Distribution of Welfare Change, Inequality & Poverty Under Consumption Rationing: ***Estimates Based on a Wartime Iranian Budget Survey***

Feridoon Koohi-Kamali, 2005*

This paper deals with the measurement of welfare gains and losses due to consumption rationing. The measure employed is a first-order approximation for welfare change based on re-pricing of rationed goods at corresponding market prices; applied to an Iranian budget survey with dual price structure. The distributional comparisons of total expenditure obtained allow the identification of gainers and losers, and the results indicate rationing has led to reduced inequality, and a more notable fall in poverty. Some caveats regarding the approximations are also examined.

This paper deals with household standard of living as affected by the distribution of welfare gains and losses due to consumption rationing. Its particular aim is to provide an approximate picture of the distributional consequences of rationing. Even excluding coupon resale and/or additional market consumption of coupon goods, one would expect gains from rationing, relative to a fixed base, for some income groups. The changes in welfare relative to the mean or median total expenditure will tend to be positive for the poor, and negative for the rich, reflecting the larger weight of rationed goods in total expenditure of the poor compared to the rich[1].

This paper attempts to examine the above issues employing the 1984-85 Iranian household budget survey conducted by the Central Bank of Iran, thenceforth the 1984-5 CBIHBS, under wartime coupon rationing in urban areas. Its distinctive features are its dual fix/market price structure for food commodities and its dual expenditure structure for all commodities, including non-food[2]. In Section 1, I discuss the measurement methods employed in this paper, section 2 deals with the method and estimation of market prices of rationed goods. The main results, employing section 2 prices, assess distributional consequences of rationing; they appear in section 3. Section 4 discusses the policy relevance of the findings, while section 5

*This paper is revised and expanded version Koohi-Kamali (1992).

[1]The gains will still be larger for the poor with partial ration take-up, though relatively less so if resale is ruled out, and provided the subsidy applies to food items with low income elasticities, as was more or less the case with the Iranian coupon system.

contains the conclusion of the paper.

1 Measurement of Welfare Under Rationing

Imagine we compare the position of a household under existing prices of rationed quantities with a position in which those same quantities are consumed at market prices. This would be one way of measuring welfare change due to rationing. Such simple comparisons can be interpreted as the compensation required to keep welfare unchanged, that is, minimum expenditures required to reach the same utility level at alternative price vectors. Money metric utility, Samuelson (1974), labelling indifference curves by their minimum costs, is commonly employed for this purpose. Suppose we want to retain the utility level u^0 reached under rationing but in terms of the vector of free market prices of rationed goods p^1. The minimum expenditure to reach u^0, defined by the cost function $c(p^1, u^0)$, can be obtained by reflating the minimum cost of u^0 at the (rationed) fix-price vector p^0, that is, $p^0 q^0$, by a price index $p(p^1, p^0; u^0)$, which compares the cost of the base utility at two different price vectors. Since this price index involves unobservable u^0, one common approximation for it is the Laspeyres index, $\frac{p^1 q^0}{p^0 q^0}$, based on observable values. As it can be shown that the true index cannot exceed the Laspeyres, we have $c(p^1 u^0) \leq (\approx) p^0 q^0 * \frac{p^1 q^0}{p^0 q^0}$

One can alternatively measure the welfare change due to rationing in terms of consumer surplus and its related concept of a Hicksian Compensation Variation (*CV*), Hicks (1955). *CV* measure the difference between the minimum costs of u^0 at price vectors $p^0; P^1$, and $c(p^1, u^0)$ is obtained by adding this difference to $p^0 q^0$. Once again, a similar approximation is given by $c(p^1, u^0) \leq (\approx) p^0 q^0 + (p^1 - p^0) q^0$. With this method, *CV* is approximated by $(p^1 - p^0) q^0$. This is

[2]For the sampling features of the survey, see section 3.

also a common method for an approximate measurement of welfare gain and loss in other similar contexts, see Mohring (1971) and King (1983) for its employment in welfare assessment of alternative tax policies.

The question is how good these approximations are. Since they refer to a fixed utility level rather than a fixed quantity vector, consumers can take advantage of substitution when prices change, and still remain on the same indifference curve. However, the above approximations assume that the scope for substitution among goods in response to relative price change is small and can be ignored, Deaton (1980)[3]. When this is not so, then the Laspeyres index provides a poor approximation. Such, for example, can be the case in computation of price indices between countries when relative prices may differ substantially, thus resulting in large substitution effects in response to long-term relative price change. Perhaps a more relevant example closer to our concerns is construction of indices for comparisons of controlled and post-liberalisation prices. In such cases too, there is some evidence suggesting that Laspeyres and Paasche indices may have a poor performance. For instance, Osband (1991) believes they are biased upward, thus exaggeration inflation in a privatising economy. However, one would not expect large substitution effects in the short run in a war economy. In a such an economy, there will typically be varying degree of shortages, particularly of non-rationed goods, even when the aggregate consumption is close to voluntary levels, hence limiting the scope for substitution compared to peacetime. In the rest of this paper, I shall work with $(p^1 - p^0)q^0$, although the issue is further examined in section 4.

2 Estimation of Market Prices of Rationed Goods

The measurement of the welfare impact of rationed goods on the household standard of

---

[3]The last inequality approximation, for instance, is obtained from a Taylor series expanded around p^0, with further terms excluded. The first term among those removed contains the matrix of substitution. Its exclusion therefore amounts to assuming *separability* between the rationed and non-rationed goods by this method of approximation.

living requires re-pricing of commodities bought at fixed prices, at corresponding market prices. There is no direct market price information for commodities available from the survey employed; what is available is the expenditure for a particular quantity of a food commodity by each household, recorded separately for each item bought at fixed prices, mostly coupon goods, and at market prices, mostly non-coupon goods. Moreover, price data is available only for the food and tobacco group. For non-food commodities, only expenditure data is recorded, albeit by type of purchase. Any attempt at re-pricing of rationed non-food goods at market prices would have to make assumptions based on the available information on prices and quantities of food commodities.

The first step in the estimation of market prices for rationed goods is to aggregate the individual household fixed and market prices into a pair of corresponding prices, one for each fix-price/rationed food commodity since we require estimates not for all food commodities but only for the limited group of fix-price/rationed food commodities. There are different ways of estimating such rationed and market prices for rationed goods, but since sample data usually contains outliers, it is obviously preferable to use robust weighted prices in order to guard against them. I employ a weighted average where each price is weighted by its corresponding ratio of quantity purchased to the total quantity of the commodity in question purchased by all households. Such an aggregate pair of rationed/market prices for each commodity is employed for two different purposes.

The first purpose is to re-price, each rationed food expenditure, at its equivalent free market price. In this regard, it should be remembered that not all households with purchases of a given rationed good have market purchases of the same good. Aggregate prices obtained from all rationed purchases for rationed goods and from all market purchases for market goods allow re-pricing of rationed purchases in such cases.

The second purpose of aggregate food prices is to employ them for re-pricing of non-food rationed commodities. Since no price information is available for the latter group, I assume that such re-pricing would increase expenditure on each non-food rationed commodity by a common factor. I assume this factor to be the average overall ratio between each pair of rationed and market food prices, or each pair of rationed and market food expenditures if quantities of rationed food goods remain constant[4]. One should employ the aggregate price ratios based on weighted aggregate prices to guard against outliers. However, before summing up the weighted prices of each good to obtain this aggregate ratio, one could weight each weighted price pair by its rationed quantity to protect the constructed price ratios even more against outlying observations. I employ this double-weighted method[5].

First, I define the weighted, market or rationed, price for each good *j* averaged over households:

$$p^{w}_{ja} = \sum_{h=1}^{H} [\, p_{jha} \cdot \frac{q_{jha}}{\sum_{h=1}^{H} q_{jha}} \,]$$

where *h*=1,2,...,H stands for number of households, and *j*=1,2, ...,N for number of goods, *w* for weighted average ***annual*** prices, which can be only one of the two rationed *a*=*r* or free market *a*=*m* types.

Second, using these average quantity-weighted prices, I accordingly define a rationed quantity-weighted aggregate ratio of market prices to ration prices, thus using quantity weights twice, averaged over goods:

[4]Aggregate commodity prices for re-pricing of rationed expenditures requires only one set of estimates, namely market prices of rationed goods. However, estimating average price difference between ration and market prices needs a pair of aggregate prices for each commodity. Thus, here we also have to estimate aggregate ration prices. Variations in ration prices will, of course, be relatively small.

[5]One can alternatively construct aggregate food price ratios based on simple aggregate prices, but weigh each simple price pair by its rationed quantity. A priori, one would expect the double quantity-weighted ratio adopted here to be more robust than such a single quantity-weighted aggregate price ratio.

$$EQ = \frac{\sum_{j=1}^{N} p^{w}_{jm} \cdot q_{jr}}{\sum_{j=1}^{N} p^{w}_{jr} \cdot q_{jr}} \qquad (1)$$

(1) provides estimates of the common factor of increased expenditure on non-food rationed commodities.

Since I present separate estimates for gains and losses from the purchase of a narrow list of coupon goods, and from that of all fixed-price purchases, I first offer a definition of each before going further. In this chapter, the "Coupon Scheme" refers to the aggregate price ratio estimated for the narrow group of coupon goods with universal entitlement[6]. The aggregate ratios were also estimated for *all* fix-price consumption, inclusive of the above "Coupon Scheme" goods. It should be mentioned that fix-price consumption of non-coupon goods is not universal, but the privilege of some groups such as public sector employees. However, some poor households are also likely to be among those with a privileged access, for instance families of the war veterans, most of which are from low income groups. I shall call this broader fix-price scheme, the "General Scheme".

Table 1 shows these estimates calculated separately for each city group for the above two categories of fixed-price schemes. The data was searched for outlying observations and their influence on the method of construction adopted for prices and price ratios were checked. The *a priori* expectation regarding the robustness of (1) is fully borne out by the examination of the outcome [7]. Based on (1), table 1 suggests that market cost of rationed food commodities is about

---

[6]Its main items are rice, vegetable oil, sugar, red meat, poultry, cigarettes, for which the survey provides price and expenditure information by type; and soap and washing-powder, with only expenditure information by type.

[7]For example, a particularly striking case of an outlier for a single market purchase of sugar was found for Tehran, and aggregate price and price ratios were estimated before and after removing this observation by other methods in addition to EQ in order to assess their robustness. The result using EQ hardly changes, but this is not so for other methods, though the simple, unweighted method is the most sensitive to this outlier: its price ratio for the 'Coupon Scheme' falls from 34 to 2 after excluding the outlier.

twice its rationed cost for the general scheme, and two and half times for the coupon scheme. On this basis, the narrow group of coupon food displays a greater "shortage", or degree of subsidy, than the group of food commodities taken as a whole.

One final issue has to be looked at before we turn to the use of market prices for re-pricing of rationed goods. Lacking separate information on fixed and market non-food prices[8], I assume that market cost of rationed goods would increase by the same amount as the average aggregate difference between the rationed cost and the market cost of rationed food commodities. I have no means of assessing the reasonableness of this assumption empirically, but it must be noted that non-food rationed goods constitute about 15-20% of all fix-price expenditure under the coupon scheme, making the proposed method more acceptable for this scheme. By contrast, since expenditure on public utilities, run by the state, are of a fix-price type, they account for a large percentage of such expenditure. If effects of rationing in a more strict sense are of interest, then the above method is not likely to be very reliable when applied to the entire non-food fix-price expenditure, a point worth remembering when examining the estimated welfare change due to rationing.

We are now in possession of basic price information to carry out our re-pricing exercise. The idea is, that evidence of actual expenditure, without allowing for different prices that households would have had to pay to obtain rationed goods at market prices underestimates the standard of living. Each rationed food commodity is re-priced using a single quantity-weighted aggregate market price for that commodity. The same procedure is followed for construction of a single rationed price for the same commodity.

Expenditures on non-food rationed commodities are raised by using estimates based on (1) from table 1 below. All coupon non-food expenditures were multiplied by the resulting

[8]Separate non-food information for market and fixed prices are collected for some goods, but I know of *no* published source, which makes this available even for a commodity like petrol.

common factors for each city group given in table 1 below, that is, by 2.492, 2.346, and 2.502. All non-food fix-price purchases, inclusive of non-food coupon goods, were raised for their respective city group by 2.097, 1.907, and 2.113. The re-pricing of expenditure on the rationed goods components of total expenditure raises the measured standard of living for anyone with recorded purchases for rationed goods. However, relative to a fixed base, poorer households, which have a greater weight of rationed goods in their total expenditure than richer households, would increase their standard of living more compared to the rich, under the re-priced distribution. Re-pricing therefore reduces inequality compared to the original distribution. The extent to which this is so is a subject of section 3. However, the question of sample weights enters into this exercise and must be dealt with before distributional estimates are examined.

3 Consumption Rationing, Inequality, and Poverty

The assessment of the distributional consequences of rationing requires comparison of actual with re-priced total expenditure. We must thus decide on an appropriate definition of total expenditure in order to carry out the exercise. A common practice is to exclude the expenditure category of 'housing investment'. This category contains infrequently purchased residential and non-residential homes, and its inclusion in the non-food group will significantly distort sample averages[9]. However, the imputed market rental value of housing is included for owner-occupied, subsidised, or free housing. The inclusion of the latter will increase non-food expenditure [10]. Moreover, the massive wartime population displacement usually makes free or cheap housing a widespread phenomenon not just confined to the poor, and total expenditure should contain imputed values in this case.

[9]For instance, the inclusion will lower the average food share by about 5.5 %.

[10]Some believe this may distort the Engel prediction on food share behaviour for the poor. At the very bottom, the food share may display an increase with income, see, Bhanoji Rao (1981). This can quite plausibly be explained by an underestimation of housing cost.

The final issue is the one of population weights. One sampling feature of the 1984-85 CBIHBS is its employment of optimal stratified sampling. In this method widely different sampling rates are deliberately adopted for strata so as to minimize the variance of the overall mean, that is, large sampling rates are used for strata where the variance among the elements is large, and small sampling rates used where it is smaller. For a fixed sample size of *n*, this is achieved by setting the sampling rate of a stratum g proportional to its standard deviation S_g: $\frac{n_g}{N_g} = k\, S_g$, Cochran (1977). Compared to self-weight sampling, optimal stratified sampling results in large gains in precision when differences between S_g are large, that is, when characteristics of the population have highly skewed distributions. With skewed within stratum distributions, the application of self-weighting would lead to over-sampling parts of the population with the largest variance. In the measurement of welfare with which this chapter is concerned, this would result in over-sampling the richer parts of the population or larger households. Disproportionate sampling would be more robust in this regard. With optimal stratified sampling, each observation must be multiplied by its corresponding probability of selection defined by its population weight before aggregate and mean values can be obtained at the level of each stratum, and then added cumulatively to obtain values for the entire sample.

The evidence in some studies supports the practical importance of these weights[11]. My own estimates partially confirm these findings. Table 2 compares the reported 84-85 CBIHBS population weighted averages for food share, total expenditure and household size, with those without employing such weights obtained from the 84-85 survey of the present study. A key

[11]Deaton (1997) has given the difference, arising from the employment of ‘inflation’ weights, in mean household per capita total expenditure for Pakistan by over 15%. Thus, there seems to be enough correlation between probability weights and PCE in this case. Coulombe and Demery (1993) report a massive decline of 24% in average household size over 1985-88 for the Living Standard Measurement Study (LSMS) panel surveys of the Cote d’Ivoire mainly because of over-sampling of larger households, in turn due to over-sampling of larger dwellings, in the earlier survey years. Demery and

value is the average budget food share. This is 40.7 percent weighted and 39 percent unweighted, thus a change of only 1.7 percent is due to population weights. However, a relatively small change in the food share can result in a large change in the average total expenditure, according to table 2 of about 6 percent, that is, about three and half times bigger than the food share difference. Although this is relatively smaller than that reported by Deaton (1997) for Pakistan, its effect on the ranking of households by total expenditure is not negligible; hence, it must be taken into account. Table 2 also checks the impact of the population weights on average household size; the difference is relatively small.

The necessity of weighting observations for assessment of distributional changes gives rise to a new problem. The assessment of distribution in relation to rationing requires estimating mean total expenditure for decile expenditure classes. To estimate decile means one would ordinarily arrange all households in increasing order of total expenditure and then divide them into ten equal-sized groups in terms of number of households. However, with a random sample, the selected households must represent not the population of the particular city from which they are drawn, but also that of any sub-group of the population which may be of interest to analysis. Since the probability of selection varies with the sub-group of the population in question, selected households would have to be given different weights to account for this change in probability of selection as the underlying sub-groups of population change. To determine weighted decile expenditure class boundaries we should have ten equal-sized groups in terms of weighted households but this usually leaves unequal number of households in each decile. The way this boundary problem is resolved is as follows:

Households are arranged in increasing order of their unweighted total expenditure, net of housing investment, with corresponding weight in front of each observation. Next to the column

---

Grootaert (1993) re-weighted the raw data, and found that mean estimates of income, expenditure, and poverty are all significantly affected by this correction.

of weights, a column of cumulative weights is constructed. Since these population weights are, in fact, probabilities of being included in the sample drawn, the column of cumulative weights should approximately add up to one for each city group over which the weights have been estimated, provided no substantial errors have been made in estimations. Assuming this is the case we can mark off 0.1, or the nearest observation to it, from the column of cumulative weight and find out the corresponding total expenditure from the expenditure column. This total expenditure defines the first expenditure class boundary. Similarly, the cumulative weight 0.2 enables us to define the second expenditure class boundary, and so on. Since households are arranged in increasing order of total expenditure, the first class will contain the poorest households, the second class the next group of poor households with total expenditure higher than the first group and so on. The appendix provides more details on the population weights employed.

The cumulative weights for Tehran, whose estimation requires no more information than number of households in each sample block and the total number of such households in all selected blocks in Tehran, add up to 0.99999. For large cities where some modification had to be made using the population of large cities, see Appendix below, the total cumulative weight adds up to 0.994916. However, for small cities where the weights were the product of three ratios, using census population figures, which sometimes contain inaccuracies, errors have inevitably crept into estimation, resulting in a total cumulative weight of 0.952010. While the total cumulative weights for Tehran and large cities are reasonable approximations to one, the gap for small cities is somewhat large. For this last group of cities I have divided each individual cumulative weight by 0.952010, which ensures that the new, adjusted cumulative weights add up exactly to one[12]. Therefore, while I use cumulative weights to mark off class boundaries for Tehran and large cities, it is the adjusted cumulative weights, which were used for this purpose in

[12]This can introduce a bias but if so, given the small difference of the total cumulative weight from one, the bias is also likely to be small. It is thus worth making the adjustment.

small cities. I then multiply each observation within a decile by its weight and add them up cumulatively. The total thus obtained is an estimate of total mean expenditure for that decile.

In order to explore the links between household demographic and economic features, I divide total expenditure by number of persons in the household and estimate decile means for per capita total expenditure taking into account family size as a first approximation to household demographic effect. In addition, I also give to each child under the age of 15 a weight equal to half of that for an adult, and divide household total expenditure by the total number of equivalent adults to obtain equivalent per capita expenditure for that household, see also Cowell and Mercader-Prats (1999). This is the second approximation to family demographic composition. Households are thus arranged in increasing order of their per capita/equivalent per capita total expenditure before being divided into deciles. We are thus looking at two different types of distributions: mean decile per capita expenditure and equivalent per capita expenditure, before and after re-pricing. The estimates are shown in tables 3, 4, and 5. In each cell, the top figure refers to estimated mean decile expenditure and the figure under it in brackets indicates corresponding standard deviations.

The following are the definitions of the column headings in tables 3, 4, and 5. The estimates under the column headed *Market* refer to mean per capita, and equivalent per capita of actual total expenditure *without* repricing the purchases made at fixed prices. Those appearing under the *General* column refer to mean per capita, and equivalent per capita total expenditure *with repricing of all fix-price* purchases, inclusive of the coupon group, at their corresponding free market prices, covering the entire food and non-food groups. The scope of re-pricing for the food group is therefore the same as that defined earlier for the "General" scheme in table 1 above, though here non-food fixed-price purchases are also re-priced in accordance with our earlier discussion. Those arranged under the column *Coupon* refer to mean per capita, and equivalent per capita, total expenditure with *re-pricing confined to the coupon group assessed at corresponding* *market prices*. Again, the scope for the coupon group of goods is the same as that defined for the "coupon" scheme in table 1 above, plus re-pricing of non-food coupon purchases by a common factor defined as the aggregate price ratio of the food coupon group detailed earlier.

In order to turn these estimates into pure numbers easily comparable across rows, we have deflated mean decile expenditure, dividing it by the *median* of the entire distribution. In the present distributional context, median as a measure of central tendency is superior to mean due to its robustness to outliers at either extreme of the distribution. However, note that since weights are applied to ranking households, the median will also be weighted. To estimate the weighted median for total expenditure, we have to find out the level so that sum of the weights for all households with lower total expenditure is half the total sum of weights[13]. Since the distribution is based on weights, this middle observation would not be the one that divides the distribution into two equal halves in terms of number of households. The middle observation in this case is the per capita total expenditure of the household with a cumulative weight of 0.5, or the nearest to it, which is roughly the expenditure defining the boundary between the fifth and sixth weighted deciles. These deflated deciles mean expenditures are reported as the bottom set of figures in each cell. In discussing the results, I am mainly concerned with these sets of estimates.

First of all, note that as a result of re-pricing of rationed goods, original mean expenditure will increase for all households, including the rich, since every household consumes some rationed commodities. This is shown by the "Market" estimates of mean total expenditure (top line in each cell), which is always higher for the "Coupon" estimates than the original and still higher for the "General" estimates than the coupon. This applies to both per capita and equivalent per capita in

[13]The weighted median $\bar{x}_w$ is defined as $\sum_{i=1}^{w} 1(x_i \leq \bar{x}_w) v_i = 0.5$ where v_i's are the normalized weights and the function 1(e) is 1 if e is true and 0 otherwise. Other percentiles can be estimated similarly by replacing 0.5, see Deaton (1997). This is close to the method I adopted above for defining decile expenditure classes.

all city groups. We can also interpret each decile standard deviation as a measure of inter-class *inequality*. It can be seen that this inequality follows the increase in mean decile expenditure after re-pricing of rationed goods throughout the three tables.

However, it is the third set of median deflated figures (bottom line in each cell) that allows much wider scope for comparison. Specifically, these deflated means provide evidence on the turning point for distributional gains from rationing, in addition to their measuring overall reduction in inequality. Take the latter for Tehran first. Table 3 shows that for per capita expenditure (the first three columns) the median deflated difference between the mean of the 1st decile and that of the 10th decile narrows from 3.9794 for the original per capita distribution to 3.6828 for the Coupon per capita and further to 3.6524 for the General per capita. Hence, distribution under the general scheme displays greater equality than that under the coupon scheme. It is also commonplace to examine the ratios of decile means. The corresponding ratio values are 14.5802, 12.8855, and 12.2178. Thus, the ratio estimates also point to greater equality after re-pricing, though on the ratio basis, the Coupon scheme shows greater equality than the General scheme.

The same pattern can be observed for the equivalent per capita distribution, as in the last three columns of table 3, where the difference between the mean of the top and the bottom deciles narrows from the original value of 3.6009 to the Coupon 3.3537 and further to the General 3.4287. Note the reversal in inequality ranking, on the median deflated differences, between the Coupon and General schemes as we move from per capita distribution to equivalent distribution. The corresponding ratio estimates are 13.0714 (Market), 11.0742 (Coupon), and 11.6185 (General). The median deflated ratios of mean expenditure suggest the same ranking of inequality as that based on the differences. Keeping in mind the auxiliary schemes and the significance of the public sector as suppliers of some code 1 purchases that make up the General scheme, one cannot assume that the latter will be the most equal even though it contains the commodities in the Coupon scheme. Whether the General scheme leads to greater equality depends on the distribution of those fix-price purchases not equally accessible to every one, over all households.

For large cities, these median deflated differences, respectively under the original "Market", Coupon and General distributions, are 3.8333, 3.4993 and 3.5013 for per capita means, and 3.6546, 3.2727, 3.2132 for equivalent per capita means, again ranking changes as we move from the Coupon scheme to the General one. This time, however, inequality is slightly greater under the General per capita distribution, but this is reversed as we move to equivalent per capita distribution where inequality becomes greater under the Coupon scheme. The corresponding median deflated per capita ratios are 12.8678, 10.8795, 11.6845; and equivalent per capita ratios are 12.1625, 10.2711; 10.8504. Thus, while there is agreement on inequality ranking on both the difference and ratio basis for equivalent per capita values, here the per capita ratio ranking, unlike the per capita difference ranking, suggests the General scheme to be resulting in greater equality.

For small cities, the corresponding median deflated differences between the top and bottom deciles are 3.5149, 3.4495 and 3.2394 for per capita distribution and 3.3116, 3.2980 and 3.2030 for equivalent distribution. The corresponding per capita ratios are 15.3348, 13.4262, and 12.5404, while those for equivalent per capita ratios are 14.6336, 12.9320, and 12.2742. Thus, in small cities the General scheme results in more equality than the Coupon under both per capita and equivalent per capita, regardless of whether median deflated differences or ratios are employed. Note that in both large and small cities each equivalent value is below its corresponding per capita, the difference being about 0.2 to 0.3 except for the General scheme in small cities. Thus, the evidence for all three city groups suggest that allowing for variation in needs by age has its own important contribution to make towards reducing overall inequality quite apart from the influence of subsidised rationed commodities.

General improvements in inequality do not tell us much about how the gains are distributed and which decile groups have experienced reductions in total expenditure under rationing. In particular, it is of considerable interest to find out how the poor households fare under rationing or to assess theoretical propositions which suggest, for example, that, under rationing, the lowest income households gain, middle income households' position remain unchanged, while top income households lose, see Sah (1987). Decile by decile comparison of mean expenditure is particularly useful in shedding light on such questions.

Compare now the original per capita distribution with the Coupon for Tehran. It can be said that the median deflated mean total expenditure is higher for re-priced distribution up to and including the seventh decile. Only the eighth, ninth, and tenth deciles have median deflated coupon mean expenditures lower than the originals, that is, gains from rationing, relative to the median, benefit not only the poor but also the middle deciles and some better off households. When we compare the original per capita with the General per capita, however, the change-over occurs earlier in the distribution, up to and including the fourth decile, with the exception of the eighth decile which has a slight gain of 0.02 under the coupon scheme. As for the equivalent distribution, the switch occurs between the eighth and ninth deciles, while for the General scheme it is even further up the scale between the ninth and the tenth. Thus, only in the case of the per capita distribution under the Coupon scheme are the gains confined to the lowest four, poor deciles. Elsewhere, gains from rationing extend well into the upper expenditure groups, though typically the gains are insignificant the further up one goes; important gains being confined to the bottom groups.

We now look at the influence of equivalent distribution on decile by decile changes in table 3. Here, one finds that the difference between the deflated decile means of the original and Coupon for equivalent per capita are larger for gainers and smaller for losers than those for per capita distribution, the exception being the poorest 1st decile where the gain is larger under per capita distribution. The changeover in such differences from gain to loss occurs between the seventh and eighth deciles under per capita, and between the eighth and ninth deciles for equivalent, corresponding to their respective decile by decile switches. However, in all cases the differences in gains and losses between the two distributions are insignificant, suggesting only a slight role for equivalent per capita in this context. Thus, using an adult equivalent weight of 0.5 for children does not make much difference to mean expenditure if we look at the decile by decile distribution. In particular, while the equivalent per capita concept makes a clear contribution to the reduction of overall inequality in Tehran, its impact on poverty and the expenditure means of poor deciles is negligible.

Identifying the decile groups gaining from rationing in large cities from table 4 we can see that deflated mean expenditure increases as we move from the original ("Market") to the re-priced distributions for the first four deciles, and thereafter decreases. The same applies to a comparison of the original equivalent per capita and the Coupon one, where gains are exclusive to the bottom four deciles. In fact this holds uniformly throughout table 4 whether we compare per capita /equivalent per capita or the Coupon/General schemes. Thus, the boundary between the fourth and fifth deciles divides the table neatly into two sections, a lower part containing gainers, and an upper part of losers.

Comparing gains and losses under per capita and equivalent per capita for large cities shows that gains and losses (differences between the original and equivalent decile means) are larger/smaller under per capita distribution than they are under equivalent distribution. The changeover from gains to losses occurs between the fourth and fifth deciles (from 0.0148 for the fourth per capita decile to -0.0040 for the fifth, and from 0.0008 for the fourth equivalent per capita decile to -0.0138 for the fifth). The switch occurs at the line dividing the table into two. More important, however, are the differences in gains and losses under the two distributions. Again, such differences are small, suggesting the adult equivalent weight of 0.5 does not significantly change the decile by decile distribution.

For small cities in table 5, the switch in distributional gains from rationing under the Coupon scheme is between the fifth and sixth deciles for per capita expenditure. Nevertheless, when we compare the original with the General scheme, only households in the 1st decile display clear gains from rationing. However, the gap in decile changeover between the two schemes may be somewhat closer as the means of the second deciles are practically identical for the original and the General; the same applies to the fifth decile means for the original and Coupon. Comparing the original equivalent per capita with the Coupon, we can see that the changeover is much further up, between the seventh and eighth deciles, and this is also the turning point for the General scheme. Thus, while distributional gains from rationing are similar to large cities for per capita expenditure, being mainly confined to the poor, in this case the very poor, the pattern for equivalent per capita distribution is similar to Tehran, where even some rich households display gains from rationing. Once again, significant gains are confined to the bottom three deciles while gains for other groups are relatively modest, as was the case in Tehran. Comparing per capita and equivalent per capita for small cities shows that differences between the original and Coupon decile means are now larger for gains and smaller for losses under equivalent than under per capita distribution. Thus, the pattern is similar to Tehran's. The changeover from gains to losses occurs between the sixth and seventh deciles for per capita, and a little further up between the seventh and eighth deciles for the equivalent distribution. However, since the magnitude of losses and gains are slight, one cannot say much about the significance of higher gains and losses under the distribution of adult equivalent total expenditure.

4Policy Relevance of Empirical Results

The above conclusion regarding the distributional consequences of rationing raises the question of their robustness to the method employed. The evident disadvantage of the above method is that everyone appeared to gain from the Iranian war rationing during 1984-85, which may not necessarily be the case when examined in a general equilibrium context. This broader distributional context is usually analysed with the help of applied general equilibrium models. Equally important is the relevance of war rationing public food distribution schemes in peace. In this regard, I shall examine some general features of a particular policy alternative based on retaining some of the advantages of the coupon system without allowing universal entitlement, a feature made unnecessary by improved food supplies in peace-time conditions. This alternative is expected to involve lower resources while it is also likely to target the poor more effectively. In this section, I examine these two issues.

It is obvious that with a single cross section budget survey one can only examine changes in demand. By necessity, supply is assumed constant. By contrast general equilibrium analysis is based on both supply and demand information, and allows for the feed-back effects of changes in demand, through the effects that changes in endogenous labour supply and output have on key variables of price and quantity with which a welfare change is measured. An important feature of the measurement of welfare change in this framework is that a policy change in the price and quantity of rationed goods may, unlike the survey based partial analysis, lead not only to welfare gain for some but also to welfare *loss* for others. Most analyses of the welfare implications of a policy employ the Computable General Equilibrium (CGE) format because of its ease of application[14]. Typically, a benchmark model, or a base year scenario, is calibrated to match the key features of the economy in that year. Then counterfactual scenarios are employed in which an exogenous factor is allowed to change in the benchmark model, and the consequences for welfare are computed[15].

---

[14]However, the implications of a change in welfare are most explicit in the Negishi (1971) format. This defines a general equilibrium as a maximum point of a social welfare function, interpreted as the choice of a particular Pareto efficient allocation represented by the preferences of a decision-maker/central planner on how to trade off individual utilities against each other, see Gunning and Keyzer (1995) for further discussion.

[15]It should be mentioned that the relative ease of CGE in application is bought at some expense, as

One area in relation to which the issue of welfare gain and loss can clearly be seen is in the analysis of the effect of changing tax policy on individuals with different marginal tax rates. For example, while the partial analysis of tax reduction may suggest welfare gains for everyone, though unequally distributed, and thus a Pareto optimal outcome, the CGE simulations usually show that some groups, in the case of a tax reduction, those with a very low marginal tax rate, may incur welfare losses from such a policy change. An example, which demonstrates this clearly, is the measurement of distortions from the 1989 U.S marginal tax rate changes on the distribution of U.S income and welfare by Altig and Carlstrom (1999, esp. discussion of fig. 6 and table 4). They define the welfare gain from a policy change as the percentage decrease in 'full wealth' (present values of labour income and of bequests) taken from an individual in 1989 in order to keep the pre-change 1986 utility unchanged. Their CGE simulations, based on 1989 as the benchmark, show the very wealthiest enjoy a welfare gain equal to 2% of their wealth, while the second poorest suffer a welfare loss of 0.33 %. However, when the experiment is repeated with partial equilibrium analysis by holding factor prices fixed, the welfare loss found in the CGE simulation fails to re-appear. Thus, the outcome is Pareto optimal with the latter method but not the former. Nonetheless, as expected from the small amount of welfare loss, the distribution of welfare due to a reduction in marginal tax rate shows very little difference in outcomes when the general and the partial methods are compared. In particular, both methods show welfare gains concentrated at the upper end of the lifetime income distribution. Thus, in this case study, the differences in estimation do not justify the extra effort involved in the employment of the CGE method. The partial method, with income treated as exogenous, provides a very good

---

the full parameter estimation of a proposed CGE model is usually impossible. First, there are identification problems arising from having a very large number of endogenous variables. Then, there is the lack of data to deal with, an especially acute problem for developing countries. This is particularly evident for input data, where one rarely has access to more than one input-output table. In practice, CGE calibrations often rely on: borrowed elasticities, adjustments in coefficients, and, more important, on parameter restriction by the imposition of theoretical assumptions, in order to reproduce the base approximation to the general equilibrium estimates[16]. One would expect a similar conclusion to emerge for the findings of this chapter. Since gains of the better off are relatively small, they may or may not show up in a broader model, while one would not expect very different measures of the much larger gains to the poor.

Mention should also be made of the budget survey-based approach to the measurement and evaluation of efficiency/equity trade-off of different schemes. This type of approach, common in the assessment of alternative tax schemes, provides a complementary method to that based on the CGE computation. The latter, due to its elaborate economy-wide structure, usually provide details for only a limited number of aggregate groups of consumers, and this may prove insufficient for reaching a clear policy judgment. The larger number of survey observations on households allow policy assessment to be based on a less aggregated grouping, see King (1983) or Ahmad and Stern (1991).

It should, however, be noted that the empirical contents of many of the CGE models are usually not a great deal superior to those reported in this paper because they also make use of the *separability* assumption. Given the elaborate nature of these models and the need to keep the number of parameters to be estimated as small as possible, this is inevitable but has its own cost. A frequently used functional form based on additive separability for consumption demand in these models is the Linear expenditure System (LES), although other functional forms based on this type of preference, such as the Cobb-Douglas, are also commonly employed[17]. Such functional

---

model. I shall return to the consequences of this practice below.

[16]Pointing out that 'welfare implications are little changed in the partial equilibrium', Altig and Carlstrom conclude that ' both the quantitative and qualitative outcomes are similar to those obtained from the full general equilibrium experiments'.

[17]The CGE models explicitly dealing with equity-efficiency comparisons of the alternative uses of government expenditure on food subsidies appear to be the most directly relevant to the assessment of the findings of this chapter. Examples are the CGE models of Narayana et al. (1988) and Parikh (1994) for India; and of Kehoe and Serra-Puche (1983) and (1986) for Mexico. Yet the former are based on the LES consumption demand function in India; the latter employs the Cobb-Douglas functional form for the same purpose for Mexico. Nor are the examples confined to the models of coupon subsidies. Just how common is the use of the separability assumption can also be seen from the CGE model of an

forms allow for substitution in a highly restrictive manner, for instance ruling out inferior goods, important in the food consumption of the poor, particularly in a developing economy. When the data comes from more normal conditions with relative absence of corner solutions, preference separability does lead to serious errors in judging the empirical results. For instance, the application of separable models suggests a proportionality between expenditure and own price elasticities where less restrictive functional forms reveal none; still worse, the empirical conclusions change radically with minor alterations in the precise definition of preference separability employed, see Deaton (1975) and (1981). The empirical results obtained with this type of model must therefore be regarded with skepticism because they are not the outcome of genuine measurement but of the theoretical assumptions incorporated into the model.

Let us now turn to the second issue, namely self-selecting public expenditure policy alternatives to wartime universal rationing as tools for the alleviation of poverty. There is clearly no justification for the state provision of subsidised food to all consumers regardless of income in peacetime conditions with normal food supplies. However, as an anti-poverty scheme for peacetime, rationing by queue appears to be more effective, as its selective mechanism tends to concentrate resources on those with a low value of time, though it is not the only such mechanism[18]. It is, however, important to realize that the queue will not necessarily result in redistribution towards the poor because the value of time, on its own, does not suffice to ensure that outcome. At least two main issues are involved here.

First and foremost, how a particular queue allocation divides fixed and variable costs to

---

influential group of writers on Africa, namely Bevan et at (1990, esp. appendix 2): the rural and urban household demand functions in this model are both based on the LES. In survey-based empirical studies on uniform commodity taxation, separability plays a prominent role in estimation of the equity/efficiency trade-off; see Ahmad and Stern (1991).

[18]It is also possible to base self-selective distribution on quality and packaging directed to items that only the poor consume. Such a scheme was has been successfully implemented in Tunisia, where, for example, subsides on cube sugar were eliminated but retained on unrefined brown sugar, see Tuck and Lindert (1996). Iran also employs this type of self-selective method on a limited scale, i.e. flour subsidy is not available for expensive types of bread and pastry items.

consumers, is important for whether it succeeds in targeting the poor. The prerequisite for this, implicitly assumed in most models of the queue, is that time expenditure is a variable cost. If consumers are allowed to purchase as much as they wish once they have reached the top of the queue, then the immediate effect would be that a minority, who can afford to undertake some fixed inventory cost, will stockpile the good, leaving little for others[19]. To prevent hoarding by the few, the government stores often restrict the amount purchased per visit.

The empirical evidence on this question is examined in Alderman (1987). He employs the Heckman two-stage method, with one equation determining the decision to enter the queue, while the second is a demand equation conditional on entry. The exogenous queue time variable represents the fixed cost of queuing in the entry equation, and its variable cost in the conditional demand equation, see also Melenberg and Soest (1996) for a similar analysis. Alderman's estimates of time elasticities indicate most of the effect, for the main commodity examined, bread, is due to fixed costs, suggesting the allocation by the queue in this case is not very successful in targeting resources on the poor. This appears to be the result of relative lack of control on the amount purchased per visit. The Iranian bread queue has the same feature[20]. By contrast, the current Iranian queue scheme for milk operates mainly based on variable cost[21]. Since the more water-tight alternative of preventing consumers from re-joining the queue may involve more costly administration, restricting the allocations to relatively small quantities per visit is an

[19]Indeed, in the extreme case of unlimited purchase per visit combined with retrading, consumers can hire the individual with the lowest queuing cost to purchase the good on their behalf, dispensing with the need to queue altogether.

[20]However, peak queuing times may see a limit imposed on the amount purchased per visit, usually 10 loaves, in order to keep down time expenditure by those consuming bread in small amounts as a palatable good.

[21]Currently, subsidised milk is available to all households on a first-come-first-served basis regardless of their demographics. However, each household can purchase only one bottle per visit, and only together with a compulsory purchase of another at market price. Those consuming entirely at market price can obtain milk without queuing. The monitoring agent is the local grocer who is supposed to refuse supply to more than one member of the same household on the same day, or to those not known to him personally coming from other localities. Although the scheme can clearly benefit from further control over the local distributors, it is a definite improvement over the wartime universal

interesting aspect of the milk scheme to mitigate the problem of fixed cost, thereby further strengthening the link between the variable cost of time and the ration obtained.

Second, Barzel (1974) argues that the queue will not target subsidised resources on those with the greatest need for them, if the marginal propensity to consume dominates the price effect, since, in that event, rationing by queuing applied to goods with high income elasticity would produce more benefit to the rich than to the poor. Therefore, if the queue is to be employed as an allocation scheme targeted to benefit the poor, it must be confined to markets where goods have low income elasticities, to health or housing rather than opera or sports facilities[22]. However, this problem arises when the queue is the only available scheme for the shortage good. In Barzel's queue model, time is the only cost, and a parallel market is ruled out. By contrast, such a market were allowed, as in the model by Nichols et al (1971), it would be possible for consumers to combine cash and time costs in order to minimize expenditure according to the value of time. In this scheme, the high income elasticity goods consumed by the rich find an alternative channel, allowing the poor to obtain comparatively more benefit from the queue. Queuing schemes for food in developing countries, most commonly for bread, operate along with a parallel free market. Indeed one of the most prevalent queuing schemes in developed economies, namely publicly funded health services such as the British NHS, usually operate together with a parallel market, often within the same unit. The evidence suggests that other issues are probably less important for a queue-based scheme of food allocation[23].

---

rationing.

[22]Frech III and Lee (1987) examine the welfare effects of disregarding this issue in quota rationing between the rural and urban sectors during the two petrol shortage crises of the 70's in the USA, and the measured loss in welfare appears quite large.

[23]Beveridge (1942), an influential document in the birth of publicly funded national health schemes, although stressed access for all regardless of income, was also quite clear on the queuing basis of its proposed health scheme. The queue was to be mainly for those who cannot afford to pay for health care. However, it was intended to operate in combination with a parallel market service for those wishing to avoid waiting for it. Successive governments of different persuasions, however, have denied the existence of such an allocational queue; the excess demand is claimed to represent a backlog of cases, which should be dealt with by additional short-term resources. As a policy, the latter has tended to

5Conclusion

To summarise, re-pricing rationed goods at market prices improves measured inequality throughout the country, regardless of city group, distribution type, or rationing scheme. Decile by decile gains for the poor, the bottom three or four expenditure classes, is significant in all cases. Gains are exclusive to the poor in large cities and, for per capita distribution, in small cities. In Tehran and, for equivalent per capita distribution, in small cities, the gains extend well into the middle and upper deciles. However, compared to the gains for the bottom deciles, gains in middle and upper deciles are negligible, suggesting that differences for small cities under the two distributions might be more apparent than real. Household distribution in equivalent rather than per capita terms is important for inequality but the influence of adult equivalent weighting on decile gains and losses is very slight. Estimates for the General scheme show greater irregularity compared to the Coupon, especially in small cities[24].

The possible policy relevance of these findings may be of some interest. It seems probable that if some aspects of rationing are to be retained in peacetime Iran, then the queue, or at least some such self-selecting scheme, provides a more suitable distributional method than quota rationing. Provided food distribution is firmly based on the variable component of queuing time, the other conditions for the relative success of distribution by through queuing in targeting the poor are usually met in food rationing schemes. It is quite possible for the state to spend significantly less on subsidizing the consumption of sugar, oil, and rice, and still achieve a

---

create additional demand, often through greater number of hospital referrals by general practitioners, without solving the problem of long-run excess demand. The other major queue type, namely, that based on a waiting list, would not work as policy tool for food allocation. This latter type of rationing is most effective when applied to goods with *unpredictable* demand, which explains its prevalence in the allocation of medical care for which individuals cannot forecast short term demand. When this is not the case, it will not be an effective distributional policy tool. Since short-term demand for food is predictable, consumers rationed by waiting list can simply order in advance causing an indefinite lengthening of the queue, thus disabling it from performing its rationing function, Lindsay and Feigenbaum (1984).

[24]A Central Bank of Iran (1989) study on poverty on urban Iran also concluded that rationing has significantly benefited the poor.

substantial reduction in poverty, by abolishing universal entitlement in favour of the queue-targeting scheme[25]. For this reason, the current Iranian scheme of milk distribution represents an intelligent attempt to modify wartime rationing to peacetime conditions and deserves to be considered for the future of food policy in Iran, and more generally for the design of public schemes for food support in other developing counties[26]. This scheme also avoids the major problem of most anti-poverty programmes, namely, how to identify the poor. It has a built-in mechanism for the identification of the poor.

Table 1-Aggregate Price Ratio of Market to Fix-price Purchase of Food Commodities						
Aggregate Price Ratio	Tehran	Large Cities	Small Cities	Tehran	Large Cities	Small Cities
	General Scheme			Coupon Scheme		
EQ	2.097	1.907	2.113	2.492	2.346	2.502

**** = General scheme***: *All* goods purchasable at fixed prices, inclusive of the narrow group of coupon goods, but also inclusive of goods not necessarily universally available to all citizens, such as beans, fresh fruits, bus tickets.
 # = Coupon scheme: Exclusive to the narrow group of goods purchasable by coupon or identity card, and universally available to all citizens, such as rice, cigarettes, washing powder.

Table 2-Comparison of Averages for the Key Values 1984-85 CBIHBS With and Without Population Weights			
	Total Expenditure **‖**	Budget Food share	Household Size
Population Weights*****	1574547	40.7	4.90
Without Weights**#**	1669942	39.0	5.02

Source: * = CBIHBS (1985, table 26, and p.4), # = Obtained from the survey of this study
‖ = *Definition of total expenditure*: Annual (in Rails), inclusive of the imputed rental value of housing, plus tax and pension contributions, but exclusive of housing investment.

[25]World Bank (2003) analysis of the removal of subsidies in Iran has pointed out that despite the inefficiency of the current self-selection rationing based on queuing, 'its rational is sensible', (p.115).

[26]There are indications that public food programmes in such countries as Iraq or Afghanistan are poorly targeted, resulting in significant leakages as evidenced by the scale of coupons resale either because the (Afghan) poor regard their food rations too good for own consumption, or that the (Iraqi) rich consider theirs being unpalatable. A distribution scheme based on queuing time, perhaps combined with smaller rations, but without a limit on re-queuing, is likely to weed out those not depended on food rations.

Table 3-Comparison of Deflated Mean Expend.,Repricing Fix-price Goods at Market rate (Weighted Monthly Exp. In Rails)						
Tehran						
Decile Class	Per Capita			Equivalent Per Capita		
	Market	General	Coupon	Market	General	Coupon
Decile 1 Mean St. Dev. Def.Mean	8203 (49) 0.2930	13279 (70) 0.3073	10070 (55) 0.3283	10379 (64) 0.2983	16948 (91) 0.3229	12601 (67) 0.3329
Decile 2 Mean St. Dev. Def.Mean	12977 (29) 0.4636	20639 (50) 0.4776	15142 (33) 0.4937	16084 (35) 0.4623	25978 (62) 0.5046	19043 (40) 0.5031
Decile 3 Mean St. Dev. Def.Mean	16762 (35) 0.5988	26688 (46) 0.6175	19527 (35) 0.6367	2051 (35) 0.5909	33163 (55) 0.6441	23978 (40) 0.6334
Decile 4 Mean St. Dev. Def.Mean	20933 (35) 0.7478	32612 (52) 0.7546	23880 (38) 0.7786	25718 (42) 0.7392	39744 (60) 0.7719	29236 (39) 0.7723
Decile 5 Mean St. Dev. Def.Mean	25626 (38) 0.9154	39441 (60) 0.9126	28651 (35) 0.9342	31393 (50) 0.9023	47159 (54) 0.9160	35134 (46) 0.9281
Decile 6 Mean St. Dev. Def.Mean	30572 (44) 1.0921	46891 (71) 1.0851	34207 (56) 1.1153	37333 (46) 1.0730	56809 (89) 1.1034	41495 (58) 1.0962
Decile 7 Mean St. Dev. Def.Mean	37376 (60) 1.3352	56733 (90) 1.3128	41064 (59) 1.3389	44089 (59) 1.2672	68672 (105) 1.3338	48769 (61) 1.2884
Decile 8 Mean St. Dev. Def.Mean	45623 (78) 1.6298	71023 (131) 1.6435	49523 (77) 1.6147	54786 (105) 1.5746	85328 (157) 1.6573	59776 (108) 1.5791
Decile 9 Mean St. Dev. Def.Mean	63927 (201) 2.2836	95007 (250) 2.1984	67823 (186) 2.2113	75536 (223) 2.1710	113423? (295) 2.2030	80223 (207) 2.1193
Decile10 Mean St. Dev. Def.Mean	119601 (1428) 4.2724	171123 (2062) 3.9597	123023 (1412) 4.0111	135665 (1430) 3.8992	193149 (2062) 3.7516	139552 (1419) 3.6866

Table 4-Comparison of Deflated Mean Expend., Repricing Fix-price Goods at Market Rate (Weighted Monthly Exp. In Rails)						
Large Cities						
Decile Classes	Per Capita			Equivalent Per Capita		
	Market	General	Coupon	Market	General	Coupon
Decile 1						
Mean	6246	8534	7678	7976	10868	9806
St. Dev.	(45)	(59)	(52)	(57)	(75)	(65)
Def.Mean	0.3230	0.3277	0.3542	0.3274	0.3262	0.3530
Decile 2						
Mean	9449	13170	11172	12079	16844	14358
St. Dev.	(21)	(30)	(24)	(31)	(37)	(29)
Def.Mean	0.4887	0.5058	0.5153	0.4959	0.5056	0.5169
Decile 3						
Mean	11830	16271	13720	15247	21116	17613
St. Dev.	(23)	(29)	(19)	(28)	(36)	(27)
Def.Mean	0.6118	0.6249	0.6328	0.6260	0.6338	0.6341
Decile 4						
Mean	14490	19723	16567	18646	25265	21286
St. Dev.	(26)	(29)	(29)	(30)	(33)	(36)
Def.Mean	0.7493	0.7574	0.7641	0.7655	0.7584	0.7663
Decile 5						
Mean	17891	24015	19972	22622	30229	25415
St. Dev.	(25)	(33)	(28)	(32)	(47)	(36)
Def.Mean	0.9252	0.9222	0.9212	0.9287	0.9074	0.9149
Decile 6						
Mean	21579	28562	24110	27028	36148	30489
St. Dev.	(33)	(41)	(37)	(39)	(47)	(46)
Def.Mean	1.1159	1.0969	1.1121	1.1096	1.0851	1.0976
Decile 7						
Mean	26289	34468	29115	33193	42805	36596
St. Dev.	(43)	(51)	(42)	(52)	(62)	(48)
Def.Mean	1.3595	1.3237	1.3429	1.3627	1.2849	1.3174
Decile 8						
Mean	32399	42901	35421	40774	53441	44157
St. Dev.	(61)	(86)	(59)	(71)	(71)	(77)
Def.Mean	1.6755	1.6475	1.6338	1.6740	1.6740	1.5896
Decile 9						
Mean	44400	55241	47327	53731	68808	57554
St. Dev.	(128)	(134)	(132)	(155)	(147)	(151)
Def.Mean	2.2961	2.1214	2.1829	2.2059	2.0654	2.0719
Decile10						
Mean	80371	99704	83549	96992	117912	100715
St. Dev.	(993)	(1099)	(1022)	(1118)	(1259)	(1175)
Def.Mean	4.1563	3.8290	3.8535	3.9820	3.5394	3.6257

Table 5-Comparison of Deflated Mean Expend., Repricing Fix-price Goods at Market Rate (Weighted Monthly Exp. In Rails)

Small Cities

Decile Classes	Per Capita			Equivalent Per Capita		
	Market	General	Coupon	Market	General	Coupon
Decile 1						
Mean	4546	7209	5736	5729	8914	7158
St. Dev.	(36)	(59)	(48)	(51)	(77)	(62)
Def.Mean	0.2452	0.2807	0.2776	0.2429	0.2841	0.2764
Decile 2						
Mean	8551	11821	10006	10625	15405	12639
St. Dev.	(24)	(31)	(24)	(33)	(37)	(33)
Def.Mean	0.4612	0.4603	0.4842	0.4505	0.4910	0.4879
Decile 3						
Mean	11765	15911	13512	14830	20401	17116
St. Dev.	(22)	(27)	(23)	(27)	(32)	(33)
Def.Mean	0.6346	0.6196	0.6538	0.6288	0.6502	0.6608
Decile 4						
Mean	14707	19681	16511	18455	24802	20607
St. Dev.	(23)	(25)	(27)	(28)	(31)	(29)
Def.Mean	0.7933	0.7664	0.7990	0.7824	0.7904	0.7956
Decile 5						
Mean	17434	23563	19432	21763	29060	24290
St. Dev.	(18)	(32)	(21)	(26)	(36)	(28)
Def.Mean	0.9404	0.9176	0.9403	0.9227	0.9261	0.9378
Decile 6						
Mean	20184	27918	22545	25394	43141	28147
St. Dev.	(28)	(32)	(30)	(29)	(46)	(35)
Def.Mean	1.0887	1.0872	1.0909	1.0766	1.0880	1.0867
Decile 7						
Mean	24662	32361	27110	29954	40555	33565
St. Dev.	(35)	(44)	(36)	(46)	(45)	(47)
Def.Mean	1.3303	1.2602	1.3119	1.2699	1.2924	1.2958
Decile 8						
Mean	30030	39801	32895	37478	48276	41048
St. Dev.	(49)	(66)	(52)	(59)	(72)	(62)
Def.Mean	1.6198	1.5499	1.5918	1.5889	1.5385	1.5847
Decile 9						
Mean	38983	50892	42035	48829	62837	52365
St. Dev.	(82)	(123)	(87)	(98)	(138)	(101)
Def.Mean	2.1028	1.9818	2.0340	2.0702	2.0026	2.0217
Decile10						
Mean	69709	90393	77024	83840	109419	92583
St. Dev.	(608)	(1412)	(1388)	(666)	(1603)	(1579)
Def.Mean	3.7601	3.5201	3.7271	3.5545	3.4871	3.5744

Appendix Population Weights

This appendix deals with the population weights employed above. However, these are better understood if we first discuss population weights for the original sample size of 6000 before turning to the type of modifications employed for the estimations presented above.

A1Original Population Weights

Since the survey is not "self-weighting", thus the fraction of total households included is not the same in all strata; sums and means cannot be obtained by simple adding up of the results for households. Each household has a population coefficient representing its probability of being included in the sample. Households belonging to the same bloc have the same coefficient. The method of estimating the coefficients can be expressed in notations as follows

$$\overline{X} = \frac{1}{R}\sum_{l=1}^{24}\sum_{g=1}^{2}\sum_{k=1}^{S_{g1}}\sum_{f=1}^{N_{kgl}}\sum_{h=1}^{R}\frac{M_{fkgl}}{\sum_{f=1}^{N_{kgl}}M_{fkgl}} * \frac{M_{kgl}}{\sum_{k=1}^{S_{gl}}M_{kgl}} * \frac{M_{gl}}{M_{l}} * \frac{M_{l}}{\sum_{l}M_{l}} * X_{hfkgl}$$

h=index of number of households in a selected block,

f=index of number of selected blocks,

k=index of selected city,

g=group index, equal to 1 for randomly selected small cities and 2 for self-representing cities, i.e. all other cities,

l=index of province (from 1 to 24),

M=population, or number of households in population,

R=number of selected households in a selected block (8 for Tehran and 4 elsewhere),

N=number of selected blocks,

S=number of selected cities,

X=household characteristic of interest, such as, housing expenditure,

$\overline{X}$ =Mean value of household characteristic of interests over the whole country, that is, all urban areas. The same holds true if a mean value of a household characteristic is required at a lower level of aggregation. For example, for all small cities we have

$$\overline{X} = \frac{1}{R}\sum_{l=1}^{24}\sum_{k=1}^{S_{1l}}\sum_{f=1}^{N_{k1l}}\sum_{h=1}^{R}\frac{M_{fk1l}}{\sum_{f=1}^{N_{k1l}} M_{fk1l}} * \frac{M_{k1l}}{\sum_{k=1}^{S_{1l}} M_{k1}} * \frac{M_{1l}}{\sum_{l=1}^{24} M_{1l}} * X_{hfk1l}$$

where g can only take the value 1.

The value of a given household characteristic is then weighted by its corresponding coefficient and the results are added cumulatively to obtain the mean value of that characteristic over all households. Only the ratio of households in a selected block of a city to all households in all selected blocks in that city can be obtained from the survey. All other ratios, such as, number of households in a selected small city over number of households in all selected small cities, must be obtained from the census. Since the results of the 1985-86 census were not available in 1984-85, the population ratios were obtained by the CBI from the 1975-76 figures, then up-dated to the 1984-85 levels to take into account demographic change.

A2Modified Population Weights

The magnetic tape made available for this study contains a large random selection the households (4246) out of the entire 60000 that constituted the 1984-85 CBIHBS. The random rule used for this purpose was discussed in Appendix A1. The removal of a fifth of households means that we have variable numbers rather than a fixed number of households per block. This means that in weighing observations by population coefficients, it is not enough to match each estimated coefficient with the corresponding household, but the weights themselves must be modified, multiplying 4/x or 8/x where x is the number of households in a given block, in order to obtain new estimates for these weights. These new estimates must be matched with their corresponding households, not the original estimates obtained from the entire sample.

The second problem was to decide what to do with eight cities, which were used in calculation of average figures for the whole country, but were not included in the categories of small or large cities. The CBI no longer publishes separate estimations for household expenditure in small and large cities as it considers its survey to be insufficient in size in these categories. However, estimation of population weights are available for each household at the national level, as well as provisional, small and large city levels, with the exception of the eight cities for which only national weights are available. If we include these weights in any of our city groups, then the original city group weights are no longer usable, since small or large city weights involve estimates of total population of large cities or that of sampled small cities. Therefore, these estimates and the weights based on them would change, if we included the population of the eight cities, used only for the original national weights, among them. Since they contain about 6.5% of the sampled households, rather than leaving these eight cities out, it was decided to include them in the small city category.

The CBI defines a city over 20,000 to be a large city. However, each group contains cities that on a size criterion alone should have been in the other group. In fact, six of these eight cities have populations just over 20,000. The separating of cities into large and small also depends to some extent on the social and economic importance of the city, either because the city is an important regional centre, or an important import/export port, even though they may have populations below 20,000. Indeed, the growing importance of most of these cities since the last census of 1975-76 has been the reason for their inclusion in national estimates. Some of these eight cities could have been included with large cities, but their allocation to one category, small cities, was decided because this as the revised estimation easier by confining it to small city groups, though this may somewhat alter the picture for small cities.

The original weights for large cities are based on the total large city population, which includes Tehran. Since I investigate Tehran separately, specific weights had to be re-estimated for Tehran and for all other large cities (excluding Tehran). For the Tehran weights, the ratios of number of households in each selected block to all households in all selected blocks were available from the original estimates for Tehran. For large cities, the weights were re-estimated using the smaller total large city population, excluding Tehran. The re-estimation provided for each household results in one revised ratio for Tehran and two revised ratios for small cities. The final multiplication of several ratios to obtain relevant weights was carried out by computer taking into account the missing households in each block resulting from the removal of one fifth of the original set.

Finally, I have two observations on the estimates used for the population of cities. The formulae for these weights are based on households rather than individuals, but only for the ratio $\frac{M_{fkgl}}{\sum_{f=1}^{N_{kgl}} M_{fkgl}}$ is the information available in terms of households. For other ratios involved in estimating weights, population in terms of individuals were used as an approximation to population in terms of households. It should be remembered that this assumes that average number of persons per household
in a city approximately equals the average number of persons per household in all cities [27].

The second issue relates to the CBI's estimated 1984-85 population for the 72 cities in the survey based on the 1975-76 population census. A few of these estimates were implausibly smaller or larger than the corresponding 1985-86 census population subsequently available, see Statistical Centre of Iran (1986). In such cases new replacement estimates were obtained by assuming the ratio of population in 1985-86 to that in 1984-85 for a city was equal to the equivalent ratio for the corresponding province, or *vice versa* if a replacement was estimated for a province in 1984-85.

---

[27] We assume: $$\frac{M_k}{\sum_k M_k} \approx \frac{\dfrac{(\frac{M_k * A_k}{\quad})}{A_k}}{\dfrac{[(\sum_{kM_k})*A]}{A}}$$ where M_k is the population of city *k* and A_k its average number of persons per household in city k and A the average number of households in all cities. The approximation of population by households assumes $A_k = A$. The extent of the inaccuracy of weights using population depends on how far apart is A_k from A, see Kuznets (1979), PP.102-108.

References:

Ahmad H, and N Stern (1991), *Theory and Practice of Taxation in Developing Counties*, Oxford: Oxford University Press.

Alderman, H. (1987), "Allocation of Goods Through Non-Price Mechanisms: Evidence on Distribution by Willingness to Wait", *Journal of Development Economics*, Vol. 25: 105-24.

Altig D. and C. Carlstrom (1999), 'Marginal Tax Rates and Income Inequality in a Life-Cycle Model', *American Economic Review*, 89, 5: 1197-1215.

Barzel, Y., (1974), "A Theory of Rationing by Waiting: Time as a Market Clearing Price", *Journal of Law and Economics*, Vol. XVII (April): 75-95.

Bevan, D, P. Collier, and J. W. Gunning (1990), *Controlled Open Economies: A Neo-classical Approach to Structuralism*, Oxford: Clarendon Press.

Beveridge, W. H. (1942), *Social Insurance and Allied Services*, London: HMSO

Bhanoji Rao, V. V. (1981), "Measurement of Deprivation and Poverty Based on the Proportion Spent on Food: An Exploratory Exercise", *World Development*, 9(4), 753-68.

Central Bank of Iran (1984-85), Definitions and Terms Used in the 1984-85 Household Budget Survey, mimeo: Tehran (in Persian).

Central Bank of Iran (1986), *The Results of Household Budget Survey of Urban Iran, 1984-5*: Tehran, (in Persian).

Central Bank of Iran (1986), *Poverty: Concept, Measurement, and Criteria for its Identification in Iran*, Internal Report, Depart. of Economic Statistics: Tehran (in Persian).

Cockran, W. (1977), *Sampling Techniques*, 3rd ed. New York: John Wiley and Son.

Coulombe, H. and L. Demery, (1993), "Household Size in Cote d'Ivoire, Sampling Bias in the CILSS", *LSMS, WP* No. 97 Washington: World Bank.

Cowell, F. and M. Mercader-Prats (1999), 'Equivalence Scales and Inequality', Discussion Paper No. DARP 27, *STICERD*, London School of Economics.

Deaton, A. S. (1975), "A Reconsideration of the Empirical Implications of Additive Preferences", *Economic Journal*, 84: 338-48.

Deaton, A. S. (1980),"The Measurement of Welfare", *LSMS ,WP* No.7, Washington: World Bank.

Deaton, A. S. (1981), " Optimal Taxes and the Structure of Preferences", *Economtrica*, 49: 1245-60.

Deaton, A. (1997), *The Analysis of Household Surveys, A Microeconometric Approach to*

Development Policy, London: Johns Hopkins University Press.

Demery, L. and C. Grootaert (1993), "Correcting for Sample Errors in the Measurement of Welfare and Poverty: the Case of Cote d'Ivoire Living Standard Survey", *World Bank Economic Review*, Vol. 7: 263-292.

Frech III H. and W. Lee (1987), “The Welfare Cost of Rationing-by-Queuing Across Markets: Theory and Estimates From the U.S. Gasoline Crises”, Quarterly Journal of Economics, Feb.: 97-108.

Gunning, J W and M A Keyzer (1995), “Applied General Equilibrium Models for Policy Analysis”, in (eds) Behrman, J and T.N. Sinivasan, *Handbook of Development Economics*, Vol. III, Elsevier Science B.V.

Hicks, J.R. (1955), *A Revision of Demand Theory*, Oxford: Clarendon Press.

Kehoe, T. J. and J. Serra-Puche (1983), “ A Computational General Equilibrium Model with Endogenous Unemployment: An analysis of the 1980 fiscal reform in Mexico”, *Journal of Public Economics*, 22:1-26.

Kehoe, T. J. and J. Serra-Puche (1986), 'A General Equilibrium Analysis of Price Controls and Subsides on Food in Mexico', *Journal of Development Economics*, 21: 65-87.

King, M. (1983), “Welfare Analysis of Tax reforms Using Household Data”, *Journal of public Economics*, 21: 183-214.

Koohi-Kamali, F. (1992), “Rationing of Commodities and its Welfare Consequences”, in The 1991-92 Lectures from Monthly Seminars: Central Bank of Iran: Tehran (in Persian).

Kuznets, S (1979), “Population Growth and Structural Shift”, in w Gelenson (ed.) *Economic Growth and Structural Change in Taiwan*, 15-131, New Heaven: Yale University Press.

Lindsay, C, M. and B. Feigenbaum (1984)," Rationing by Waiting Lists", *American Economic Review*, Vol. 74, 3: 404-417.

Melenberg, B. and A Soest (1996), “Parametric and Semi-parametric Modelling of Vacation Expenditures”, *Journal of Applied Econometrics*, Vol. 11: 59-76.

Mohring, H. (1971), “Alternative Welfare Gain and Loss Measures”, *Western Economic Journal*, 9: 349-68.

Narayana, M.S.S, K.S. Parikh, and T.N. Srinivasan (1988), “Rural Works Programs in India: Costs and Benefits”, *Journal of Development Economics*, 29:131-56.

Nichols, D., E. Smolensky and T. Tideman, (1971), "Discrimination By Waiting Time in Merit Goods", *American Economic Review*, Vol. 61: 312-323.

Negishi, T. (1972), *General equilibrium and international trade*, Amsterdam: North-Holland.

Osband, K.(1991), "Index Number Biases During Price Liberalization", *IMF Working Papers*, WP/91/76.

Parikh, K. S. (1994), 'Agricultural Policy in India: Some Explorations', in (eds) J. Mercenier and T.N. Srinivasan, *Applied General Equilibrium and Economic Development: Present Achievements and Future Trends*, Ann Arbor: Michigan U P.

Sah, R.J. (1987), "Queues, Rations, and Market: Comparisons of Outcomes For the Poor and the Rich", *American Economic Review*, Vol. 77, no. 1: 69-77.

Samuelson, P.A. (1974), "Complementarity-An Essay on the 40th Anniversary of the Hicks-Allen Revolution in Demand Theory', *Journal of Economic Literature*, Vol. 12:1255-89.

Tuck, L. and K. Lindert (1996), “From Universal Food Subsidies to a Self-Targeted Program: A Case Study in Tunisian Reform”, *World Bank Discussion Paper No. 351*, Washington, D.C.

World Bank (2003), *Iran, Medium Term Framework for Transition: Converting Oil Wealth to Development*, Social and Economic Development Group, Report No 25848-Iran: Washington, D.C.