Discussion Paper ERU/2008 - 05
April 15, 2008

Estimating Equivalence Scales Through Engel Curve Analysis

Amita Majumder♣
Indian Statistical Institute, Kolkata

and

Manisha Chakrabarty♣
Indian Institute of Management Calcutta

Abstract

This paper proposes a simple two-step estimation procedure for Equivalence scales using Engel curve analysis based on a single cross section data on household level consumer expenditure. It uses Quadratic Logarithmic (QL) preferences with the maintained hypothesis of Generalized Equivalence Scale Exactness (GESE) (Donaldson and Pendakur, 2004). The novelty of the proposed procedure is that it neither requires any assumption on the form in which demographic attributes enter into the system of demands, nor any algebraic specification of the functions that appear in the budget share equations. More importantly, it does not require a computationally heavy estimation of complete demand systems. As an illustrative exercise the methodology is applied to Indian consumer expenditure data.

Keywords: Equivalence scales, Equivalence Scale Exactness (ESE), Generalized Equivalence Scale Exactness (GESE).

JEL classification numbers: C13, C21, D12, J16.

1. Introduction

Equivalence scale is defined as the relative cost of maintaining the same level of utility under different demographic regimes. It is, therefore, an inevitable element of welfare comparison between households. There are several theoretical and structural problems

♣ The authors thank Professor Dipankor Coondoo of the Indian Statistical Institute for helpful suggestions. The authors would also like to thank Professor Gauthier Lanot of Keele University, UK for his immense help and constructive suggestions. The usual disclaimer applies.

in the calculation and interpretation of *equivalence scales* and it is well known that equivalence scales are identifiable only under explicit assumptions. Muellbauer (1974) recognised one of the fundamental problems of identification of *equivalence scale.* He argued that welfare comparison across households require unconditional *equivalence scales* which is based on utility derived from both goods and household's fertility-decision on having children. But traditional budget data allow us to calculate only the conditional *equivalence scales*, where different equivalence scales can be consistent with the same preferences. The issue of identifiability of household *equivalence scale* has been discussed in many studies, which include Pollak and Wales (1979, 1981, 1992), Deaton and Muellbauer (1986), Fisher (1987), Lewbel (1989), Deaton, Ruiz-Castillo and Thomas (1989), Blundell and Lewbel (1991, 1994), Dickens, Fry and Pashardes (1993), Blackorby and Donaldson (1994), Lewbel (1997), Pendakur (1999) and Lewbel and Pendakur (2006).

Functional specification of the demographic vector augmented demand system plays an important role in the identification issue. *Equivalence scales* can not be recovered from demand behaviour in a single cross-section study (where there is no price variation) in case of a *rank-two* demand system with budget shares linear in logarithm of expenditure. Examples are PIGLOG[1] systems such as the Almost Ideal Demand System or the Translog demand system [Muellbauer (1974), Blackorby and Donaldson (1994), Pashardes (1995), Phipps (1998)]. Introduction of price variation also cannot solve this problem due to limited covariance between prices and demographic characteristics because while prices vary across period, variation in household characteristics occurs within a period [Dickens et al. (1993), Ray (1983)].

[1] The Price Independent Generalized Log-Linear (PIGLOG) systems are characterized by the cost function of the form $C(u,p) = \{b(p)\}^u a(p),$ where *p* is the price vector, *b(p)* is homogeneous of degree zero and *a(p)* is linear homogeneous in prices.

Deaton, Castillo and Thomas (1989) suggested parameter restriction such as Demographic Separability[2] (DS) as a remedial measure which imposes zero demographic substitution effect; but this restriction can yield biased estimates of equivalence scales. On the other hand, a *rank-three* demand system or a *rank-two* model that allows for non-linear log-expenditure effects on the budget share enables estimation of identifiable scales, where scales are invariant to the utility level at which the welfare comparisons are made, without the restriction of DS [Pashardes (1995)].

The property of invariance of equivalence scales to the utility level has been termed Independent of Base (IB) by Lewbel (1989) and Equivalence Scale Exactness (ESE) by Blackorby and Donaldson (1994). Formally, equivalence scales satisfy IB/ESE if and only if the cost function is separable in the utility level and the household attributes (Lewbel, 1989; Blackorby and Donaldson, 1993), implying that the equivalence scales depend only on prices and demographic composition. Although this property is frequently used in the literature, there is no rationale for assuming that the results of comparison should be equal for 'rich' and 'poor' households (Szulc, 2003; Donaldson and Pendakur, 2004).[3] Donaldson and Pendakur argue that there may be two reasons why equivalence scales should depend on total expenditure. "First, because economies of household formation are associated with sharable commodities such as housing whose expenditure share decreases as total expenditure rises, it is reasonable to expect expenditure-dependent equivalence scales for multi-person households to increase with expenditure. Second, because the consumption of many luxuries, such as eating in good

[2] An item-group is said to be demographically separable from a demographic group, if changes in the demographic structure within the demographic group exert only income-like effects on the goods in the item-group.

[3] In a cross-country study of equivalence scales by Lancaster, Ray and Valenzuela (1999) wide variation in equivalence scales across countries that span a wide range of per capita GNP has been observed.

restaurants or attending the theatre, are more enjoyable when done in groups, we may expect equivalence scales for households with more than one member to decrease with expenditure".[4] They propose a generalisation of ESE, which they call Generalised Equivalence Scale Exactness (GESE) that allows the scales to be different for rich and poor. They also show that if GESE is a maintained hypothesis, and the reference expenditure function is not PIGLOG, the 'equivalent expenditure function'[5] can be identified from demand behaviour.

In this paper we propose an estimation procedure for Equivalence scales using Engel curve analysis based on a single cross section data on household level consumer expenditure, the underlying system being Quadratic Logarithmic (QL) (Lewbel, 1990) in a GESE set up. The novelty of our procedure is that we work with the general form of QL system, without an explicit functional form involving the parameters. To be precise, the proposed methodology does not require any assumption on the form in which demographic attributes enter the system. Briefly, the estimation involves two steps. In the first step, the set of item-specific Engel curves relating budget shares to the logarithm of income is estimated for different demographic groups in a single equation framework using household level consumer expenditure data.[6] In the second step the equivalence scale for each demographic group is estimated using the coefficients of the item-specific Engel curves, estimated in the first step, considering commodities as

---

[4] Recent works of Koulovatianos et al. (2005a, 2005b) based on survey data also report evidence that equivalence scales are decreasing in income.

[5] Equivalent expenditure for a household is the expenditure level which would make the reference household as well off as the members of the household. Thus, equivalence scale is actual expenditure divided by equivalent expenditure.

[6] In fact, the proposed method does not require a computationally heavy estimation of complete demand systems. *Equivalence scales* in a system framework have been estimated by Pashardes (1995), Lancaster and Ray (1998), Szulc (2003), Majumder and Chakrabarty (2003), Lyssiotou and Pashardes (2004) and Donaldson and Pendakur (2006).

observations in a pooled regression of demographic groups and commodities. The validity of ESE assumption is then tested under this general set up.[7]

The paper is organized as follows: Section 2 sets out the estimation procedure for the equivalence scales; Section 3 describes the data used for the illustrative exercises done and presents the results; and finally, Section 4 concludes the paper.

2. The Proposed Procedure

The cost function underlying the Quadratic Logarithmic (QL) systems, namely, the Quadratic Almost Ideal Demand System (QUAIDS) of Banks, Blundell and Lewbel (1997) and the Generalized Almost Ideal Demand System (GAIDS) of Lancaster and Ray (1998), is of the form

$$C(u,p) = a(p).exp\left(\frac{b(p)}{(1/\ln u) - \lambda(p)}\right) , \qquad (1)$$

where $a(p)$ is homogeneous of degree one in prices, $b(p)$ and $\lambda(p)$ are homogeneous of degree zero in prices and u is the level of utility.

From (1), the demographic vector augmented Quadratic Logarithmic Indirect Utility Function can be written as:

$$V(p,y,z) = \left[\left(\frac{\ln y - \ln a(p,z)}{b(p,z)}\right)^{-1} - \lambda(p,z)\right]^{-1} , \qquad (2)$$

where y is income and z is the vector of demographic characteristics.

[7] It may be pointed out that as per the existing literature the test of the ESE property is conclusive only in case of rejection as suggested by Blundell and Lewbel (1991), Blackorby and Donaldson (1993, 1994). Murthi (1994) tested the restriction implied by exactness in the context of different parametric forms of engel curves on Sri Lankan data and in most of the cases exactness was not rejected. Pashardes (1995), on the other hand, found rejection of the hypothesis on UK data for the model he proposed. Gozalo (1997) and Pendakur (1994) proposed different nonparametric tests of the IB restriction on engel curves. Gozalo statistically rejected IB while Pendakur did not reject.

Donaldson and Pendakur (2004) showed that GESE with QL preference implies the following relations:

$$\ln a(p,z) = K(p,z)\ln a^0(p) + \ln G(p,z) \quad (3)$$
$$b(p,z) = K(p,z)b^0(p) \quad (4)$$
$$\lambda(p,z) = \lambda^0(p) \quad (5)$$

where *a(.)* is homogeneous of degree one in *p*, *b(.)* and $\lambda(.)$ are homogeneous of degree zero in *p*, $a^0(p) = a(p,z^0)$, $b^0(p) = b(p,z^0)$, and $\lambda^0(p) = \lambda(p,z^0)$, *0* being the reference household. It is evident from the above relationships that *K(p,z)* is homogeneous of degree zero in prices.

The logarithm of Equivalence scale under GESE is given by:

$$ln\,S(p,y,z) = \frac{(K(p,z)-1)\ln y + \ln G(p,z)}{K(p,z)}. \quad (6)$$

S(.) is increasing (decreasing) in *y* if *K(p,z)>1 (K(p,z)<1).*[8] ESE implies $K(p,z) = 1$, so that equivalence scale is independent of income, and in that case

$$ln\,S(p,z) = ln\,a(p,z) - ln\,a^0(p). \quad (7)$$

Now, applying Roy's identity to (2), the budget share equations are given by

$$w_i(p,y,z) = \frac{\partial\, ln\,a(p,z)}{\partial\, ln\,p_i} + \frac{\partial\, ln\,b(p,z)}{\partial\, ln\,p_i} ln\frac{y}{a(p,z)} + \frac{\partial \lambda(p,z)}{\partial\, ln\,p_i}\frac{1}{b(p,z)}(ln\frac{y}{a(p,z)})^2$$

or, $$w_i = \alpha_i(p,z) + \beta_i(p,z)\ln(\frac{y}{a(p,z)}) + \frac{\lambda_i(p,z)}{b(p,z)}(\ln(\frac{y}{a(p,z)}))^2$$

(8)

[8] A possible practical problem could be that for *K(p,z)<1*, the equivalence scale *S(p,z)* may turn out to be less than 1 for high values of *y* when *lnG(p,z)* is small.

where $\alpha_i(p,z) = \frac{\partial \ln a(p,z)}{\partial \ln p_i}$, $\beta_i(p,z) = \frac{\partial \ln b(p,z)}{\partial \ln p_i}$, $\lambda_i(p,z) = \frac{\partial \lambda(p,z)}{\partial \ln p_i}$ and *y* is the total expenditure.

Now, given household level consumer expenditure data, one can define specific demographic groups and classify each household as a member of certain demographic group. Thus, for commodity group *i* and demographic group *j* the household-level budget share equations (8) can be written as:

$$w_{ih}^j = \alpha_i(p,z^j) + \beta_i(p,z^j)\ln(\frac{y_h^j}{a(p,z^j)}) + \frac{\lambda_i(p,z^j)}{b(p,z^j)}(\ln(\frac{y_h^j}{a(p,z^j)}))^2 ; \qquad (9)$$

i=1,2,…,n; *j=0,1,2,….,J*; $h = 1,2,..., H^j$; where z^j is the demographic vector and H^j is the number of households in group *j,* respectively.

Rearranging the terms, equation (9) can be written as

$$\begin{aligned} w_{ih}^j = {} & [\alpha_i(p,z^j) - \beta_i(p,z^j)\ln a(p,z^j) + \frac{\lambda_i(p,z^j)}{b(p,z^j)}(\ln a(p,z^j))^2] \\ & + [\beta_i(p,z^j) - 2\frac{\lambda_i(p,z^j)}{b(p,z^j)}(\ln a(p,z^j)] y_h^{*j} + \frac{\lambda_i(p,z^j)}{b(p,z^j)} y_h^{*j^2} \end{aligned} \qquad (10)$$

where $y_h^{*j} = \ln(y_h^j)$.

Note that, for a single cross section data prices may be assumed fixed. Hence, equation (10) can be written as

$$w_{ih}^j = [\alpha_i^j - \beta_i^j \pi_j + \lambda_i^{*j}\pi_j^2] + [\beta_i^j - 2\lambda_i^{*j}\pi_j] y_h^{*j} + \lambda_i^{*j} y_h^{*j^2} \qquad (11)$$

where $\pi_j = \ln a(p,z^j)$.

Equivalently, $w_{ih}^j = a_i^j + b_i^j y_h^{*j} + c_i^j y_h^{*j^2}$ (12)

where $a_i^j = \alpha_i^j - \beta_i^j \pi_j + \lambda_i^{*j}\pi_j^2$
(13)

$$b_i^j = \beta_i^j - 2\lambda_i^{*j}\pi_j$$

(14)

$$c_i^j = \lambda_i^{*j}.$$

(15)

Thus, using the cross-section data, the following budget share equation for item *i* and demographic group *j* can be estimated (first stage estimation) taking households belonging to the demographic group as observations:

$$w_{ih}^j = a_i^j + b_i^j y_h^{*j} + c_i^j y_h^{*j^2} + \varepsilon_{ih}^j. \qquad (16)$$

In order to estimate the equivalence scales from (6) we need to have estimates of $K(p,z)$ and $lnG(p,z)$, which can be obtained from the parameter estimates of equation (16) and equations (3)-(5) as follows.

Note from equations (8), (11) and (15) that

$$c_i^j = \lambda_i^{*j} = \frac{\partial\lambda(p,z^j)}{\partial ln\, p_i}\frac{1}{b(p,z^j)}$$

or, $c_i^j = \dfrac{\partial\lambda^0(p)}{\partial ln\, p_i}\dfrac{1}{b(p,z^j)}$, since by GESE $\lambda(p,z^j) = \lambda^0(p)$ (from equation (5)).

Again,

$$c_i^0 = \lambda_i^{*0} = \frac{\partial\lambda^0(p)}{\partial ln\, p_i}\frac{1}{b^0(p)}$$

Hence, $\dfrac{c_i^0}{c_i^j} = \dfrac{b(p,z^j)}{b^0(p)} = K(p,z^j)$ by Equation (4)

or, $c_i^0 = K^j c_i^j$.

(17)

The estimates of K^j can be obtained by regressing $\hat{c}_i^0$ on $\hat{c}_i^j$ (without intercept) for each *j*, taking items as observations, *i=1,2,…,n.*[9] Hence ESE can be tested by testing the hypothesis that the slope coefficient =1 in this regression.

We now propose a simple method for estimating $ln G^j (= \pi_j - K^j \pi_0)$ under the additional assumption that $\beta_i^j = \beta_i + \gamma_j$, say. Now note from (14) and (15) that

$$b_i^j - b_i^0 = (\gamma_j - \gamma_0) - 2c_i^j \pi_j + 2c_i^0 \pi_0 . \qquad (18)$$

Given the estimates $\hat{b}_i^j$, $\hat{c}_i^j$, equation (18) is written as

$$\hat{b}_i^j - \hat{b}_i^0 = \gamma_j^* + 2\hat{c}_i^j (K^j \pi_0 - \pi_j) + e_i^j, \quad i = 1,2,...,n; \quad j = 1,2,...,J \qquad (19)$$

using equation (17), where e_i^j is a composite error term. Here again, it may be pointed out that although the relationship in (18) is exact, replacement of the variables by their estimated values yields a regression set-up. e_i^j is a linear combination of the individual errors of estimation of. $b_i^j, c_i^j, b_i^0, c_i^0$. Using arguments similar to those used in estimation of K^j, estimates of $ln G^j (= \pi_j - K^j \pi_0)$ and γ_j^* are obtained from a pooled regression of demographic groups and commodities.

Finally, given income $\tilde{y}$ and demographic group *j*, equivalence scale under GESE and a given price level, can be estimated using the following expression:

$$ln S(\tilde{y}, z^j) = \frac{(\hat{K}^j - 1) ln \tilde{y} + ln \hat{G}^j}{\hat{K}^j},$$

where $\hat{K}$ and $ln \hat{G}$ are the estimated values obtained from (17) and (19).

To obtain the standard error of this generalized expenditure dependent equivalence scale, for which the analytical expression is not possible to derive, we use bootstrap method to obtain the approximate standard errors.

[9] See Appendix A1 for an explanation for a regression set-up although the relationship in (17) is exact.

For estimation of equivalence scales under ESE, we proceed as follows. Note that under ESE, $c_i^0 = c_i^j \; \forall \; j$. After having obtained $\hat{c}_i^0$ from estimation of equation (16) for the reference group, the budget shares for the other demographic groups are now estimated by putting in this restriction for each *j*. This yields estimates of $\hat{a}_i^j$'s and $\hat{b}_i^j$'s under ESE. Estimates of $(\pi_j - \pi_0)$ are then obtained from the following regression equation[10]

$$\hat{b}_i^j - \hat{b}_i^0 = 2\hat{c}_i^0(\pi_0 - \pi_j) + e_i^j, \;\; i = 1,2,...,n; \quad j = 1,2,...,J \,. \tag{20}$$

3. Data and Results

The data for the present analysis have been taken from the data collected by the National Sample Survey Organization (NSSO), India, in its 61st round enquiry on Employment-Unemployment during July, 2004 – June, 2005. The data provide information on household characteristics, demographic particulars and employment status at the individual level within each household surveyed. In addition, this survey also provides data on consumption expenditure on several detailed items and total expenditure. Since the estimation is based on a single cross section data, prices are assumed fixed. We also assume that all demographic groups face the same price. Data for only the urban sector have been used to illustrate the estimation procedure described in Section 2. The all India urban data we consider here consist of 5959 households comprising only three types of households based on demographic composition[11]. The

[10] There will be no constant term here in view of the fact that under ESE $b(p,z) = b^0(p)$, which implies $\beta_i^j = \beta_i$ for all *j*.

[11] Here All-India refers to 15 major states, viz., Andhra Pradesh, Assam, Bihar, Gujarat, Harayana, Punjab, Karnataka, Kerala, Madhya Pradesh, Maharashtra, Orissa, Rajasthan, Tamil Nadu, Uttar Pradesh and West Bengal.

reference households are taken to be those consisting of 2 adults only. The two other household-types are households consisting of (i) 2 adults plus 1 male child (0-17 years) and (ii) 2 adults plus 1 female child (0-17 years). These groups consist of 3321, 1513, 1125 households, respectively. We consider 10 commodity groups, namely, (i) Cereals and cereals substitutes, (ii) Milk and milk products, (iii) Edible oils, (iv) Meat, fish & egg, (v) Sugar & salt, (vi) Other food, (vii) Pan, tobacco & intoxicants, (viii) Clothing & footwear, (ix) Services and (x) Other non-food[12].

The estimation procedure involves first estimating equation (16) for 10 commodity groups $(i = 1,2,,....10)$ and for the three household types $(j = 0,1,2)$ mentioned earlier. The variable y_h^{*j} denotes logarithm of total expenditure of household h belonging to the demographic group j. The regression results obtained from estimating equation (16) are presented in Table 1. It is observed that except in cases of 'Other food', 'Pan, tobacco & intoxicants' and 'Clothing & footwear', for all other items most of the coefficients turn out to be significant.

The estimated values of K for two household types, viz., households with 2 adults plus 1 male child and households with 2 adults plus 1 female child, taking the two-adult household as numeriare, are reported in Table 2. The results of the test for ESE are also presented. It is evident from the results that ESE is rejected at 5% level of significance for this data set.

Our next step of estimation involves estimating equation (19), from which estimate of $log\, G^j = -(K^j \pi_0 - \pi_j)$ can be obtained directly as the coefficient of $2\hat{c}_i^j$. The estimated values of $log\, G^j$ turn out to be 2.946 and 2.876 for demographic groups 1 and 2, respectively.

[12] "Other food" includes beverages, processed foods, vegetables and fruits. "Other non food" includes fuel and light, entertainment, education, medical, transport, rent & tax, personal care, toilet article, sundry

Finally, we calculate log equivalence scale for demographic group *j* at income level $\tilde{y}$ from the expression $ln\,S(\tilde{y},z^j) = \frac{(K^j - 1)ln\,\tilde{y} + ln\,G^j}{K^j}$ by using the estimated values of both K^j and $\log G^j$. The equivalence scales at different levels of income for the two household types are presented in Table 3. The minimum value for income has been chosen to be a value close to the sample minimum. We report equivalence scale up to the income level of Rs.10,000/- basically for two reasons. First, only 2% of the sample fall beyond this level; and second, the value of the equivalence scale starts to become implausible (less than one) at this level. However, as pointed out in footnote 8, this could be due to a problem with the GESE set up itself.

Note that given $K^j < 1$, the equivalence scale is a decreasing function of income by construction as observed in Table 3. This corroborates the findings from the studies of Donaldson and Pendakur (2004, 2006) using Canadian data. Similar results have been obtained through a subjective (survey) method for evaluating equivalence scales using data from Germany and France (Koulovatianos et al., 2005a) as well as from Cyprus (Koulovatianos et al., 2005b). The result implies that the cost of raising a child *relative to the income level* is much higher for a poorer household than for a richer household, a scenario that fits well into the Indian context. The fact that a child is indeed a ‘burden’ for a ‘poor’ household in India, is reflected through the high values of equivalence scales at the lower end of the income distribution. The bootstrapped estimates of standard errors (from 2000 re-samples) reveal that except for the lowest income group, almost all values are significant.

For comparability with other Indian studies, the ESE equivalence scales are also presented in Table 3. The values 0.319 for boys and 0.376 for girls indicate that boys cost less than girls in an overall sense. This observation is in line with the finding by

---

article.These items have been merged to avoid too many zero observations.

Lancaster, Ray and Valenzuela (1999) who obtain equivalence scales (averaged over three children groups, viz., 0-4 years, 5-14 years, 15-17 years) to be 0.171 for boys and 0.192 for girls under a Rank 3 demand system for India. Similar pattern has also been noted by Chakrabarty (2000) for the state of Maharashtra (India). Here the Engel equivalence scale for a boy (0-14 years) turns out to be 0.502 and that for a girl (0-14 years) turns out to be 0.569, and the corresponding Rothbarth scales turn out to be 0.047 and 0.069 under a QL budget share curve.

4. Conclusion

In this paper we have proposed a simple estimation procedure for Equivalence scales in a GESE set up, using Engel curve analysis based on a single cross section data on household level consumer expenditure where the budget shares are Quadratic Logarithmic (QL) in income. The novelty of our procedure is that no explicit algebraic form for the coefficients of the Engel curves (which are functions of demographic variables[13]) is required. In other words, the proposed method, which is a two-step procedure for estimating equivalence scales, does not require any assumption on the form in which demographic attributes enter the system of demands. More importantly, the proposed method does not require a computationally heavy estimation of complete demand systems. As an illustrative exercise the methodology is applied to a limited number of demographic groups where children of 0-17 years of age have been clubbed into one group. The procedure is, however, extendable to any number of groups, subject to availability of data in each demographic group.

From the test of validity of ESE assumption it emerges that ESE is rejected on Indian data and the generalized equivalence scale is found to be inversely related to income, a result that corroborates the findings of other studies on developed and underdeveloped countries. It is also observed that boys cost less than girls.

References

Banks, J., R. Blundell and A. Lewbel (1997): Quadratic Engel Curves and Consumer Demand, *Review of Economics and Statistics*, 79, 527-539.

Blackorby, C. and D. Donaldson (1993): Adult-equivalence Scales and the Economic Implementation of Interpersonal Comparisons of Well-being, *Social Choice and Welfare*, 10, 335-361.

Blackorby,C. and D. Donaldson (1994): Measuring the Cost of Children: A Theoretical Framework in Blundell, R., I. Preston and I. Walker (eds.) The Measurement of Household Welfare, Cambridge University Press, Cambridge, 51-69.

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

Chakrabarty, M. (2000): Gender-Bias in Children in Rural Maharashtra - An Equivalence Scale Approach, *Journal of Quantitative Economics*, 16(1), 51-65.

Deaton, A. and J. Muellbauer (1986): On Measuring Child cost: With Application to Poor Countries, *Journal of Political Economy*, 79, 481-507.

Deaton, A., J. Ruiz-Castillo and D. Thomas (1989): The Influence of Household Composition on Household Expenditure Patterns: Theory and Spanish Evidence, *Journal of Political Economy*, 97, 179-200.

Dickens, R., V. Fry and P. Pashardes (1993): Nonlinearities, Aggregation and Equivalence Scales, *Economic Journal*, 103, 359-368.

Donaldson, D. and K. Pendakur (2004): Equivalent-expenditure Functions and Expenditure Dependent Equivalence Scales, *Journal of Public Economics*, 88, 175-208.

Donaldson, D. and K. Pendakur (2006): The Identification of Fixed Costs from Consumer Behaviour, *Journal of Business and Economic Statistics*, 24, 255-265.

Fisher, F. M. (1987): Household Equivalence Scales and Interpersonal Comparisons, Review of Economic Studies, 54, 519-524.

Gozalo, P. (1997): Nonparametric Bootstrap Analysis with Applications to Demographic Effects in Demand Functions, *Journal of Econometrics*, 81, 357-393.

Koulovatianos, C., C. Schröder and U. Schmidt (2005a): On the Income Dependence of Equivalence Scales, *Journal of Public Economics*, 89, 967-996.

Koulovatianos, C., C. Schröder and U. Schmidt (2005b): Properties of Equivalence Scales

[13] As we are dealing with cross section data, prices are assumed fixed.

in Different Countries, *Journal of Economics*, 86,19-27.

Lancaster, G. and R. Ray (1998): Comparison of Alternative Models of Household Equivalence Scales: The Australian Evidence on Unit Record Data, *The Economic Record*, 74, 1-14.

Lancaster, G., R. Ray and M. R. Valenzuela (1999): A Cross-Country Study of Equivalence Scales and Expenditure Inequality on Unit Record Household Budget Data, *Review of Income and Wealth*, 45, 455-482.

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

Lewbel, A. (1990): Full Rank Demand Systems, *International Economic Review*, 31, 289-300.

Lewbel, A. (1997): 'Consumer Demand Systems and Household Equivalence Scales' in Pesaran, M.H. and M.R. Wickens (eds.) *Handbook of Applied Econometrics*, Vol. II (Microeconomics), 167-201.

Lewbel, A. and K. Pendakur (2006): Equivalence Scales: Entry for *The New Palgrave Dictionary of Economics*, 2nd Edition.

Lyssiotou, P. and P. Pashardes (2004): Comparing the True Cost of Living Indices of Demographically Different Households, *Bulletin of Economic Research*, 56, 21-39.

Majumder, A. and M. Chakrabarty (2003): Relative Cost of Children : The Case of Rural Maharashtra, India, *Journal of Policy Modeling*, 25, 61-76.

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

Murthi, M. (1994): Engel Equivalence Scales in Sri Lanka: Exactness, Specification, Measurement Error in Blundell, R., I. Preston and I. Walker (eds.) The Measurement of Household Welfare, Cambridge University Press, Cambridge.

Pashardes, P. (1995): Equivalence Scales in a Rank-3 Demand System, *Journal of Public Economics*, 58, 143-158.

Pendakur, K. (1999): Estimates and Tests of Base-independent Equivalence Scales, Journal of Econometrics, 88, 1-40.

Phipps, S. (1998): What Is the Income "Cost of Child"? Exact Equivalence Scales for Canadian Two-Parent Families, Review of Economics and Statistics, 80, 157-164.

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

Pollak, R. and T. Wales (1981): Demographic Variables in Demand Analysis, *Econometrica,* 49, 1533-1551.

Pollak, R. and T. Wales (1992): Demand System Specification and Estimation, Oxford

University Press, London.

Ray, R. (1983): Measuring the Costs of Children: An Alternative Approach, *Journal of Public Economics*, 22, 89-102.

Szulc, A. (2003): Is It Possible to Estimate Reliable Household Equivalence Scales?, *Statistics in Transition*, 6, 589-611.

Table 1: Parameter estimates of Engel curves

Commodity groups	Reference Group (2 adults)			Household type 1 (2 adults + 1 male child)			Household type 2 (2 adults+1 female child)		
	$\hat{a}_0$	$\hat{b}_0$	$\hat{c}_0$	$\hat{a}_1$	$\hat{b}_1$	$\hat{c}_1$	$\hat{a}_2$	$\hat{b}_2$	$\hat{c}_2$
Cereals & cereals substitute	1.903*** (0.111)	-0.358*** (0.027)	0.017*** (0.002)	2.354*** (0.197)	-0.454*** (0.047)	0.022*** (0.003)	2.313*** (0.221)	-0.444*** (0.053)	0.022*** (0.003)
Milk	-1.072*** (0.102)	0.298*** (0.025)	-0.019*** (0.002)	-0.906*** (0.173)	0.246*** (0.042)	-0.015*** (0.003)	-1.535*** (0.166)	0.399*** (0.040)	-0.025*** (0.002)
Edible oil	0.263*** (0.036)	-0.037*** (0.009)	0.001** (0.0005)	0.152*** (0.053)	-0.013 (0.013)	-0.0004 (0.001)	0.207*** (0.075)	-0.022 (0.018)	0.0002 (0.001)
Egg, fish & meat	-0.181* (0.076)	0.070*** (0.019)	-0.005*** (0.001)	-0.461*** (0.132)	0.135*** (0.032)	-0.009*** (0.002)	-0.117 (0.136)	0.053 (0.033)	-0.004** (0.002)
Salt & sugar	0.299*** (0.032)	-0.045*** (0.008)	0.002*** (0.0005)	0.146*** (0.053)	-0.008 (0.013)	-0.001 (0.001)	0.284*** (0.066)	-0.042*** (0.016)	0.001 (0.001)
Other food products	0.097 (0.141)	0.033 (0.035)	-0.004* (0.002)	-0.027 (0.201)	0.060 (0.049)	-0.005* (0.003)	-0.031 (0.139)	0.022 (0.034)	-0.002 (0.002)
Pan, tobacco & intoxicants	-0.274** (0.125)	0.087*** (0.032)	-0.006*** (0.002)	-0.197 (0.203)	0.067 (0.051)	-0.005 (0.003)	0.018 (0.183)	0.011 (0.045)	-0.001 (0.003)
Sevices	0.055 (0.154)	-0.042 (0.040)	0.005** (0.003)	0.643*** (0.215)	-0.185*** (0.054)	0.014*** (0.003)	0.519** (0.264)	-0.153** (0.066)	0.012*** (0.004)
Clothing & footwear	-0.042 (0.059)	0.035** (0.015)	-0.003*** (0.001)	-0.010 (0.111)	0.030 (0.027)	-0.003 (0.002)	0.072 (0.113)	0.008 (0.028)	-0.001 (0.002)
Other non-food	-0.685** (0.338)	0.179** (0.088)	-0.006 (0.006)	-1.611*** (0.552)	0.406*** (0.138)	-0.020** (0.009)	-1.442** (0.584)	0.359** (0.147)	-0.017* (0.009)

Note: Standard errors are reported in parenthesis.

*, ** and *** indicate significance at 10%, 5% and 1% levels, respectively.

Table 2: Estimated values of K for two household types $(j = 1,2)$

Household type 1 (2 adults + 1 male child (0-17 years))	Household type 2 (2 adults+1 female child (0-17 years))
0.6605 (Standard error = 0.1150) R^2 =0.778	0.6842 (Standard error = 0.0809) R^2 =0.884
H_0: K=1 $\lvert t_9 \rvert = 2.95$, p-value: 0.016	H_0: K =1 $\lvert t_9 \rvert = 3.91$, p-value: 0.004

Table 3: Equivalence scales

Income Level (Rs.)	**GESE** Equivalence Scales	
	Household type 1 (2 adults + 1 male child (0-17 years))	Household type 2 (2 adults + 1 female child (0-17 years))
1000	2.504 (4.314)	2.696 (2.534)
1500	2.034 (1.314)	2.230** (0.942)
2500	1.566*** (0 .349)	1.757 *** (0.286)
5000	1.098*** (0.261)	1.271 *** (0. .211)
10000	0.770** (0.299)	0.919*** (0. 264)
ESE scale	1.319	1.376

Note: A two-adult household has a value 1. Bootstrapped standard errors are in parentheses.
*, ** and *** indicate significance at 10%, 5% and 1% levels, respectively.

Appendix

A1

From equation (17) we have $c_i^0 = K^j c_i^j$.

To estimate K^j we replace c_i^j 's by their estimated values. Let $\hat{c}_i^j = c_i^j + \delta_i^j$, where δ_i^j 's are the errors.

Then, $\hat{c}_i^0 - \delta_i^0 = K^j(\hat{c}_i^j - \delta_i^j)$

Or, $\hat{c}_i^0 = K^j \hat{c}_i^j + (\delta_i^0 - \delta_i^j K^j)$

Or, $\hat{c}_i^0 = K^j \hat{c}_i^j + \delta_i^{j^*}$, say. (*)

Note that the regression error is assumed to be present only because of estimation errors in the first stage. Since the first stage estimates are unbiased and consistent, asymptotically equation (*) would hold exactly. Now, as the observations here are over *items*, the itemwise errors can be assumed to be uncorrelated with the regressor as the estimation errors originate from estimation of itemwise budget shares separately.