the review of income and wealth

Review of Income and Wealth
Series 59, Number 1, March 2013
DOI: 10.1111/roiw.12008

ESTIMATION OF EQUIVALENCE SCALES UNDER CONVERTIBLE RATIONING

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This paper provides estimates of equivalence scales based on three Quadratic Almost Ideal type models of the food share with coupon resale, and on their extended versions. A chief feature of the models is the method of dealing with infrequency of zero expenditure on coupon goods through imputation. The models are applied to an Iranian wartime budget survey. The results indicate little scope for measurement error. However, they also reveal a strong *cross-section* price effect, which proves robust to a variety of checks. The effect of price heterogeneity proves critical in identification of the scale estimates, providing relatively rare empirical evidence consistent with the Base Independence (BI) hypothesis. The BI food share estimates of cost of children have plausible values.

JEL Codes: D13, I31, J13

Keywords: base independence, cost of children, equivalence scales, food prices, welfare comparison

1. INTRODUCTION

The natural context of consumption rationing is a wartime economy, yet little is known empirically about consumption in such an economy. In particular, in order to estimate adult equivalence scales from a demand function, or an Engel curve, for such an economy, one would first have to establish whether a conditional function is appropriate. The available wartime estimates for such scales (e.g. Nicholson, 1949), provide no justification for the employment of a conventional function. The latter depends on whether rationing seriously restricts consumption, for example when additional consumption cannot be secured through the market, and if so, whether an Engel curve, or a demand function is required. In addition, estimating a demand function requires data on price variations, generally unavailable from a single cross-section survey. The present paper provides estimates of adult equivalence scales based on a cross-section budget survey whose unique features, especially with regard to price heterogeneity, provide answers to the above issues.

I employ three models with food share as a welfare indicator for this purpose. The models are based on a specific functional form that assumes the general absence of corner solutions. The welfare indicator can be interpreted in terms of

Note: I wish to express my special debt to John Muellbauer for his suggestions. Comments on the earlier drafts of this paper by Frank Cowell, Assad Monajemi, and Hashem Pesaran are also gratefully acknowledged. I also wish to thank Ran Liu for help in preparing the final draft. They bear of course no responsibility for remaining errors or shortcomings, which are mine alone.

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Published by Blackwell Publishing, 9600 Garsington Road, Oxford OX4 2DQ, UK and 350 Main St, Malden, MA, 02148, USA.

the Engel, Barten, or Prais–Houthakker models of equivalence scales identified from the assumption that the scale for food is the same as the average scale for the rest of household consumption. An unusual feature of the main model, in the context of *cross-section* household data, is its incorporation of differences in prices faced by different households. In the field of consumption analysis, the Barten model and its various extensions have provided the rationale for the existence of such cross-sectional "price," or more precisely quasi-price effects exerted on consumption by differences in household demographics. The hypothesis of different cross-section quasi-prices faced by different households has been extensively tested and rejected by Muellbauer (1977). By contrast, the household-specific price effects presented in the model below do not impose the restriction of the Barten model, and the evidence presented suggests they have significant impact on the estimates, and bear some affinity to the literature on income-dependent prices, (e.g. Rao, 2000).

The estimates are based on microdata from a household expenditure survey conducted by the Central Bank of the Islamic Republic of Iran during 1984–85 (henceforth CBIHBS 84–5), a year that belongs to the 1980–88 period of the Iran–Iraq war during which an extensive system of mainly food coupon rationing was implemented in Iran, which, in a scaled-down version, continues to this day. The dataset consists of some 4300 households surveyed in 72 selected cities throughout the country over an 11-month period,[1] employing multistage sampling techniques (see Koohi-Kamali, 2004). Two of its features are of particular relevance to this paper. First, the survey employs two-day and one-month recall periods for food, e.g. meat and rice; and one-month and one-year recalls for non-food, e.g. clothes and durables. Note that a two-day recall is too short to capture adequately the voluntary cycles of purchase, let alone the longer frequency cycles of ration distribution. It is therefore likely to be a major source of measurement error. Second, the distinctive feature of the survey is its recording of household expenditure in terms of fixed-price and/or market-price purchases separately. In addition, for the food and tobacco group alone, this dual structure also includes separate recording of quantities purchased, thus effectively providing a *dual price structure* for the entire food and tobacco group. There are two sources of fixed-price quantity and expenditure: that from purchases available to *all* consumers, e.g. rice, by *coupon*; and that available to only some without coupons, e.g. fresh fruit and vegetables from employee co-ops, stores for the veterans, etc. The distributional network for the latter tends to favor *segments* of the poor, while the better off mainly account for market-price purchases of this group (see Supplementary Appendix 1, Table A2).

I start with an examination of equivalence scales independence from base utility, and with the question of corner solutions in consumption in Section 2. Econometric modeling of adult equivalence scales under wartime conditions is addressed in Section 3; Section 4 gives the estimation results. A conclusion (Section 5) sums up the main findings of the paper.

[1]The CBIHBS 84–5 was the first to be conducted annually under the current Republic. Logistic problems prevented its commencement on time; for this reason the first month of the survey contains twice as many households as the remaining 10 months.

2. Basic Issues

2.1. *Identification of Equivalence Scales*

I mentioned the identifying assumption of equality of scales across commodity groups for food share scales. However, there are additional problems facing the identification of equivalence scales. These stem from estimation with demand data when applied to households varying in demographics (see Pollak and Wales, 1979; see also Deaton and Muellbauer, 1980; Blundell and Lewbel, 1991), and from the standard Almost Ideal (AI) functional form, due to its approximate linearity (see Blackorby and Donaldson, 1993). This study employs non-linear AI as a solution to the latter.

When dependence of equivalence scales on income is not critical, it would be a desirable property of the scale to be independent of the base level utility at which demographic comparison is being made, and hence applicable at all levels of income. This property, known as Base Independence (BI), or Equivalence Scale Exactness (ESE), requires the cost function to be multiplicatively decomposable into two functions: one depending only on prices and utility; the other only on prices and household demographics (Lewbel, 1989; Blackorby and Donaldson, 1993; Donaldson and Pendakur, 2003). BI has testable implications that can be examined by introducing suitable functional forms for total expenditure and demographics into the food share equation.

There are, however, some functional forms for which the scales remain unidentified even when BI is assumed, or imposed as an empirical restriction. In particular, Blackorby and Donaldson (1993, sections 5 and 7) have argued that the imposition of BI on the expenditure function of the AI model cannot identify equivalence scales because that model's functional form is PIGLOG, that is log-linear in utility (or income). This has resulted in testing BI for scales obtained from extended AI forms that can deal with the identification problem. Banks *et al.* (1997) proposed an extended AI as a quadratic polynomial generalization of the Working–Leser model, henceforth QUAIDS, that nests the AI model and preserves its convenient flexible functional form. Similarly Blundell and Lewbel (1991) proposed testing BI by the inclusion of an interactive term between household size (number of children), and the log-linear income term in budget share equations. However Dickens *et al.* (1993) have demonstrated that the effect of the omitted second-order income term, when significant, is likely to be picked up by the interactive terms, resulting in a false rejection of BI. To overcome this problem, they propose to test the restriction by the inclusion of a log-linear income, its second-order term, and an interactive term. To be consistent with BI, the interactive term should be insignificant, after controlling for changes in the intercept due to differences in household size (see Blundell *et al.*, 1998). These studies, except for Pendakur (1999), find the BI hypothesis is strongly rejected. Indeed, rejection of BI is very common and its acceptance a rather rare outcome; see however Section 4.3. A further implication of BI is the invariance of the budget share equations across households of different composition. When BI changes in scales are independent of income, all households respond identically to a proportional change in income. Pendakur (1999) calls this condition *shape invariance*. This suggests that the shape of the Engel curves of the reference and non-reference households are just scaled version of each other, but they retain the same curvature, i.e. they have different intercepts.[2]

The above discussion suggests that QUAIDS can provide the basis for testing BI, provided we identify vertical from horizontal shifts in the curvature of the food share equations, that is, testing should check whether an interactive term between income and scale is statistically significant, given the inclusion of a size-interactive intercept term. I adopt this approach below, modified for income adjusted for coupon subsidies, as discussed below, in place of total expenditure. The strategy I follow is to employ QUAIDS as my baseline estimates for three different models of the demand for food; the model that results in the most plausible adult equivalence scales is then further tested for BI, in addition to its robustness to measurement errors, and for its correction to quality-adjustment in the price data employed.

2.2. *Corner Solutions*

I take up the issue of corner solutions for food consumption in this section. Assume "food" is rationed and available to all at quantities q_1 at the fixed price of p_0 up to a maximum amount $\bar{q}_1$ beyond which it can only be bought at market prices higher than p_0, such as p_1 (Deaton, 1986). "Non-food," representing all other commodities, is freely available at market prices p_2 in quantities q_2. Given corner solutions, the budget constraint would be

$$z = p_0 q_1 + p_2 q_2; \quad q_1 \le \bar{q}_1, \tag{1}$$

where z is total expenditure (or income). However, if ration resale is even semi-legal, then a parallel market for the sale of the unused rations $(\bar{q}_1 - q_1)$, trading at p_1, will develop. The demand function would then be unconditional, with the budget constraint

$$z = p_0 \bar{q}_1 + p_1 (q_1 - \bar{q}_1) + p_2 q_2 \quad or \quad z + (p_1 - p_0)\bar{q}_1 = p_1 q_1 + p_2 q_2. \tag{2}$$

Unlike (1), under (2) the definition of income is inclusive of ration subsidy (see Deaton, 1986). The question is which one is relevant to the survey under examination. The Iranian system of rationing broadly tolerates the market for resale of coupons, and although this had never been officially acknowledged, there is a *de facto* market in coupons, and the dealers operate conspicuously near distribution centers. Being conducted in commodity-specific coupon-currencies rather than in goods further facilitates coupon trading.[3] The annual per capita rationed averages for some key food commodities for three groups of cities (relative to their values in the capital, Tehran) were examined for this purpose (see Table A1). The rise of quantity averages with income is the dominant pattern,

[2]However, the importance of non-linear Engel curves to the identification of adults equivalence scales has long been recognized; see Muellbauer (1975b) on food scales; further discussed in Muellbauer (1987).

[3]The only subsidized food item without coupon was fresh milk, distributed only to households having a member under 2 or over 60, confined to Tehran, and had uneven distribution.

suggesting the averages are take-up, not consumption. Therefore, given resale, we observe $q_1 < \bar{q}_1$, but not $\bar{q}_1$ itself. The rich will of course be recorded as having spent, per capita, more on rationed coupon goods than poor households. The reason for this is, as income rises the quantity consumed comes closer to the (full entitlement) quantity purchased. Overall, the dominant pattern appears to suggest a general *absence* of corner solutions. Note also that the coupon food price rise under convertibility would ensure that the aggregate food market is relatively free from corner solutions if some of the shortages of the non-coupon segment spill into the coupon goods markets.

An important consequence follows from the above, namely that demand for food can be analyzed by a conventional function, provided that income is corrected for subsidies received on coupon goods.

3. Models

3.1. *The Food Share Models*

Section 2.2 examined a two-commodity budget constraint. More realistically, suppose good 1 is a composite food commodity, "rice." It is rationed at $\bar{q}_1$, and its price is $\bar{p}_1$. Good 2 is the other food commodity, which is not rationed. Good 3 is non-food, also not rationed. Goods 1 and 2 are necessities. Then we have

$$\bar{p}_1\bar{q}_1 + p_2q_2 + p_3q_3 = x. \tag{3}$$

Let there be an open market for the sale or purchase of the "rationed" rice; so that the market purchase or sale of rice for $\bar{q}_1$ take place at the market price p_1. Thus, the budget constraint is

$$p_1\bar{q}_1 + p_2q_2 + p_3q_3 = x + (p_1 - \bar{p}_1)\bar{q}_1 = z, \tag{4}$$

i.e., similar to (2). Assuming no transaction costs, consumers will just behave as if the price of rice is p_1 and the ration endowment $(p_1 - \bar{p}_1)\bar{q}_1$ is added to the budget. In terms of true consumption, perfectly conventional Marshallian demand functions result:

$$q_j = g_j(z, p_1, p_2, p_3). \tag{5}$$

The budget share equation for rice and for food in general will have a perfectly conventional form at given prices; assuming z is correctly measured.

Our dependent variable is the food share

$$w = \frac{(p_1q_1 + p_2q_2)}{z}. \tag{6}$$

However, one probably observes $w = \frac{(p_1\bar{q}_1 + p_2q_2)}{z}$ for many poor households where $q_1 < \bar{q}_1$. This is because the survey does not explicitly record resale of goods, or explicitly asks for *net* purchases.

In general, for a given household h,

$$(7) \qquad w_h = f\left(\frac{z_h}{m_h}, p_1, p_2, p_3\right) + \varepsilon_h,$$

where ε_h is a random error, and it is assumed that systematic differences between households are reflected in the equivalence scale m_h. The Engel model is a specialized version of (7) when prices are constant and can therefore be dropped from (7). This is the second food share model. Finally, it is further assumed that households may face different prices, so modifying (7) as

$$(8) \qquad w_h = f\left(\frac{z_h}{m_h}, p_{1hlt}, p_{2hlt}, p_{3hlt}\right) + \varepsilon_h,$$

where subscripts indicate that such prices may vary for each household h, and, additionally, for each location l and each month t. This is the third food-share model.

Since the food share is the welfare indicator in all three models above, food equivalence scales of each are implicitly defined by the equalization of the food share of a household whose characteristics define the base with that of a selected household with different characteristics. Furthermore, in the application to follow, the QUAIDS functional form of Banks *et al.* (1997) and its extensions will be adopted as an approximation for $f(\,)$. A number of important measurement issues will now be discussed. In the next section, I shall take up the practical solutions adopted.

There are three main problems here. One is that the ration subsidy $(p_1 - \bar{p}_1)\bar{q}_1$ needs to be imputed and added to x in (4); similarly for other subsidized goods sold below market prices. The other is the need to recognize explicitly that different households do not face the same free market prices. In other words, the conventional assumption in Engel curve or budget share analysis that households face the same prices is not, in fact, valid. Finally, there is the bias in the measurement of q_1 for poorer households to deal with.

$\bar{q}_1$ could be derived from averages, at the most disaggregated commodity level available, of ration quantities, varying by month, location, and household characteristics. One can then impute the average values to households in place of the actual observations in the observation period. This method should provide a good approximation for food coupon consumption when ration take-up is full. Moreover, additional market price consumption of rationed food also poses no problem as the survey records market price purchases separately.

A related issue is the problem of measurement error in z. The two-day recall period for some food commodities is bound to induce large measurement errors in total expenditure, especially under rationing. If there is some kind of regular income measure y, it may well be better to use it instead of total expenditure; then, $z = y + (p_1 - \bar{p}_1)\bar{q}_1$ where, of course, all imputed ration levels and corresponding price differences are incorporated in $(p_1 - \bar{p}_1)\bar{q}_1$. Since the survey contains income as well as expenditure information, I shall employ income data for y in my definition of z given below.

p_1 can be derived by averaging prices paid by different households, at a given location and time. We can adopt exactly the same procedure for p_2. Note, however, that these average values also contain quality effects since consumers pay different prices for different varieties of the same good. For this reason, such averages are referred to as unit values (see Deaton, 1988; Deaton *et al.*, 1994). For p_3, the present household budget survey does not contain non-food price information, but at the national level, the monthly official *non*-food CPI can be used as a non-food price index, though it is noted that this is not available by region (see Section 3.3).

The main component of z is monthly income after tax and pension payments, but inclusive of imputed rent on property for owner-occupier households. To income thus obtained must be added all the subsidies consumers receive on a variety of goods through various fixed-price channels, i.e. the sum of each fixed-price purchase valued at their corresponding price difference.[4]

For non-food commodities with no price information, I have scaled their fixed-price expenditure by the food group average ratio of market prices to fixed-prices (−1) in order to obtain a measure of their subsidies (see Appendix 2 for details).

One would also have to subtract sales of any goods to derive consumption. If sales are under-recorded or not recorded, there could be a bias in the relationship for the demand for the rationed good, that is, for poor households where $q_1 < \bar{q}_1$, $\bar{q}_1$ may be recorded instead of q_1. This would tend to bias upward the estimated z. In order to avoid this bias, one would have to limit the estimation sample to those households for which resale is likely to be negligible. I shall return to this issue in Section 3.4. I shall also assume that transactions costs are small enough not to make much difference to ration resale.[5]

The other main problem is the question of price differences across households, taken up in Section 3.2. Note that cross-section price variation by income group causes a specification problem for Engel curves, hence the employment of (8). However, (8) is needed anyway if prices vary by month and/or location.

There are two main components to the definition of food expenditure required to obtain food share as the dependent variable: coupon and non-coupon expenditures. Food share as the dependent variable in (6) can now be rewritten as

$$w_h = \frac{\sum_{j=1}^{k_1} p_{1j} q_{1jh} + \sum_{j=1}^{k_2} p_{2j} q_{2jh}}{z_h}. \tag{9}$$

[4]I also note that location/month coupon averages used for imputation must be defined over all observations, not just non-zero observations, to avoid imputing to households grossly inflated ration levels; and all purchases, zero and positive, are replaced by their corresponding imputed averages to avoid overstating the averages.

[5]The above can be contrasted with, for example, the Blundell and Meghir (1987) model of infrequency of purchase defined as a function of income. However, such a model would not be adequate for our purposes since an additional source of zero due to non-purchase is equally important. One would require a double-hurdle model in the present context, such as Deaton and Irish (1984). The latter has not led to tractable results, despite its simplifying assumption that infrequency purchases are independent of income. By contrast, the much simpler approach adopted in this study has proved very effective in dealing with infrequency of zeros (see Table 1, column 5), provided the sample selection rule employed adequately addresses the coupon resale issue.

The numerator of (9) defines total food expenditure, exclusive of tobacco, at market prices, the first term representing the coupon group and the second the non-coupon group. The imputation results in one observing q_{1jh} as the imputed ration rather than the imputed ration minus sales. Finally, (9) remains the definition of the food share for the model regardless of whether prices in (9) are of monthly/location or income-group specific type.

3.2. *Market Prices*

Consumer goods prices paid by households are influenced by variation over time and location, but also by variation in quality. In the Fisher and Shell view of "simple repackaging" quality variations, higher quality is just like having more of a good.[6] Then, in a cross-section the rich and poor face the same quality corrected prices, though the rich pay more per unit for higher quality. This would suggest using prices varying by time and location, but not by income. Variation by income in this view reflects quality differences, and the averages are unit values, not prices.

There is an alternative to this quality-based view, which, however, is only relevant to the non-coupon segment of the wartime food market. In a war economy with unequal purchasing power, the rich can more effectively adapt consumption to a price rise of non-rationed food, due to a shortfall of, or excess demand for, superior commodities by either *involuntary* substitution of cheaper goods, and/or by quality shading. In the former case, the non-quality adjustment will spill shortages into these additional markets, some of which will be absorbed by the market clearing prices of the convertible coupon goods, without leading to curtailment of food expenditure by the rich. By contrast, the poorer households, with a larger share of food in their budget, are expected to be highly sensitive to food price changes, less able to substitute for non-coupon food, and more likely to cut consumption, or drop out of the markets for some inferior food commodities altogether. Thus, for (non-coupon) normal goods, one may observe increasing demand, and hence, rising prices with income.[7] In this view, quintile averages are a combination of price and quality effects, reflecting changes in both shortage and quality for non-coupon goods.[8] This (non-coupon group) price hypothesis seems

[6]In this approach, a quality improvement is the same as having a larger package of the old good at the same price, or having to pay less per unit of the old good, and the required price adjustment for quality change is "simple," that is, independent of utility and all quantities consumed. These effects result in scaling each price, entered directly in the utility function, by its quality parameters defining varieties of each good (see Fisher and Shell, 1971; Muellbauer, 1975a). The application of this model, however, requires time-series of cross section data.

[7]A shortfall, perhaps even a modest one, coming from either the supply or demand side, may be the immediate cause, but inequality of purchasing power will exacerbate this by triggering a price spiral. There is no doubt that wartime profiteering by speculation and hoarding also contribute to such price rises, perhaps even leading to a huge market gap for the good. However, speculation usually feeds on shortages but is often not their principal cause and cannot provide an explanation of how shortages are generated and sustained over a longer period. In this regard, see Sen's (1981) analysis of the Great Bengal famine, separating its first, defining phase, based on falling purchasing power, from its second, speculative phase.

[8]However, note that voluntary non-consumption for coupon goods cannot be considered as being due to shortages; given trade in rations, coupon prices must be regarded as market clearing prices.

to draw some support from the discussion, in Appendix 1, of the evidence on the pattern of zeroes observed for non-coupon food and mostly accounted for by the two poor bottom total expenditure quintiles.

I employ three types of indices in order to examine the quality and price views. These are constructed as follows. To avoid a large number of "zero" or, more accurately, "missing" price observations, due to none and/or infrequent purchases, the list of food items used to construct p_1 and p_2 is based on food commodities purchased by at least 10 percent of the sampled households. I shall employ Laspeyres price indices for p_1 and p_2. All prices referred to below are market prices. I shall also employ household-specific prices, not just for the non-coupon group, but also for the coupon group. I deal with missing values for quintile prices by imputing values from non-missing observations; however, this requires justifying that missing values have a random pattern, i.e. unrelated to the variable (see Little and Rubin, 2002). Figure A1 in Appendix 3 tests for whether the missing data are independent of the quintile price variables. Given the random pattern of missing price values confirmed in Figure A1, all missing values have been filled by interpolation from the nearest observations.

The first type of price index assumes all income groups pay the same unit price for a good. These location/monthly indices are defined over 33 observations. Thus, for good j, in location l, at month t; with the base as $l = 1$, and $t = 1$, we have:

$$p_{1lt} = \frac{\sum_{j=1}^{k_1} p_{1jlt} \cdot \bar{q}_{1j1}}{\sum_{j=1}^{k_1} p_{1j11} \cdot \bar{q}_{1j1}} \tag{10}$$

$$p_{2lt} = \frac{\sum_{j=1}^{k_2} p_{2jlt} \cdot q_{2j1}}{\sum_{j=1}^{k_2} p_{2j11} \cdot q_{2j1}}. \tag{11}$$

Quantity weights in (10) are defined as the annual per capita ration level for each coupon good, whilst those in (11) are defined over total annual per capita consumption of non-coupon goods. The estimates based on (10) and (11) are examined in Section 4.1.

The second index I employ is of a total expenditure quintile type, using annual quantity weights for each quintile separately. Thus, for quintile i, with the base as $l/t/i = 1$:

$$p_{1lti} = \frac{\sum_{j=1}^{k_1} p_{1jlti} \cdot \bar{q}_{1j11}}{\sum_{j=1}^{k_1} p_{1j111} \cdot \bar{q}_{1j11}} \tag{12}$$

(13)
$$p_{2lti} = \frac{\sum_{j=1}^{k_2} p_{2jlti} \cdot q_{2j11}}{\sum_{j=1}^{k_2} p_{2j111} \cdot q_{2j11}}$$

Equations (12) and (13), each with 165 observations, are the main price indices employed in this study; the estimates based on them appear in Section 4.3.

Finally, I shall also employ a version of (12) and (13) corrected for quality variation. Prais and Houthakker (1955, ch. 8) suggested that a rough measure of quality effect can be obtained from the regression of the logarithm of prices on the logarithm of household per capita income. I follow a similar procedure carried out in two stages. First, I regress the logarithm of quintile "prices" (12) and (13) on the logarithm of household per capita income, and household demographics. The logarithms of quintile per capita income, using income and size values aggregated to quintile levels, were then multiplied by these estimates, and the results subtracted from the corresponding quintile unit values as defined in (12) and (13). Therefore, I ran

(14)
$$\ln P_{1lti} = v + \delta_1 \ln\left(\frac{z_h}{n_h}\right) + \sum_1^{n-1} \gamma_d \left(\frac{n_d}{n_h}\right)$$

(15)
$$\ln P_{2lti} = \upsilon + \delta_2 \ln\left(\frac{z_h}{n_h}\right) + \sum_1^{n-1} \gamma_d \left(\frac{n_d}{n_h}\right),$$

where d age groups for $n = 4$, defined as proportions of size, are: small child (0–6 years), older child (7–15 years), "middle-aged" adult (16–54 years), leaving adults above 54 as the omitted group. This results in "quality elasticity" of about $\hat{\delta}_1 = 0.03558$ (t-ratio = 6.39) for the coupon, and $\hat{\delta}_2 = 0.09822$ (t ratio = 14.02) for the non-coupon group. Then quality-adjusted prices are obtained from:

(16)
$$\ln p^*_{1lti} = \ln p_{1lti} - \hat{\delta}_1 \ln\left(\frac{z_{lti}}{n_{lti}}\right)$$

(17)
$$\ln p^*_{2lti} = \ln p_{2lti} - \hat{\delta}_2 \ln\left(\frac{z_{lti}}{n_{lti}}\right)$$

where superscript * denotes quality-corrected prices.

However, there may be an identification problem with the above method as it assumes that unit values increase to the extent of the full amount of the price increase. Deaton (1988) proposed a method to resolve this identification problem by exploiting regional unit value/price variation in budget surveys, since rural prices vary between isolated village markets, but at a village level, they remain constant; only quality (unit value) varies. However, the result of applying the proposed method to different developing countries has shown the impact of the correction, to the naïve method estimates, to be insignificant in all of them, even when an additional source of bias due to measurement error is also corrected (Deaton, 1990;

1997, p. 303). Although it would be difficult to apply such an approach to the urban survey of this study, the "prices" obtained from the naïve method remain nevertheless unidentified, due to a quality component, to an unknown degree, and one cannot safely work with unit values as though they were prices.

3.3. *Price and Demographic Specifications*

In order to obtain a clean estimate for size effect with the QUAIDS models, size parameters need to be incorporated into the quadratic term. For (8), this suggests the z_h terms should be written so as to ensure that the expression inside the brackets is the same for both the log-linear and its quadratic terms, that is,

$$\text{(18)} \quad w_f = \alpha_0 + \alpha_1\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right] + \alpha_2\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]^2 + \beta_1 \ln\left(\frac{P_{1lti}^f}{P_{3t}^{nf}}\right) + \beta_2 \ln\left(\frac{P_{2lti}^f}{P_{3t}^{nf}}\right) + \gamma_a\left(\frac{n_a}{n_h}\right) + \gamma_{bc}\left(\frac{n_{bc}}{n_h}\right) + \gamma_{sc}\left(\frac{n_{sc}}{n_h}\right) + \delta v + \varepsilon_f,$$

where ε_f is the error term, ζ_n is the size parameter, and $-\alpha_1\zeta_n$ is the compound coefficient of $\ln n_h$. However, unlike the linear AI in Deaton (1997), the non-linear estimation by QUAIDS employed for (18) returns 0–1 bounded estimates for the elementary parameter ζ_n, so various reported coefficients for size (ζ_n) are direct estimates of economies of scale in consumption, and values close to one indicate little scope for economies, those closer to zero substantially more so. The superscripts *f*, *nf*, and *g* stand for the food, the non-food, and the general price indices. Non-food prices are from published sources, so I have to assume $P_{3lti}^{nf} \approx P_{3t}^{nf}$; $P_{3lti}^{g} \approx P_{3t}^{g}$ for the aggregate prices (Central Bank of the Islamic Republic of Iran, 1984–85). v, a vector of auxiliary variables, such as regional or employment dummies, is also included in (18) in order to take into account factors whose influences are specific to the sample. Finally, note that I deflate the two food price indices in (18) by the CPI of non-food, P_{3lt}^{nf}, or more precisely by P_{3t}^{nf}. Some such deflation is required by the homogeneity restriction of a demand function (Deaton and Muellbauer, 1980). Moreover, (18) employs only two price indices, since the non-food index is not significant.[9] Equation (18) provides my main baseline QUAIDS estimates with quintile prices defined by (12) and (13).

I shall also employ two other versions of (18) to examine, in the light of my discussion in Section 3.2, the alternative price effect formulations. The first version, without price terms, is the extended Engel model based on the quadratic version of the commonly employed semi-logarithmic Working–Leser curve:

$$\text{(18}'\text{)} \quad w_f = \alpha_0 + \alpha_1\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right] + \alpha_2\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]^2 + \gamma_a\left(\frac{n_a}{n_h}\right) + \gamma_{bc}\left(\frac{n_{bc}}{n_h}\right) + \gamma_{sc}\left(\frac{n_{sc}}{n_h}\right) + \delta v + \varepsilon_f.$$

[9]With four prices, there should be three price ratios in (18). However, the non-food price term was dropped, as it was statistically insignificant.

The location/monthly-price version is

$$(18'') \quad w_f = \alpha_0 + \alpha_1\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right] + \alpha_2\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]^2 + \beta_1 \ln\left(\frac{P_{1lt}^f}{P_{3t}^{nf}}\right) + \beta_2 \ln\left(\frac{P_{2lt}^f}{P_{3t}^{nf}}\right) + \gamma_a\left(\frac{n_a}{n_h}\right) + \gamma_{bc}\left(\frac{n_{bc}}{n_h}\right) + \gamma_{sc}\left(\frac{n_{sc}}{n_h}\right) + \delta v + \varepsilon_f.$$

As shown below, a comparison of estimates from (18)–(18″) would suggest that (18) offers the most plausible adult equivalence scale values. I shall then further examine my results from (18) for BI; for robustness to measurement errors in income, and to correction to quality variation in price data. I now turn to these issues.

The discussion in Section 2.1 suggests that testing for BI with (18) should also be checked by separately identifying vertical from horizontal shifts in the curvature of (18), that is, by adding $\alpha_0 \cdot (\zeta_n \ln n_h)$ and $\alpha_3\left\{(\varsigma_n \ln n_h)\cdot\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]\right\}$ as interactive terms (see, for example, Blundell *et al*., 1998, equation 32); a significant estimate for the latter term would indicate rejection of the BI hypothesis. Accordingly, I employ an extended version of (18) with interactives as

$$(19) \quad w_f = \alpha_0 \cdot (\zeta_n \ln n_h) + \alpha_1\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right] + \alpha_2\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]^2 + a_3\left\{(\zeta_n \ln_h)\cdot\left[\ln\left(\frac{z_h}{p_t^g}\right) - \zeta_n \ln n_h\right]\right\} + \beta_1 \ln\left(\frac{P_{1lti}^f}{P_{3t}^{nf}}\right) + \beta_2 \ln\left(\frac{P_{2lti}^f}{P_{3t}^{nf}}\right) + \gamma_a\left(\frac{n_a}{n_h}\right) + \gamma_{bc}\left(\frac{n_{bc}}{n_h}\right) + \gamma_{sc}\left(\frac{n_{sc}}{n_h}\right) + \delta v + \varepsilon_f$$

I will also present estimates by a version of (19) by IV, using non-food total expenditure, and a few other ownership of durable instruments, to obtain 2SLS predicted values of income in order to check robustness of the scale estimates to any remaining measurement error income. Finally, I also estimate a version of (19) with quality-adjusted prices to check robustness to quality variation in price data based on (16) and (17) price indices, where superscript * indicates quality-corrected prices, that is,

$$(20) \quad w_f = \alpha_0 \cdot (\zeta_n \ln n_h) + \alpha_1\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right] + \alpha_2\left[\ln\left(\frac{z_h}{P_t^g}\right) - \zeta_n \ln n_h\right]^2 + a_3\left\{(\zeta_n \ln_h)\cdot\left[\ln\left(\frac{z_h}{p_t^g}\right) - \zeta_n \ln n_h\right]\right\} + \beta_1 \ln\left(\frac{P_{1lti}^{*f}}{P_{3t}^{nf}}\right) + \beta_2 \ln\left(\frac{P_{2lti}^{*f}}{P_{3t}^{nf}}\right) + \gamma_a\left(\frac{n_a}{n_h}\right) + \gamma_{bc}\left(\frac{n_{bc}}{n_h}\right) + \gamma_{sc}\left(\frac{n_{sc}}{n_h}\right) + \delta v + \varepsilon_f$$

I take a childless couple as the reference household, $n_0 = 2$ with income z_0; adding one adult/child to a two-adult household containing adults aged 16–54 (and keeping prices the same). Equations (18)–(18″) and (19)–(20) make the calculation of the scales dependent on z_0, here chosen to be the sample mean income $\bar{z}$. The quadratic equation, resulting from $w_0 = w_h$, has two solutions for $\ln z_h$ whose exponential values give z_h. The smaller of the two roots corresponds to the negatively sloped portion of the quadratic function. Dividing this solution by z_0 would then give a cost estimate of an additional person in each age group. However, calculating scales by setting $w_h = w_0$ does not mean that the influence of prices on them is omitted, as with the Engel model. If the referenced and selected households both face the same prices, then prices drop out. If they do not, then price terms will simply be added to the scale calculations. The adult equivalence costs, thus obtained, are examined in Sections 4.1 and 4.3.[10]

3.4. *Resale Sample Selection*

Finally, in order to avoid estimation bias resulting from imputing full rations to poor households who are likely to have some ration resale, I have tested whether the estimates differ across the sub-samples of the rich and poor. Here, the latter is defined as the bottom 30 percent of households in terms of income since this is where the effect of resale on estimates is most pronounced.[11] The estimates obtained from the applications of (18)–(18″), (19), and (20) are then made subject to a parameter stability pseudo F-test, similar to that employed for testing pooled samples, to find out if the 70 percent and 30 percent samples are likely to have different parameter estimates. The test is explained in the note to Table 1.

4. Estimation Results

I first examine the estimates based on the assumption that households face the same prices, given location and time. Throughout, I employ a sample of the semi-nuclear family consisting of parents and, if any, their immediate children, regardless of child age. This sample selection rule allows some variation on adult age group required to obtain adult costs while avoiding the undue estimation effect of non-nuclear households that have some features of their own, for example such households often have more employed adults.

4.1. *Households Facing Same Market Prices*

The estimation results, the corresponding equivalence scale, and economies of scale ("*lnn*") estimates by (18′) and ((18″) appear in Table 1, columns 1 and 2. The quadratic income terms are insignificant; the coupon price term is mildly significant but that for non-coupon food is insignificant, and there is little evidence for economies of scale. Moreover, the hypothesis of parameter stability across the 70

[10]The apparent income dependence of the scale in (18) is artificial and due to the linear and squared terms having different demographic deflators, but nothing fundamentally hinges on this specification. See Section 4.3 on testing for income-dependent scales.

[11]Early primary estimates with different sample divisions, at 25 or 40 percent of the bottom households, produced similar results.

TABLE 1

Estimates from Equations (18)–(18″) and (20) for Iran, 1984–85 (abs. t-values in parentheses)

	1	2	3	4	5	6
Incep	2.1274	3.5042	7.4702	–	–	–
	(2.87)	(3.94)	(4.00)			
Int*sz†	–	–	–	2.6494	4.7778	2.5856
				(0.89)	(0.49)	(0.87)
lninc	–0.2200	–0.2366	–1.0956	0.1208	0.1473	0.1621
	(1.53)	(1.65)	(3.37)	(2.69)	(2.03)	(3.55)
lninc^2	0.0054	0.0061	0.0409	–0.0085	–0.0109	–0.0096
	(0.77)	(0.88)	(2.87)	(2.25)	(1.75)	(2.52)
lnp_1	–	0.1157	0.1924	0.1923	0.2040	0.1750
		(2.73)	(5.01)	(4.99)	(5.03)	(4.46)
lnp_2	–	0.0342	0.4237	0.4137	0.4304	0.3834
		(0.90)	(13.26)	(12.97)	(11.76)	(11.52)
lnn	0.9693	0.9673	0.6308	0.3231	0.1694	0.3237
	(5.99)	(6.08)	(4.68)	(1.10)	(0.54)	(1.06)
Inc*sz†	–	–	–	–0.2128	–0.3660	–0.2065
				(0.95)	(0.51)	(0.92)
adults	0.0636	0.0644	0.0371	0.0361	0.0346	0.0398
	(2.51)	(2.54)	(1.15)	(1.09)	(1.01)	(1.19)
ch615	0.0583	0.0593	0.0072	0.0008	0.0127	0.0020
	(1.42)	(1.44)	(0.14)	(0.01)	(0.23)	(0.04)
ch0–6	0.0461	0.0468	0.0024	–0.0003	0.0086	–0.0021
	(1.02)	(1.04)	(0.04)	(0.01)	(0.15)	(0.04)
R^2	0.1453	0.1482	0.2301	0.2264	0.1813	0.2096
rmse	0.0395	0.0394	0.1862	0.1867	0.1921	0.1887
N	2788	2788	1936	1936	1936	1936
F-tst*	1.3799	1.3541	2.0580	1.6199	3.0117	2.4740
+1ad	1.48	1.48	1.29	1.30	1.27	1.33
+1cb	1.46	1.46	1.21	1.20	1.22	1.21
+1sc	1.40	1.40	1.20	1.20	1.21	1.20

Notes:

Column 1: *Full* sample with (18′), regressors include city dummies; Inc*sz estimate for addition of an interactive term, with Int*sz control, as defined below, is –0.0229/*t-ratio* = 5.82.

Column 2: *Full* sample with (18″), regressors include city dummies; Inc*sz estimate for addition of an interactive term, with Int*sz control, is –0.0280/*t-ratio* = 4.60.

Column 3: *70%* sample with (12), (13), and (18), regressors include city dummies.

Column 4: *Col. 3* equation replaced with its (19), with (12) and (13) prices.

Column 5: *Col. 4* equation estimated by 2S LS with non-food market-priced exp. as main instrument with 1st stage estimate as 0.2420 with t-ratio = 18.42.

Column 6: *Col. 4* replaced by (20) based on quality adjusted prices as defined in (16) and (17).

†Interactive terms are defined as $\alpha_0 \cdot \{(\varsigma_n \ln n_h)$, and $\alpha_3 \left\{ (\varsigma_n \ln n_h) \cdot \left[\ln\left(\frac{z_h}{P_t^g} \right) - \zeta_n \ln n_h \right] \right\}$; "Int*sz" stands for size-interacted intercept $\hat{\alpha}_0$, and "Inc*sz" stands for size-interacted income $\hat{\alpha}_3$ in columns 4–6.

*$F(q,(n_{70\%} + n_{30\%} - 2q)) = [(SSE_F - SSE_{70\%} - SSE_{30\%})/q]/[(SSE_{70\%} + SSE_{30\%})/(n_{70\%} + n_{30\%} - 2q)]$, where q is the number of parameters and F stands for the 100% sample; critical value for $F^{5\%}_{(28\ or\ 30\ or\ 33;\infty)} = 1.459$.

Samples confined to the head, spouse, and, if any, their immediate children regardless of age; except for variables indicated in the table, the same equation employed throughout. Sample mean of the food share: columns 1–2 = 0.424; columns 3–6 = 0.443.

vs 30 percent samples is *accepted*, suggesting robustness to coupon resale and imputation. Note that the equivalence scales are identical with and without location/monthly prices included in the food share equation. Moreover, household size and adults effects are significant, but the scale values (columns 1 and 2) are close to the head-count ratio of each household to that of the reference household, and this is reflected in the (close to 1) estimates for economies. These equivalence scale values must be regarded as implausibly large, typical of the magnitudes usually produced with the Engel model for developing countries, and attributed to the scale overestimation inherent to that model (see Nicholson, 1976; Deaton and Muellbauer, 1986). Nevertheless, according to Table 1, at least small children have smaller effects than older children, and adults.[12]

Two conclusions emerge. First, (18′) specification results in large estimates of child cost close to the head-count ratio. Second, including prices makes little difference to the scale estimates if the demand function is (18″); the Engel curve (18′) provides nearly identical approximations to (18″). However, it is possible to obtain plausible estimates that are BI, and prove robust to measurement error in income, and to quality correction to prices, from an alternative modeling of price effects incorporated in (18) and its extensions by (19) and (20). A brief examination of the pattern of such price effects will prove helpful to the estimation results based on (18), (19), and (20).

4.2. *Pattern of Quintile Market Prices*

An example of the regional graphs of the logarithm of (12) and (13) to non-food CPI, exactly as the price terms that will be used to estimate the food share equations (18) and (19) in the next section,[13] are shown in Figures 1 and 2 for Tehran; on non-food CPI, see Sections 3.1 and 3.3. In assessing the price patterns in these plots, it should be remembered that the explanation suggested for the existence of household-level prices does not apply to the coupon food group, as convertibility will ensure uniform prices across income classes, given time and location. The employment of quintile prices, and their graphical presentation for both coupon and non-coupon food, allows this contrast to be examined more clearly. The Tehran non-coupon group shows the curves for the top two quintiles are distinctly above those for the three bottom quintiles. By contrast, for the Tehran coupon group prices, there is hardly any distinct pattern between the rich and poor quintiles. This is of course what one would expect, given convertibility. More generally, the non-coupon food index for the top quintiles also tends to be higher than those for other quintiles in the large and small city groups. For instance, taking 0.1 (in Figure 2) as a dividing value for the top two quintiles, only three observations in the non-coupon group are below 0.1 in both large and small cities. This type of price behavior seems common in wartime economies. Brittain (1960) shows that the British consumer price index increased monotonically with

[12]The adult equivalence scales obtained by extending (18′) and (18″) into equations similar to (19) (inclusive of interactive terms) still result in large scale values, especially regarding the cost of an additional adult. Nonetheless, these are somewhat smaller than those reported in Table 1. They are available, but not reported here for two reasons. First, unlike the subsequent estimates below, the estimates by extended versions of (18′) and (18″) strongly reject the BI hypothesis (the precise estimates for α_3 obtained with equations inclusive of the interactive intercept, however, are reported at the bottom of Table 1, corresponding to those in columns 1 and 2 with the interactives added). Second, the economies of scale values obtained with the interactives added fall outside the expected upper bound of 1.

[13](12) and (13) for the bottom poor quintiles are not relevant to the estimation results in the next section as these results are obtained with a sample confined to the top 70 percent of income category.

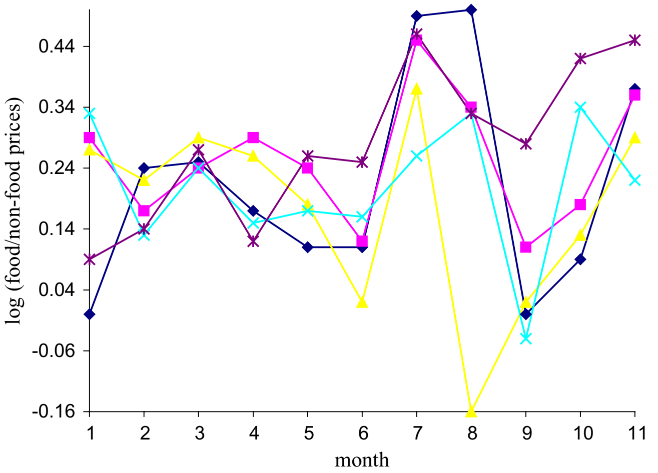

Figure 1. Tehran Coupon Food Price Index (Base: month = 1; quintile = 1)

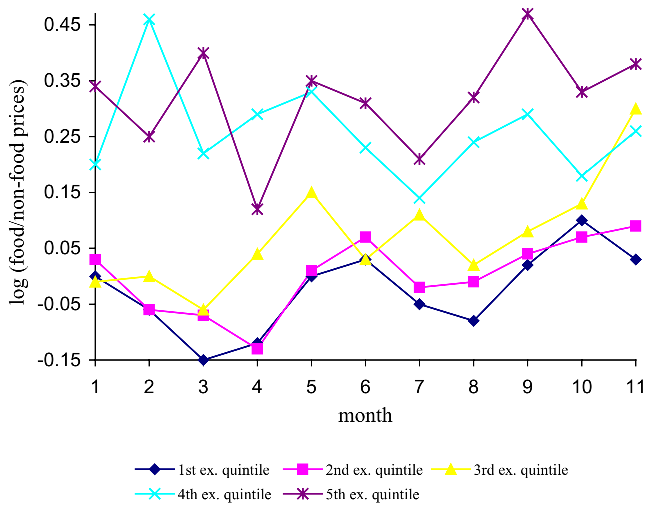

Figure 2. Tehran Non-Coupon Food Price Index (Base: month = 1; quintile = 1)

income during World War II.[14] More similar to the above are Henderson (1949–50, table X), in that the evidence displays, separately for the food group and in the aggregate, price increase by expenditure group; the aggregate prices following closely the changes in the food group prices. However, the pattern of coupon prices appears easier to explain by the market clearing prices due to convertibility.

[14]Brittain (1960, table 5) reports the following consumer price increases, with 1953 as base, in Britain for nine (weekly) income groups (in Pounds Sterling) during in 1938–46. Under 3: *49*; 3–6: *53*; 6–8: *59*; 8–10: *60*; 10–14: *61*; 14–20: *63*; 20–30: *66*; 30–50: *67*; 50 and over: *71*.

4.3. *Households Facing Different Market Prices*

Columns 3–6 represent the main estimation results of this paper based on the quintile prices of Section 3.2, column 3 retaining the same food share equation as in columns 1–2; columns 4–6 are with food share equations (19) and (20) inclusive of the interactive terms. All estimates are based on the top 70 percent sample, excluding the bottom 30 percent of poor households. I adopt this truncation in the light of the reported F-values, columns 3–6, *rejecting* the hypothesis of parameter similarity across the two sub-samples.

Column 3 is the (18) baseline QUAIDS (without the interactives). Both income terms α_1 and α_2 ("lninc" and "lninc^2" in Table 1, column 3) are now significant. Note, however, that the coupon price parameters β_1 ("$\ln p_1$"in Table 1), monthly *and* quintile, are similar in columns 2 and 3; it is the difference regarding the non-coupon price estimates β_2 ("$\ln p_2$"), which changes dramatically in magnitude and significance when monthly prices are replaced with quintile prices in the food share model; 0.0342 (0.90) in column 2, compared to 0.4237 (13.26) here. These estimates suggest food price elasticity not far from −1, but less close to −1 than those in column 2, as $\beta_i \approx 0$ imply −1; and larger in absolute value than (uncompensated) own price elasticities from developing countries.[15] The household size effect and age effects are all insignificant. Note, however, that the scale estimate of 0.63 (4.68) has now more plausible value. Note also that the F-statistic for parameter similarity across the 30 percent poor and 70 percent non-poor samples rejects the hypothesis, given the 5 percent critical value of 1.459 for $F_{(28,1936)}$. Finally, the costs of an additional person in each of the three age groups, at 1.29, 1.21, and 1.20, are much more plausible and very different from those reported in columns 1 and 2. The remaining columns of Table 1 are therefore devoted to a further scrutiny of column 3 estimates for identification and robustness to measurement error and to quality variation.

Column 4, the scale estimates, obtained from (19) inclusive of the interactive terms to test for BI, are quite close to those in column 3. The major differences relate to the additional estimates for the interactives, included in (19) in order to test whether the scales reported in column 3 are independent of income (or utility). These show that, after controlling for a size-interacted intercept ("int*sz") in column 4, the interaction term of logarithms of household size and household income ("Inc*sz") turns out *insignificant*, hence *failing to reject the BI hypothesis*. The income terms in column 4 are both significant. Once again, note that β_1 ("$\ln p_1$") estimates, from columns 2 and 4, are similar; it is the estimates β_2 ("$\ln p_2$") which differ sharply between the two columns. This suggests the main impact on equivalence scale values comes from the employment of the quintile-specific non-coupon price indices. Finally, note that the F-statistic for parameter similarity between the two (bottom and top income) samples is rejected yet again.

In column 5 I have repeated the estimation of the column 4 equation with the generalized IV method to check for robustness to any remaining measurement error in income, with an estimate of total non-food expenditure as the main

[15]These tend to be negative but often quite close to zero (see Pitt, 1983); especially so with regard to the aggregate food price elasticity, e.g. between −0.26 and −0.32 for India (Deaton, 1997).

instrument for income reported at the bottom of Table 1.[16] The results are similar to those reported in columns 3 and 4; in particular, the interactive term, of logarithms of income and size, at –0.366 (t-ratio = –0.51), remains insignificant, suggesting the column 5 scales are identified. This suggests that the imputation method employed has been very effective in dealing with infrequency measurement error. Note also that the hypothesis of parameter similarity between the 30 vs. 70 percent samples is rejected once again.

In column 6 I check for robustness of column 4 estimates to quality-adjusted prices, employing equation (20) with prices as defined in (16) and (17). Column 6 of Table 1 shows the results and they are similar to those in columns 4 and 5. The scales are identical to those in columns 3–5 for children, but slightly larger for adults. The term for logarithm of income squared ("lninc_2" in Table 1) remains significant; and once again the earlier scale identification results reappear with an insignificant size-income term (–0.2065, t-ratio = –0.92); evidence of independence of estimated scales from income.

This limited impact of quality change on the child costs, and on the estimates of estimates of scale, is to be expected, given the quality effect estimates of Section 3.2 and the evidence from budget surveys of other developing countries. Quality variation, at any rate, seems unlikely to provide the main component of an explanation for the magnitude of price effect encountered in this work. Bearing in mind the cautionary comments concerning the identification issue above (Deaton, 1997, p. 291) extensive examination concludes that the role of quality variation is "modest."

I end this section by noting that most studies report rejection of BI; failure to reject BI reported in this study appears to be a relatively rare outcome, particularly with parametric tests. A notable exception in support of the BI hypothesis is Pendakur (1999). He attributes such evidence mainly to the application of semi-parametric tests, and notes the inability of parametric tests in detecting the evidence in support of the BI; Gozalo (1997) also finds more limited support for the BI, but mostly from the parametric part of his evidence.[17]

5. Conclusion

In this paper, I obtained equivalence scale estimates for wartime Iran by employing a quadratic extension of the AI type demand function for food, and compared baseline QUAIDS estimates obtained from three models of equivalence scale; those based on quintile (non-coupon) food prices have the most plausible values. A key feature of the model employed is the correction of the budget constraint for coupon subsidies through imputation. Bearing in mind the limitations of this study on unavailability of non-food dual price information similar to those for food, and the possible shortcoming of the method employed to correct

[16]The full first-stage estimates of instruments used are available, but excluded to save space.

[17]Gozalo reports BI for the Engel food share scales (independent of prices). Pendakur's cross equation semi-parametric tests provide more limited support for BI than his single-equation results. He explains some of the differences in parametric and non-parametric test results by differences in consumption of child goods, but test results among household with varying numbers of children cannot be similarly explained. He also mentions that "noise in the data may be a factor." Note in this regard, the robustness of my estimates to measurement error in income discussed above with respect to Table 1, column 5 estimates.

for quality in price data, the main, quintile-price-based model has successfully dealt with measurement error in total expenditure, not a small gain, given the complexities of the issues related to multiple sources of zero expenditure. The model also explicitly recognizes that in a wartime economy, households are unlikely to face the same prices, and the modeling of this effect reveals an unusually strong cross-section price effect on the food share. The effect remains robust in terms of its scale estimates (given in Table 1, columns 3–6); the final resulting scales are obtained allowing for both vertical and horizontal demographic shifters. The quintile price effect proves important for overcoming the equivalence scales under-identification so common in the literature, adding another piece of relatively rare evidence consistent with BI; to my knowledge, the first piece of evidence on BI for a developing country. By contrast, BI is rejected for estimates based on a food share model with uniform prices, or on the Engel model without price effects.

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Supporting Information

Additional Supporting Information may be found in the online version of this article:

Appendix 1: Background Evidence
Table A1: Per Capita Ration Quantities of Principal Coupon Goods (Base: City = 1, quintile = 3)
Table A2: Frequency of Market Non-Purchase Out of 55 (11 Months by 5 Quintiles) for Each Food Item Consumed by At Least 10% of the Sample
Appendix 2: Non-Food Coupon Subsidies
Appendix 3: Testing of Missing Completely At Random Data for Market Food Prices
Figure A1: Stem-and-Leaf Display of t-Statistics Testing MCAR for Missing Market Food Prices; Top Three Quintiles