}),$$

it is clear that

$$\Delta M = \Delta P_{21} + \Delta R_1 = \Delta P_{12} + \Delta R_2 = \Delta P_{1t} + \Delta P_{2t} + \Delta R_t \text{ for all } t \neq \tau.$$

Let us agree that, in an intertemporal context, we would prefer that real inequality goes down, that is to say, that $\Delta R_t < 0$ for all t. On the other hand, let us presume that we would rather have relative prices evolving over time in a manner more favorable (or less unfavorable) for the poor than the rich. Assume that individual true cost-of-living indexes indicate that the price level keeps rising over time for all consumers, and consider first the case in which $1 < t < 2$. Then

$$x_1^h = c^h(u_1^h, p_1) < x_{1t}^h = c^h(u_1^h, p_t) \text{ for all h.}$$

But if p_t is less damaging to the poor, the increase in the minimum total expenditure necessary to sustain u_1^h will be greater for the rich than for the poor. Consequently, the inequality of the distribution $\{x_{1t}^h\}$ will be greater than for the distribution $\{x_1^h\}$, and hence ΔP_{1t} will be positive. Similarly, we would have that

$$x_2^h = c^h(u_2^h, p_2) > x_{2t}^h = c^h(u_2^h, p_t) \text{ for all h,}$$

but more so for the rich, so that the inequality of the distribution $\{x_2^h\}$ will be greater than that of $\{x_{2t}^h\}$, and hence ΔP_{2t} will be positive too. For analogous normative reasons, in the two limiting cases in which $t = 1$ or 2, we would prefer to have $\Delta P_{21} > 0$, and $\Delta P_{12} > 0$, respectively. Finally, notice that if $1 < 2 < t$, we would rather have $\Delta P_{1t} > 0$ but $\Delta P_{2t} < 0$, while if $t < 1 < 2$ it would be $\Delta P_{1t} < 0$ and $\Delta P_{2t} > 0$.

In all cases, since the desired sign of at least one of the price terms and the real inequality component go in the opposite direction, we cannot envision a prefered sign for ΔM: changes in money inequality are not a good indicator of whether the situation is improving or not from a normative point of view. In an intertemporal context the decomposition we have discussed is unavoidable.

I. 2. Household composition effects

In the seminal paper already quoted, Muellbauer (1974a) observed that the usual conditions of S-concavity and symmetry (or anonimity) of the social welfare function, made sense only in a society of consuming units with identical needs. Then he went on to suggest a simultaneous treatment of the effects on inequality of both price and household composition structures in the context of index number theory and the duality approach to consumer demand, to whose development he was contributing so much at that time(3).

As long as we want to take into account the demographic heterogeneity of any society, we must confront two issues: what is the unit of analysis, that is, what do we mean by "individuals"; and how do we deal with the fact that they have equity relevant non-income related different needs, that is, how do we make them comparable in "income" space.

As far as the first problem is concerned, the choice is among the household, the family, or the person. There is no question that the level and trend of inequality, and the ranking of specific groups, depend crucially on this decision(4). Following the generally accepted practice, we recommend the person for two types of reasons. Firstly, according to the individualistic position of the profession in moral matters, it is the economic welfare of each person in society that should count. Secondly, Western societies, Spain included, are witnessing a process of growth of independent living arrangements which results from a long-established movement from the extended to the nuclear family and, at different speeds, a more recent trend toward the splitting of the nuclear family itself. Any analysis which does not take the person as the unit, would tend to obscure these changes. Thus, no other alternative will be pursued here.

However, expenditure data comes typically aggregated at the level of the household, and even when we have information on personal incomes we have to contend with non earners. There is evidence that inequalities inside the household are empirically relevant(5). But, like most of the literature, we will accept the equal-sharing assumption, according to which all persons share equally from the household scale variable x_{τ}^{h}, conveniently adjusted to permit interpersonal welfare comparisons among households of different size and composition.

Following Muellbauer and the rest of the profession thereafter, the adjustment of the household income variable will be carried on by means of equivalence scales. For that purpose, let households be characterized by household income x, a vector a in a set A of demographic characteristics, and a set of so-called unconditional preferences for commodities and characteristics. If we let $c^h(u, p, a)$ be the corresponding unconditional cost function, the equivalence scales for household h are defined by

$$d^h(a, a^r; u, p) = c^h(u, p, a)/c^h(u, p, a^r),$$

which gives the cost of attaining a utility level u at prices p by a household of characteristics a, relative to the cost of attaining this utility at those prices by a reference household of characteristics a^r. The function $d^h(\cdot)$ provides the numbers with which one would deflate the income distribution in order to adjust it to a needs-corrected basis, according to *this* household's unconditional preferences. As Pollak and Wales (1979) and Pollak (1991) insist, we are still lacking a theory of interpersonal welfare comparisons in the presence of several unconditional preferences. But then, whose preferences should be selected for the income adjustment?

These difficulties are usually dealt with assuming the existence of an unconditional preference ordering common to all consuming units. That is,

$$d^h(a, a^r; u, p) = d(a, a^r; u, p) \text{ for all } h = 1,\dots,H.$$

Armed with this single deflator we can construct the distribution of adjusted or equivalent income per person $\{z^{ir}_{\tau t}\}$, where, for each person i in household h,

$$z^{ir}_{\tau t} = x^h_{\tau t}/d(a^h, a^r; u^h_\tau, p_t) = c(u^h_\tau, p_t, a^r).$$

This means that each person i in period t is assigned the income that a household of characteristics a^r would need, at prices p_t, to achieve the utility level u^h_τ attained in period τ by the household to whom the person belongs.

Having selected an equivalence scale $d(\cdot)$ and a reference type a^r, let us define the change in money inequality for a pair of distributions of equivalent income per person by

$$\Delta M(d(\cdot), a^r) = I(\{z^{ir}_2\}) - I(\{z^{ir}_1\}).$$

If we define the corresponding change in real inequality and the price effects for any year t in the obvious manner, we will have as before

$$\Delta M(d(\cdot), a^r) = \Delta P_{1t}(d(\cdot), a^r) + \Delta P_{2t}(d(\cdot), a^r) + \Delta R_t(d(\cdot), a^r).$$

Exactly as in price space, we must face an index number problem: for each reference type we will have a different decomposition of this sort. Clearly, to avoid this problem we must have

$$d(a, a^r; u, p) = d(a, a^r; p) \text{ for all } u.$$

In this case,

$$d(a, a^{r'}; p) = d(a, a^r; p)\, d(a^{r'}, a^r; p),$$

and for all τ and t

$$I(\{z^{ir'}_{\tau t}\}) = I(\{z^{ir}_{\tau t} / d(a^{r'}, a^r; p)\}) = I(\{z^{ir}_{\tau t}\})$$

if the inequality measure is scale invariant[(6)].

Finally, given an equivalence scale d(·) and a reference type a^r, it would be convenient to establish the relationship between the inequality of unadjusted income per household, $\{x^h_\tau\}$, and of equivalent income per person, $\{z^{ir}_\tau\}$. Let us consider, for instance, the change in money inequality, and let us denote by $\{z^{hr}_\tau\}$ the distribution in which each household is assigned the adjusted income $x^h_\tau / d(a^h, a^r; p_\tau)$. Then, for each τ, the change in money inequality attributable to the change from unadjusted to equivalent income will be denoted by

$$\Delta EQ_\tau = I(\{z^{hr}_\tau\}) - I(\{x^h_\tau\}),$$

and the variation attributable to the change in the unit of analysis from the household to the person by

$$\Delta UA_\tau = I(\{z^{ir}_\tau\}) - I(\{z^{hr}_\tau\}).$$

It is clear that

$$\Delta M = I(\{x^h_2\}) - I(\{x^h_1\})$$

$$= [I(\{x^h_2\}) - I(\{z^{hr}_2\})] + [I(\{z^{hr}_2\}) - I(\{z^{ir}_2\})] + [I(\{z^{ir}_2\}) - I(\{z^{ir}_1\})] +$$

$$[I(\{z^{ir}_1\}) - I(\{z^{hr}_1\})] + [I(\{z^{hr}_1\}) - I(\{x^h_1\})]$$

$$= -\Delta EQ_2 - \Delta UA_2 + \Delta M(d(\cdot), a^r) + \Delta EQ_1 + \Delta UA_1.$$

I. 3. The social welfare aspects of the exercise

One of the main features of Sen (1976)'s review of the field of real national income measurement, was the explicit recognition of the difficulties of interpretation entailed in comparisons involving different groups of people at two moments in time and/or two points in space. This is indeed a harder question than the usual comparison of alternative positions faced by the same population as in traditional welfare economics, social choice theory or the standard theory of national planning.

To begin with, suppose one is trying to compare the income distributions of two communities of the same size. In the present context of intertemporal comparisons for a single country, there is less reason for requiring two different set of value judgements. Thus, in spite of Sen's remark that "a community of Benthamites may turn Rawlsian", we will assume the stationarity of social judgements underlying specific inequality measures.

According to Sen, there are two types of questions to be raised in making intertemporal real income comparisons, which we apply here to inequality judgements. The first one is: 'Is Spain better off with the inequality exhibited by the 1980-81 distribution than it *would have been* with the one exhibited by the 1973-74 distribution at constant prices?' The difficulty lies in interpreting the meaning of the Spanish 1980-81 community getting the 1973-74 income distribution, a problem even more acute when dealing with two different communities. We have to admit that we don't have a clear way to establish a correspondence between the Spaniards at the two moments in time out of the S! different one-to-one correspondences between two groups of S persons each. This is one reason why it is helpful to investigate a number of relevant partitions, each entailing a particular way of aggregating individuals into a tractable number of types for which a natural correspondence between the two dates can be proposed.

The second question would be: 'Is Spain better off in 1980-81 than it *was* in 1973-74?' This is a harder issue, for which the constancy of the underlying indifference map is quite insufficient for a meaningful answer. In our case, the analysis of inequality would have to be completed with efficiency considerations. Also, as we shall see, any empirical especification of the "income" variable will manifest a variety of shortcomings as a measure of the economic position of the people. As pointed out by Muellbauer (1974a), it will typically ignore utility and disutility aspects from work and leisure, or intergenerational issues not apparent from two cross-sections of the same population. Besides, as Sen also notices, there are non economic dimensions pertaining to welfare comparisons which could have changed between periods 1 and 2 without affecting the social preference map defined over income distributions -for example, the degree of participation in decision making. But this, of course, escapes for the moment from even the richest conceptual framework at our disposal.

On the other hand, when the two populations differ in size we must face an additional difficulty of interpretation. The way out sanctioned by professional practice, is to accept an axiom which makes inequality measurement invariant to identical replicas of the population. In intertemporal comparisons for the same country with large size samples, this final difficulty does not seem to be the more damaging to the standard procedures we will be following here.

In view of the above, the properties characterizing the different inequality measures in applied work, are best seen as a purely personal expression of some values concerning society. But which other properties should they posessed? There are several criteria about the meaning of inequality that almost all types of approaches in the literature have in common. Together with the invariance to identical replicas of the population, these are a condition on symmetry, scale-invariance and the Pigou-Dalton Principle of Transfers as captured by S-convexity.

As it is well known(7), these four properties completely characterize the usual Lorenz quasi-ordering. Also, since we will be examining different partitions of the population, it will be convenient to work with additively separable indicators. But combined with the rest of the axioms already mentioned, this last assumption has drastic implications(8): the class of relative inequality measures gets reduced to the following parametric specification, known as the Generalized Entropy family of indices of relative inequality:

$$I_c = (1/H)\,(1/c^2 - c)\,\Sigma_h\,\{(z^h/\mu(z^h)^c - 1\}, \qquad c \neq 0, 1;$$
$$I_0 = (1/H)\,\Sigma_h\,\log\{\mu(z^h)/z^h\}\,;$$
$$I_1 = (1/H)\,\Sigma_h\,\{z^h/\mu(z^h)\}\,\log\{z^h/\mu(z^h)\}\,,$$

where $\mu(\cdot)$ is the mean of the distribution, I_1 is the Theil index, I_0 is the mean logarithmic deviation, and the larger and more positive (the more negative) c is, the more sensitive the measure I_c is to income differences at the top (bottom) of the distribution.

II. STATISTICAL PROCEDURES AND EMPIRICAL IMPLEMENTATION

Given data on the distributions of household net expenditures and demographic characteristics for a population in two moments of time, to actually obtain empirical results on the evolution of inequality one needs to have: 1) household specific price indices to express the original distributions at common prices; 2) a single set of equivalence scales to permit welfare comparisons of households of different size and/or composition.

From the beginning of the empirical analysis of demand, some practitioners have always been tempted by the possibility of estimating equivalence scales along with the usual price and income effects. In this way, observed behavior, welfare and household characteristics would be linked in a systematic manner -a feature of the econometric approach which makes it appealing to many economists.

For that purpose, one needs to assume the existence of a single unconditional preference ordering over commodities and demographic attributes, common to all consuming units. This is an assumption plagued with a number of well known difficulties(9), beginning with the normative objections raised by Fisher (1987) to the claim that equal utility implies equal welfare.

Special attention should be paid to the fundamental identification problem -first raised by Pollak and Wales (1979)- according to which observed consumption patterns, conditional on demographic characteristics, do not permit the recovery of unconditional preferences over commodities and demographic attributes. As established by Blundell and Lewbel (1991), conditional demands determine only 'relative' equivalence scales, which are ratios of true cost of living indices for different demographic groups. Equivalence scales at a given point in time in a single price regime remain unidentified, although commodity demands serve to determine the way the scales change over time in response to price changes.

The standard practice has consisted of arbitrarily selecting intuitively attractive assumptions about equivalence scales to narrow the range of possible specifications for preferences(10). Econometrically, this amounts to selecting a particular cardinalization of the cost function which rationalizes the data on conditional commodity demands(11). When they exist, the empirical implications of such a procedure are by no means always tested, and when they are they have been rejected(12). Other identification issues, which require new questionable assumptions, often make difficult and/or dubious the estimation of these models(13). In addition, the crucial assumption of a common utility function for all consuming units is seldom tested, and the available evidence does not support it(14). A fact to be expected if, as pointed out by Pollak and Wales

(1979), the distribution of unconditional preferences among the population is not independent on the distribution of demographic attributes.

Besides the usual data problems, most of these econometric models share still other understandable limitations which reduce their value for applied welfare analysis: they are static, cover only expenditures on private goods, and do not deal with the allocation of time within and outside the household. Finally, we must ask whether the normative conclusions on inequality measurement and their trend are unduly sensitive to the demand-functional form one chooses to estimate price and demographic effects(15), or to the specification of the function relating equivalence scales to characteristics(16).

As Coulter *et al.* (1992a) conclude, the host of assumptions underlying the econometric models are not overwhelmingly persuasive, at least for income distribution assessment purposes. On the contrary, many of them rest on potentially controversial normative judgements. Moreover, it comes as no surprise that different assumptions lead to different scales. Thus, it is fair to say that there is no "correct" set of scales and that searching for some would possibly be misguided.

The question is that other approaches are not convincing either, and do not generate robust empirical results, as documented in Buhmann *et al.* (1988) and Coulter *et al.* (1992a). Thus, a range of scales is not only inevitable but also legitimate. This is a serious problem, since thanks to the work of Coulter *et al.* (1992b), we are aware of the impact on the measurement of relative inequality of different views about the generosity of the scale.

What is to be done? There are two immediate alternatives(17). In the first place, if one insists in pooling people of different characteristics adjusting total household income or expenditures by means of some equivalence scales, then robustness should be checked by estimating inequality for different values of the key parameters which determine the scales.

In the second place, we can always study each homogeneous household type separately. Following Blundell and Lewbel (1991), one would estimate distinct cost of living indices for each household type. These would suffice to construct relative equivalence scales to express the available money distributions at comparable money units of selected years in order to distinguish price from real effects on inequality measurement. Then, following Coulter *et al.* (1992a), one would only use measures that are additively decomposable by population subgroup to minimize the impact of "inappropriate" scale relativities which, under the assumptions that will be revised presently, will contaminate only the between-group component.

Unfortunately, Blundell and Lewbel's suggestion is a non-trivial and expensive project which we may attempt at a later date. In the meanwhile we thought interesting to explore the distinction between money and real inequality using Laspeyres type statistical price indices rather than true cost-of-living constructions from an explicit behavioral model. Of course, the main advantages of such an approximation exercise are that it is valid for a wide class of individual preferences, and can be carried on at a relatively low computational cost while allowing for detailed commodity disaggregation.

The empirical implementation of this double strategy requires the discussion of the following three points: (i) which is the best variable to represent the household's standard of living; (ii) how to compare the money distributions of different time periods; and (iii) how to deal with the demographic heterogeneity of the population.

(i) The scale variable

Banks *et al.* (1991), in the context of a theoretical framework allowing a high degree of generality in the intertemporal structure of tastes and prices, found that single-period measures, such as income and consumption, will require relatively strong assumptions to act as suitable indicators of a household living standard over the complete life-cycle. However, since we do not have appropriate life-cycle welfare measures, we must make a decision on the basis of the following arguments.

On the one hand, as these authors point out, there seems to be a consensus that current consumption is preferable to current income as a measure of the household's permanent economic position(18). On the other hand, budget studies are designed to provide an accurate measurement of all types of expenditures, while declared income is often seriously underreported. In particular, the expenditures of the self-employed, those working in the agricultural sector, and the suppliers of the "underground" or the "irregular economy", present no special measurement problems, which is surely not the case for their incomes. At any rate, in the Spanish case, since more than 60 % of all households report greater expenditures than total income, until this circumstance is explicitly modeled we recommend concentrating all the attention on household total expenditure, as an estimate of private total consumption, rather than total income.

In our surveys, the concept of total expenditure includes transfers made by the household, as well as a number of imputations for consumption and wages in kind, subsidized meals at work, and a market rental value, estimated by the owner, for owner-occupied housing. However, our experience with the 1980-81 EPF(19) indicates that discontinuous household expenditures on some durables, whose occurrence may distort heavily the total, are best considered investment rather than consumption. These include current acquisitions of cars, motorcycles and other means of private transportation, as well as house repairs financed by either tenants or owner-occupiers. Thus, our estimate of household current consumption, x^h_τ, will be total household expenditures, net of these investment items.

This is of course a measure of private consumption of goods and services, which does not include neither leisure nor the impact of the public sector via taxes or publicly provided goods and services. The possible effect on the standard of living of asset ownership or liquidity constraints will be absent also from the analysis.

(ii) Repricing the scale variable

In order to express the various expenditure distributions in comparable money units of the same time period, we use household specific Laspeyres price indices whose construction deserves some explanation.

The present system of official price indices for Spain is based on the year 1983. Being impossible to extend backwards this system beyond 1978, we decided to use the previous system which has 1976 as the base year. Since we have monthly price data from 1976 onwards, and we know the quarter during which each household of the second survey was interviewed, it is possible to select one of them, namely Winter 1981, as situation 2. Unfortunately this is not the case for the first survey: we only have annual price data from 1960 to 1975, and we do not have information about the time structure of the survey during the span July 1973 to June 1974. Therefore, the benchmark corresponding to situation 1 will have to be the average between 1973 and 1974. Moreover, we will consider other time periods: 1978, within the interval (1973-74 , Winter 81), and 1984 outside of it.

As reported in Higuera and Ruiz-Castillo (1991), to compare a price vector in a given year t with the prices in the base year 1976 for a household h, we constructed individual indices of the type

$$I^h(p_t, p_{76}; w^h_\tau) = \Sigma_j w^h_{j\tau} I_{jt},$$

where $w^h_{j\tau}$ is the share of total expenditure devoted to commodity j by household h in the survey year τ, I_{jt} is the official price index for commodity j in year t, and j = 1, ... , 58.

To express a given distribution -for instance the distribution $\{x^h_1\}$ of net total expenditures in situation 1- in money terms of a year t, we need individual Laspeyres type indices based on situation 1. These are easily constructed as follows:

$$L^h(p_t, p_1 ; w_1^h) = I^h(p_t, p_{76} ; w_1^h) / I^h(p_1, p_{76} ; w_1^h)$$

where $$p_1 = (1/2)\, p_{73} + (1/2)\, p_{74} .$$

Then, the repriced distribution will be

$$y_{1t}^h = x_1^h \, L^h(p_t, p_1 ; w_1^h).$$

for h = 1,..., 24.151 and t = 1978, Winter 1981, and 1984. Similarly, the repriced distributions for the second survey data will be

$$y_{2t}^h = x_2^h \, L^h(p_t, p_2 ; w_2^h).$$

for h = 1,..., 23.952 , p_2 = Winter 81, and t = 1973-74, 1978, and 1984.

Of course, a statistical price index provides only an upper bound to the true cost-of-living construction. Therefore, we will have

$$y_{\tau t}^h = x_\tau^h \, L^h(p_t, p_\tau ; w_\tau^h) \geq x_\tau^h \, L^h(p_t, p_\tau ; u_\tau^h) = x_{\tau t}^h .$$

Hence, the nature of our approximations will depend on the substitution bias incurred with the use of statistical indices. In particular, if the bias is greater for the rich, as can be expected, then for any τ and t we will have

$$I(\{y_{\tau t}^h\}) > I(\{x_{\tau t}^h\}).$$

Thus, the expressions

$$\Delta Y_L = I(\{y_{21}^h\}) - I(\{x_1^h\}),$$

and $$\Delta Y_U = I(\{x_2^h\}) - I(\{y_{12}^h\})$$

will provide a lower and an upper bound, respectively, to the theoretical expressions $\Delta R_1, \Delta R_2$ when the latter are negative. Unfortunately, for t different from τ nothing can be said *a priori* about the relationship between $\Delta Y_t = I(\{y_{2t}^h\}) - I(\{y_{1t}^h\})$ and ΔR_t.

(iii) The treatment of differences in household size

Having renounced to an explicit behavioral model for equivalence scales, how should we deal with the demographic heterogeneity of the population? We will follow Coulter *et al.* (1992b) who have studied how measures of inequality and poverty defined over distributions of equivalent income change if equivalence scale relativities are changed. They start from a simple characterization of equivalence scales as a function of household size s^h and one parameter Θ, independently of prices and utility levels. In the notation of Section II:

$$d^h(a, a^r; u, p) = d(s^h, \Theta),\ \Theta > 0,\ \partial M_s/\partial s > 0, \text{ and } \partial M_s/\partial \Theta > 0,$$

where the reference type a^r is taken to be a household consisting of one adult. For tractability reasons, they accept Buhmann *et al.* (1988) suggestion of working with

$$d(s^h, \Theta) = (s^h)^\Theta,\ \Theta \in [0,1],$$

because it provides a good approximation to virtually all the different scales currently used in empirical studies of income distribution.

Following this approach, the objects of study will be the distributions $\{z^i(\Theta)\}$ in which each person i receives the equivalent total expenditure of the household h to which she belongs,

$$z^i(\Theta) = y^h/(s^h)^\Theta, \text{ for all i in h},$$

for several appropriate values of Θ, including $\Theta = 1$ which assigns to each person the household per capita total expenditure.

The main finding so far is that there is a systematic relationship between equivalence scale generosity and the extent of relative inequality and poverty. For a given income distribution, and most measures, the extent of equivalent income inequality and poverty first falls and then rises, as relativities are increased from their minimum level. Moreover the changes induced by changing scale relativities are considerable whichever measure is used. In an intertemporal context, we are only aware of Jenkins (1991) work on the evolution of money inequality in the United Kingdom, where the sensitivity of the conclusions to the scale generosity was systematically studied.

Coulter *et al.* (1992a) second suggestion of working with measures that are decomposable by population subgroup in order to minimize the impact of "inappropriate" scale relativities, deserves some elaboration. Let us take the General Entropy family I_c of measures of relative inequality, and consider any partition of the distribution $\{z(\Theta)\}$ into $\{z^k(\Theta)\}$, k = 1,...,K disjoint subgroups. Then we know that

$$I_c(z(\Theta)) = \Sigma_k\, [v^k(\Theta)]^c\, (n^k)^{1-c}\, I_c(z^k(\Theta)) + I_c(\mu(z^k(\Theta)) = W_c^{(k)}(\Theta) + B_c^{(k)}(\Theta),$$

where:

$v^k(\Theta)$ = share of aggregate equivalent income held by group k's members;

n^k = group k's population share, independent of Θ;

$I_c(z^k(\Theta))$ = inequality of the distribution of equivalent income per person within k's group;

$\mu(.)$ = mean of the corresponding distribution;

$W_c^{(k)}(\Theta)$ = *within*-group component of total inequality;

$B_c^{(k)}(\Theta)$ = *between*-group component of total inequality, calculated as if each person within a given group received that group's mean income.

Consider the following two questions for a given partition: 1) by how much would total inequality go down if income differences between members of the partition were the only ones to exist?, or 2) by how much would total inequality go down if we were to eliminate income differences between groups, maintaining existing inequality within each of them? It is well known[20] that the only member of the General Entropy family for which this two questions have the same answer -namely, the between-group component in the above decomposition- is the mean logarithmic deviation.

In the particular case in which the population gets partitioned by household size into j = 1,...,J groups, we will have that all persons i in a given group j will be assigned the equivalent income

$$z^i(\Theta) = y^h / j^\Theta$$

of the household h to which she belongs. Therefore, since $I_c(.)$ is an index of relative inequality, we will have that, for each j,

$$I_c(z^j(\Theta)) = I_c(z^j) \text{ for all } \Theta \in [0,1].$$

Nevertheless, since the terms $v^j(\Theta)$ will still be dependent on Θ, in general the within-group component will be dependent on Θ also. Only in the case c = 0, we will have

$$I_0(z(\Theta)) = \Sigma_j (n^j)\, I_0(z^j) + I_0(\mu(z^j(\Theta)) = W_0^{(j)} + B_0^{(j)}(\Theta);$$

that is, only in this case using the wrong equivalence scale contaminates exclusively the between-group component.

If we want to study any other partition k = 1,...,K, we should apply the decomposability of the mean logarithmic deviation to the within-group component of the partition by household size which, as we have seen, is independent of the equivalence scale. Then

$$I_0(z(\Theta)) = \Sigma_{jk} (n^{jk}) I_0(z^{jk}) + \Sigma_j (n^j) I_0(\mu(z^{j1}), \ldots, \mu(z^{jK})) + B_0^{(j)}(\Theta)$$

$$= W_0^{(jk)} + B_0^{(k \to j)} + B_0^{(j)}(\Theta)$$

where

$W_0^{(jk)}$ = within-group component of total inequality in the partition by household size *and* the characteristic k;

$B_0^{(k\to j)}$ = demographically weighted average of the impact of characteristic k on each of the members of the partition by household size, or 'true' (independent of Θ) between-group component of total inequality due to the effect of characteristic k.

This convenient new concept is not to be confused with

$$B_0^{(k)}(\Theta) = I_0(\mu(z^1(\Theta),...,\mu(z^K(\Theta))$$

or

$$B_0^{(kj)}(\Theta) = I_0(\mu(z^{1}1),..., \mu(z^{1}J),...,\mu(z^{K}1),..., \mu(z^{K}J))$$

which are, respectively, the standard between-group components in the partitions by characteristic k and characteristics j and k simultaneously. As a matter of fact, it is easy to see that

$$B_0^{(k\to j)} = B_0^{(kj)}(\Theta) - B_0^{(j)}(\Theta) = B_0^{(k)}(\Theta) + [B_0^{(j\to k)} - B_0^{(j)}(\Theta)],$$

where

$$B_0^{(j\to k)} = \Sigma_k(n^k)I_0(\mu(z^{k}1),..., \mu(z^{k}J)).$$

III. EMPIRICAL RESULTS

We will answer the following questions for the country as a whole: **1.** How does the measurement of relative inequality in situations 1 and 2 change with the variation in the two parameters Θ and c representing, respectively, the generosity of the equivalence scale and the aversion to inequality? **2.** How does the change in money inequality vary with Θ and c? **3.** How good are our lower and upper approximations to the change in real inequality, and how varies the relationship between money and real inequality as a function of Θ and c? **4.** In particular, how can we explain the difference between the evolution of inequality for the distribution of household expenditure per household -the case $\Theta = 0$- and the evolution of inequality for the distribution of equivalent expenditure per person for values of Θ greater than 0? **5.** Which is the distributional incidence at the national level of the change in relative prices in the period 1973-74 to 1984?

6. The next question has to do with the evolution of money and real inequality for each household type with comparable needs in the partition by household size. For other partitions into Autonomous Communities, municipality size, and educational attainment of the household head, the questions will be: **7.** Which particular groups gain and loose in money and real inequality for different values of c and Θ? **8.** Which partition explains better total inequality in both survey years, that is, which partition generates a larger between-group component in the two situations?

1. For the two surveys, Table 1 shows the distribution of persons by household size, as well as mean total expenditure for each group relative to the mean. The parameter Θ, representing the weight given to household size, takes on the values considered in Buchman *et al.*(1988): 0.00, 0.25, 0.36, 0.55, 0.72, and 1.00. Moreover, we have considered the so-called Oxford equivalence scale, widely used internationally, including the Spanish INE. It gives a unit weight to the first adult -a person 14 or more years old- 0.7 to each additional adult, and 0.5 to every person less than 14 years old. It will be referred to by the symbol EQ.

It is important to note that when size plays no role, that is, when $\Theta = 0$, we always use the distribution of household total expenditure per household. However, when $\Theta \neq 0$, all estimates refer to a distribution of equivalent expenditure per person. In all cases we have used the information on blowing up factors provided by INE. Thus, ours are not sample estimates but blown up estimates for the total population.

The patterns are very similar for the two situations: about 60 per cent of individuals live in 3-, 4- or 5-person households. Overall mean

TABLE 1. Mean equivalent expenditure by household size as a function of Θ

EPF 1973-74

Household size	Number of persons as % of total pop.	Mean equivalent expenditure as % of overall mean Θ = 0.00	0.25	0.36	0.55	0.72	1.00	EQ
1	2.2	37.4	47.7	56.2	74.2	94.7	139.4	91.7
2	10.9	66.8	71.7	78.2	90.6	102.7	124.6	117.3
3	15.7	95.0	92.0	96.1	103.0	108.9	118.0	115.3
4	23.9	112.2	101.0	102.2	103.8	104.5	104.4	106.6
5	19.8	124.0	105.7	104.4	101.5	97.1	98.5	95.7
6	13.2	133.8	108.9	105.3	99.0	93.1	83.0	86.8
7 or +	14.3	156.8	120.5	112.9	100.4	84.2	90.0	78.4
Overall mean exp.		254.608	199.722	169.491	128.330	100.645	68.319	103.879
Cov(z, log s)/z		63.3	41.9	32.0	14.1	- 2.7	- 32.3	- 11.3

EPF 1980-81

Household size	Number of persons as % of total pop.	Mean equivalent expenditure as % of overall mean Θ = 0.00	0.25	0.36	0.55	0.72	1.00	EQ
1	2.1	41.9	54.1	63.5	83.5	105.9	154.9	101.5
2	11.4	70.1	76.1	82.8	95.4	107.6	129.6	121.6
3	15.1	96.1	94.3	98.2	104.6	110.2	118.5	116.1
4	25.5	113.2	103.4	104.3	105.3	105.6	104.7	107.6
5	20.1	122.4	105.7	104.0	100.6	97.1	90.5	94.2
6	12.5	129.0	106.4	102.6	95.9	89.8	79.5	86.3
7 or +	13.3	145.6	112.9	105.8	94.0	84.2	69.4	72.8
Overall mean exp.		854.082	661.107	562.887	428.522	337.645	230.871	352.469
Cov(z, log s)/z		146.3	81.3	51.4	- 2.3	- 52.7	-141.6	- 95.6

expenditure declines substantially as Θ increases. However, among those from small-sized households, mean equivalent total expenditure raises with Θ, while the reverse occurs for those from large households. These regularities are reflected in the normalized covariance between log household size and equivalent expenditure, which is positive at low Θ but negative at high Θ. This seems to confirm similar findings by Coulter *et al.* (1992b) for the U.K., and is consistent with a U-form for inequality measurement in a single cross-section as Θ varies from 0 to 1. Notice that, judging by equivalent expenditure means, the Oxford scale seems to be located somewhere in the interval Θ = [0.72,1].

Table 2 presents information on inequality at the national level for the following members of the General Entropy family: the first Theil measure (c = 1), the mean logarithmic deviation (c = 0), a top-sensitive

TABLE 2. How relative inequality varies as the aversion to inequality, the equivalence scale generosity, and the time period change

EPF 1973-74

c	Money units of period	Θ =	0.00	0.25	0.36	0.55	0.72	1.00	EQ
2	73-74		0.2934	0.2351	0.2284	0.2249	0.2312	0.2659	0.2386
	1978		0.3042	0.2444	0.2378	0.2346	0.2415	0.2776	0.2489
	Winter81		0.3172	0.2569	0.2504	0.2475	0.2550	0.2931	0.2626
	1984		0.3228	0.2618	0.2552	0.2522	0.2598	0.2984	0.2668
1	73-74		0.2316	0.1833	0.1781	0.1740	0.1763	0.1932	0.1802
	1978		0.2380	0.1889	0.1837	0.1797	0.1819	0.1989	0.1855
	Winter81		0.2457	0.1967	0.1916	0.1879	0.1915	0.2082	0.1943
	1984		0.2493	0.1999	0.1948	0.1911	0.1936	0.2111	0.1969
0	73-74		0.2498	0.1842	0.1775	0.1713	0.1717	0.1851	0.1746
	1978		0.2570	0.1899	0.1832	0.1769	0.1772	0.1905	0.1797
	Winter81		0.2642	0.1970	0.1904	0.1844	0.1849	0.1986	0.1875
	1984		0.2675	0.1999	0.1933	0.1872	0.1876	0.2010	0.1897
-1	73-74		0.3930	0.2410	0.2268	0.2124	0.2094	0.2248	0.2121
	1978		0.4095	0.2507	0.2361	0.2211	0.2178	0.2329	0.2200
	Winter81		0.4195	0.2596	0.2450	0.2301	0.2270	0.2427	0.2292
	1984		0.4253	0.2637	0.2489	0.2336	0.2304	0.2458	0.2321

EPF 1980-81

c	Money units of period	Θ =	0.00	0.25	0.36	0.55	0.72	1.00	EQ
2	73-74		0.2141	0.1694	0.1656	0.1656	0.1737	0.2081	0.1877
	1978		0.2211	0.1756	0.1717	0.1717	0.1798	0.2144	0.1938
	Winter81		0.2279	0.1823	0.1786	0.1790	0.1877	0.2238	0.2018
	1984		0.2344	0.1877	0.1839	0.1841	0.1926	0.2284	0.2063
1	73-74		0.1857	0.1455	0.1419	0.1405	0.1447	0.1646	0.1543
	1978		0.1907	0.1502	0.1466	0.1451	0.1493	0.1690	0.1587
	Winter81		0.1950	0.1545	0.1511	0.1498	0.1543	0.1744	0.1636
	1984		0.1997	0.1586	0.1550	0.1536	0.1578	0.1776	0.1668
0	73-74		0.2024	0.1497	0.1450	0.1418	0.1447	0.1619	0.1536
	1978		0.2080	0.1547	0.1499	0.1467	0.1495	0.1665	0.1582
	Winter81		0.2119	0.1586	0.1539	0.1509	0.1538	0.1712	0.1625
	1984		0.2168	0.1626	0.1578	0.1546	0.1573	0.1743	0.1657
-1	73-74		0.2948	0.1899	0.1806	0.1730	0.1750	0.1965	0.1866
	1978		0.3061	0.1979	0.1884	0.1805	0.1824	0.2040	0.1940
	Winter81		0.3104	0.2021	0.1926	0.1849	0.1870	0.2090	0.1985
	1984		0.3190	0.2082	0.1984	0.1903	0.1921	0.2138	0.2033

measure (c = 2) -which is half the square of the coefficient of variation- and a bottom-sensitive measure (c = -1). Each of the two distributions is expressed in money units of situation 1 (1973-74), situation 2 (Winter of 1981), as well as the additional time periods 1978 and 1984.

Figure 1 graphs how inequality varies with Θ and c when both distributions are expressed at the prices of their respective time periods, that is, in money units of 1973-74 and Winter of 1981, respectively.

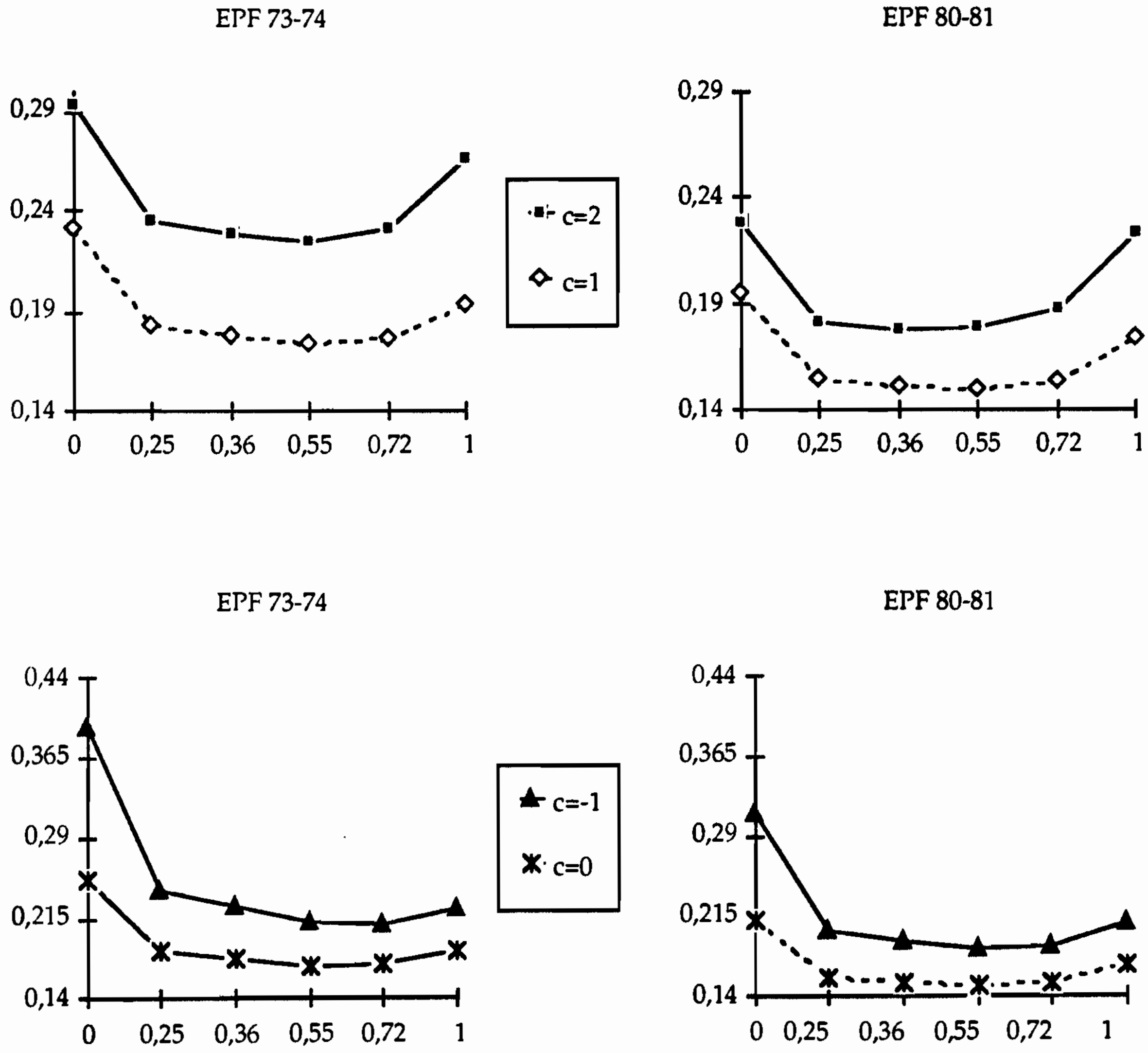

Figure 1

It appears that, for both surveys, a U-form is more pronounced for higher values of c. However, when we compute the range for each c as [(max/min) - 1] 100, excluding either the polar case Θ = 0 or both Θ = 0 and 1, we obtain:

Range of variation in percentage terms

c =	2	1	0	-1
EPF 1973-74:				
$\Theta \in [0.25, 1]$	18.2	11.0	8.1	15.1
$\Theta \in [0.25, 0.72]$	6.1	5.3	7.5	15.1
EPF 1980-81:				
$\Theta \in [0.25, 1]$	25.3	16.4	13.4	13.0
$\Theta \in [0.25, 0.72]$	13.0	9.2	7.7	9.3

All of these values are rather large and, as a matter of fact, greater than the ones reported in Coulter *et al.* (1992b) for c = 1 and 0. When we restrict ourselves to the per person distributions, in both surveys the maximum inequality value is reached at c = 2 and $\Theta = 1$, and the minimum at c = 0 and $\Theta = 0.55$. The ranges of variation across all values of c are 55.2 % and 48.1 % for 1973-74 and 1980-81, respectively.

Finally, notice that the inequality estimates for the Oxford scale are again within those of the interval [0.72, 1] for Θ, but much closer to the lower bound in 1973-74 and to the upper bound in 1980-81.

2. Figure 1 illustrates another important fact: for all values of c, the curve showing inequality in situation 2 is below the one in situation 1. That is, as we knew from other studies money inequality has improved in Spain during this period, or $\Delta M(\Theta) < 0$ for all Θ, independently of the inequality index used. In order to work with positive numbers, numerical estimates of

$$-\Delta M(\Theta)/I(z_1^i(\Theta)) = -[I(z_2^i(\Theta)) - I(z_1^i(\Theta))]/I(z_1^i(\Theta))$$

for selected values of Θ are provided in Table 3 under the heading ΔM. For all values of c, the improvement in money inequality decreases continuously as we keep giving more weight to household size. This effect is more pronounced as we decrease c. Thus, in the case c = - 1, ΔM falls by more than 50 % as Θ varies in the interval [0.25, 1].

The range of variation is large: for per person distributions, the maximum value of ΔM is 22.5 %, reached when c = 2 and $\Theta = 0.25$, while the minimum is 7 % for c = -1 and $\Theta = 1$. For the important case c = 0, the change in money inequality when $\Theta = 0.25$ or EQ, for example, is 12.0 % and 6.9 %, respectively.

3. The change in money inequality has a very limited interest in itself. What matters is the change in real inequality and the distributional role of changes in relative prices at different dates. Because of the index number problem, we know that the change in real inequality at situation 1

TABLE 3. Percentage change of money and real inequality, relative to inequality in 1973-74, as the parameters c and Θ change

c		Θ= 0.00	0.25	0.55	1.00	EQ
	ΔM	22.3	22.5	20.4	15.8	15.4
	ΔY_L	27.0	28.0	26.4	21.7	21.3
	ΔY_U	30.4	31.8	30.5	26.0	25.5
2	$(\Delta Y_U - \Delta Y_L)/\Delta Y_L$	12.6	13.6	15.5	19.8	19.7
	$\Delta Y_L/\Delta M$	1.211	1.244	1.294	1.373	1.383
	$\Delta Y_U/\Delta M$	1.363	1.413	1.494	1.645	1.656
	$(\Delta Y_U - \Delta Y_L)/\Delta M$	15.2	16.9	20.0	27.2	27.3
	ΔM	15.8	15.7	13.9	9.7	9.2
	ΔY_L	19.8	20.6	19.3	14.8	14.3
	ΔY_U	21.9	23.0	21.9	17.4	17.0
1	$(\Delta Y_U - \Delta Y_L)/\Delta Y_L$	10.6	17.5	13.5	17.6	18.9
	$\Delta Y_L/\Delta M$	1.253	1.312	1.388	1.526	1.554
	$\Delta Y_U/\Delta M$	1.386	1.465	1.575	1.794	1.848
	$(\Delta Y_U - \Delta Y_L)/\Delta M$	13.3	15.3	18.7	26.8	29.4
	ΔM	15.2	13.9	12.0	7.5	6.9
	ΔY_L	19.0	18.7	17.2	12.6	12.0
	ΔR_{78}	19.6	19.1	17.6	12.9	12.3
0	ΔR_{84}	20.3	20.3	19.0	14.5	13.8
	ΔY_U	21.0	20.9	19.6	14.8	14.3
	$(\Delta Y_U - \Delta Y_L)/\Delta Y_L$	10.5	11.8	13.9	17.5	19.2
	$\Delta Y_L/\Delta M$	1.250	1.345	1.433	1.680	1.739
	$\Delta Y_U/\Delta M$	1.382	1.504	1.633	1.973	2.072
	$(\Delta Y_U - \Delta Y_L)/\Delta M$	13.2	15.9	20.0	29.3	33.3
	ΔM	21.0	16.1	12.9	7.0	6.4
	ΔY_L	25.0	21.2	18.5	12.6	12.0
	ΔY_U	27.7	23.9	21.3	15.0	14.4
-1	$(\Delta Y_U - \Delta Y_L)/\Delta Y_L$	10.8	12.7	15.1	19.5	20.0
	$\Delta Y_L/\Delta M$	1.190	1.317	1.434	1.800	1.875
	$\Delta Y_U/\Delta M$	1.319	1.484	1.651	2.143	2.250
	$(\Delta Y_U - \Delta Y_L)/\Delta M$	12.9	16.7	21.7	34.3	38.5

prices, ΔR_1, need not be equal to the same concept at situation 2 prices, ΔR_2, nor at other different dates, say ΔR_t. Nevertheless, for the sake of robustness we hope that all these magnitudes are not too distant in practice.

We do not have a direct estimate of these concepts. Rather, we have performed an approximation exercise to establish a lower bound ΔY_L for ΔR_1, as well as an upper bound ΔY_U for ΔR_2 when the theoretical concepts are negative. We have also estimated ΔR_t for t = 1978 and 1984. The results, always relative to inequality in situation 1 and expressed in positive figures, are also in Table 3.

Now we should ask: how good is our approximation? To begin with, for all c and Θ the lower bound is always smaller than the upper bound. As a matter of fact, as shown by way of example only for c = 0, it is also always the case that

$$\Delta Y_L < \Delta R_{78} < \Delta R_{84} < \Delta Y_U,$$

with ΔR_{78} very close to ΔY_L, and ΔR_{84} very close to ΔY_U in all cases. What we may call the approximation error to the change in real inequality, ($\Delta Y_U - \Delta Y_L$), is roughly invariant with c, grows slowly with Θ, and seems to be of a tolerable order of magnitude: between 10-20 % of the value of ΔY_L, or 15-35 % of the value of ΔM, depending on Θ.

On the other hand, as can be seen in Table 3, both ΔY_L and ΔY_U decrease with Θ, but less so than ΔM. Interestingly enough, for almost all values of Θ our estimates of the change in real inequality are larger for top-sensitive and bottom-sensitive members of the General Entropy family than for the mean logarithmic measure. At any rate, the central finding is that the improvement in real inequality during this period is always greater than the improvement in money inequality.

Concentrating the attention to important selected cases, we have:

	c	Θ	ΔM	ΔY's	$(\Delta Y - \Delta M)/\Delta M$
Minimum:	-1	1	7.0 %	12.6-15.0 %	80-114 %
Intermediate:	0	0.55	12.0 %	7.6-19.0 %	43-63 %
Maximum:	2	0.25	22.5 %	28.0-31.8 %	24-41 %

Relative to situation 1, while money inequality has improved in a range that goes from 7 to 22.5 %, real inequality has improved from about 13 to 30 %. Thus, the improvement in real inequality is between 24 and 114 %

greater than the improvement in money inequality, depending on our value judgements about c and Θ and the approximation error in our estimates. In particular, when we choose the Oxford scale, real inequality has improved between 38 to 125 % more than money inequality, depending on the parameter c.

4. Let us focus on our lower bound estimate of real inequality change in the distribution of total expenditure per household when c = 0, that is

$$\Delta Y_L(0) = I_0(\{y_{21}^h\}) - I_0(\{x_1^h\}),$$

where $$y_{21}^h = x_2^h L^h(p_1, p_2; w_2^h)$$

is our estimate of household h total expenditure in situation 2 at prices of situation 1. In the previous section we saw that, for any other $\Theta' > 0$,

$$\Delta Y_L(0) = \Delta Y_L(\Theta') + [\Delta EQ_1(\Theta') + \Delta UA_1(\Theta')] - [\Delta EQ_2(\Theta') + \Delta UA_2(\Theta')],$$

where, for each τ, $\Delta EQ_\tau(\Theta')$ is the change in real inequality attributable to the move from unadjusted to equivalent income, and $\Delta UA_\tau(\Theta')$ is the variation attributable to the change in the unit of analysis from the household to the person. The estimates for these terms when $\Theta' = 0.55$ and 1.00 are as follows:

	$\Theta' = 0.55$	$\Theta' = 1.00$
$\Delta EQ_1(\Theta')$	−0.05711	−0.04690
$-\Delta EQ_2(\Theta')$	0.04568	0.02708
$\Delta UA_1(\Theta')$	−0.02130	−0.01773
$-\Delta UA_2(\Theta')$	0 .01490	0.01345
$\Delta Y_L(\Theta')$	−0.02953	−0.02326
$\Delta Y_L(0)$	−0.04736	−0.04736

Observe that $\Delta EQ_\tau(\Theta) < 0$ and $\Delta UA_\tau(\Theta) < 0$ for all τ and Θ; that is, when we go both from unadjusted to equivalent expenditure, or from the household to the person, there is a reduction on measured inequality. However, we see that the "equivalent expenditure" effect is always greater in absolute terms than the "unit of analysis" effect. Also, since situation 1 figures are always greater than those for situation 2, we have that the improvement in real inequality when $\Theta = 0$ is greater than when Θ' is 0.55 or 1.00.

5. Clearly, changes in relative prices must have played a positive redistributive role. Recall that the decomposition of the change in money inequality into a real and a price effect in the two polar cases requires that

$$\Delta M = \Delta P_{21} + \Delta R_1 = \Delta P_{12} + \Delta R_2,$$

where situation 1 is 1973-74 and situation 2 is Winter of 1981. Also, for any $t \neq 1$ or 2,

$$\Delta M = = \Delta P_{1t} + \Delta P_{2t} + \Delta R_t.$$

As we have seen, we have estimated this last expression for t = 1978 and 1984 -situations 3 and 4, respectively- so that we can learn about the distributional impact of relative prices along the sequence 1973-74, 1978, Winter 1981, and 1984. This can be done on the basis of either of the two budget surveys, as in Table 4 where we report on the cumulative sequences $(\Delta P_{21} - \Delta P_{23})$, ΔP_{21}, $(\Delta P_{21} + |\Delta P_{23}|)$, and ΔP_{13}, ΔP_{12}, ΔP_{14}, corresponding to the second and the first situation, respectively. Only the case c = 0 is presented here, since the interested reader can obtain most of the remaining information from Table 3.

TABLE 4. Percentage change of price effects, relative to inequality in period 1973-74, as the equivalence scale parameter change. Case c = 0

Impact of price changes from 1973-74 to:	$\Theta =$	0.00	0.25	0.55	1.00	EQ
1978	$\Delta P_{21} - \Delta P_{23}$	2.2	2.7	2.8	2.5	2.6
	ΔP_{13}	2.9	3.1	3.3	2.9	2.9
Winter-81	ΔP_{21}	3.8	4.8	5.3	5.0	5.1
	ΔP_{12}	5.8	7.0	7.6	7.3	7.4
1984	$\Delta P_{21} + \|\Delta P_{23}\|$	5.8	7.0	7.3	6.7	6.9
	ΔP_{14}	7.1	8.5	9.2	8.6	8.9
ΔM		15.2	13.9	12.0	7.5	6.9
$\Delta P_{21}/\Delta M$		0.249	0.345	0.440	0.671	0.737
$\Delta P_{12}/\Delta M$		0.381	0.501	0.636	0.969	1.070

According to both surveys, for all values of Θ relative prices have evolved in all periods less favorably for the rich, thus contributing to make the improvement in real inequality greater than for money inequality. Price effects in all periods follow an inverted U pattern as the generosity of the scale increases in the interval [0,1], reaching a maximum at the value $\Theta = 0.55$. The impact is considerably greater during the shorter period 1978-Winter 1981: about twice as large as the initial period 1973-74 to 1978. The positive effect is somewhat weaker during the period Winter 1981-1984.

In particular, in the decomposition of money inequality the price component remains in the range 4 to 7 %. Since the improvement in money inequality decreases with Θ, the importance of price effects relative to ΔM grows considerably: from about 30 to almost 100 % as Θ moves from 0 to 1.

6. Traditionally, in applied work with a single cross-section, given a partition of the population by any non-income characteristic there has been two issues of interest: the contribution to inequality by each individual group, and the magnitude of the between-group component as a measure of the explanatory power of overall inequality by the characteristic in question. Since we have two cross-sections, we can also study the evolution of money and real inequality for each of the groups considered.

In the previous section we reviewed the advantages of the mean logarithmic deviation as an additively decomposable measure of relative inequality. For the partition by household size, the incidence of each group's inequality will be measured by π^j, the ratio of its contribution to within-group inequality to its own demographic weight, that is,

$$\pi^j = [I_0(z^j)/W_0^{(j)}]/n^j .$$

A value of π^j greater than 1, for instance, means that this household size contributes to overall inequality -corrected by the fact that within-group inequality is not the only source of inequality for the population as a whole- more than what it could be expected from its demographic importance. Notice, of course, that this measure is independent of Θ.

Table 5 presents the empirical evidence. The main result is that, in both surveys, small sized households and those with 7 or more members register the largest contributions relative to their demographic importance. Four person households contribute the least to total inequality.

TABLE 5. Inequality within the partition by household size. Case c = 0

	EPF 1973-74			EPF 1980-81		
Number of people	(1) % n^j	(2) % $I_0(z^j)/W_0^{(j)}$	π^j = (2)/(1)	(1) % n^j	(2) % $I_0(z^j)/W_0^{(j)}$	π^j = (2)/(1)
1	2.2	4.5	2.04	2.1	4.1	1.96
2	10.9	15.9	1.46	11.4	15.7	1.38
3	15.7	14.4	0.92	15.1	15.5	1.03
4	23.9	19.7	0.83	25.5	21.6	0.85
5	19.8	18.0	0.91	20.1	18.1	0.90
6	13.2	12.3	0.93	12.5	10.5	0.84
7 and +	14.3	14.9	1.04	13.3	14.5	1.09
Total	100.0	100.0		100.0	100.0	

How much of total inequality can be attributed to differences in household size? The standard answer to this question is in terms of the between-group component when c = 0. Unfortunately, as we saw in the previous section, this statistic is contaminated by the generosity of the scale one cares to choose. The evidence for our two surveys and different members of the General Entropy family is in Table 6. Even excluding the case $\Theta = 0$, the percentage of total inequality attributable to differences between the means of the various household groups, varies widely with Θ for all selected values of c.

TABLE 6. Between-group inequality as a percentage of overall inequality in the partition by household size

	c	Θ = 0.00	0.55	1.00	EQ
	1	22.9	0.90	7.78	4.92
EPF 1973-74	0	24.4	0.97	8.22	5.20
	-1	19.1	0.84	6.92	4.39
	1	21.3	0.86	11.8	7.75
EPF 1980-81	0	22.4	0.87	12.3	8.01
	-1	18.3	0.72	10.3	6.93

Table 7 contains important evidence, independent of the generosity of the scale, on the decomposition of money inequality into a real and a price effect at prices of situations 1 and 2 for households classified by their size.

TABLE 7. Change in money and real inequality within the partition by household size

Number of people	ΔM^j	=	ΔR^j_1	+	ΔP^j_{21}	=	ΔR^j_2	+	ΔP^j_{12}
c = 1									
1	-18.4	=	-23.3	+	4.9	=	-28.2	+	9.8
2	-19.9	=	-24.5	+	4.6	=	-28.4	+	8.5
3	-5.2	=	-11.7	+	6.6	=	-13.2	+	8.1
4	-10.9	=	-16.6	+	5.7	=	-19.2	+	8.3
5	-15.6	=	-20.8	+	5.1	=	-23.3	+	7.7
6	-24.9	=	-28.9	+	4.0	=	-32.9	+	8.0
7 and +	-11.8	=	-17.5	+	5.7	=	-18.9	+	7.1
c = 0									
1	-16.1	=	-20.9	+	4.7	=	-24.8	+	8.7
2	-17.3	=	-21.9	+	4.6	=	-25.4	+	8.1
3	- 1.8	=	- 8.3	+	6.5	=	- 9.7	+	7.9
4	-10.2	=	-15.7	+	5.4	=	-18.2	+	8.0
5	-13.4	=	-18.5	+	5.0	=	-20.8	+	7.4
6	-20.2	=	-24.6	+	4.4	=	-27.8	+	7.6
7 and +	- 9.3	=	-14.8	+	5.5	=	-16.0	+	6.7
c = -1									
1	-20.9	=	-25.9	+	5.0	=	-29.3	+	8.4
2	-19.8	=	-24.7	+	4.9	=	-29.1	+	9.2
3	3.5	=	- 3.9	+	7.4	=	-5.2	+	8.8
4	-12.6	=	-18.3	+	5.7	=	-21.7	+	9.1
5	-13.2	=	-18.5	+	5.3	=	-21.2	+	8.0
6	-22.0	=	-27.0	+	5.0	=	-30.0	+	7.9
7 and +	- 4.5	=	-10.4	+	5.9	=	-11.9	+	7.4

The pattern does not change qualitatively with parameter c: except 3-person households for the bottom sensitive measure c = -1, money inequality improves for all groups, but less so than real inequality. From this point of view, the evidence for the progressive role played by relative prices over this period is overwhelming.

Nevertheless, the improvement in real inequality is not the same for all groups: in absolute terms, 1-, 2- and 6-person households, which represent about 26 % of the population, experience improvements of more than 20 % relative to inequality in 1973-74. The important group of 3-person households experience the minimum improvement, 4-12 %. However, the size of the price effect appears to be of the same order of magnitude for all groups, in the range of 5-9 %, depending on whether we take the lower or the upper estimate and on the value of c.

7. When inequality is measured by the mean logarithmic deviation, a value greater (lower) than 1 for the expression

$$\pi^k(\Theta) = [I_0(z^k(\Theta))/W_0^{(k)}(\Theta)]/n^k$$

indicates whether group k contributes to within-group inequality more (less) than its demographic weight leads us to expect. The results for the partitions by Autonomous Communities (CCAA), the size of the municipality of residence (MUN), and the educational level attained by the household head (EDC) are presented in Table 8.

TABLE 8. Demographic distribution and contribution to inequality by individual groups
Case c = 0 and Θ = 0.55

	1973-74		1980-81			1973-74		1980-81	
CCAA	n^k	π^k	n^k	π^k	MUN	n^k	π^k	n^k	π^k
Andalucía	17.2	1.18	17.2	1.15	< 2.000	11.3	1.16	10.3	1.14
Aragón	3.3	1.23	3.2	0.98	2-10.000	20.2	1.02	19.3	1.05
Asturias	3.0	0.85	3.0	1.06	10-50.000	23.2	0.92	21.7	0.93
Baleares	1.7	0.93	1.7	1.17	Capitals	45.3	0.99	48.8	0.98
Canarias	3.5	1.08	3.6	1.06					
Cantabria	1.4	0.95	1.4	0.97					
Cast.-León	7.3	1.18	6.8	1.11	EDUCATION				
C.-La Mancha	4.6	1.19	4.4	1.05	Illiterate	6.5	1.33	6.3	1.34
Cataluña	15.9	0.76	15.9	0.84	No studies	18.7	1.09	25.1	1.14
C. Valenciana	9.4	0.90	9.8	0.95	Primary	60.6	0.96	48.5	0.94
Extremadura	3.1	1.22	2.8	1.05	Secondary	4.9	0.93	6.7	0.91
Galicia	7.4	0.97	7.5	1.08	High School	3.6	0.91	4.8	0.86
Madrid	11.9	1.02	12.4	0.97	Professional	0.7	0.80	1.6	0.77
R. Murcia	2.4	0.88	2.5	1.04	Short Collg.	2.2	0.83	3.5	0.81
Navarra	1.3	0.77	1.4	0.85	University	2.8	0.97	3.5	0.88
País Vasco	5.9	0.80	5.7	0.77					
La Rioja	0.7	0.80	0.7	0.73					

A seven year period is not long enough to observe dramatic changes in the frequency distribution by Autonomous Communities and municipal size. The two *Castillas* and *Extremadura* loose some population, while *Comunidad Valenciana* and *Madrid* gain some. Also, rural Spain sees its demographic weight reduced by little more than 3 percentage points in favor of Provincial capitals. In this context, the switch observed between the categories "No studies" and "Primary School" is worrisome since it has no known explanation; however, the 1980-81 figures are closer to reality according to other statistical sources. The greater importance in the second survey of Short College degrees and University attainment levels is to be expected.

Those individual groups for which their contribution to money inequality decreases over the period, must have experienced an improvement in real inequality and/or must have received a favorable impact from the change in relative prices. The full information on this issue for c = 0 and 1, and Θ = 0.55 and 1.00 is in Tables A, B, and C in the Appendix. Table 9 presents a joint picture -including the previous results by household size- of an average for those two values of Θ in the case c = 0 for our lower bound estimate for real inequality change at situation 1 prices.

A number of comments are in order. In the first place, it seems that the approximation we have attempted is relative immune to the index number problem: except two cases -*La Rioja* and municipalities of less than 2.000 inhabitants- when there is an improvement in real inequality $|\Delta Y_L^k| < |\Delta Y_U^k|$, that is, our lower bound estimate for ΔR_1 is below the upper bound estimate for ΔR_2. Notice that when ΔR_τ is positive, under the hypothesis that the substitution bias in the statistical price indices is greater for the rich than the poor, ΔY_L^k and ΔY_U^k become upper and lower bounds for ΔR_1 and ΔR_2, respectively. In those cases we observe, as expected, that $\Delta Y_L^k > \Delta Y_U^k$. At any rate, the difference between these two quantities is rarely greater than 2-3 percentage points, although such gap represents more than 20 % of the lower bound for 5 Autonomous Communities, one municipal size, and one educational group.

In the second place, there is an impressive evidence in favor of an improvement in real inequality across all partitions: only 3 -*Baleares, Asturias* and *Murcia*- out of the 29 new groups register an increase in real inequality. Taking from here on the data for the case c = 0, we observe that improvements in real inequality for a sizable part of the population are above 20 %, relative to inequality in situation 1: 1-, 2-, and 6-person households (about 26 % of the population), 4 Autonomous Communities -*Aragón, Extremadura, Castilla-La Mancha* and *País Vasco*- representing more than 16 %, and all households whose head has attained a High

TABLE 9 . Evolution of inequality for different partitions. Case c = 0

1973-74 : Π^k		$\% \Delta Y_L$		1980-81: Π^k		
2.05	1 person	30	Aragón			
1.46	2 person		Profess.	1.96	1 person	(1, 1)
1.33	Illiterate	24	6 person	1.38	2 person	(2, 2)
1.23	Aragón		Extremadura			
1.22	Extremadura	23	H. School		Illiterate	(3, 3)
1.19	C. La Mancha	22	2 person	1.17	Baleares	(4, 21
1.18	Castilla-León		University	1.15	Andalucía	(5, 8)
	Andalucía	21	1 person	1.14	< 2.000	(6, 9)
1.16	< 2.000		Sch. Collg.		No studies	(7, 1
------	------	20	C. La Mancha	------	------	------
1.09	No studies		País Vasco	1.11	Cast.-León	(8, 7)
1.08	Canarias	------	------	1.09	7 or +	(9, 12)
		18.5	5 person	1.08	Galicia	(10, 16
1.04	7 or +		Madrid	1.06	Asturias	(11, 29
1.02	Madrid		La Rioja		Canarias	(12, 11
	2.000-10.000					
------	------	18	Secondary	1.05	Extremadura	(13, 5)
0.99	Prov. capitals	16.5	Cast.-León		C-La Mancha	(14, 6
		16	Primary		2.000-10.000	(15, 13
0.97	Galicia	15.7	4 person	1.04	R. Murcia	(16, 2
	University	14.8	7 or +	1.03	3 person	(17, 24
0.96	Primary S.	14	Prov. Cap.	0.98	Aragón	(18, 4)
0.95	Cantabria	13	Canarias		Prov. Cap.	(19, 15
0.93	6 person	11.5	Andalucía	0.97	Cantabria	(20, 19
	Baleares		< 2.000		Madrid	(21, 13
	Secondary		10.000-50.000	0.95	C. Valenciana	(22, 27
0.92	3 person	------	------	0.94	Primary	(23, 18
		10	Illiterate 0.93		10.000-50.000	(24, 25
	10.000-50.000	8.3	3 person	0.91	Secdry..	(25, 2
0.91	5 person	8	No studies	0.90	5 person	(26, 23
	High School					
0.90	C. Valenciana	7	Navarra	------	------	------
------	------	6.5	2.000-10.000	0.88	University	(27,17
0.88	R. Murcia	4.5	Cantabria	0.86	High Sch.	(28, 2
0.85	Asturias		Valencia	0.85	4 person	(29, 3
0.83	4 person	4	Cataluña		Navarra	(30, 3
	Short Collg.	1	Galicia	0.84	6 person	(31, 2
0.80	País Vasco	------	------		Cataluña	(32, 3
	La Rioja	+ 6	R. Murcia	0.81	Sch. Collg.	(33, 31
	Professional	+ 11	Asturias	0.77	País Vasco	(34, 32
0.77	Navarra	+ 15	Baleares		Profes.	(35, 34
0.76	Cataluña			0.73	La Rioja	(36, 3

School degree or more, which are more than 10 %. At the other extreme, a number of groups had improvements in real inequality below 10 %: 3-person households (15 %), households headed by an illiterate person or without formal studies (24-30 %), small municipalities of 2.000 to 10.000 inhabitants -but with the rest of rural Spain not doing much better- some richer Autonomous Communities -*Navarra, Cataluña* and *Comunidad*

Valenciana- and some not so rich ones -*Cantabria, Galicia, Murcia* and *Asturias.*

In the third place, the evidence in favor of the redistribution induced by changes in relative prices is conclusive: for all specifications of c and Θ, all price effects are positive. Furthermore, as we saw for household size, the range of variation of, say, ΔP^{k}_{21} is small for all partitions: *Murcia*, households headed by an illiterate or a person without studies plus *Canarias*, experience only a 3-3.5 % price effect, while *Baleares, Galicia* and *Cataluña* -which did not do particularly well in real inequality- had an effect close to 7 %. However, the vast majority of the population -in fact, the whole of it when classified by municipality size- exhibits a positive price effect in the interval 4.3-5.6 %.

What about money inequality? Notice that a positive price effect, which reflects a pro-poor change in relative prices, when added up to an already positive change in real inequality -as in the case of *Baleares* or *Asturias*- leads to an alarmingly large deterioration in money inequality, while added up to groups with slight reductions in real inequality may lead up to a misleading increase in money inequality, as in the case of *Cataluña, Comunidad Valenciana,* illiterates or municipalities with 2.000-10.000 inhabitants.

Our last point is that if price effects are relatively neutral across groups, then the variability in money inequality is essentially due to the variability in real inequality. The transition from the first to the second list of π^{k}'s in Table 9 captures the evolution in money inequality. The last column in that Table informs about the rankings occupied by each group in terms of π^{k} in 1980-81 and 1973-74, respectively. Some groups who were doing badly at the beginning of this period, end up appearing 9 or more positions below at the end of it, like *Aragón, Extremadura* or *Castilla-La Mancha*. Others who were doing fine end up doing much better, like households headed by a person with a University degree or 6-person ones, or, on the contrary, end up much higher in the ranking by π^{k}: 3-person households, *Baleares, Asturias,* and *Murcia*. Finally, some high in the ranking in situation 1, like 1- and 2-person households, remain equally high but with a much reduced contribution to overall inequality above their demographic weight.

8. Recall that only when we use the mean logarithmic deviation, the within-group component of total inequality for the partition by household size is independent of the generosity of the scale. That is,

$$I_0(z(\Theta)) = \Sigma_j (n^j)\, I_0(z^j) + I_0(\mu(z^j(\Theta)) = W^{(j)}_0 + B^{(j)}_0(\Theta).$$

For any other partition by characteristic k, we would have

$$I_0(z(\Theta)) = \Sigma_k (n^k) I_0(z^k(\Theta)) + I_0(\mu(z^k(\Theta)) = W_0^{(k)}(\Theta) + B_0^{(k)}(\Theta).$$

Therefore, to understand the role of characteristic k = 1,...,K in explaining overall inequality, in the previous section we suggested the partition of the term $W_0^{(j)}$ into

$$W_0^{(j)} = \Sigma_{jk} (n^{jk}) I_0(z^{jk}) + \Sigma_j (n^j) I_0(\mu(z^{j1}),..., \mu(z^{jK})) = W^{(jk)} + B_0^{(k\to j)}$$

where $B_0^{(k\to j)}$ was called the true (independent of Θ) between-group component.

Thus, the expression $[B_0^{(k\to j)}/W_0^{(j)}]100$ would provide a Θ-independent measure of the importance of such between-group component. The evidence is as follows:

	CCAA	MUN	EDC
1973-74	12.3	12.1	24.9
1980-81	8.5	9.1	25.2

We observe that the explanatory power of the socioeconomic variable EDC is twice as large as the geographic characteristics CCAA and MUN which, in any case, appears to be decreasing during this period, as reported also by Bosch *et al.* (1989) using the measure $B_0^{(k)}$, which depends on Θ. To compare these two statistics we will use the expressions

$$m^{(k\to j)}(\Theta) = 100\, [B_0^{(k\to j)}/I_0(z(\Theta)]$$

and

$$m^k(\Theta) = 100\, [B_0^{(k)}(\Theta)/I_0(z(\Theta)].$$

The evidence is as follows:

	1973-74				1980-81			
$\Theta =$	0.00	0.55	1.00	EQ	0.00	0.55	1.00	EQ
CCAA								
$m^{(k\to j)}(\Theta)$	-	12.2	11.3	11.8	-	8.4	7.4	7.8
$m^{(k)}(\Theta)$	8.1	11.2	11.8	12.0	5.9	8.0	7.9	8.1
MUN								
$m^{(k\to j)}(\Theta)$	-	12.0	11.1	12.0	-	9.0	8.0	8.7
$m^{(k)}(\Theta)$	9.2	11.5	11.3	12.1	6.8	8.8	8.2	8.9
EDC								
$m^{(k\to j)}(\Theta)$	-	24.6	22.7	24.9	-	25.0	22.2	24.8
$m^{(k)}(\Theta)$	21.0	24.2	24.2	24.4	20.9	25.1	22.5	25.8

We observe only a slight difference between the two. Since

$$B_0^{(k\to j)} = B^{(k)}(\Theta) + [B_0^{(j\to k)}(\Theta) - B^{(j)}(\Theta)],$$

it is clear that the expression in brackets is close to zero for all Θ, that is, the distributions where each person receives the mean of her household size in her Autonomous Community, say, or the mean of her household size in the country as a whole, exhibit very similar inequality.

So far, we have examined a cross-section at a time. We are also interested in the explanatory role of the between-group component in the evolution of money or real inequality. Let us denote by $\Delta_M W_0^{(j)}$ the change in the within-group component of overall money inequality for the partition by household size. Using the corresponding notation, the decomposition of this term into a real and a price effect will be

$$\Delta_M W_0^{(j)} = \Delta_{R_1} W_0^{(j)} + \Delta_{P_{21}} W_0^{(j)}.$$

For any other partition by characteristic k, we will have

$$\Delta_M W_0^{(jk)} + \Delta_M B_0^{(k\to j)} = \Delta_{R_1} W_0^{(jk)} + \Delta_{R_1} B_0^{(k\to j)} + \Delta_{P_{21}} W_0^{(jk)} + \Delta_{P_{21}} B_0^{(k\to j)}$$

and $$\Delta_M B_0^{(k\to j)} = \Delta_{R_1} B_0^{(k\to j)} + \Delta_{P_{21}} B_0^{(k\to j)}.$$

This last expression provides a Θ–independent decomposition of the change in money inequality between groups into a real and a price effect. Relative to $W_0^{(j)}$ in situation 1, we have in percentage terms:

$$[\Delta_M B_0^{(k\to j)}/W_0^{(j)}]100 = [\Delta_{R_1} B_0^{(k\to j)}/W_0^{(j)}]100 + [\Delta_{P_{21}} B_0^{(k\to j)}/W_0^{(j)}]100$$

CCAA	- 4.8	=	- 5.2	+	0.4
MUN	- 4.1	=	- 4.7	+	0.6
EDC	- 2.6	=	- 4.6	+	2.0

Once more, we observe that changes in money inequality are not a good indicator of changes in real terms: the improvement in real inequality between groups is of a similar order of magnitude for the three partitions; however, the pro-poor effect of prices has a noticeable diminishing impact on between-group inequality only for EDC.

IV. CONCLUSIONS

Any welfare or inequality comparison of a pair of income or expenditure distributions at different points in time and/or space, requires a solution to two well known problems. The first is how to express the two distributions in comparable money units. The second problem arises because we use information on household total expenditure as the best proxy for a household's standard of living, but we choose the person as the unit of analysis. The question is how to treat the heterogeneity of a population consisting of persons who belong to households of different demographic composition and, hence, of different needs.

The paper´s aim is the evolution of relative inequality in Spain, making use of two large household budget surveys collected in 1973-74 and 1980-81. The decomposition of the change in money inequality into a real and a price effect occupies the center of the analysis. Inequality is measured with the help of several members of the General Entropy family of additively decomposable indices of relative inequality.

We do not work with an explicit behavioral model, in which one could conceivably estimate income, price and demographic effects, so as to construct true cost-of-living indices and equivalence scales in order to solve the above mentioned problems. Instead, we have attempted an approximation exercise which does not rely on too specific assumptions on individual preferences and does not require its recovery by means of expensive econometric procedures.

For the first problem, we use statistical Laspeyres type price indices which are household specific. The two available distributions are expressed in common money units at the dates the samples were taken, as well as at other years within and outside this period. Under the assumption that substitution bias in statistical price indices are greater for the rich than for the poor, we estimate bounds for the change in real inequality at 1973-74 and Winter of 1981 prices.

For the second problem, we take solely into account household size, and study how measured inequality varies with a parameter Θ which captures the weight one is prepared to give to household size in the definition of equivalent expenditure per person. We analyze the robustness of our conclusions on the evolution of inequality at the national level with changes in Θ for several members of the General Entropy family, and we contrast this approach with a study of the partition by household size where each group consists of households which are presumed to have identical needs.

We find useful to concentrate most of the attention to the mean logarithmic deviation, since only for this member of the General Entropy family results about the within-group component of total inequality in the partition by household size are not contaminated by an inappropriate specification of the generosity of the equivalence scale. This opens up the way for a treatment of other household characteristics in the explanation of overall inequality through their between-group component in the corresponding decomposition.

A summary of empirical results should illustrate the usefulness of this exercise in descriptive statistics.

1. In agreement with Coulter *et al.* (1992b)'s results for the U.K., we find that as we give more weight to household size, inequality in both surveys first declines and then increases until we reach the per capita expenditure distribution. Excluding the polar case in which size is given no weight, the range of variation for different specifications of the parameter c, which identifies members of the General Entropy family, is 5-15 % for the 1973-74 survey and 8-13 % for the 1980-81 one. By appropriately choosing parameters c and Θ, inequality for a single cross-section can vary as much as 50 %.

As we knew from previous studies, money inequality at the national level has improved for all specifications of c and Θ. However, such an improvement decreases continuously as we give more weight to household size, an effect more pronounced for high values of c, that is, for members of the General Entropy family more sensitive to inequality at the top of the distribution. Relative to inequality in situation 1, the maximum and minimum estimates of the improvement in money inequality are 7 % and 22.5 %, respectively.

2. For all values of c and Θ, our lower bound estimate for the change in real inequality in situation 1, ΔY_L, is always smaller than our upper bound estimate of that change at situation 2 prices, ΔY_U. This convenient finding is confirmed for 35 groups out of a total of 36 arising from the four partitions studied. On the other hand, our estimates for 1978 and 1984 are always contained in the interval $[\Delta Y_L , \Delta Y_U]$. The size of such interval is 10-20 % of ΔY_L, or 15-35 % of the change in money inequality, depending on Θ. All of which suggests that our approximation to the "true" change in real inequality might be appropriate.

3. The central finding is that the improvement in real inequality is always greater than the improvement in money inequality: when the latter achieves its minimum at 7 %, the former is 12.6-15 %; at the maximum of 22.5 % for money inequality, the change in real inequality is 28.0-31.8 %. For the important case of the mean logarithmic deviation when c = 0, the relationship is 12 % *versus* 7.6-19.0 % at an intermediate value for Θ.

4. Here, as in many other studies, we find that as we go from household total expenditure per household to equivalent expenditure per person, relative inequality decreases. We provide an explanation in terms of the movement from total to equivalent expenditure, on the one hand, and from the household to the person, on the other hand. Both effects reduce measured inequality, but the first is stronger than the second.

5. There is no doubt that changes in relative prices from 1973-74 to Winter of 1981 in Spain have been less damaging to the poor than to the rich. When c = 0, as Θ varies in the interval [0, 1] money inequality improves by 15-7 % during this period. The proportion of this change explained by the price effect varies with Θ from about 30% to almost 100 %. Having computed also real and price effects at 1978 and 1984, we observe that during the intermediate period, 1978 to Winter 1981, the impact of price changes is greater than during the two larger periods, 1973-74 to 1978 and Winter 1981 to 1984.

6. The partition by household size leads to a fundamental result independent of the generosity of the scale: except for 3-person households according to a measure more sensitive to inequality at the bottom of the distribution, money inequality improves for all groups, but less so than real inequality. At the top of the list, 1-, 2-, and 6-person households, which represent about 26 % of the population, experience improvements in real inequality of more than 20 % relative to inequality in situation 1. However, the size of price effects are in the range of 5-9 % for all groups.

7. When we partition the population by other household characteristics, we find that for c = 1 or 0 and several values of Θ, only 3 out of 29 new groups experience a decrease in real inequality. At the same time, price effects always induce a pro-poor redistribution for all specifications. Furthermore, for the majority of the population price effects, from the point of view of situation 2, for example, represent 4.3-5.6 % of inequality in 1973-74. Therefore, the variability observed in money inequality is essentially due to the variability in real inequality.

8. To evaluate the importance of inequality between groups of different partitions, we propose a measure independent of Θ. In both surveys, we find that the explanatory power of overall inequality provided by the educational level attained by the household head, is about twice as large as the one provided by the Autonomous Community or the size of the municipality where one lives: the between-group component in the case of the socioeconomic characteristic explains 24 % of total inequality, while the geographic factors explain only 8.5-12.3 % of the total. We observe also that there is an improvement in real inequality between groups of a similar order of magnitude for the three partitions, but the pro-poor price effect is only noticeable for the education variable.

La Rioja	$\Theta = 0.55$	-21.2	=	-26.4	+	5.2	=	-30.5	+	9.3
	$\Theta = 1.00$	- 7.9	=	-14.7	+	6.8	=	-12.9	+	5.0

$c = 0$

Andalucía	$\Theta = 0.55$	-10.7	=	-15.3	+	4.6	=	-17.3	+	6.6
	$\Theta = 1.00$	- 4.5	=	- 8.8	+	4.3	=	-10.5	+	6.0
Aragón	$\Theta = 0.55$	-27.3	=	-33.2	+	5.9	=	-36.6	+	9.3
	$\Theta = 1.00$	-20.1	=	-26.2	+	6.1	=	-29.3	+	9.2
Asturias	$\Theta = 0.55$	13.4	=	8.1	+	5.3	=	3.1	+	10.4
	$\Theta = 1.00$	21.1	=	14.6	+	5.5	=	10.2	+	9.9
Baleares	$\Theta = 0.55$	15.4	=	9.0	+	6.4	=	6.1	+	9.3
	$\Theta = 1.00$	29.9	=	23.2	+	6.7	=	21.1	+	8.8
Canarias	$\Theta = 0.55$	-10.9	=	-14.3	+	3.5	=	-15.9	+	5.0
	$\Theta = 1.00$	- 9.1	=	-13.0	+	3.8	=	-13.5	+	4.3
Cantabria	$\Theta = 0.55$	- 6.4	=	-11.3	+	4.9	=	-15.0	+	8.6
	$\Theta = 1.00$	6.9	=	2.0	+	4.9	=	- 1.2	+	8.1
Castilla-León	$\Theta = 0.55$	-14.5	=	-19.5	+	4.9	=	-22.3	+	7.8
	$\Theta = 1.00$	- 9.8	=	-14.3	+	4.5	=	-17.2	+	7.4
C.-La Mancha	$\Theta = 0.55$	-19.2	=	-24.4	+	5.2	=	-24.4	+	5.2
	$\Theta = 1.00$	-10.5	=	-15.3	+	4.9	=	-15.9	+	5.4
Cataluña	$\Theta = 0.55$	1.1	=	- 6.2	+	7.3	=	- 7.7	+	8.8
	$\Theta = 1.00$	4.7	=	- 2.2	+	6.9	=	- 3.7	+	8.4
C. Valenciana	$\Theta = 0.55$	- 3.1	=	- 9.3	+	6.2	=	-10.8	+	7.7
	$\Theta = 1.00$	6.4	=	0.3	+	6.1	=	- 1.2	+	7.6
Extremadura	$\Theta = 0.55$	-21.4	=	-25.7	+	4.2	=	-28.3	+	6.8
	$\Theta = 1.00$	-19.36	=	-22.9	+	3.6	=	-26.0	+	6.7
Galicia	$\Theta = 0.55$	1.3	=	- 6.0	+	7.3	=	- 9.3	+	10.6
	$\Theta = 1.00$	10.0	=	3.2	+	6.8	=	0.0	+	10.0
Madrid	$\Theta = 0.55$	-13.4	=	-19.2	+	5.8	=	-22.4	+	9.0
	$\Theta = 1.00$	-12.8	=	-18.1	+	5.4	=	-21.2	+	8.4
R. Murcia	$\Theta = 0.55$	6.9	=	3.4	+	3.5	=	- 0.4	+	7.3
	$\Theta = 1.00$	10.9	=	8.3	+	2.6	=	4.2	+	6.7
Navarra	$\Theta = 0.55$	- 0.8	=	- 5.8	+	6.6	=	- 7.1	+	7.9
	$\Theta = 1.00$	- 2.0	=	- 7.9	+	5.9	=	- 9.6	+	7.5
País Vasco	$\Theta = 0.55$	-16.0	=	-21.1	+	5.1	=	-24.2	+	8.2
	$\Theta = 1.00$	-13.9	=	-18.7	+	4.8	=	-22.0	+	8.1
La Rioja	$\Theta = 0.55$	-16.5	=	-22.1	+	5.7	=	-22.7	+	6.2
	$\Theta = 1.00$	- 9.1	=	-15.3	+	6.1	=	-14.0	+	4.8

APPENDIX

TABLE A. Change in money and real inequality within Autonomous Communities

	$c = 1$	ΔM^k	=	ΔY^k_L	+	ΔP^k_{21}	=	ΔY^k_U	+	ΔP^k_{12}
Andalucía	$\Theta = 0.55$	-13.4	=	-18.2	+	4.8	=	-20.5	+	7.1
	$\Theta = 1.00$	- 6.9	=	-11.5	+	4.6	=	-13.5	+	6.6
Aragón	$\Theta = 0.55$	-26.6	=	-32.8	+	6.2	=	-36.1	+	9.4
	$\Theta = 1.00$	-18.2	=	-24.8	+	6.6	=	-27.6	+	9.4
Asturias	$\Theta = 0.55$	6.4	=	1.3	+	5.1	=	- 3.6	+	10.0
	$\Theta = 1.00$	13.6	=	8.4	+	5.2	=	- 3.8	+	9.8
Baleares	$\Theta = 0.55$	20.9	=	13.2	+	7.7	=	9.4	+	11.4
	$\Theta = 1.00$	38.9	=	30.5	+	8.4	=	28.6	+	10.3
Canarias	$\Theta = 0.55$	- 8.0	=	-12.4	+	4.4	=	-13.3	+	5.3
	$\Theta = 1.00$	- 8.4	=	-12.8	+	4.4	=	-13.3	+	4.9
Cantabria	$\Theta = 0.55$	-18.8	=	-22.8	+	3.9	=	-25.8	+	6.9
	$\Theta = 1.00$	- 7.3	=	-11.5	+	4.2	=	-13.9	+	6.6
Castilla-León	$\Theta = 0.55$	-11.1	=	-16.2	+	5.1	=	-19.4	+	8.3
	$\Theta = 1.00$	- 3.7	=	- 8.5	+	5.0	=	-11.3	+	7.6
C.-La Mancha	$\Theta = 0.55$	-18.2	=	-23.9	+	5.7	=	-24.1	+	5.8
	$\Theta = 1.00$	- 7.8	=	-13.0	+	5.2	=	-14.1	+	6.3
Cataluña	$\Theta = 0.55$	- 2.6	=	- 9.7	+	7.1	=	-11.4	+	8.8
	$\Theta = 1.00$	1.3	=	- 5.5	+	6.8	=	- 7.3	+	8.6
C. Valenciana	$\Theta = 0.55$	- 3.9	=	-10.3	+	6.4	=	-11.2	+	7.3
	$\Theta = 1.00$	3.7	=	- 2.4	+	6.0	=	- 3.6	+	7.2
Extremadura	$\Theta = 0.55$	-24.6	=	-28.7	+	4.1	=	-31.9	+	7.3
	$\Theta = 1.00$	-24.6	=	-28.0	+	3.4	=	-32.7	+	8.1
Galicia	$\Theta = 0.55$	1.9	=	- 5.5	+	7.4	=	- 8.9	+	10.8
	$\Theta = 1.00$	9.8	=	3.1	+	6.7	=	- 0.5	+	10.3
Madrid	$\Theta = 0.55$	-21.2	=	-26.4	+	5.2	=	-30.5	+	9.3
	$\Theta = 1.00$	-19.7	=	-24.5	+	4.9	=	-28.5	+	8.9
R. Murcia	$\Theta = 0.55$	1.3	=	- 1.8	+	3.1	=	- 6.3	+	7.5
	$\Theta = 1.00$	4.4	=	2.4	+	2.0	=	- 2.5	+	6.9
Navarra	$\Theta = 0.55$	-12.2	=	-18.0	+	5.6	=	-18.1	+	5.9
	$\Theta = 1.00$	-23.6	=	-28.0	+	4.5	=	-28.3	+	4.7
País Vasco	$\Theta = 0.55$	-18.3	=	-23.3	+	5.0	=	-26.7	+	8.4
	$\Theta = 1.00$	-14.3	=	-19.1	+	4.8	=	-22.6	+	8.3

TABLE B. Change in money and real inequality within municipalities by size

	$c = 1$	ΔM^k	=	ΔY^k_L	+	ΔP^k_{21}	=	ΔY^k_U	+	ΔP^k_{12})
< 2.000	$\Theta = 0.55$	- 7.5	=	-12.5	+	5.0	=	-11.1	+	3.6
	$\Theta = 1.00$	2.0	=	- 3.1	+	5.1	=	- 1.7	+	3.8
2.000-10.000	$\Theta = 0.55$	- 6.7	=	-11.9	+	5.2	=	-14.4	+	7.7
	$\Theta = 1.00$	1.6	=	- 3.2	+	4.8	=	- 5.8	+	7.4
10.000-50.000	$\Theta = 0.55$	- 9.7	=	-15.2	+	5.5	=	-17.5	+	7.8
	$\Theta = 1.00$	- 6.1	=	-11.1	+	5.0	=	-13.4	+	7.2
Prov. capitals	$\Theta = 0.55$	-14.4	=	-19.9	+	5.5	=	-23.0	+	8.5
	$\Theta = 1.00$	-11.0	=	-16.2	+	5.2	=	-19.2	+	8.1
	$c = 0$									
< 2.000	$\Theta = 0.55$	- 11.1	=	-15.7	+	4.6	=	-14.6	+	3.5
	$\Theta = 1.00$	- 5.0	=	- 9.3	+	4.3	=	- 8.6	+	3.5
2.000-10.000	$\Theta = 0.55$	- 6.6	=	-11.5	+	4.9	=	-14.0	+	7.4
	$\Theta = 1.00$	2.7	=	- 2.0	+	4.6	=	- 4.3	+	7.0
10.000-50.000	$\Theta = 0.55$	- 8.4	=	-13.6	+	5.2	=	-15.9	+	7.5
	$\Theta = 1.00$	- 4.5	=	- 9.3	+	4.8	=	-11.4	+	7.0
Prov. capitals	$\Theta = 0.55$	-10.2	=	-15.8	+	5.6	=	-18.5	+	8.3
	$\Theta = 1.00$	- 7.1	=	-12.4	+	5.3	=	-14.9	+	7.8

TABLE C. Change in money and real inequality for household head's educational levels

c = 1		ΔM^k	=	ΔY^k_L	+	ΔP^k_{21}	=	ΔY^k_U	+	ΔP^k_{12}
Illiterates	Θ = 0.55	-13.4	=	-16.7	+	3.3	=	-18.6	+	5.2
	Θ = 1.00	- 0.2	=	- 3.1	+	2.9	=	- 3.8	+	3.6
No studies	Θ = 0.55	- 8.8	=	-12.6	+	3.8	=	-14.4	+	5.5
	Θ = 1.00	- 2.1	=	- 5.6	+	3.5	=	- 7.2	+	5.1
Primary School	Θ = 0.55	-14.6	=	-19.6	+	5.0	=	-22.2	+	7.7
	Θ = 1.00	- 9.6	=	-14.3	+	4.7	=	-16.6	+	7.1
Secondary School	Θ = 0.55	-14.8	=	-19.9	+	5.1	=	-22.7	+	7.9
	Θ = 1.00	-10.4	=	-15.5	+	5.1	=	-18.1	+	7.7
High School	Θ = 0.55	-22.8	=	-28.0	+	5.2	=	-30.3	+	7.4
	Θ = 1.00	-25.2	=	-29.1	+	3.8	=	-31.8	+	6.5
Professional	Θ = 0.55	-22.2	=	-27.8	+	5.6	=	-27.9	+	5.7
	Θ = 1.00	-31.0	=	-36.5	+	5.5	=	-36.8	+	5.8
Short College	Θ = 0.55	-18.0	=	-22.6	+	4.6	=	-26.3	+	8.3
degree	Θ = 1.00	-21.8	=	-25.7	+	3.9	=	-29.4	+	7.7
University	Θ = 0.55	-18.3	=	-22.8	+	4.5	=	-25.7	+	7.5
	Θ = 1.00	-17.3	=	-21.1	+	3.8	=	-25.5	+	8.2
c = 0										
Illiterates	Θ = 0.55	-12.1	=	-15.1	+	3.0	=	-17.3	+	5.2
	Θ = 1.00	- 0.2	=	- 2.8	+	3.0	=	- 3.5	+	3.7
No studies	Θ = 0.55	- 9.0	=	-12.7	+	3.7	=	-14.9	+	5.9
	Θ = 1.00	- 3.0	=	- 6.4	+	3.4	=	- 8.3	+	5.4
Primary School	Θ = 0.55	-14.6	=	-19.4	+	4.8	=	-22.0	+	7.4
	Θ = 1.00	-10.5	=	-14.9	+	4.5	=	-17.4	+	6.9
Secondary Sch.	Θ = 0.55	-14.6	=	-19.5	+	4.9	=	-22.8	+	8.2
	Θ = 1.00	-10.9	=	-15.5	+	4.6	=	-18.7	+	7.8
High School	Θ = 0.55	-17.8	=	-22.9	+	5.1	=	-25.1	+	7.4
	Θ = 1.00	-19.5	=	-23.5	+	4.1	=	-25.9	+	6.4
Professional	Θ = 0.55	-16.5	=	-22.9	+	6.4	=	-23.5	+	7.0
	Θ = 1.00	-25.8	=	-31.9	+	6.0	=	-32.8	+	6.9
Short College	Θ = 0.55	-15.3	=	-20.2	+	4.9	=	-23.6	+	8.4
degree	Θ = 1.00	-17.8	=	-22.0	+	4.2	=	-25.3	+	7.5
University	Θ = 0.55	-21.0	=	-25.1	+	4.1	=	-28.5	+	7.6
	Θ = 1.00	-17.0	=	-20.5	+	3.5	=	-25.0	+	8.0

utility, permit the identification of the so-called general translog household equivalence scales. In Jorgenson and Slesnick (1984b) measures of absolute and relative inequality are then derived from a social welfare function defined over the distribution of price-dependent indirect utility functions. These are estimated with 1958-78 data for the U.S.; the differences between any two years in these series give an estimate of the corresponding change in money inequality. In Slesnick (1990), two time series of relative inequality measures were estimated with U.S. data for the period 1947-85: one dependent on prices, whose rate of change provides again a measure of the evolution of money inequality; and another series estimated at constant prices with 1947 as the reference time period, whose rate of change provide a measure of the evolution of real inequality. Of course, the difference between the two series for each given year t provide an estimate of the impact on inequality of the change in prices from 1947 to the year in question.

Also relevant for our topic, in Muellbauer (1974c) money expenditure distributions for 1964 and 1970 in the U.K. were simply adjusted for differences in household size by means of some equivalence scales borrowed from Stark (1972). Then, based on some estimates of the Linear Expenditure System of demand equations for 9 commodity groups, true cost-of-living indices were used to express the 1970 money expenditure distribution for adult equivalents into 1964 prices in order to compute the change in real inequality between these two dates. Muellbauer (1974a) contains an estimate of the impact of price changes on inequality during the period 1970-72.

(12) On this issue, see Blundel and Lewbel (1991)

(13) See the discussion in Coulter *at al.* (1992a).

(14) See Barnes and Gillingham (1984) and Nicol (1989).

(15) Ray (1985) found that both the inequality estimates for childless couples and for couples with one child and, more disturbingly, their evolution over time, were highly sensitive to the two demand systems used.

(16) This issue has been stressed by Browning (1991) and by Coulter *et al.* (1992a).

(17) Coulter *et al.* (1992a) mention also the "Sequential Dominance" approach suggested by Atkinson and Bourguignon (1987) which requires as a necessary condition that relative numbers of each different household type are the same in the distributions to be compared, a feature not present in our data.

(18) On this issue, see also the discussion in Atkinson (1990).

NOTES

(1) For early studies on India, see Iyengar (1967) and Mahalanobis (1972); for U.K. studies, see Prais (1959), Nicholson (1975), Lesser (1976) and the references quoted in Muellbauer (1974c); for the U.S., see Michael (1970), Hollister and Palmer (1972), and Hagemann (1982); and for Spain, see Abadía (1986a).

(2) Alternatively, with the help of true cost-of-living indices of the Laspeyres type, $L^h(p_t, p_\tau; u^h_\tau)$, we can reprice the original total money expenditure distributions as follows

$$x^h_\tau L^h(p_t, p_\tau; u^h_\tau) = c^h(u^h_\tau, p_\tau)\,[c^h(u^h_\tau, p_t)/c^h(u^h_\tau, p_\tau)] = c^h(u^h_\tau, p_t) = x^h_{\tau t}.$$

(3) See Muellbauer (1974b).

(4) See, for example, Kuznets (1976), Danziger and Taussing (1979), Johnson and Webb (1979), Datta and Meerman (1980), Cowell (1984), and Ruiz-Castillo (1987).

(5) See Haddad and Kambur (1990) and the references quoted there.

(6) The same argument goes through for any other class of mean-invariant inequality measures by appropriately modifying the definition of the equivalence scale.

(7) See, for example, Foster (1985). Notice that accepting the transfer principle when analyzing distributions of equivalent income implies the acceptance of a value judgement which has been questioned very forcefully by Cowell (1980), Jenkins and O'Higgins (1989), and Glewwe (1991). For a useful discussion, see Coulter *et al.* (1992a).

(8) See Shorroks (1980).

(9) For a discussion see Deaton and Muellbauer (1980) and Ruiz-Castillo (1991), and Coulter *et al.* (1992a).

(10) Good recent examples are provided by Blackorby and Donaldson (19??) and Lewbel (1989), which explore the implications of assuming that equivalence scales are independent of utility levels.

(11) A good example for our purposes is provided by the work of Jorgenson and Slesnick (1983, 1984a, 1987), where the assumption of independence of the scales on the utility levels, combined with conditions for exact aggregation in the context of the translog model of indirect

REFERENCES

Abadía, A. (1986a), "Composiciones de demandas, precios relativos y variaciones de capacidad de compra bajo indiciación de rentas," *Investigaciones Económicas*, : 69-81.

Abadía, A. (1986b), "Inflation, Relative Prices and Welfare Redistribution in Spain, 1976-1984," *Economic Letters*, **20**: 387-390.

Abadía, A. (1987), "Indice de Precios de Consumo, Coste de Vida y distribución del bienestar: 1976-1985," *Investigaciones Económicas*, **11**: 179-190.

Atkinson, A. (1970), "On the Measurement of Inequality," *Journal of Economic Theory*, **2**: 244-263.

Atkinson, A. B. (1990), "Comparing Poverty Rates Internationally: Lessons from Recent Studies in OCDE Countries," London School of Economics, The Welfare State Programme, Discussion Paper WSP/53.

Atkinson, A. B. and F. Bourguignon (1987), "Income Distribution and Differences in Need," in G. R. Feiwell (ed), *Arrow and the Foundations of the Theory of Economic Policy* , London: Macmillan.

Banks, J., R. Blundell and I. Prescot (1991), "Lyfe-Cycle Expenditure Allocations and the Consumption Costs of Children," Institute for Fiscal Studies, Working Paper Series, No. W91/12.

Barnes, R. and R. Gillingham (1984), "Demographic Effects in Demand Analysis: Estimation of the Quadratic Expenditure System Using Microdata," *Review of Economic Studies and Statistics*, **66**: 591-601.

Blundell, R. and A. Lewbell (1991), "The Information Content of Equivalence Scales," *Journal of Econometrics*, **50**: 49-68.

Bosch, A., C. Escribano and I. Sánchez (1989), *Evolución de la desigualdad y la Pobreza en España*. Madrid: Instituto Nacional de Estadística.

Browning, M. (1990), "Modeling the Effects of Children on Household Economic Behavior," MaMaster University, Working Paper Series, No. 90-121.

Buhmann, B., L. Rainwater, G. Schmauss 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**: 115-142.

(19) Ruiz-Castillo (1987).

(20) See Shorrocks (1980).

Iyengar, N. S. (1967), "Study of Differential Price Movements and Consumer Behavior: An Application of Fractile Graphical Analysis," *Indian Economic Review*, **II**, No.2.

Iyengar, N. S. and N. Bhattacharya (1965), "On the Effect of Differentials in Consumer Price Index on Measures of Inequality," *Sankhya, Series B*, **27**: 47-56.

Jenkins, S. P. (1991), "Income Inequality and Living Standards: changes in the 70's and 80's," *Fiscal Studies*, **12**: 1-28.

Jenkins, S. and M. O'Higgins (1989), "Inequality Measurement Using Norm Incomes," *Review of Income and Wealth*, **35**: 265-282.

Johnson, P. and S. Webb (1989), "Counting People witth Low Incomes," *Fiscal Studies*, **10**: 66-82.

Jorgenson, D. W. and D. Slesnick (1983), "Individual and Social Cost of Living Indexes," in W. E. Diewert and C. Montmarquette (ed), *Price Level Measurement* , Ottawa: Statistics Canada.

Jorgenson, D.W. and D.T. Slesnick (1984a), "Aggregate Consumer Behavior and the Measurement of Inequality," *Review of Economic Studies*, **51**: 369-392.

Jorgenson, D. W. and D. T. Slesnick (1984b), "Inequality in the Distribution of Individual Welfare," in R. L. Basmann and G. F. Rhodes (ed), *Advances in Econometrics, Vol. 3* , Greenwich: JAI Press.

Jorgenson, D. W. and D. T. Slesnick (1987), "Aggregate Consumer Behavior and Household Equivalence Scales," *Journal of Business and Economic Statistics*, **5**: 219-232.

Kuznets, S. (1976), "Demographic Aspects of the Size Distribution of Income: An Exploratory Survey," *Economic Development and Cultural Change*, **25**: 1-94.

Lesser, C. E. V. (1976), "Income, Household Size, and Price Changes 1953-1973," *Oxford Bulletin of Economics and Statistics*, **38**: 1-10.

Lewbel, A. (1989), "Household Equivalence Scales and Welfare Comparisons," *Journal of Public Economics*, **39**: 377-391.

Mahalanobis, P. C. (1972), "Disparities in the Level of Living," in S. L. N. Simha (ed), *Economic and Social Development* , Vora and Co.

Michael, R. T. (1979), "Variations Across Households in the Rate of Inflation," *Journal of Money, Credit, and Banking*, **11**: 32-46.

Muellbauer, J. (1974a), "Inequality Measures, Prices, and Household Composition," *Review of Economic Studies*, **41**: 493-504.

Muellbauer, J. (1974b), "Household Composition, Engel Curves, and Welfare Comparisons Between Households," *European Economic Review*, **5**: 103-122.

Muellbauer, J. (1974c), "Prices and Inequality: the United Kingdom Experience," *Economic Journal*, **84**: 32-55.

Nicholson, J. L. (1975), "Whose Cost of Living," *Journal of the Royal Statistical Society, Part 4*, **138**: 540-542.

Nicol, C. (1989), "Testing a Theory of Exact Aggregation," *Journal of Business and Economic Statistics*, **7**: 259-265.

Pollak, R. (1991), "Welfare Comparisons and Situation Comparisons," *Journal of Econometrics*, **50**: 31-48.

Pollak, R. and T. Wales (1979), "Welfare Comparisons and Equivalent Scales," *American Economic Review, Papers and Proceedings*, **69**: 216-221.

Prais, S. J. (1959), "Whose Cost of Living?," *Review of Economic Studies*, **26**(February): 126-134.

Ray, R. (1985), "Prices, Children and Inequality: Further Evidence for the United Kingdom, 1965-1982," *Economic Journal*, **95**: 1069-1077.

Ruiz-Castillo, J. (1987), "La medición de la pobreza y la desigualdad en España, 1980-81," Banco de España. Servicio de Estudios, Estudios Económicos, número 42.

Ruiz-Castillo, J. (1991), "Difficulties in the Use of Equivalence Scales for Normative Purposes," Universidad Carlos III de Madrid, Working Paper, 91-03.

Rusell, R. (1983), "Comments on "Individual and Social Cost of Living Indexes" by D.W. Jorgenson and D. Slesnick," in W. E. Diewert and C. Montmarquette (ed), *Price Level Measurement* , Ottawa: Statistics Canada.

Sen, A. (1973), *On Economic Inequality*. Oxford: Clarendon Press.

Sen, A. (1976), "Real National Income," *Review of Economic Studies*, **43**: 19-39.

Shorrocks, A. (1980), "The Class of Additively Decomposable Inequality Measurements," *Econometrica*, **48**: 613-625.

Slesnick, D. (1990), "Inflation, Relative Price Variation, and Inequality," *Journal of Econometrics*, **43**: 135-151.

Stark, T. (1972), *The Distribution of Personal Income in the UK, 1949-1963*. Cambridge: Cambridge University Press.