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Understanding Consumption Patterns[1]

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Abstract: The analysis of consumer demand is one of the major successes of economics as it represents the near perfect marriage of theory and econometrics. This paper reviews, distills and systematises some of the major empirical findings on consumption patterns, concentrating in particular on the more recent (and, in some cases, more controversial) evidence. One of the key conclusions of the paper is that on the basis on new methods, the hypotheses of homogeneity, symmetry and preference independence are not at such wide variance with the data as was once thought to be the case.

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1 Introduction

The study of consumption patterns is important for a number of reasons. First, as total consumption absorbs more than 70 percent of GDP in most countries, it is the largest of the macroeconomic aggregates, thus having great significance for the state of the economy as a whole and business conditions. Second, the pattern of consumption contains a wealth of useful information regarding economic welfare and living standards. Closely allied to this is that as consumption (both current and future) is the ultimate objective of all economic activity and economic systems (mercantilists notwithstanding), in a fundamental sense consumption patterns are an objective way of measuring and assessing economic performance. Finally, an understanding of the price-responsiveness of consumption is of crucial importance for a host of microeconomic policy issues including public-utility pricing, the measurement of distortions, optimal taxation and the treatment of externalities.

It is partly for these reasons that the analysis of consumption has attracted the attention of some of the best minds in economics and econometrics. Additional reasons which account for the extent of sophisticated econometric analy-

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sis of consumption patterns include advances in econometric methodology and computing technology, as well as the availability of large-scale databases, both time series and cross sectional (Theil, 1980). But perhaps the most important is the near perfect marriage of theory and econometrics offered by consumer demand, a situation almost unparalleled in any other field of economics. The utility-maximising theory of the consumer gives rise to demand equations which can be aggregated over individuals to yield market demand curves which, under certain conditions (much weaker than usually believed), have more or less the same properties, enabling them to be applied to aggregate data. The hypotheses derived from utility theory such as homogeneity and symmetry can then be tested econometrically, so that there is a smooth transition from theory to application.

Major book-length works on the theory and measurement of consumer demand include Barnett (1981), Bewley (1986), Deaton (1975), Deaton and Muellbauer (1980a), Goldberger (1987), Lluch, Powell and Williams (1977), Phlips (1974), Pollak and Wales (1992), Powell (1974), Theil (1975/76), Theil, Chung and Seale (1989), Theil and Clements (1987) and Theil and Suhm (1981). Survey articles of this area are provided by Barten (1977), Blundell (1988), Brown and Deaton (1972) and Deaton (1986).

The objective of this paper is to review, distill and systematise some of the major empirical findings on consumption patterns, concentrating in particular on the more recent (and, in some cases, more controversial) evidence. In order to make the paper accessible to nonspecialists, the exposition is kept as nontechnical as possible. We start by using price-quantity data for the OECD to introduce indexes of consumption patterns. Then follows a discussion of Engel's law (and its modern extension) and recent evidence on homogeneity and symmetry. Later parts of the paper deal with preference independence, the issue of the constancy of tastes and aspects of functional form of demand equations. Most of the discussion deals with results from time-series/cross-country data. We thus do not cover the interesting recent developments with microdata (see, e.g., Blundell, 1988). This choice reflects considerations of space, as well as our view that the analysis of aggregate time-series data is still not a well-understood and settled area.

2 OECD Consumption

In this and the next section, which draw on S. Selvanathan (1988a) and Stening (1985), we introduce some basic concepts by using consumption data from most of the OECD countries. Table 1 gives some broad charactistics of the database. There are 18 countries with a bit less than 20 annual observations in most.

Let p_{it} be the price and q_{it} the per capita quantity consumed of good i during year t. Let there be n goods, so that total expenditure is $M_t = \sum_{i=1}^{n} p_{it} q_{it}$ and the

Table 1. Characteristics of the OECD database

Country (1)	Sample period (2)	Sample size (3)	Per capita GDP in 1975: International dollars (4)	Per capita GDP in 1975: (4) with US = 100 (5)	Number of commodity groups (6)
1. US	1960–1981	21	7132	100	10
2. Canada	1960–1981	21	6788	95	10
3. Sweden	1964–1981	17	6749	95	10
4. Switzerland	1960–1981	21	6082	85	9
5. Denmark	1966–1981	15	5969	84	10
6. Australia	1960–1981	21	5919	83	10
7. France	1964–1981	17	5864	82	10
8. Germany	1960–1981	21	5758	81	8
9. Belgium	1960–1981	21	5554	78	10
10. Norway	1964–1981	17	5419	76	10
11. Netherlands	1952–1977	25	5321	75	10
12. Iceland	1960–1973	13	5201	73	10
13. Finland	1960–1977	17	5192	73	10
14. Austria	1964–1981	17	4994	70	10
15. Japan	1970–1981	11	4905	69	8
16. UK	1964–1981	17	4601	65	10
17. Spain	1964–1977	13	4032	57	10
18. Italy	1964–1981	17	3870	54	10

The GDPs in international dollars are from Summers and Heston (1984) and are computed on the basis of the purchasing powers of national currencies, not prevailing exchange rates. Sample size is after lagging.

proportion of total expenditure devoted to commodity i is $w_{it} = p_{it}q_{it}/M_t$. This w_{it} is called the budget share of good i. For brevity, we shall refer to M as "income". Table 2 presents the budget shares at sample means for each commodity in the 18 countries. For example, looking at the sixth row we see that, on average, Australians spend 19 percent of their income on food, 9 percent on beverages, 9 percent on clothing, 16 percent on housing and so on. The last row of the table presents the budget share of each commodity averaged over the 18 countries. As can be seen, the average OECD consumer spends 24 percent of income on food; and food, housing and transport together occupy about 50 percent of the total in most countries.

The log-change in the price of i is defined as $Dp_{it} = \log p_{it} - \log p_{i,t-1}$. When multiplied by 100, this log-change is approximately the annual percentage change. Columns 2–11 of Table 3 contain the mean price log-changes. The last row reveals that, on average, all prices grow between 6 and 9 percent per annum. The Divisia price index is a budget-share-weighted average of the price log-changes,

$$DP_t = \sum_{i=1}^{n} \bar{w}_{it} Dp_{it} \ , \tag{2.1}$$

Table 2. Budget shares of 10 commodities for 18 countries (Means × 100)

Country (1)	Food (2)	Beverages (3)	Clothing (4)	Housing (5)	Durables (6)	Medical care (7)	Transport (8)	Recreation (9)	Education (10)	Miscellaneous (11)
1. US	14.38	4.24	8.03	19.29	7.28	9.46	15.93	6.25	1.81	13.33
2. Canada	16.42	6.21	8.41	18.85	8.73	3.84	15.02	5.96	2.33	14.22
3. Sweden	20.72	7.96	8.41	22.79	7.66	2.27	13.65	9.18	.18	7.18
4. Switzerland	22.73	9.42	6.77	18.06	7.80	6.52	10.99	9.36		8.35
5. Denmark	19.18	9.09	6.97	20.92	8.98	1.92	15.04	7.88	.83	9.19
6. Australia	19.48	8.96	9.13	15.77	8.19	6.04	15.13	5.71	.69	10.89
7. France	21.20	4.69	8.22	14.41	10.23	10.50	11.99	6.05	.29	12.42
8. Germany	30.08		10.36	15.79	12.04	2.68	13.06	7.44		8.55
9. Belgium	23.11	6.95	7.53	15.72	14.14	6.74	10.67	3.98	.23	10.93
10. Norway	23.45	7.56	9.97	14.54	8.86	4.20	13.31	7.68	.54	9.89
11. Netherlands	26.20	6.38	14.78	10.81	12.48	6.67	6.67	3.41	3.10	9.50
12. Iceland	23.50	7.86	10.66	19.06	10.00	6.07	12.13	5.24	.44	5.05
13. Finland	28.24	8.62	8.26	14.36	7.51	2.54	13.60	5.92	1.44	9.49
14. Austria	22.06	7.29	11.37	11.99	8.89	3.49	13.36	5.67	.35	15.54
15. Japan	27.45		7.76	16.78	6.76	8.92	8.97	9.05		14.32
16. UK	18.92	6.26	8.40	18.20	7.78	.85	13.13	7.88	1.98	16.50
17. Spain	34.08	3.51	10.58	13.99	8.55	4.62	9.50	4.28	2.02	8.87
18. Italy	30.85	5.69	9.41	12.94	6.79	3.73	10.48	7.05	.43	12.62
Mean	23.45	6.92	9.16	16.35	9.04	5.05	12.37	6.56	1.11	10.94

Table 3. Prices of 10 commodities and divisia price index for 18 countries (Mean log-changes × 100)

Country (1)	Food (2)	Beverages (3)	Clothing (4)	Housing (5)	Durables (6)	Medical care (7)	Transport (8)	Recreation (9)	Education (10)	Miscellaneous (11)	Divisia price index (12)
1. US	5.00	4.85	3.29	4.67	4.25	5.81	4.87	3.56	5.91	5.37	4.78
2. Canada	6.15	4.71	3.87	5.18	4.49	4.96	4.72	3.89	8.43	6.73	5.27
3. Sweden	7.17	7.27	4.82	8.50	7.30	6.14	7.66	6.57	9.13	9.10	7.43
4. Switzerland	4.00	4.06	3.39	5.43	3.86	6.49	4.55	4.10		5.47	4.55
5. Denmark	8.69	6.68	6.57	11.43	8.55	8.91	8.79	7.62	9.54	9.30	8.90
6. Australia	6.05	6.58	6.12	7.43	4.33	8.42	6.31	6.97	8.38	7.53	6.58
7. France	7.30	6.47	6.74	8.80	6.76	6.99	8.13	6.13	9.02	8.28	7.48
8. Germany	3.23		3.67	5.57	3.04	5.58	4.13	3.74		4.61	3.96
9. Belgium	4.56	4.61	3.89	5.76	4.36	6.02	5.28	5.00	4.54	6.88	5.13
10. Norway	6.96	6.99	6.74	7.20	7.02	8.25	7.32	5.36	7.85	7.91	7.05
11. Netherlands	3.71	2.91	3.48	5.08	2.29	7.73	3.38	5.51	5.91	4.80	4.07
12. Iceland	13.74	11.25	9.69	11.66	10.13	14.10	11.25	12.66	12.78	11.48	11.79
13. Finland	7.75	6.87	6.09	5.90	6.25	6.22	7.16	6.32	9.44	8.17	7.02
14. Austria	4.57	2.93	3.67	7.49	3.59	8.94	6.12	4.21	4.20	6.53	5.28
15. Japan	7.61		7.99	7.58	6.71	7.36	8.70	7.35		8.88	7.85
16. UK	9.00	8.49	6.88	11.28	8.68	9.36	10.25	8.85	10.73	9.99	9.56
17. Spain	9.79	7.57	11.24	8.58	11.07	11.69	7.89	12.54	12.82	10.87	10.06
18. Italy	9.67	7.07	10.37	10.89	10.37	8.72	10.83	9.25	10.03	11.65	10.13
Mean	6.94	6.21	6.03	7.69	6.28	7.87	7.08	6.64	8.58	7.98	7.05

Table 4. Per capita quantities consumed of 10 commodities and divisia volume index for 18 countries (Mean log-changes × 100)

Country (1)	Food (2)	Beverages (3)	Clothing (4)	Housing (5)	Durables (6)	Medical care (7)	Transport (8)	Recreation (9)	Education (10)	Miscellaneous (11)	Divisia volume index (12)
1. US	.60	1.41	2.33	2.97	1.64	4.49	2.80	4.10	3.33	1.77	2.43
2. Canada	.91	2.57	3.12	3.17	3.02	.22	4.17	6.27	4.20	2.79	2.87
3. Sweden	.50	.80	2.66	2.35	1.18	3.02	2.06	4.53	3.17	−.95	1.71
4. Switzerland	1.72	1.57	.75	2.08	.62	2.85	3.63	2.98		2.79	2.13
5. Denmark	.17	1.43	.59	3.23	−1.01	1.50	1.49	3.34	8.53	.75	1.42
6. Australia	1.35	1.03	.71	3.54	4.45	1.23	3.49	4.33	−2.40	1.95	2.37
7. France	1.62	1.37	1.71	4.89	3.31	6.68	4.65	5.37	2.43	3.20	3.57
8. Germany	2.02		2.36	3.24	3.51	1.45	5.69	3.90		3.77	3.15
9. Belgium	1.53	2.58	2.76	2.80	4.20	5.13	4.52	4.04	1.98	2.98	3.11
10. Norway	1.35	2.08	1.40	3.43	3.21	2.14	4.06	5.56	1.26	2.17	2.68
11. Netherlands	2.22	4.39	3.19	3.93	6.91	5.53	7.52	4.07	5.15	5.20	4.34
12. Iceland	1.64	5.11	5.65	2.54	8.12	7.20	8.58	5.92	4.54	8.56	5.07
13. Finland	1.77	4.89	1.62	3.60	5.41	6.43	6.48	6.18	−1.60	5.58	3.86
14. Austria	1.90	2.34	4.26	4.29	3.61	2.62	5.58	4.35	1.94	1.98	3.27
15. Japan	1.50		1.47	4.93	1.58	6.07	4.10	3.26		3.40	3.13
16. UK	.15	1.48	2.52	1.87	1.58	.90	2.81	3.65	2.10	1.70	1.76
17. Spain	2.82	4.01	3.00	3.79	3.44	9.24	10.66	5.31	4.24	5.72	4.40
18. Italy	2.20	3.15	2.81	2.80	3.90	6.82	5.71	3.98	1.61	3.65	3.33
Mean	1.44	2.51	2.39	3.30	3.26	4.08	4.89	4.51	2.70	3.17	3.04

where $\bar{w}_{it} = 1/2(w_{it} + w_{i,t-1})$ is the arithmetic average of the budget share of commodity i in years $t-1$ and t. Column 12 of Table 3 gives the mean of (2.1) for each country.

Columns 2–11 of Table 4 present the means of $Dq_{it} = \log q_{it} - \log q_{i,t-1}$, the log-change in per capita consumption of i. Note from the last row that on average transport and recreation grow the fastest (4.9 and 4.5 percent per annum), while food has the lowest growth rate (1.4 percent). The index (2.1) measures the overall growth in prices. The analogous Divisia volume index is defined as

$$DQ_t = \sum_{i=1}^{n} \bar{w}_{it} Dq_{it} \ . \tag{2.2}$$

This measures the overall growth in per capita consumption and is given in column 12 of Table 4. The last entry in this column indicates that per capita consumption as a whole increases by 3.0 percent per annum on average in the OECD.

3 Divisia Variances

The Divisia price and volume indexes, defined by equations (2.1) and (2.2), are budget-share-weighted first-order moments of the price and quantity log-changes. The corresponding second-order moments are the Divisia variances,

$$\Pi_t = \sum_{i=1}^{n} \bar{w}_{it}(Dp_{it} - DP_t)^2 \ , \qquad K_t = \sum_{i=1}^{n} \bar{w}_{it}(Dq_{it} - DQ_t)^2 \ .$$

These variances measure the degree to which the prices and quantities of the individual goods change disproportionately. When all the prices and quantities change proportionately, these two variances vanish.

Tables 5 and 6 present these variances. Comparing these two tables, we see that the quantity variances systematically exceed the corresponding price variances. This pattern agrees with the results of Clements (1982, 1983), Meisner (1979b), E. A. Selvanathan (1987), Theil (1967) and Theil and Suhm (1981) and seems to have the status of an empirical regularity. Another empirical regularity is that higher inflation (i.e., *DP*) is associated with a larger dispersion of relative prices (Π); for a survey of the literature on inflation and relative price variability, see Marquez and Vining (1984).

The Divisia price-quantity covariance and correlation are

$$\Gamma_t = \sum_{i=1}^{n} \bar{w}_{it}(Dp_{it} - DP_t)(Dq_{it} - DQ_t) \ , \qquad \rho_t = \frac{\Gamma_t}{\sqrt{\Pi_t K_t}} \ .$$

These measure the co-movement of prices and quantities. As demand curves slope down, we would expect the consumer to move away from those goods

Table 5. Divisia price variances in 18 countries

Year (1)	US (2)	Canada (3)	Sweden (4)	Switzerland (5)	Denmark (6)	Australia (7)	France (8)	Germany (9)	Belgium (10)	Norway (11)	Netherlands (12)	Iceland (13)	Finland (14)	Austria (15)	Japan (16)	UK (17)	Spain (18)	Italy (19)
1953	—	—	—	—	—	—	—	—	—	—	.34	—	—	—	—	—	—	—
1954	—	—	—	—	—	—	—	—	—	—	1.18	—	—	—	—	—	—	—
1955	—	—	—	—	—	—	—	—	—	—	.21	—	—	—	—	—	—	—
1956	—	—	—	—	—	—	—	—	—	—	.93	—	—	—	—	—	—	—
1957	—	—	—	—	—	—	—	—	—	—	.49	—	—	—	—	—	—	—
1958	—	—	—	—	—	—	—	—	—	—	.84	—	—	—	—	—	—	—
1959	—	—	—	—	—	—	—	—	—	—	.09	—	—	—	—	—	—	—
1960	—	—	—	—	—	—	—	—	—	—	.86	—	—	—	—	—	—	—
1961	.03	.08	—	.25	—	.72	—	.45	1.82	—	.10	1.39	.28	—	—	—	—	—
1962	.07	.07	—	.12	—	.95	—	.17	.59	—	.08	3.56	.38	—	—	—	—	—
1963	.03	.15	—	.24	—	.89	—	.18	.96	—	.18	3.18	.47	—	—	—	—	—
1964	.82	.17	—	.26	—	.27	—	.21	.18	—	.67	2.29	1.39	—	—	—	—	—
1965	.04	.19	.17	.32	—	.47	.17	.18	1.53	.19	.37	1.00	.87	.35	—	.38	.97	.15
1966	.22	.35	.32	.15	—	.13	.18	.23	.65	.20	.64	2.03	.39	.24	—	.13	.20	.08
1967	.26	.34	.32	.50	1.15	.19	.32	.41	.22	.37	.95	.26	.56	.50	—	.11	.72	.18
1968	.11	.31	.15	.46	.64	.22	.38	.90	.15	.12	.45	.92	3.36	.58	—	.20	.91	.19
1969	.12	.22	.51	.29	.38	.63	.15	.29	.14	.16	.57	6.08	.05	.19	—	.08	.72	.05
1970	.07	.40	.19	.25	.62	.52	.11	.09	.49	.32	.79	2.45	.22	3.16	—	.17	.88	.13
1971	.16	.34	.21	.22	.68	.56	.10	.19	.36	.22	.97	2.16	.17	.51	.39	.29	.37	.23
1972	.15	.35	.42	.41	.50	.16	.14	.08	.45	.08	.67	3.79	.33	.34	.41	.47	.84	.34
1973	1.48	1.71	.39	.59	1.25	.96	.28	.04	.51	.09	.49	2.69	.25	.69	1.68	.99	.59	.52
1974	.62	1.07	2.19	.23	1.09	1.54	.74	.43	.76	.45	1.02	—	1.84	1.04	2.82	.46	.88	1.77
1975	.33	.47	.34	.25	.24	1.65	.20	.05	.88	.90	.90	—	1.87	.37	.75	1.01	.54	.91
1976	.44	.95	.27	.72	.79	.66	.25	.06	.54	.07	.55	—	.91	.52	.47	.36	1.11	1.13
1977	.30	.83	.42	.02	.34	.44	.28	.11	.51	.07	.64	—	1.45	.35	.40	.18	1.56	.67
1978	.40	.98	.62	.41	.30	.77	.06	.08	.27	.32	—	—	—	.45	.81	.40	—	.43
1979	.47	.27	.64	.85	1.83	.42	.24	.43	.76	.25	—	—	—	.40	.24	.54	—	.46
1980	.96	.18	.47	.25	.79	.35	1.01	.16	1.76	.55	—	—	—	.51	.42	1.17	—	1.00
1981	.40	.53	.42	.28	.43	.29	.20	.23	3.45	1.10	—	—	—	.69	.03	2.20	—	.16
Mean	.36	.47	.47	.34	.73	.61	.28	.24	.81	.32	.60	2.45	.87	.64	.77	.54	.79	.49

All entries are to be divided by 1000.

Table 6. Divisia quantity variances in 18 countries

Year (1)	US (2)	Canada (3)	Sweden (4)	Switzerland (5)	Denmark (6)	Australia (7)	France (8)	Germany (9)	Belgium (10)	Norway (11)	Netherlands (12)	Iceland (13)	Finland (14)	Austria (15)	Japan (16)	UK (17)	Spain (18)	Italy (19)
1953	—	—	—	—	—	—	—	—	—	—	2.10	—	—	—	—	—	—	—
1954	—	—	—	—	—	—	—	—	—	—	1.94	—	—	—	—	—	—	—
1955	—	—	—	—	—	—	—	—	—	—	2.21	—	—	—	—	—	—	—
1956	—	—	—	—	—	—	—	—	—	—	1.86	—	—	—	—	—	—	—
1957	—	—	—	—	—	—	—	—	—	—	1.16	—	—	—	—	—	—	—
1958	—	—	—	—	—	—	—	—	—	—	2.35	—	—	—	—	—	—	—
1959	—	—	—	—	—	—	—	—	—	—	1.42	—	—	—	—	—	—	—
1960	—	—	—	—	—	—	—	—	—	—	.60	—	—	—	—	—	—	—
1961	.63	8.64	—	.89	—	.55	—	.62	1.45	—	.60	2.69	2.07	—	—	—	—	—
1962	.72	.57	—	.65	—	2.35	—	.81	.25	—	.68	7.51	1.29	—	—	—	—	—
1963	.44	.63	—	.28	—	3.22	—	.51	.52	—	.48	11.38	1.74	—	—	—	—	—
1964	.34	.47	—	.66	—	.48	—	.56	1.23	—	3.35	5.52	3.86	—	—	—	—	—
1965	.30	.63	.83	.30	—	.19	.52	.81	.50	.54	2.18	1.41	4.49	.84	—	.66	1.72	.50
1966	.46	.87	.71	.30	—	.57	.41	.18	.31	.28	1.98	4.23	.76	.90	—	.16	4.23	.52
1967	.15	.50	.49	.21	.42	.65	.33	.34	.42	.52	1.71	2.69	.93	.30	—	.32	.70	1.06
1968	.59	.61	1.31	.15	.54	.85	.29	.64	.78	.52	1.33	8.39	.75	.31	—	.30	1.75	.10
1969	.55	.35	.76	.21	1.19	1.10	.50	1.91	.52	2.32	2.41	3.33	3.80	.56	—	.24	.97	.73
1970	.69	1.82	.53	.11	.94	.28	.57	1.10	.24	2.61	1.88	19.19	.95	1.60	—	.43	.40	.59
1971	.92	2.51	1.38	.22	1.09	.37	.63	.28	.74	1.02	1.06	8.40	1.06	2.36	.66	.98	.88	.40
1972	.65	1.51	.83	.17	.86	.38	.55	.26	.99	.94	2.86	4.32	1.79	2.39	.49	1.94	2.67	.60
1973	1.31	2.35	1.06	.73	1.17	2.52	.75	.95	2.10	.43	1.61	1.80	2.43	.60	1.29	1.02	1.11	1.03
1974	1.55	1.96	1.45	.93	3.82	1.49	1.19	.62	1.81	.51	1.74	—	2.52	1.16	3.31	1.11	.89	.77
1975	.68	.23	.32	1.78	2.08	1.54	.75	1.21	2.04	.73	4.38	—	2.95	.45	1.59	.21	.48	.97
1976	.54	.35	.40	.47	1.64	.58	.99	.69	.75	1.39	.52	—	3.15	.61	1.06	.16	.28	.32
1977	.31	.67	.73	.35	.55	.56	.44	.88	1.07	1.30	1.49	—	2.04	2.32	.63	.44	.39	.56
1978	.54	.61	.58	.46	.88	.99	.69	.18	.57	2.85	—	—	—	2.21	.97	.71	—	.15
1979	.36	.27	.13	.21	.33	.25	.32	.16	.88	.72	—	—	—	.82	.87	.32	—	.66
1980	1.21	.38	.46	.13	1.96	.44	.55	.26	1.75	.68	—	—	—	.50	1.41	.41	—	.31
1981	.26	.23	.61	.34	.44	.16	.41	.43	.75	1.21	—	—	—	.45	1.07	.52	—	.69
Mean	.63	1.25	.74	.45	1.19	.93	.58	.64	.94	1.09	1.78	6.22	2.15	1.08	1.21	.58	1.27	.59

All entries are to be divided by 1000.

Table 7. Divisia price-quantity variances in 18 countries

Year (1)	US (2)	Canada (3)	Sweden (4)	Switzerland (5)	Denmark (6)	Australia (7)	France (8)	Germany (9)	Belgium (10)	Norway (11)	Netherlands (12)	Iceland (13)	Finland (14)	Austria (15)	Japan (16)	UK (17)	Spain (18)	Italy (19)
1953	—	—	—	—	—	—	—	—	—	—	−.45	—	—	—	—	—	—	—
1954	—	—	—	—	—	—	—	—	—	—	−.88	—	—	—	—	—	—	—
1955	—	—	—	—	—	—	—	—	—	—	.21	—	—	—	—	—	—	—
1956	—	—	—	—	—	—	—	—	—	—	−.95	—	—	—	—	—	—	—
1957	—	—	—	—	—	—	—	—	—	—	.33	—	—	—	—	—	—	—
1958	—	—	—	—	—	—	—	—	—	—	.13	—	—	—	—	—	—	—
1959	—	—	—	—	—	—	—	—	—	—	−.64	—	—	—	—	—	—	—
1960	—	—	—	—	—	—	—	—	—	—	−.33	—	—	—	—	—	—	—
1961	.45	.01	—	−.32	—	−.11	—	−.40	−.97	—	−.67	−.27	−.57	—	—	—	—	—
1962	.33	−.69	—	−.61	—	.06	—	−.69	−.38	—	−.01	−.65	−.06	—	—	—	—	—
1963	.10	−.69	—	−.55	—	−.22	—	.74	.59	—	−.82	−.70	.30	—	—	—	—	—
1964	−.08	−.45	—	−.39	—	−.87	—	−.71	.31	—	−.57	−.49	−.72	—	—	—	—	—
1965	.10	−.31	−.38	−.59	—	−.01	.42	−.73	−.04	.05	−.38	−.35	−.57	−.78	—	−.21	−.58	.08
1966	−.30	−.57	.15	.13	—	.61	.15	.10	−.31	−.16	−.43	−.62	.22	.54	—	.15	−.58	−.15
1967	−.06	.22	−.23	.19	−.12	−.29	.63	.12	−.21	−.36	−.14	.17	.50	.11	—	.27	−.08	−.13
1968	−.18	−.37	−.19	.61	.69	.05	.20	.11	.29	−.87	−.25	−.46	−.53	.15	—	.41	−.78	.01
1969	−.22	.07	−.47	−.07	−.52	−.74	.44	−.62	−.58	−.67	−.02	−.80	−.25	−.20	—	−.01	.34	−.66
1970	.45	−.06	−.36	−.47	−.10	−.26	.43	−.10	−.25	−.26	−.51	−.16	−.07	−.36	—	−.02	−.30	−.36
1971	.67	.17	−.66	−.06	.19	−.48	−.09	−.07	.27	−.23	−.44	.30	.05	−.42	−.04	.23	.24	.34
1972	−.82	−.65	−.70	−.08	−.71	−.50	−.87	−.89	−.48	−.28	.16	−.61	−.25	−.39	−.27	−.60	−.61	−.60
1973	−.97	−.78	−.64	.37	−.75	−.65	−.77	−.57	−.56	−.45	−.66	−.55	−.43	−.34	−.34	−.90	−.69	−.58
1974	−.59	−.77	−.20	−.26	−.35	−.53	−.62	−.20	−.46	−.08	−.37	—	−.46	−.10	−.75	−.76	−.54	−.63
1975	−.59	−.62	−.56	−.09	−.13	−.67	.04	−.06	.02	−.87	−.49	—	−.32	.38	.34	−.05	−.19	−.52
1976	.08	−.55	−.43	.29	−.59	−.31	−.11	.14	−.38	−.32	−.21	—	−.65	.27	−.74	−.28	.57	.23
1977	.04	−.76	−.61	−.41	−.46	.34	−.58	−.77	.65	.45	−.80	—	−.37	−.55	.29	−.29	−.01	−.31
1978	−.88	−.88	−.16	−.72	−.53	−.58	−.05	−.22	−.11	−.23	—	—	—	.03	−.19	−.36	—	−.30
1979	−.70	−.69	−.44	−.59	.28	.00	−.22	−.10	−.31	−.04	—	—	—	.25	.07	.06	—	−.04
1980	−.70	−.46	−.56	−.42	−.07	−.77	−.48	−.69	−.82	−.49	—	—	—	.00	−.67	−.26	—	.42
1981	−.19	−.67	−.14	−.77	−.28	.04	−.60	−.35	−.70	−.85	—	—	—	−.83	.22	−.26	—	−.03
Mean	−.19	−.45	−.39	−.23	−.23	−.28	−.12	−.28	−.21	−.33	−.37	−.40	−.25	−.13	−.19	−.17	−.25	−.19

which have above-average price increases and vice versa. Thus we expect ρ to be negative. Table 7 presents the Divisia correlations for the OECD. Of the 322 correlations, about 75 percent are negative. The last row of the table shows that, on average, the correlations are negative for each country, as expected.

4 Engel's Law

One of the major empirical regularities in consumption economics is *Engel's law*, i.e., the budget share of food falls with increasing income, or, in other words, the income elasticity of demand for food is less than unity. The classic reference in this area is Houthakker (1957) who estimated Engel curves for a large number of household surveys; in each case food was found to be a necessity.

More recently, Theil et al. (1989) have shown that it is now possible to make a more precise statement regarding the dependence on income of the food budget share. Working (1943) observed that the budget share for food seems to be a linear function of the logarithm of income. This observation was generalized by Leser (1963) to a complete set of n commodities to yield

$$w_i = \alpha_i + \beta_i \log M \ , \qquad i = 1, \ldots, n \ . \tag{4.1}$$

Choosing the income unit such that $M = 1$ for some household, α_i is then interpreted as the budget share of i for that household. The coefficient β_i is interpreted as 100 times the change in the budget share of i resulting from a 1 percent increase in income. We shall refer to (4.1) as Working's model.

Theil (1987a), using data from Kravis et al. (1982) for 34 countries, plots the budget share of food against the log of real per capita total consumption. Figure 1 gives such a plot (with consumption scaled such that $M = 1$ for the poorest country). The solid line is the LS regression line. As can be seen, the points are all scattered around a downward-sloping line which gives strong visual support for Working's model for food. The slope of the LS line is $-.15$ (standard error .01), which is an estimate of β_i for food. This estimate does not take account of cross-country differences in the relative price of food; however, approximately the same value of β_i emerges when these differences are allowed for (Theil, 1987a). The estimate of β_i of $-.15$ is in remarkable agreement with other estimates of the food β_{ij} see Table 8. Thus $\beta_i = -.15$ seems to be more or less a natural constant.

To analyze the implications of this β_i-estimate, consider moving from one country to another. Assume that the first country's per capita income is M, while that of the second is $2M$. Therefore,

$$\Delta(\log M) = \log 2M - \log M = \log 2 = .69.$$

From (4.1) we have

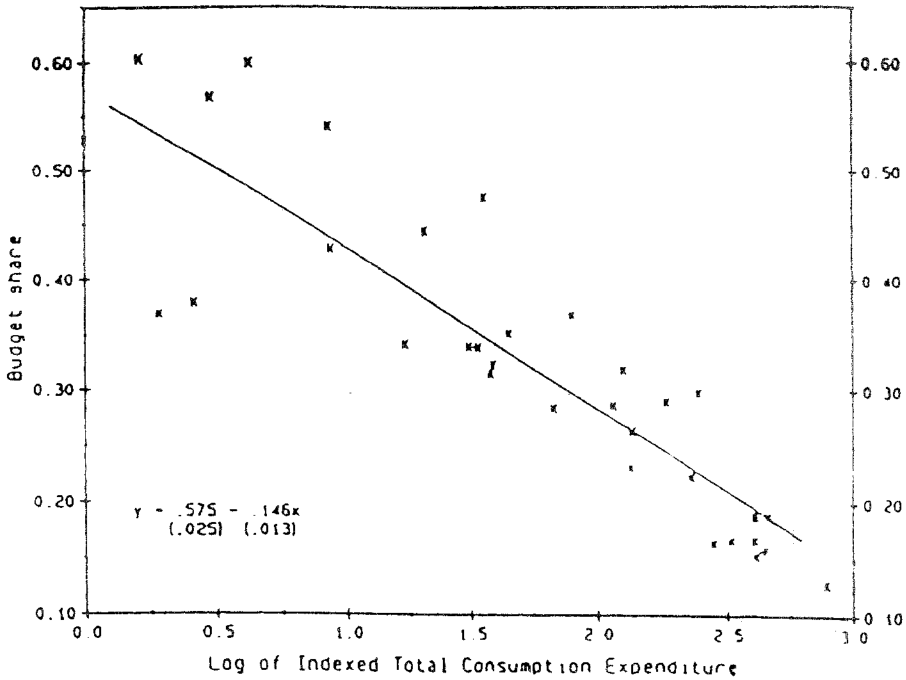

Fig. 1. Budget share of food against scaled total consumption expenditure per capita in 1975 for 34 countries

Table 8. Previous estimates of working's income coefficient for food

Author(s)	Country	Estimate
Aasness and Rodseth (1983)	Norway	−.17
Blanciforti and Green (1983)	US	−.13
Chung and Lopez (1988)	Spain	−.18 & −.16
Deaton and Muellbauer (1980b)	Britain	−.16
Finke et al. (1984)	Japan	−.15
Musgrove (1985)	Dominican Republic	−.14
Theil (1987a)	Cross country	−.14 to −.16
Theil and Finke (1984)	Netherlands	−.13
Theil et al. (1987)	China	−.13

Source: Chung and Lopez (1988, Table 1).

$$\Delta w_i = \beta_i \, \Delta \log M \ .$$

Consequently, when $\beta_i = -.15$ the effect of doubling income on the food budget share is

$$\Delta w_i = -.15 \times .69 = -.10 \ .$$

That is, when income doubles, the budget share declines by 10 percentage points. Theil et al. (1989) refer to this as *the strong version of Engel's law.*

One use of this law is to make real income comparisons across consumers or countries on the basis of the food share. As this share is obviously a pure number, it is directly comparable over time, countries etc., independent of inflation, exchange rates and so on. Consequently, such indirect income comparisons could be an attractive short-cut.

5 Homogeneity and Symmetry

The Marshallian demand equation for good i is

$$q_i = q_i(M, p_1, \ldots, p_n) \ . \tag{5.1}$$

If income and prices all increase proportionately, then the quantity demanded of each good should remain unchanged. That is, for $\alpha > 0$, $q_i(\alpha M, \alpha p_1, \ldots, \alpha p_n) = q_i(M, p_1, \ldots, p_n)$, or equation (5.1) is homogeneous of zero degree. This is known as the absence of money illusion or *demand homogeneity* and can be expressed as

$$\eta_i + \sum_{j=1}^{n} \eta_{ij} = 0 \ ,$$

where $\eta_i = \partial(\log q_i)/\partial(\log M)$ is the income elasticity of good i; and $\eta_{ij} = \partial(\log q_i)/\partial(\log p_j)$ is the $(i, j)^{\text{th}}$ uncompensated price elasticity.

The system of n demand equations given by (5.1) for $i = 1, \ldots, n$ also satisfies *Slutsky symmetry*, viz.

$$\frac{\partial q_i}{\partial p_j} = \frac{\partial q_j}{\partial p_i} \qquad i, j = 1, \ldots, n \ , \tag{5.2}$$

where the derivatives hold real income constant. In words, the effect on beer drinking of a \$1-increase in the price of a bottle of wine is exactly equal to the response of wine consumption to the same rise in the price of a bottle of beer. Slutsky symmetry can be formulated in terms of elasticities by defining $\eta'_{ij} = \eta_{ij} + w_j \eta_i$ as the $(i, j)^{\text{th}}$ compensated price elasticity, where w_j is the budget share of j. Restriction (5.2) then becomes

$$w_i \eta'_{ij} = w_j \eta'_{ji} \qquad i, j = 1, \ldots, n \ .$$

Homogeneity and symmetry would seem to constitute highly plausible hypotheses, hypotheses that would not be expected to be at great variance with the data. Surprisingly, however, when these restrictions are tested the bulk of the evidence points towards rejection; see Barten (1977) and Keuzenkamp and Barten (1991) for surveys of these tests.

These rejections represented a major puzzle. The response of Theil, and some others, was not to take the "evidence" too literally. He argued in *Theory and Measurement of Consumer Demand* that these findings reflect

data imperfections rather than unorthodox consumer behavior. What is needed in such a situation is an explicit recognition on the part of the analyst that his data are imperfect. He should guide these data so that they yield a sensible picture. It is unavoidable that this picture of the real world will be biased toward simplicity, but it is appropriate to realize that this kind of bias is inherent in every model. (Theil, 1975/6, pp. 250–1 of Volume 1).

As a consequence, Theil felt comfortable simply imposing the constraints, especially homogeneity. Others have responded to the rejection puzzle by emphasizing dynamics (Anderson and Blundell, 1983), simultaneity (Attfield, 1985) and the use of non-stationary variables (Bewley and Elliott, 1989).

Theil's skepticism about the rejections was not shared by everyone. Some took the findings seriously and concluded that consumption theory should be rejected (see, e.g., Christensen et al., 1975). Apparently, this troubled Theil sufficiently to encourage two of his students, Laitinen and Meisner, to study the rejections by means of Monte Carlo experiments. The idea is to simulate data sets under the null, use these data to estimate the demand equations and then test homogeneity and symmetry. As the hypotheses are true by construction, the rejection rates should coincide with the nominal significance levels if the tests are working satisfactorily.

Table 9 shows Laitinen's (1978) and Meisner's (1979a) results. As can be seen from columns 2 and 3, the rejections of homogeneity are far too high, with the rejection rate increasing with n, the number of commodities. The results for symmetry (columns 4 and 5) also indicate a substantial bias against the null, but the problems are now not quite so bad for $n = 5$ or 8. The source of the problem is that these tests involve the replacement of the inverse of the disturbance covariance matrix $\sum^{-1}$ with its estimator S^{-1}, the inverse of the residual moment matrix. This replacement has an asymptotic justification (when the number of observations $T \to \infty$), but this becomes problematic when the number of commodities in the model is large relative to T; in this case, S is near singular

Table 9. Percentage rejections of homogeneity and symmetry

	Rejections of homogeneity		Rejections of symmetry	
Number of commodities, n (1)	5 percent (2)	1 percent (3)	5 percent (4)	1 percent (5)
5	14	6	9	3
8	30	16	26	8
11	53	35	50	37
14	87	81	96	91

and S^{-1} explosive. See Theil (1987b) for further details and Bera et al. (1981) and Bewley (1983, 1986) for related contributions.

If the standard tests are defective, how does one proceed? There are two ways of avoiding the problems associated with asymptotics. First, there is now available an exact finite-sample test for homogeneity (Laitinen, 1978). Second, Monte Carlo testing (Theil, 1987b), which is distribution-free, can be employed. This involves simulating a large number of values of the test statistic under the null hypothesis to construct its empirical distribution. The data-based value of the test statistic is then compared with this distribution, rather than its asymptotic counterpart.

S. Selvanathan (1987a) tests homogeneity and symmetry with the Monte Carlo technique for 18 OECD countries. She employs a variant of the Rotterdam model (with intercepts to take account of residual trends) and Table 10 summarizes the results. There are 100 values of each test statistic, the observed (or data-based) value plus 99 simulated values. Consequently, the hypothesis is rejected at the 5 percent level if the observed value is among the largest 5.

Columns 2–11 of Table 10 give the ranks of the homogeneity test statistic for each commodity. Looking at the last two rows, it can be seen that durables tends to be a troublemaker. Column 14 gives the rank of the test statistic for the homogeneity of the n equations jointly. As can be seen, the hypothesis is acceptable for all 18 countries except Belgium at the 5 percent level; and at the 1 percent level homogeneity cannot be rejected for any country. Column 15 of Table 10 shows that symmetry is rejected at the 5 percent level for Switzerland only. The hypothesis is acceptable in all cases at the 1 percent level.

On the basis of these new tests, the conclusion is that, on the whole, homogeneity and symmetry are not grossly incompatible with the data. The previous rejections reported in the literature reflect faulty econometric procedures (the breakdown of asymptotic theory, in particular), rather than a flaw in the theory of the utility-maximizing consumer.

6 Preference Independence

When the consumer's tastes can be described by means of a utility function which can be written as the sum of n sub-utility functions, each involving one good only, then tastes are said to exhibit *preference independence*. Formally, the utility function is of the preference independent form if

$$u(q_1, \ldots, q_n) = \sum_{i=1}^{n} u_i(q_i) \ , \qquad (6.1)$$

so that the marginal utility of good i is independent of the consumption of j, $i \neq j$. The specification (6.1) is also known as additive preferences.

Table 10. Ranks of test statistics for homogeneity and symmetry

Country	Homogeneity													
	Food	Beverages	Clothing	Housing	Durables	Medical care	Transport	Recreation	Education	Miscellaneous	Percent significant at 5% level	Percent significant at 1% level	All goods	Symmetry
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
1. US	76	89	53	40	1	77	98	3	70	25	30	10	80	71
2. Canada	65	18	68	68	37	100	8	49	47	39	10	10	29	32
3. Sweden	91	74	95	24	35	49	2	58	48	37	10	0	75	14
4. Switzerland	16	45	17	92	29	40	16	96	—	94	0	0	82	96
5. Denmark	77	55	14	29	51	23	84	78	13	31	0	0	25	41
6. Australia	37	35	81	6	100	63	34	9	7	72	10	10	87	2
7. France	83	60	31	67	95	76	2	58	28	34	10	0	61	80
8. Germany	23	—	26	83	52	77	29	72	—	55	0	0	14	90
9. Belgium	10	37	70	3	100	93	25	79	55	29	20	10	99	30
10. Norway	83	1	77	76	61	12	70	63	51	27	10	10	64	11
11. Netherlands	39	38	92	91	2	55	55	26	59	45	10	0	37	56
12. Iceland	68	17	48	69	83	57	12	98	44	5	10	0	52	38
13. Finland	61	48	41	65	66	34	62	70	10	23	0	0	3	42
14. Austria	40	2	81	72	40	58	78	33	47	58	10	0	18	21
15. Japan	85	—	9	50	2	84	81	16	—	78	13	0	89	56
16. UK	19	81	74	25	85	41	20	64	50	83	0	0	43	30
17. Spain	63	56	73	35	51	58	37	76	32	25	0	0	2	87
18. Italy	20	52	41	32	73	95	87	55	32	55	0	0	29	49
Percent significant at 5% level	0	13	0	6	28	6	17	11	0	0			6	6
Percent significant at 1% level	0	6	0	0	17	6	0	0	0	0			0	0

If the commodities are fairly broad groups, such as food, housing, clothing and so on, then (6.1) could be a reasonable working hypothesis as it conveys the idea that total utility is obtained from the utility derived from food *and* utility from housing *and* from clothing *and* so on. These broad commodity groups can be interpreted as representing the "basic wants" of the consumer and if these are truely basic wants, they could be expected to exhibit little interaction in the utility function.

To set out the operational implications of the assumption of preference independence, we use a double-log approximation to the general demand equation (5.1):

$$\log q_i = \alpha_i + \eta_i \log M + \sum_{j=1}^{n} \eta_{ij} \log p_j \ , \tag{6.2}$$

where η_i is the income elasticity of i and η_{ij} is the uncompensated elasticity of demand for good i with respect to the price of j. Let $\log P = \sum_{i=1}^{n} w_i \log p_i$ be the Divisia price index in levels, $\log Q = \log M - \log P$ be (the logarithm of) real income and η'_{ij} the $(i, j)^{\text{th}}$ compensated price elasticity. Then, using the Slutsky equation $\eta_{ij} = \eta'_{ij} - w_j\eta_i$, we can formulate (6.2) in terms of real income and compensated price elasticities,

$$\log q_i = \alpha_i + \eta_i \log Q + \sum_{j=1}^{n} \eta'_{ij} \log p_j \ . \tag{6.3}$$

We write ϕ for the reciprocal of the income elasticity of the marginal utility of income ("the income flexibility" for short) and $\theta_i = w_i\eta_i = \partial(p_iq_i)/\partial M$ for the i^{th} marginal share, with $\sum_{i=1}^{n} \theta_i = 1$. This θ_i answers the question "If income rises by one dollar, how much of this is spent on commodity i?" The utility function (6.1) then implies that the price elasticities take the form (see, e.g., Clements, 1987)

$$\eta'_{ij} = \phi\eta_i(\delta_{ij} - \theta_j) \qquad i, j = 1, \ldots, n \ , \tag{6.4}$$

where δ_{ij} is the Kronecker delta ($\delta_{ij} = 1$ if $i = j$, $=0$ otherwise).

We define $\log P' = \sum_{i=1}^{n} \theta_i \log p_i$ as the Frisch price index, which has marginal shares as weights. Using (6.4), the substitution term in (6.3) then becomes

$$\sum_{j=1}^{n} \eta'_{ij} \log p_j = \phi\eta_i \sum_{j=1}^{n} (\delta_{ij} - \theta_j) \log p_j = \phi\eta_i(\log p_i - \log P') \ ,$$

so that the demand equation for good i takes the form

$$\log q_i = \alpha_i + \eta_i \log Q + \phi\eta_i \log\left[\frac{p_i}{P'}\right] . \tag{6.5}$$

As can be seen, only the own relative price, $\log(p_i/P')$, appears in the demand equation. The "unrestricted" system of demand equations, (6.3) for $i = 1, \ldots, n$, contains n^2 price elasticities η'_{ij}; by contrast, (6.5) for $i = 1, \ldots, n$ contains only one additional parameter ϕ.

The own-price elasticity of demand for good i in (6.5) is

$$\frac{\partial(\log q_i)}{\partial\left[\log \frac{p_i}{P'}\right]} = \phi\eta_i \ .$$

Strictly speaking, because the Frisch deflator is used in the relative price term of (6.5), this is a Frisch elasticity which holds constant the marginal utility of income. It can be converted to the more conventional real-income-constant elasticity by simply taking out the term $\log p_i$ from the Frisch index $\log P' = \sum_{i=1}^{n} \theta_i \log p_i$; this elasticity then takes the form $\partial(\log q_i)/\partial(\log p_i) = \phi\eta_i(1 - \theta_i)$, which is approximately equal to its Frisch counterpart $\phi\eta_i$ if the marginal share θ_i is small (as it is likely to be since $\sum_{i=1}^{n} \theta_i = 1$).

It is clear from (6.1) that preference independence is the *simplest* possible specification of tastes. The simplicity and transparency of the assumption is also clear from the demand equation (6.5) which focuses exclusively on the role of income and the own-relative price. As a general principle, simplicity is usually to be preferred to complexity in the sense that if the world can be understood with a simple (not to be confused with simplistic!) model, then there is no point to pursuing more complicated alternatives. This is merely a matter of scientific efficiency.

7 The Validity Preference Independence

In the previous section we argued that the assumption of preference independence leads to an attractive simplification of demand equations. There are two distinct justifications for this assumption (Clements, 1987): (i) The economic justification in terms of preference independence being plausible when the commodities in question are broad aggregates; and (ii) the statistical justification that the assumption reduces the number of unknown coefficients to be estimated from the order of n^2, which for even moderate-sized systems can be too large for precise estimates, to something much less. However, whether it is truly legitimate to invoke the assumption of preference independence is largely an empirical question. In this section we discuss some empirical evidence on this topic.

We return to (6.5), the demand equation for good i under preference independence, and write it as

$$\log q_i = \alpha_i + \eta_i \log Q + \gamma_i \log\left[\frac{p_i}{P'}\right], \tag{7.1}$$

where

$$\gamma_i = \phi\eta_i \tag{7.2}$$

is the own-price elasticity of demand. As (7.2) holds for $i = 1, \ldots, n$, it is clear that preference independence implies that price elasticities are proportional to income elasticities with ϕ the (negative) factor of proportionality. In other words, luxuries are more price elastic than necessities (Deaton, 1974, Pigou, 1910).

The proportionality between income and price elasticities may, at first, seem counter intuitive. The income elasticity reflects the luxuriousness of the good, while the size of the price elasticity is an indication of the availability of substitutes. As these are different dimensions of the good, one may be tempted to argue that there should be no *a priori* relationship between the elasticities. On the other hand, however, it is common practice to identify as necessities those goods which the consumer cannot easily do without (e.g., food); while luxuries involve discretionary expenditure for which there are many competing uses and thus can be easily foregone or postponed (e.g., durables). This use of language points in the direction of a relationship between the two types of elasticities (Deaton, 1987).

To analyse the evidence on the proportionality relationship (7.2), S. Selvanathan (1988a) estimates (7.1) in first-difference form for 18 OECD countries (with the Frisch price index replaced with its Divisia counterpart). Table 11 summarises the results in the form of cross-country frequency distributions of the elasticities. As can be seen, all the income elasticities for food are less than one which supports Engel's law. Also, housing is always a necessity and durables a luxury. Finally, most of the own-price elasticities lie between 0 and -1.

The income and price elasticities reported in Table 11 are unrestricted in that they are not constrained by (7.2). Table 12 gives the joint frequency distribution of the income and price elasticities. As can be seen, 55 percent of commodities

Table 11. Frequency distributions of income and price elasticities of 10 commodities in 18 countries (Percentages)

Range (1)	Food (2)	Beverages (3)	Clothing (4)	Housing (5)	Durables (6)	Medical care (7)	Transport (8)	Recreation (9)	Eduction (10)	Miscellaneous (11)	All Goods (12)
Income Elasticity (η_i)											
$(-\infty, -1]$	0	0	0	0	0	0	0	0	0	0	0
$(-1, 0]$	6	0	0	11	0	6	0	0	13	0	3
$(0, 1]$	94	75	6	89	0	78	11	28	73	67	52
$(1, \infty)$	0	25	94	0	100	17	89	72	13	33	45
Price Elasticity (γ_i)											
$(-\infty, -1]$	0	6	28	0	22	0	33	28	13	6	14
$(-1, 0]$	100	94	67	72	72	66	62	72	67	83	75
$(0, 1]$	0	0	5	28	6	28	5	0	20	11	10
$(1, \infty)$	0	0	0	6	0	6	0	0	0	0	1

Table 12. Joint frequency distribution of income and price elasticities of 10 commodities in 18 countries (Percentages)

Income elasticity η_i	Absolute value of price elasticity $\lvert\gamma_i\rvert$ $\leq\frac{1}{2}$	$>\frac{1}{2}$	Total
≤ 1	34	21	55
>1	15	30	45
Total	49	51	100

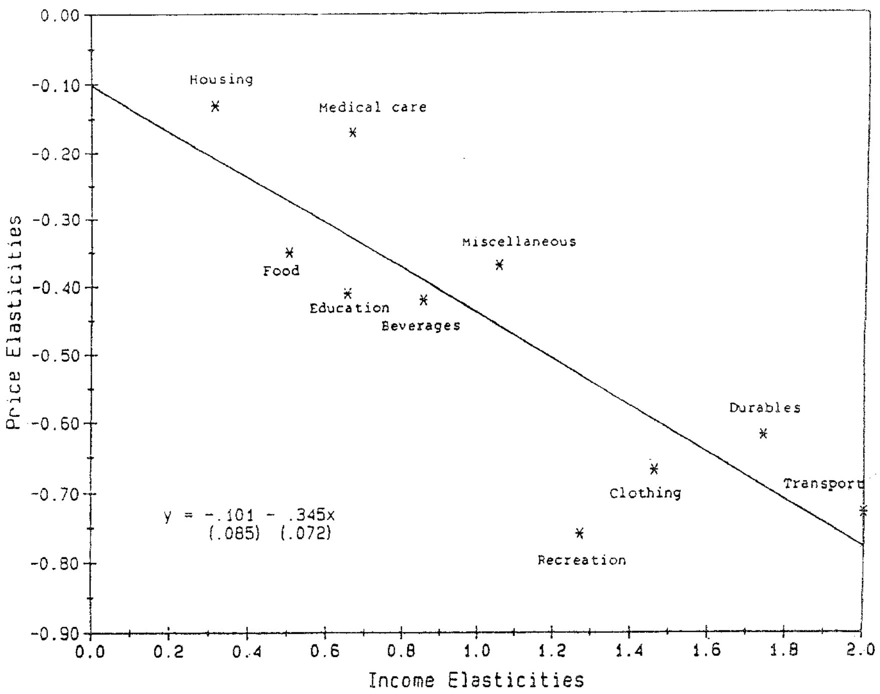

Fig. 2. Price elasticities against income elasticities for 10 commodities

are necessities ($\eta_i \leq 1$) and $34/55 = 62$ percent of these have own-price elasticities less than one half (in absolute value). Regarding the luxuries, $30/45 = 67$ percent possess price elasticities larger than one half. Consequently, there is a distinct tendency for those commodities with lower income elasticities to also have (absolutely) smaller price elasticities and vice versa. This is *prima facie* evidence in favour of (7.2).

Figure 2, from S. Selvanathan (1988a), provides another perspective on the relationship between the income and price elasticities. This plots each pair of estimates $\{\gamma_i, \eta_i\}$ averaged over the 18 countries and the solid line is the LS regression line. Here again luxuries tend to be more price elastic. See S. Selvanathan (1988a) for evidence on the proportionality relationship within each of the

18 countries. In a widely-quoted paper, Deaton (1974) carries out a similar analysis of the the elasticities derived from UK data for $n = 37$ and 8 commodities. He finds no relationship between income and price elasticities and concludes "that *the assumption of additive preferences is almost certain to be invalid in practice and the use of demand models based on such an assumption will lead to severe distortion of measurement*." (Deaton, 1974, p. 346, his emphasis)

Moreover, Deaton's rejection of preference independence on the basis of indirect evidence (the proportionality of the elasticities) is consistent with a number of more direct tests; see Barten (1977) for a survey. However, these tests have only an asymptotic justification and in light of the poor performance of the asymptotic tests of homogeneity and symmetry (discussed in Section 5 above), it is appropriate to exercise caution in taking them at face value.

S. Selvanathan (1987a, 1987b) pursues this matter by developing a Monte Carlo test of preference independence. This test works in a similar manner to the Monte Carlo tests of homogeneity and symmetry discussed in Section 5. She again employs a variant of the Rotterdam model with and without intercepts. Table 13 summarizes Selvanathan's results for 18 OECD countries. For some

Table 13. Ranks of test statistic for preference independence

	Based on S		Based on $\sum^*$	
Country (1)	No intercepts (2)	With intercepts (3)	No intercepts (4)	With intercepts (5)
1. US	100	100	58	70
2. Canada	98	96	23	15
3. Sweden	—	—	22	89
4. Switzerland	91	86	95	82
5. Denmark	—	—	47	58
6. Australia	86	70	93	92
7. France	—	—	91	86
8. Germany	95	50	84	30
9. Belgium	99	100	43	49
10. Norway	—	—	97	93
11. Netherlands	98	99	39	25
12. Iceland	—	—	7	24
13. Finland	—	—	91	59
14. Austria	—	—	81	79
15. Japan	—	—	50	79
16. UK	—	—	43	5
17. Spain	—	—	56	36
18. Italy	—	—	84	77
Percent significant at 5% level	57	57	6	0
Percent significant at 1% level	14	29	0	0

S is the residual moment matrix; and $\sum^* = \lambda^2(W - ww')$, where λ is an unknown parameter, $W = \text{diag}[w]$, $w = [\overline{\overline{w}}_i]$ and $\overline{\overline{w}}_i$ is the sample mean of the arithmetic average of the i^{th} budget share.

countries, the sample is undersized so that S, the residual moment matrix, cannot be used as the estimator of the disturbance covariance matrix. Consequently, a more parsimonious estimator is also used, $\sum^*$, which is defined in the notes to Table 13. Columns 2 and 3 of the table reveal that preference independence is rejected at the 5 percent level in 4 out of 7 countries when S can be used. When $\sum^*$ is used, for all 18 countries the hypothesis cannot be rejected at the 5 percent level when intercepts are included in the model.

Some qualifications to these results are required, such as the change in the ranking of the test statistic with the covariance estimator and the limited size of the underlying sample in some cases. The results do nonetheless provide at least some tentative indications that the status of the preference independence hypothesis is not as low as was once thought. This conclusion is also supported by Selvanathan's results on the proportionality of income and price elasticities, discussed above, which contradict Deaton's (1974) earlier findings for the UK.

8 Are Tastes Constant?

Utility maximization theory usually postulates that tastes are fixed, so that it is only the observable variables income and prices that explain consumption. There is a compelling reason for treating tastes (an unobservable) as constant: As the necessarily *ad hoc* auxiliary assumptions about the evolution of tastes are avoided, utility theory thereby has more applicability, generality and explanatory power. In a highly-influential paper, Stigler and Becker (1977) advocate treating tastes as fixed along these lines. They show that a number of examples of apparently capricious behavior (addiction, custom, tradition, fashion etc.) can in fact be reconciled with the assumption of stable preferences. Stigler and Becker (p. 89) argue that

> no significant behavior has been illuminated by the assumption of differences in tastes. Instead, they, along with the assumption of unstable tastes, have been a convenient crutch to lean on when the analysis has bogged down. They give the appearance of considered judgement, yet really have only been *ad hoc* arguments that disguise analytical failures.

The hypothesis of stable preferences can be tested by estimating demand equations for different groups of consumers and then analysing the extent to which the parameters (such as the income and price elasticities) differ across consumers. S. Selvanathan (1987c) carries out such a test with the OECD database by estimating (i) country-specific demand equations; and (ii) common demand equations for all countries whereby the data are pooled across countries. The test then involves an analysis of the deterioration in the fit of the demand equations when the data are pooled. In other words, this amounts to

Table 14. Quality of budget share predictions in OECD countries and Australian states

		RMS percentage prediction error	
Country/state (1)	Weight × 100 (2)	Individual country/ state model (3)	Pooled country/ state model (4)
OECD countries			
1. US	51.34	1.65	1.72
2. Canada	5.14	3.16	3.32
3. Sweden	1.84	1.88	1.91
4. Denmark	1.01	2.16	2.16
5. Australia	2.74	2.41	2.46
6. France	10.30	1.39	1.39
7. Belgium	1.81	2.59	2.74
8. Norway	.72	2.07	2.35
9. Netherlands	2.42	3.20	3.58
10. Iceland	.04	4.70	4.99
11. Finland	.82	3.68	3.67
12. Austria	1.25	2.34	2.50
13. UK	8.57	1.77	1.79
14. Spain	4.78	2.33	2.29
15. Italy	7.20	1.88	2.13
16. Unweighted mean		2.48	2.60
17. Weighted mean		1.87	1.95
Australian states			
18. NSW	38.53	2.16	2.14
19. Victoria	27.31	2.49	2.59
20. Queensland	14.79	3.07	2.97
21. SA	8.29	2.47	2.51
22. WA	8.47	2.54	2.46
23. Tasmania	2.61	3.42	3.56
24. Unweighted mean		2.69	2.70
25. Weighted mean		2.47	2.48

The column 2 weights for the OECD countries are proportional to GDPs in 1975 in international dollars; and the weights for the Australian states are proportional to total consumption expenditures in 1981. The RMS percentage prediction errors in columns 3 and 4 are 100 times the square root of the budget-share-weighted mean of the squared relative prediction errors of the budget shares; as an approximation, these are computed as 100 times the square root of twice the information inaccuracies.

investigating the extent to which the *same* demand equations can explain consumption in the different OECD countries. The upper part of Table 14 contains the results using the root-mean-squared (RMS) percentage prediction error as the goodness-of-fit criterion. (The 15 OECD countries with 10 commodity groups are used here; see Table 1). Although the RMSEs for the country-specific models (given in column 3) are in general a bit lower than those for the pooled model (column 4), the differences are not substantial. In addition, the countries have similar rankings with respect to the two sets of RMSEs.

Looking at the entries in row 17 of columns 3 and 4 of Table 14, the weighted means of the RMSEs are 1.87 percent for the individual country models and 1.95 percent for the pooled model. Accordingly, the average "cost" of taking tastes to be identical is 1.95 − 1.87 = .08 percentage points. This is clearly quite modest and points in the direction that tastes are not too dissimilar across countries. (For a different finding, however, see Pollak and Wales, 1987.)

There are many non-economic differences between some of the OECD countries. For example, language, culture and climate differ substantially in some cases. It may thus come as a surprise to some (not Stigler and Becker!) that tastes in these countries are more or less similar. In an attempt to control for some of the non-economic factors, S. Selvanathan (1988b, Chap. 8) conducts a similar analysis with data from different regions of the same country, the six states of Australia. The lower part of Table 14 summarizes the findings and the result is that tastes are even more similar within a country, as expected. However, as the sample size for the states is small, it is appropriate to exercise some caution here.

9 The Frisch Conjecture

In Section 6 we showed that preference independence implies that the compensated price elasticities η'_{ij} take the form (6.4). We write that equation as

$$\eta'_{ij} = \phi\eta_i(\delta_{ij} - w_j\eta_j) \ , \tag{9.1}$$

where ϕ is the income flexibility (i.e., the reciprocal of the income elasticity of the marginal utility of income); η_i is the i^{th} income elasticity; δ_{ij} is the Kronecker delta ($\delta_{ij} = 1$ if $i = j$, $=0$ otherwise); and w_j is the budget share of good j. In rewriting equation (6.4) as (9.1), we have used the definition of the income elasticity $\eta_i = \theta_i/w_i$, θ_i being the i^{th} marginal share.

If we accept the assumption of preference independence, equation (9.1) means that we can then compute the $n \times n$ matrix of the price elasticities $[\eta'_{ij}]$ from the value of the scalar ϕ, the n income elasticities and n budget shares. It is the attraction of being able to compute price elasticities in this manner that has lead equation (9.1), or a variant thereof, to be used in at least two types of applications. First, in many developing countries time-series data of sufficient length and quality are not available, so it is impossible to estimate price elasticities. But as income elasticities and budget shares are usually available from household surveys, equation (9.1), together with a ϕ-value, can then be used to generate price elasticities. The second application of this equation is in computable general equilibrium models which have such a large number of commodities that the price elasticities cannot be directly estimated. For example, Dixon et al. (1982, Sec. 29.5) use this approach in a CGE model which distinguishes over 100 consumer goods.

What value of the income flexibility should be used in equation (9.1)? Our starting point for this question is Frisch's famous conjecture that ϕ increases in absolute value as the consumer (or country) becomes more affluent. (Frisch discusses the behavior of $1/\phi$, which he calls the "money flexibility"; for clarity, in what follows we make the necessary translations from $1/\phi$ to ϕ). Frisch (1959, p. 189) provides the following numerical conjectures for the dependence of ϕ on real income:

We may, perhaps, assume that in most cases the income flexibility has values of the order of magnitude given below.

−.1	for an extremely poor and apathetic part of the population.
−.25	for the slightly better off but still poor part of the population with a fairly pronounced desire to become better off.
−.5	for the middle income bracket, "the median part" of the population.
−1.4	for the better off part of the population.
−10	for the rich part of the population with ambitions towards "conspicuous consumption".

It would be a very promising research project to determine the income flexibility for different countries and for different types of populations. A universal "atlas" should be constructed. It would serve an extremely useful purpose in demand analysis.

Frisch's conjecture has been tested by a number of authors. Brown and Deaton (1972) report estimates of the income flexibility from various studies and conclude that they show no evidence of a dependence on real income. Also, Brown and Deaton suggest that it would be fair to use a value of −.5 for ϕ. DeJanvry et al. (1972) collect estimates of the income flexibility from several studies and find a statistically significant relationship between ϕ and income, which supports Frisch. As noted by Theil (1980), however, the validity of their conclusion is based on the uncritical acceptance of the previous estimates.

Lluch et al. (1977) use time-series data for 14 countries (developed and underdeveloped) at various levels of commodity aggregation to estimate a variant of the linear expenditure system. They find strong evidence in favor of Frisch at the four-commodity level of aggregation, but less support at the two-and one-commodity levels. S. Selvanathan (1987c,d) uses time-series data for the OECD countries and finds that ϕ is unrelated to income. Selvanathan also recommends a value of about −.5 for ϕ.

Using data for 30 countries, Theil (1987a) concludes that the sample cannot support Frisch in spite of the large income variation. Theil and Brooks (1970/71) use a modified version of the Rotterdam model in which the income flexibility is a function of income. They obtain the result that the income flexibility is again unrelated to income.

This brief review indicates a lack of unanimity about the status of Frisch. It seems, however, that the broad thrust of the literature has not been too kind to

Frisch's conjecture. This is possibly not unexpected given the nature of the conjecture: The income elasticity of the marginal utility of income relates to a second-order derivative of the (indirect) utility function, so that its dependence on income involves a third derivative. It is not possible for most economic data to reveal much information about such higher-order derivatives. Consequently, it is probably best to treat ϕ as a constant equal to around $-.5$. It should be noted that the constancy of the income flexibility is not inconsistent with the idea of constant tastes discussed in the previous section.

10 Some Further International Consumption Data

In the earlier sections of the paper we presented consumption data for 18 OECD countries. In this section we use a different database which contains both developed and developing countries. Relative to the OECD, these data exhibit more cross-country variability in income. The material for this and the next two sections is mainly from S. Selvanathan (1988b).

In a pioneering book, Lluch, Powell and Williams (1977, hereafter LPW) use data for 8 commodity groups in 17 countries. Table 15 presents an overview of the data for 13 of the 17 countries with countries ordered in terms of declining per capita GNP. (See S. Selvanathan, 1988b, for the reasons for not considering 4 of the LPW countries.) As can be seen, the US has the highest per capita GNP, while Korea has the lowest with only 4 percent of the US value. Table 16

Table 15. Characteristics of the LPW database

		Per capita GNP at sample midpoint	
Country (1)	Sample period (2)	In 1970 US dollars (3)	(3) with US = 100 (4)
1. US	1955–68	3669	100
2. Sweden	1955–68	2962	81
3. Australia	1955–66	2192	60
4. UK	1955–68	1900	52
5. Israel	1959–68	1468	40
6. Italy	1955–68	1207	33
7. Puerto Rico	1955–67	1023	28
8. Ireland	1955–68	1014	28
9. Greece	1958–68	676	18
10. South Africa	1955–68	596	16
11. Panama	1960–68	564	15
12. Thailand	1960–69	148	4
13. Korea	1955–68	142	4

Source: Lluch et al. (1977, Table 3.2).

Table 16. Budget shares of 8 commodities for 13 countries (Means × 100)

Country (1)	Food (2)	Clothing (3)	Housing (4)	Durables (5)	Personal care (6)	Transport (7)	Recreation (8)	Other services (9)
1. US	26.7	9.5	22.7	7.3	8.1	15.2	5.5	4.9
2. Sweden	36.6	11.5	15.9	7.0	3.7	14.2	8.9	2.4
3. Australia	33.3	11.0	12.7	7.8	5.7	13.1	4.3	12.1
4. UK	39.7	10.7	18.3	6.6	2.3	10.9	7.5	4.0
5. Israel	31.9	9.4	19.2	7.4	6.6	7.3	8.1	10.0
6. Italy	46.3	10.1	16.6	3.2	6.3	8.1	7.5	1.8
7. Puerto Rico	35.6	10.6	15.3	6.9	6.9	12.3	8.8	3.7
8. Ireland	49.2	10.0	12.8	5.7	1.3	8.8	6.3	5.9
9. Greece	46.8	12.2	18.3	3.9	3.6	6.9	6.1	2.3
10. South Africa	36.8	11.8	16.7	7.8	4.8	13.2	4.7	4.1
11. Panama	45.5	7.4	16.7	6.4	4.7	9.4	7.5	2.3
12. Thailand	57.6	8.4	8.0	3.2	5.6	7.9	7.5	1.8
13. Korea	59.9	10.4	11.4	2.8	4.2	4.7	4.4	2.2
14. Mean	42.0	10.2	15.7	5.8	4.9	10.2	6.7	4.4

Source: Lluch et al. (1977, Table 3.3).

Table 18. Divisia moments in 13 countries

Country (1)	Price index (2)	Volume index (3)	Price variance (4)	Quantity variance (5)	Price-quantity correlation (6)
1. US	1.89	2.47	.57	.97	.20
2. Sweden	3.52	2.83	.56	2.41	−.42
3. Australia	2.51	1.75	1.28	1.32	.09
4. UK	2.88	2.52	1.33	2.15	.05
5. Israel	5.85	6.30	2.89	4.40	−.31
6. Italy	2.82	4.94	.79	2.53	−.30
7. Puerto Rico	2.40	2.49	1.32	5.17	−.30
8. Ireland	3.11	2.97	.73	1.95	−.42
9. Greece	1.93	5.71	.50	2.86	−.80
10. South Africa	2.23	1.67	1.24	.96	−.47
11. Panama	1.21	2.76	.65	.73	−.15
12. Thailand	1.92	3.37	3.63	4.51	−.38
13. Korea	13.36	2.93	1.27	5.11	.05
14. Mean	3.51	3.29	1.29	2.70	−.24

All entries in columns 2 and 3 are to be divided by 100; and those in columns 4 and 5 are to be divided by 10,000.

presents the budget shares at sample means of the 8 commodities in each country. Note the strong tendency for the food budget share to decline with increasing GNP which is in accordance with Engel's law. The upper half of Table 17 presents the average annual price log-changes, while the lower half presents the corresponding per capita quantity log-changes.

Next, we summarize these data with Divisia indexes. Let $\overline{w}_{ic}$ be the budget share of i for country c at sample means (given in Table 16); and Dp_{ic} and Dq_{ic} be the mean price and quantity log-changes (Table 17). The Divisia price and volume indexes for country c are then

$$DP_c = \sum_{i=1}^{8} \overline{w}_{ic} Dp_{ic} , \qquad DQ_c = \sum_{i=1}^{8} \overline{w}_{ic} Dq_{ic}$$

and are presented in columns 2 and 3 of Table 18. As can be seen from the last row of the table, on average, prices in these countries increase by 3.5 percent per annum while per capita consumption grows at a slightly lower rate. Columns 4–6 of the table contain the corresponding second-order moments defined in Section 3.

11 The Linear Expenditure System

Stone's (1954) linear expenditure system is probably the most popular demand system. Notable studies using this model include Deaton (1975), Goldberger

Table 17. Prices and per capita quantities consumed of 8 commodities for 13 countries (Mean annual log-changes × 100)

Country (1)	Food (2)	Clothing (3)	Housing (4)	Durables (5)	Personal care (6)	Transport (7)	Recreation (8)	Other services (9)
Prices								
1. US	1.85	1.62	1.70	.55	3.04	1.52	2.35	4.31
2. Sweden	4.00	2.25	3.85	1.79	2.85	3.46	4.22	3.70
3. Australia	2.32	1.39	4.87	.48	3.15	1.96	2.92	3.02
4. UK	2.11	1.70	4.68	1.69	3.09	2.93	4.62	3.80
5. Israel	4.85	4.65	7.81	1.94	6.46	8.06	5.97	7.22
6. Italy	2.52	2.17	3.65	−.16	4.24	2.27	4.07	4.19
7. Puerto Rico	2.62	1.44	1.27	.83	5.76	2.83	3.20	2.46
8. Ireland	3.18	.89	3.98	3.15	1.33	3.13	3.63	4.24
9. Greece	2.57	1.20	1.59	.67	2.48	1.87	.55	.69
10. South Africa	2.27	.20	3.50	.32	3.90	2.40	2.49	3.40
11. Panama	1.95	.59	.54	.11	2.24	.04	.97	.00
12. Thailand	3.24	.19	−2.63	−1.64	1.33	.06	2.43	2.02
13. Korea	13.28	14.75	11.27	12.54	16.70	12.69	13.91	14.80
14. Mean	3.60	2.54	3.54	1.71	4.35	3.29	3.95	4.14
Quantities								
15. US	.99	2.42	2.79	3.70	3.90	2.78	2.86	3.63
16. Sweden	1.51	1.76	2.77	5.25	5.62	5.32	2.28	4.03
17. Australia	.78	.66	2.34	3.14	4.09	3.12	.10	1.89
18. UK	1.62	2.44	1.98	3.08	3.75	5.53	1.37	6.48
19. Israel	4.62	6.87	5.41	11.67	5.96	9.24	8.58	5.17
20. Italy	4.09	4.62	4.62	8.51	5.96	9.46	4.28	4.47
21. Puerto Rico	−.28	3.91	3.25	4.45	1.40	5.13	4.89	5.89
22. Ireland	1.91	4.88	2.51	5.85	2.71	5.32	3.72	2.47
23. Greece	4.14	7.92	5.99	6.41	5.67	8.94	8.08	6.48
24. South Africa	1.04	2.20	.57	3.21	1.81	3.32	1.22	2.48
25. Panama	2.59	3.23	1.92	5.06	2.55	2.01	4.36	2.84
26. Thailand	2.15	5.30	1.06	10.16	3.70	6.55	5.59	7.16
27. Korea	1.94	2.35	2.27	6.88	4.95	11.55	5.69	3.31
28. Mean	2.08	3.74	2.88	5.95	4.01	6.02	4.08	4.33

Source: Derived from Lluch et al. (1977, Table 3.4 and 3.5); see S. Selvanathan (1988b, Chap. 1) for details.

and Gamaletsos (1970), Kravis et al. (1982), Lluch and Powell (1975), Lluch et al. (1977), Parks (1969), Pollak and Wales (1969, 1992) and Yoshihara (1969). In this section we set out the linear expenditure system and then present some estimates from LPW.

Our starting point is the well-known Klein-Rubin (1948) utility function,

$$u(q_1, \ldots, q_n) = \sum_{i=1}^{n} \theta_i \log(q_i - \gamma_i) , \tag{11.1}$$

where θ_i and γ_i are constants satisfying $\theta_i > 0$, $\sum_{i=1}^{n} \theta_i = 1$ and $q_i > \gamma_i$ for each i. Maximizing (11.1) subject to the budget constraint gives the corresponding demand equations. It is convenient to express these in expenditure form,

$$p_i q_i = p_i \gamma_i + \theta_i \left[M - \sum_{j=1}^{n} p_j \gamma_j \right] , \qquad i = 1, \ldots, n . \tag{11.2}$$

This model is known as the linear expenditure system (LES).

The linearity of LES is attractive in its simplicity. When the γ_i's are all positive, the model has the following intuitive interpretation: The consumer first purchases the "subsistence" quantities $\gamma_1, \ldots, \gamma_n$ at a cost of $\sum_{j=1}^{n} p_j \gamma_j$. This leaves $M - \sum_{j=1}^{n} p_j \gamma_j$ of unspent income which can be called "supernumerary" income. Then, a fraction θ_i of this supernumerary income is spent on good i. Note also that $\theta_i = \partial(p_i q_i)/\partial M$ is the i^{th} marginal share.

Two other aspects of LES should be noted. First, it cannot be used to test the homogeneity and symmetry hypotheses as these are built in or maintained hypotheses in this model. Second, equation (11.1), the utility function underlying LES, is of the preference independent form. As discussed in Section 6, this imposes certain restrictions on the price elasticities. But, as argued in Section 7, these restrictions may not be at great variance with the data if the model is applied to broad commodity groups.

We return to the LPW database described in the previous section. LPW use time-series data to estimate LES (or a variant thereof, Lluch's, 1973, extended linear expenditure system) for the 13 countries listed in Table 15. The model is estimated for each country independently of the others. The upper part of Table 19 presents the estimates of the marginal shares θ_{ic} ($i = 1, \ldots, 8$ goods; $c = 1, \ldots, 13$ countries), while the lower part gives the implied income elasticities,

$$\eta_{ic} = \frac{\theta_{ic}}{\bar{w}_{ic}} , \tag{11.3}$$

with the mean budget shares from Table 16. As can be seen from the last row of Table 19, on average food and housing are necessities, clothing is a borderline case, while the remaining five other goods are luxuries.

Notwithstanding its popularity, LES has its drawbacks. Perhaps the most important is the parameterization whereby the marginal shares are treated as constants. It is clear from (11.3) that the constancy of the marginal share implies that the income elasticity is inversely proportional to the corresponding budget share. This can give rise to problems when income is subject to large changes.

Table 19. First set of marginal shares and income elasticities of 8 commodities for 13 countries

Country (1)	Food (2)	Clothing (3)	Housing (4)	Durables (5)	Personal care (6)	Transport (7)	Recreation (8)	Other services (9)
Marginal shares								
1. US	.090	.108	.206	.106	.137	.174	.065	.113
2. Sweden	.278	.071	.145	.079	.052	.253	.097	.024
3. Australia	.143	.050	.220	.082	.133	.224	.009	.138
4. UK	.120	.067	.258	.076	.030	.277	.066	.106
5. Israel	.210	.103	.172	.119	.066	.116	.117	.097
6. Italy	.401	.087	.171	.069	.066	.117	.070	.019
7. Puerto Rico	.177	.112	.144	.068	.114	.176	.137	.072
8. Ireland	.315	.134	.108	.114	.014	.170	.075	.070
9. Greece	.341	.168	.175	.050	.048	.107	.083	.028
10. South Africa	.295	.164	.066	.115	.049	.206	.046	.059
11. Panama	.418	.081	.113	.113	.044	.085	.127	.019
12. Thailand	.482	.101	.013	.052	.052	.123	.150	.028
13. Korea	.434	.069	.085	.077	.074	.146	.078	.038
14. Mean	.285	.101	.144	.086	.068	.167	.086	.062
Income elasticities								
15. US	.337	1.137	.907	1.452	1.691	1.145	1.182	2.306
16. Sweden	.760	.617	.912	1.129	1.405	1.782	1.090	1.000
17. Australia	.429	.455	1.732	1.051	2.333	1.710	.209	1.140
18. UK	.302	.626	1.410	1.152	1.304	2.541	.880	2.650
19. Israel	.658	1.096	.896	1.608	1.000	1.589	1.444	.970
20. Italy	.866	.861	1.030	2.156	1.048	1.444	.933	1.056
21. Puerto Rico	.497	1.057	.941	.986	1.652	1.431	1.557	1.946
22. Ireland	.640	1.340	.844	2.000	1.077	1.932	1.190	1.186
23. Greece	.729	1.377	.956	1.282	1.333	1.551	1.361	1.217
24. South Africa	.802	1.390	.395	1.474	1.021	1.561	.979	1.439
25. Panama	.919	1.095	.677	1.766	.936	.904	1.693	.826
26. Thailand	.837	1.202	.163	1.625	.929	1.557	2.000	1.556
27. Korea	.725	.663	.746	2.750	1.762	3.106	1.773	1.727
28. Mean	.654	.994	.893	1.572	1.346	1.712	1.253	1.463

Source: The marginal shares are from Lluch et al. (1977, Table 3.6). The income elasticities are derived using equation (11.3).

Consider the case of food, the dominant commodity in most countries. By Engel's law, food is a necessity (i.e., $\eta_i < 1$ for $i = 1$, representing food, the first commodity). Consequently, if prices remain constant, a rise in income causes consumption of food to increase less than proportionately so that its budget share falls. It then follows from $\eta_1 = \theta_1/w_1$ that as the consumer becomes more affluent, the food income elasticity rises when θ_1 is specified as a constant. That is, food becomes less of a necessity or more of a luxury with increasing income. This behaviour of the elasticity under LES is clearly implausible, a criticism made by Theil (1983).

To illustrate, suppose we take tastes to be the same internationally (following the advice of Section 8), pool the LPW data across countries and then apply LES. This model would have the same parameters for each country. From column 2 of Table 19, the mean of the individual-country estimates of the food marginal share is .29. It is sufficient for purposes of illustration to use this as the cross-country value of this share. The observed food budget shares are given in Table 16: For the US, consumers spend 27 percent of their income on food, while the Koreans spend 60 percent. Thus we have:

	US	Korea
Food marginal share θ_1	.29	.29
Food budget share $\bar{w}_{1c}$	.27	.60
Food income elasticity $\eta_{1c} = \theta_1/\bar{w}_{1c}$	1.07	.48

The above calculation reveals a spectacular disparity in the income elasticity of food. This elasticity for the US (the richest country) is more than twice that for Korea (the poorest). In fact, if we take the results literally, food is a luxury in the US! Note also that these elasticities are nothing like the individual-country estimates given in the lower part of column 2 of Table 19. Clearly, LES fails in this cross-country application.

12 More on Working's Model

In the previous section we saw that the behaviour of income elasticities implied by LES is unattractive. It should be added that the Rotterdam model (Barten, 1964, Theil, 1965), in which the marginal shares are also specified as constants, suffers from the same defect. In this section we show that Working's (1943) model circumvents the problem.

Working's model is given by equation (4.1) which we reproduce here:

$$w_i = \alpha_i + \beta_i \log M \ , \qquad i = 1, \dots, n \ . \tag{12.1}$$

Using $w_i = p_i q_i / M$ in (12.1) and multiplying both sides by M, we get

$$p_i q_i = \alpha_i M + \beta_i M \log M \ .$$

Differentiating both sides of this equation with respect to M and using (12.1), we obtain the marginal share of good i implied by Working's model,

$$\theta_i = w_i + \beta_i \ . \tag{12.2}$$

Since the budget share w_i is not constant, neither is the marginal share under Working's model.

The income elasticity is the ratio of the marginal share to the corresponding budget share, $\eta_i = \theta_i/w_i$. It then follows from (12.2) that the income elasticity of i implied by Working's model is

$$\eta_i = 1 + \frac{\beta_i}{w_i} \ . \tag{12.3}$$

This shows that a commodity is a necessity (luxury) if β_i is negative (positive). As the budget share of a necessity falls with increasing income, this expression implies that the income elasticity of such a good falls when income increases. By a similar argument, the income elasticities of luxuries also fall as income rises. That is, as the consumer becomes more affluent, all goods become less luxurious under Working's model, which is plausible. This result is in contrast to the behaviour of the income elasticities from LES and the Rotterdam model, in which the marginal shares are constants. Therefore, Working's model is a plausible alternative to taking the marginal shares as constants.

Like all models, Working's is still not perfect however. Equation (12.1) implies that for sufficiently low or high values of income, the budget share will stray outside the [0, 1] interval. Consequently, the model is only locally valid. It could be argued that as the same local qualification must also apply with equal force to LES, we have been a little too harsh in condemning the constancy of the marginal shares of that model. Although the income elasticities implied by LES always move in the wrong direction, it may be that this is insufficient to cause major problems.

From equation (12.2), in Working's model the marginal share differs from the corresponding budget share by a constant,

$$\beta_i = \theta_i - w_i \ .$$

Thus, if we have a number of estimates of the marginal share for commodity i, θ_{ic}, $c = 1, \ldots, N$, as well as w_{ic}, $c = 1, \ldots, N$, we can obtain a simple estimate of β_i by averaging,

$$\hat{\beta}_i = \frac{1}{N} \sum_{c=1}^{N} (\theta_{ic} - w_{ic}) \ . \tag{12.4}$$

To implement (12.4) for food, we use the $N = 13$ countries from LPW, the estimates of the θ_{ic}'s given in Table 19 and the corresponding budget shares from Table 16. (Note that as the estimates of the marginal shares are obtained country-by-country by LPW, each value is independent of the other.) Table 20 contains the results for food ($i = 1$). The standard errors given in the last row indicate that the differences $(\theta_{1c} - w_{1c})$ are less variable than the marginal shares

Table 20. Marginal shares and budget shares of food for 13 countries

Country (1)	Marginal share (2)	Budget share (3)	Excess of marginal share over budget share (4)
1. US	.090	.267	−.177
2. Sweden	.278	.366	−.088
3. Australia	.143	.333	−.190
4. UK	.120	.397	−.277
5. Israel	.210	.319	−.109
6. Italy	.401	.463	−.062
7. Puerto Rico	.177	.356	−.179
8. Ireland	.315	.492	−.177
9. Greece	.341	.468	−.127
10. South Africa	.295	.368	−.073
11. Panama	.418	.455	−.037
12. Thailand	.482	.576	−.094
13. Korea	.434	.599	−.165
14. Mean	.285	.420	−.135
15. Standard error of the mean	.034	.026	.018

Source: Columns 2 and 3, Lluch et al. (1977, Tables 3.6 and 3.3).

Table 21. Estimates of working's income coefficients from LPW data (Standard errors in parentheses)

Commodity (1)	Income coefficient β_i (2)
1. Food	−.135 (.018)
2. Clothing	−.001 (.009)
3. Housing	−.013 (.014)
4. Durables	.028 (.005)
5. Personal care	.019 (.007)
6. Transport	.066 (.012)
7. Recreation	.019 (.008)
8. Other services	.018 (.006)

θ_{1c}, which points in the direction of parameterizing on β_1, rather than θ_1. Recall from Section 4 that previous estimates of Working's income coefficient for food (β_1) were closely clustered around the value −.15. The new estimate of this coefficient given in row 14 of column 4 of Table 20 is −.14 with standard error .02, which is obviously in close agreement with the other estimates. The emergence of more or less the same value of β_i in a wide variety of applications seems to give reassurance that some fundamental behavioural response is being measured here, a response which is more or less constant across consumers, time and countries.

LPW consider the following relation between the food marginal share, θ_1, and GNP per capita, Y:

$$\theta_1 = a + b \log Y \,, \tag{12.5}$$

where a and b are constants. To interpret this equation, we combine (12.1) and (12.2) for $i =$ to yield

$$\theta_1 = \lambda + \beta_1 \log M \ , \tag{12.6}$$

where $\lambda = \alpha_1 + \beta_1$ is a constant and β_1 is Working's income coefficient for food. A comparison of (12.5) and (12.6) reveals that under the not unreasonable assumption that total expenditure M is proportional to Y, the coefficient b in (12.5) is interpreted as β_1.

LPW (1977, Table 3.11, p. 52) apply (12.5) to their 17 countries and obtain the LS estimate of b of $-.179$ with standard error .045. Using the 13 countries listed in Table 15, GNP of that table and the food marginal shares of Table 19, yields a b-estimate of $-.103$ with standard error .023. Here again, these are insignificantly different from the centre-of-gravity value of $-.15$.

We return to (12.4), the estimator of Working's income coefficient for good i, and use the LPW data to implement it for $i = 1, \ldots, 8$ commodities. The procedure is exactly the same as that used for food and Table 21 contains the results. The estimates of β_i for food and housing are negative; that for clothing is close to zero; and the remaining ones are positive. This shows that food and housing are necessities, clothing is a borderline case, while the other five goods are luxuries. This is in agreement with the average income elasticities from LPW given in Table 19.

To explore further the implications of the cross-country β_i-estimates, we compute the implied marginal shares and income elasticities and Table 22 contains the results. These are to be compared with the LPW values given in Table 19. We can regard the values in Table 22 as "fitted" and those in Table 19 as "observed". The correlation coefficients between the two sets of marginal shares are:

Food	.86	Personal care	.79
Clothing	.31	Transport	.65
Housing	.64	Recreation	.73
Durables	.61	Other services	.82

As can be seen, except for clothing, the two sets of marginal shares agree well which tends to support the idea that Working's model can be satisfactorily applied to the world as a whole (the "world" being the 13 LPW countries).

13 A World Demand System

The results of the previous section are based on income responses of demand which hold constant prices. Those results are sufficiently encouraging to pro-

Table 22. Second set of marginal shares and income elasticities of 8 commodities for 13 countries

Country (1)	Food (2)	Clothing (3)	Housing (4)	Durables (5)	Personal care (6)	Transport (7)	Recreation (8)	Other services (9)
Marginal shares								
1. US	.132	.094	.214	.101	.100	.218	.074	.067
2. Sweden	.231	.114	.146	.098	.056	.208	.108	.042
3. Australia	.198	.109	.114	.106	.076	.197	.062	.139
4. UK	.262	.106	.170	.094	.042	.175	.094	.058
5. Israel	.184	.093	.179	.102	.085	.139	.100	.118
6. Italy	.328	.100	.153	.060	.082	.147	.094	.036
7. Puerto Rico	.221	.105	.140	.097	.088	.189	.107	.055
8. Ireland	.357	.099	.115	.085	.032	.154	.082	.077
9. Greece	.333	.121	.170	.067	.055	.135	.080	.041
10. South Africa	.233	.117	.154	.106	.067	.198	.066	.059
11. Panama	.320	.073	.154	.092	.066	.160	.094	.041
12. Thailand	.441	.083	.067	.060	.075	.145	.094	.036
13. Korea	.464	.103	.101	.056	.061	.113	.063	.040
14. Mean	.285	.101	.144	.086	.068	.168	.086	.062
Income elasticities								
15. US	.494	.989	.943	1.384	1.235	1.434	1.345	1.367
16. Sweden	.631	.991	.918	1.400	1.514	1.465	1.213	1.750
17. Australia	.595	.991	.898	1.359	1.333	1.504	1.442	1.149
18. UK	.660	.991	.929	1.424	1.826	1.606	1.253	1.450
19. Israel	.577	.989	.932	1.378	1.288	1.904	1.235	1.180
20. Italy	.708	.990	.922	1.875	1.302	1.815	1.253	2.000
21. Puerto Rico	.621	.991	.915	1.406	1.275	1.537	1.216	1.486
22. Ireland	.726	.990	.898	1.491	2.462	1.750	1.302	1.305
23. Greece	.712	.992	.929	1.718	1.528	1.957	1.311	1.783
24. South Africa	.633	.992	.922	1.359	1.396	1.500	1.404	1.439
25. Panama	.703	.986	.922	1.438	1.404	1.702	1.253	1.783
26. Thailand	.766	.988	.838	1.875	1.339	1.835	1.253	2.000
27. Korea	.775	.990	.886	2.000	1.452	2.404	1.432	1.818
28. Mean	.662	.990	.912	1.547	1.489	1.724	1.301	1.578

The marginal shares are computed from equation (12.2) with the observed budget shares given in Table 16 and the β_i-estimates given in Table 21. The income elasticities are computed from (12.3) using the the same approach.

ceed further and extend the approach to deal simultaneously with the effects of income and prices. In this section we ask the question, how much of the cross-country variation in consumption patterns is explained by differences in incomes and prices? To do this we employ a simple extension of Working's model which contains a price substitution term.

Our starting point is equation (6.3), the double-log demand equation for good i, which we reproduce here:

$$\log q_i = \alpha_i + \eta_i \log Q + \sum_{j=1}^{n} \eta'_{ij} \log p_j \ ,$$

where α_i is an intercept; η_i is the income elasticity of i; Q is real income; and η'_{ij} is the $(i, j)^{\text{th}}$ compensated price elasticity. To apply this equation to the LPW cross-country data, we write it in terms of changes over time; add to the variables and elasticities a country subscript c $(c = 1, \ldots, 13)$; and set $n = 8$ commodities. This yields

$$Dq_{ic} = \eta_{ic} DQ_c + \sum_{j=1}^{8} \eta'_{ijc}\, Dp_{jc} \ , \qquad (13.1)$$

where D is the log-change operator; and DQ_c is the Divisia volume index of the change in real income. Note that as log-changes are unit free, all variables in (13.1) are directly comparable across countries, independent of the different currency units.

The demand system (13.1) for $i = 1, \ldots, 8$ contains an 8×8 matrix of price elasticities $[\eta'_{ijc}]$ for each of the 13 countries. To reduce the number of unknowns, we shall invoke the assumption of preference independence. As discussed in Section 6, the price elasticities for country c then take the form

$$\eta'_{ijc} = \phi \eta_{ic}(\delta_{ij} - \theta_{jc}) \ , \qquad i, j = 1, \ldots, 8 \ , \qquad (13.2)$$

where ϕ is the income flexibility (the reciprocal of the income elasticity of the marginal utility of income); δ_{ij} is the Kronecker delta ($\delta_{ij} = 1$ if $i = j$, $=0$ otherwise); and θ_{jc} is the j^{th} marginal share in c. The use of the assumption of preference independence is probably not too bad here as the 8 commodities in the LPW database are broad aggregates. Note that as ϕ in equation (13.2) does not have a country subscript, it as specified as a constant; this is in agreement with the discussion of Section 9 above.

Let Dq_{ic} be observed value of the log-change in consumption of i in c; and $\eta_{ic} DQ_c + \sum_{j=1}^{8} \eta'_{ijc} Dp_{jc}$ be the corresponding fitted value when numerical values are used for the elasticities. The values of the variables Dq_{ic} and Dp_{ic} are given in Table 17, while those of DQ_c are in column 3 of Table 18. If e_{ic} is the difference between the observed and fitted values, then one way to measure the quality of the predictions for commodity i is the root-mean-squared error,

$$\text{RMSE}_i = \sqrt{\frac{1}{13} \sum_{c=1}^{13} e_{ic}^2} \ .$$

The performance of the system as a whole can then be measured by the budget-share-weighted average of the RMSE_i,

$$\text{RMSE} = \sum_{i=1}^{8} \bar{w}_1 \text{RMSE}_i ,$$

where $\bar{w}_i$ is the cross-country average of the budget share of i, given in the last row of Table 16.

We first use the LPW estimates of the income elasticities given in the lower half of Table 19, the marginal shares of the upper half of that table and set the income flexibility $\phi = -.5$, the value recommended in Section 9. Column 2 of Table 23 contains the results. As can be seen, food and housing share the lowest value of the RMSE of about .5 percent; the largest RMSE is 1.9 percent for durables; while the weighted average is .7 percent. Overall, these RMSEs are very low. However, it should be kept in mind that underlying these predictions are $n - 1 = 7$ free income elasticities (one is constrained by $\sum_{i=1}^{8} w_i \eta_i = 1$) for each of the 13 countries, plus a value of ϕ. Consequently, the column 2 results are based on $7 \times 13 + 1 = 92$ free parameters. As the income elasticities are unconstrained across countries, in a certain sense tastes are not constant in the demand equations underlying the results of column 2.

Next, we specify that all income elasticities are unity, so that the marginal shares are equal to the corresponding budget shares, and $\phi = -1$. This is a naive extrapolation corresponding to no-change of the budget shares (Theil, 1975/76, p. 219 of Volume 1), which serves as a standard of reference. The results in column 3 of Table 23 show that most of the RMSE_i increase substantially in relation to those of column 2 . The weighted-average RMSE of 1.9 percent in column 3 is about 2.5 times larger than its column 2 counterpart.

Finally, we use the income elasticities and marginal shares implied by Working's model (given in Table 22) and $\phi = -.5$. Here tastes are taken to be con-

Table 23. Quality of predictions of demand equations for 8 commodities in 13 countries (Root-mean-squared errors × 100)

Commodity (1)	η_{ic} = LPW, $\phi = -.5$ (2)	$\eta_{ic} = 1, \phi = -1$ (3)	$\eta_{ic} = WM, \phi = -.5$ (4)
1. Food	.51	1.41	.64
2. Clothing	.61	1.20	1.02
3. Housing	.51	2.30	1.23
4. Durables	1.89	1.63	1.70
5. Personal care	1.15	2.33	2.41
6. Transport	1.01	3.43	1.73
7. Recreation	.62	1.79	1.28
8. Other services	.88	2.42	2.37
9. Weighted mean	.71	1.86	1.15

As the demand equations are formulated in log-changes, the RMSEs × 100 are (approximately) RMS percentage prediction errors. The notation η_{ic} = LPW denotes that the income elasticities used for the predictions are the LPW estimates given in the lower part of Table 19; and $\eta_{ic} = WM$ denotes that the income elasticities are from Working's model, given in the lower part of Table 22.

stant as the income elasticities (and the marginal shares) are based on coefficients from Working's model, $\beta_1, \ldots, \beta_8$, which are the same for all countries. Column 4 of Table 23 contains the results. These RMSEs are mostly substantially lower than those of column 3, where all elasticities are unitary. At the same time however, the column 4 values are in general higher than the LPW values given in column 2; the weighted-average RMSE in column 4 of 1.2 percent is about 65 percent higher than its column 2 counterpart. Nevertheless, the average RMSE of 1.2 percent is an excellent result keeping in mind the diversity of countries involved. In addition, the predictions of column 4 are based on only 8 free parameters (7 β_i's, as one is constrained by $\sum_{i=1}^{8} \beta_i = 0$, plus ϕ). Consequently, the increase in the average RMSE from .7 percent (in column 2) to 1.2 percent (column 4) is perfectly acceptable given that the number of free parameters falls from 92 to 8.

It is thus fair to make the strong claim that this approach allows world consumption patterns to be adequately described with only 8 parameters. Of course, the qualifications to this claim are obvious: Here we are dealing with broad commodity groups and the "world" is made up of only 13 representative countries. (For a more elaborate analysis of the LPW data, see Chen, 1991, and S. Selvanathan, 1988b.)

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