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Journal of Applied Econometrics, Vol. 12, No. 5, Special Issue: The Experiment in Applied Econometrics. (Sep. - Oct., 1997), pp. 593-608.

JOURNAL OF APPLIED ECONOMETRICS, VOL. 12, 593–613 (1997)

A COMPARATIVE STUDY OF MODELLING THE DEMAND FOR FOOD IN THE UNITED STATES AND THE NETHERLANDS

HAIYAN SONG,[a] XIAMING LIU[b] AND PETER ROMILLY[b*]

[a]*Department of Management Studies, University of Surrey, Guildford, GU2 5XH, UK*
[b]*Division of Economics, School of Social Sciences, University of Abertay Dundee, Dundee DD1 1HG, UK*

SUMMARY

This paper provides time-series and cross-sectional budget survey analyses of the demand for food in the United States and the Netherlands according to the tasks set by Jan Magnus and Mary Morgan (MM). Various econometric methods, including weighted least squares (WLS), cointegration, error correction, the almost ideal demand system (AIDS), and time-varying parameter (TVP) techniques, are used and the estimated demand elasticities compared across country and over time.

J. Appl. Econ., **12**, 593–613 (1997)

No. of Figures: 1. No. of Tables: 10. No. of References: 20.

1. INTRODUCTION

Tobin's (1950) paper on food demand analysis makes systematic use of both cross-section and time-series data to derive a demand function for food in the USA. Tobin first estimates the income elasticity of demand for food from the budget survey data and the income elasticity is then used to estimate a reduced form time-series model.

Tobin's method of pooling the cross-section and time-series data was subsequently criticized by Chetty (1968), Maddala (1971), and Izan (1980). Chetty (1968) proposes a simultaneous data pooling system estimated using the Bayesian approach. Chetty's proposal is questioned by Maddala (1971) who suggests that a pre-estimation test should be considered to test the hypothesis that the income elasticities in the two types of models are equal before the data is pooled. Izan (1980) argues that the failure to account for the existence of outliers in the budget survey data, and the presence of serially correlated residuals in the time-series data, invalidates both the Chetty and Maddala conclusions. Izan uses a weighted least squares (WLS) method to correct for the presence of the outlier problem and finds that the cross-section and time-series income elasticities are not different from each other, so that the pooling of the two types of data is appropriate.

In Section 2 we undertake Task 1 set by MM and extend the study by Tobin (1950) using the budget survey data of 1950, 1960, and 1972. A WLS method suggested by Izan (1980) is used to estimate the demand models and obtain unbiased efficient estimates of the income elasticities. Section 3 relates to Task 2 of MM and focuses on the estimation of aggregate time-series demand for food models for the USA. Using the general to specific approach we examine the long-run

* Correspondence to: P. Romilly, Division of Economics, School of Social Sciences, University of Abertay Dundee, Dundee DD1 1HG, UK. E-mail: p.romilly@tay.ac.uk

CCC 0883–7252/97/050593–21$17·50

Received 1 September 1996
Revised 23 April 1997

equilibrium relationship between consumers' expenditure on food and related variables for the USA and the corresponding error correction models. Section 4 carries out the remaining part of Task 2 by estimating similar models for the Netherlands. A disaggregate time-series model is also constructed for the Netherlands. Section 5 tackles our own selected task (i.e. Task 5 of MM), that of an alternative modelling strategy using the time-varying parameter (TVP) approach to examine changes in the various demand elasticities over time. The forecasting exercise (Task 3) specified by MM is undertaken in Section 6. Section 7 provides conclusions. A brief summary of the logbook is provided in the Appendix.

2. THE MODEL OF DEMAND FOR FOOD IN THE USA: ESTIMATES FROM THE BUDGET SURVEY DATA

This and the following section deal with Task 1(a) specified by MM, namely to estimate the income elasticity of food demand in the USA using the 1950, 1960, and 1972 budget survey data sets and compare the results with that of Tobin (1950) for the 1941 budget survey data. There are a number of household-related variables which could be added to our model specification, but we retain Tobin's model for the following reasons.

First, although the composition of households (such as average number of children and average number of persons over 65 years per household) may affect food consumption, data are not complete for all the years 1950, 1960, and 1972. Furthermore, even if we reassess the size of households using so-called 'equivalent adult scales' (e.g. the 'Amsterdam scale' assigned a weight of 0·9 to an adult female) there can be no guarantee that households plan their budgets on the basis of such an assumption. Second, food consumption may also be influenced by other household variables such as the percentage of homeowning households (*HOMEOWN*). Inclusion of the homeowner variable in the model, however, shows that its coefficient is significant for 1960, but not for 1950 and 1972. Third, adopting the same model as Tobin allows direct comparisons to be made with our results.

The Tobin (1950) demand for food specification based on the budget survey data is:

$$\log \bar{c}_i = k + \alpha \log \bar{Y}_i + \delta \log n_i + u_i \tag{1}$$

where $\bar{c}_i$ is the average consumption of food for the ith group of families ($i = 1, 2, \ldots, N$), $\bar{Y}_i$ is the average disposable income, n_i is the size of family, and u_i is an error term with $u_i \sim \mathrm{ND}(0, \sigma_{u_i})$.

We re-estimate the Tobin model (1) using his data and the budget survey data for 1950, 1960, and 1972. The results are presented in Table I. The table shows problems of heteroscedasticity in all the regressions. Scatter diagrams (contained in the full logbook) of income against consumer expenditure on food indicate that the data sets for 1950, 1960, and 1972 appear to have outlier observations. Moreover, the diagrams show that the pattern of the relationship between income and food consumption in 1972 appears to be different from that of 1950 and 1960. This may be an indication of structural change in the food demand function.

One option is to eliminate the outliers from the data. But although the outlier observations are undesirable they may still contain valuable information, so that their complete elimination may result in biased estimates of the elasticities. In order to remedy the outlier problem without losing too much information, a WLS method suggested by Izan (1980) is used to obtain the estimates of the demand elasticities. The variables in the demand model are weighted by $(|\hat{u}_i|_{1/2} + \gamma)^{-1}$ where $|\hat{u}_i|$ is the absolute value of the residual obtained from the

Table I. Parameter estimates based on US budget survey data

Parameter	Tobin	1950	1960	1972
k	0·823	2·867	2·262	3·958
	(8·57)[a]	(17·96)	(10·660)	(31·422)
α	0·561	0·485	0·531	0·330
	(18·88)	(25·70)	(21·510)	(23·881)
δ	0·254	0·261	0·308	0·455
	(6·93)	(10·28)	(9·174)	(27·100)
R^2	0·93	0·94	0·910	9·950
S.E.	0·109	0·120	0·165	0·091
χ^2 white (4)[b]	9·811[c]	22·69[c]	11·22[c]	13·790[c]
Obs.	37	54	59	72

[a] The figures in parentheses are t-statistics.
[b] The White test for heteroscedasticity, where the value in parentheses is the degree of freedom. The critical value of χ^2 (4) at the 5% level is 9·488.
[c] The statistic is significant at the 5% level.

Table II. Parameter estimates after correcting for the outlier problem

		Parameters				
	γ	k	α	δ	R^2	F
1941	0·5	0·981	0·658	0·212	0·950	328·0
		(0·167)	(0·027)	(0·041)		
	0·7	0·834	0·630	0·220	0·940	292·0
		(0·124)	(0·028)	(0·040)		
	1·0	0·662	0·607	0·228	0·940	272·0
		(0·089)	(0·028)	(0·038)		
1950	0·5	1·308	0·697	0·386	0·987	1970·0
		(0·178)	(0·014)	(0·038)		
	0·7	1·267	0·671	0·373	0·983	1565·7
		(0·123)	(0·015)	(0·035)		
	1·0	1·177	0·641	0·358	0·981	1283·5
		(0·097)	(0·015)	(0·032)		
1960	0·5	0·841	0·728	0·304	0·982	1495·7
		(0·175)	(0·015)	(0·040)		
	0·7	0·872	0·704	0·309	0·976	1151·4
		(0·150)	(0·016)	(0·039)		
	1·0	0·866	0·673	0·312	0·970	903·4
		(0·121)	(0·017)	(0·036)		
1972	0·5	1·538	0·646	0·470	0·952	694·1
		(0·241)	(0·019)	(0·048)		
	0·7	1·580	0·608	0·463	0·937	513·9
		(0·207)	(0·021)	(0·046)		
	1·0	1·558	0·559	0·456	0·921	403·0
		(0·164)	(0·022)	(0·041)		

Notes:
1. Figures in parentheses are standard errors.
2. Numbers of observations for the estimation of the four models are the same as those given in Table I.

estimated cross-section regression model without considering the outlier problem. γ is a scalar greater than 0, and is included because some values of the residuals are close to zero. The estimated results based on different γ values are shown in Table II. The table shows that for a given year the demand elasticities with the outlier corrections are generally higher than those without the corrections, although Tobin's δ is lower. The heteroscedasticity problem is eliminated in all cases. The values of the elasticities appear sensitive to the choice of γ, however, and there is some variation in them over time.

3. DEMAND FOR FOOD MODELS IN THE USA: ESTIMATES FROM THE TIME-SERIES DATA

In this section we use general to specific methodology in conjunction with cointegration analysis to estimate the short- and long-run relationships between income and food consumption. The MM information pack provides the four revised time-series used in Tobin's (1950) study, which are PCFC (Tobin's S), PCY (Tobin's Y), FP (Tobin's P) and NFP (Tobin's Q), respectively. Since Tobin indicates that his quantity series (food and income) are not strictly comparable to those used in budget studies, three alternative series, *AGGEXPF*, *AGGEXP* and *AGGY*, are supplied by MM. These represent aggregate consumers' expenditure on food, aggregate total consumers' expenditure and aggregate personal disposable income, respectively, all measured in current US dollars. This data has to be transformed into variables measured in real terms. Since the information pack does not supply an overall price index such as the retail price index, we use the existing indexes to create a price index which approximates the general price index. A new index log P^* is created based on Stone's (1953) weighted index:

$$\log P_t^* = (w_t^{\mathrm{f}} \times \log p_t^{\mathrm{f}}) + (w_t^{\mathrm{nf}} \times \log p_t^{\mathrm{nf}}) \tag{2}$$

where w_t^{f} and p_t^{f} are the share of food expenditure and the price index of food, respectively, and w_t^{nf} and p_t^{nf} are the share of other expenditure and the non-food price index, respectively. *AGGEXPF* and *AGGY* in log form are then deflated by log P^* creating the real variables $\log TS_t^*$ ($= \log(AGGEXPF_t) - \log P_t^*$) and $\log TY_t^*$ ($= \log(AGGY_t) - \log P_t^*$), respectively. The corresponding per capita data are calculated as $\log S_t^* = \log TS_t^* - \log(POP_t)$ and $\log Y_t^* = \log TY_t^* - \log(POP_t)$, where POP_t is the estimated total population at time t.

The integration order of the time-series is determined prior to modelling the demand for food relationships. Applying the Dickey–Fuller type tests and the Perron (1989) outlier test for unit roots, we find that $\log PCFC_t$, $\log TS_t^*$, $\log S_t^*$, $\log PCY_t$, $\log TY_t^*$, $\log Y_t^*$, $\log FP_t$ and $\log NFP_t$ are all I(1) variables. The average family size variable $\log \bar{N}_t$ is obtained from $\log \bar{N}_t = \log(POP_t) - \log(NOH_t)$, where NOH_t is the number of households at time t. This series is found to be trend stationary. Details of these test procedures are given in the full logbook.

The demand for food relationship is modelled in terms of the following (log-linear) specifications:

$$\log PCFC_t = f(\log FP_t,\ \log PCY_t,\ \log NFP_t,\ \log \bar{N}_t) \tag{3}$$

$$\log S_t^* = f(\log FP_t,\ \log Y_t^*,\ \log NFP_t,\ \log \bar{N}_t) \tag{4}$$

$$\log TS_t^* = f(\log FP_t,\ \log TY_t^*,\ \log NFP_t,\ \log \bar{N}_t) \tag{5}$$

Table III. Cointegration regression results

Variable	log $PCFC_t$ (3i)	log $PCFC_t$ (3ii)	log S_t^* (4i)	log S_t^* (4ii)	log TS_t^* (5i)	log TS_t^* (5ii)
Sample	(1913–76)		(1929–89)		(1929–89)	
Constant	4·547	3·902	−2·810	−4·199	2·667	−0·688
log FP_t	−0·091	−0·155	0·620	0·637	0·596	0·638
log PCY_t	0·131	0·195	—	—	—	—
log Y_t^*	—	—	0·530	0·685	—	—
log TY_t^*	—	—	—	—	0·640	0·755
log NFP_t	−0·408	−0·026	−0·987	−0·779	−0·923	−0·737
log $\bar{N}_t$	−0·276	—	−1·683	—	−1·774	—
R^2	0·943	0·925	0·969	0·958	0·992	0·987
CRDW[a]	0·741	0·595	0·324	0·210	0·410	0·258
S.E.	0·019	0·021	0·053	0·061	0·0470	0·057
ADF(p)[b]	−3·66(1)	−3·321(0)	−5·001(2)[c]	−3·382(2)	−5·197(2)[c]	−3·890(2)

[a] *CRDW* is the cointegration regression Durbin–Watson statistic.
[b] ADF(p) is the Dickey–Fuller test for unit roots in the cointegration residuals.
The values in parentheses are the lag lengths of the augmented terms. The critical values of ADF(p) are from MacKinnon (1991).
[c] Significance at the 5% level.

The reasons for using a log-linear model specification are discussed in Tobin (1950). The variables in equation (3) are those used by Tobin, while those in equations (4) and (5) are defined above. Per capita consumption and income data are used in equations (3) and (4), while total income and consumption data are used in equation (5). For all three equations a cointegration relationship is tested for both with and without the inclusion of the family size variable, giving six equations in total. The Engle–Granger (1987) two-step method is initially used to test for cointegration, and the results are shown in Table III. The results from Table III are mixed. Equation (3) has negative own-price elasticities, but the corresponding values from (4) and (5) are positive.[1] All the cross-price elasticities are negative, implying complementarity between the food and non-food categories. The income elasticities are all positive, and the values in equations (4) and (5) are reasonably close to the cross-section estimates given in Table II. All the family size elasticities are negative.

Interpreting the equations as long-run relationships, it seems reasonable to test for the non-existence of money illusion, i.e. to test for homogeneity of degree zero. Imposing the appropriate restrictions gives Chi-square values (with one degree of freedom) of 0·323, 0·827, 3·242, 225·549, 23·922, and 1510·230, respectively for each of the six equations. These results show that we cannot reject the assumption of no money illusion at the 5% significance level for the first three equations, but that the assumption is clearly rejected for the remaining three. The finding of homogeneity in three of the six aggregate times series is encouraging, since this finding tends to be

[1] In fact, the own and cross-price coefficient values in Tables III–V are not elasticities corresponding to those in the Tobin model, and need to be transformed according to a suggestion of Professor Wickens. This transformation delivers the expected negative own-price elasticities: details are given in our reply to the comments following this paper (see Table R1). We retain the untransformed values in our paper in order to keep within the spirit of the experiment.

Table IV. The cointegration vectors from the Johansen procedure

Variable	Cointegration vectors: log $PCFC_t$		log S_t^*	log TS_t^*
Constant	4·629	4·105	−6·469	−3·232
log FP_t	0·013	−0·035	0·563	0·778
log PCY_t	0·171	0·260	—	—
log Y_t^*	—	—	0·253	—
log TY_t^*	—	—	—	0·583
log NFP_t	−0·245	−0·302	−0·799	−0·580
log $\bar{N}_t$	−0·249	—	−1·587	1·172

Normalized coefficient values. The lag length is 2 for all cases and the sample sizes are the same as those in Table III.

the exception rather than the rule in many demand models (see, for example, Deaton and Muellbauer, 1980a, p. 78). The ADF statistics in Table III show that there is a cointegration relationship for (4i) and (5i). The ADF statistic for (3i) is only just insignificant, however. Thus (4i) is the only equation which exhibits a long-run equilibrium relationship and satisfies the homogeneity requirement. The Johansen (1991) cointegration test procedure is then applied to the above equations. The test detects one cointegration relationship for equation (3) with and without the family size variable. But for equations (4) and (5) the test shows that there is no cointegration relationship when the family size variable is excluded. The relevant estimated cointegration vectors are given in Table IV.

The Engle–Granger and Johansen approaches both detect a cointegrating relationship in (4i) and (5i), but give conflicting results for the other equations (although (3i) is close to the borderline). The corresponding error correction models are now specified in order to examine the short-run dynamics. For comparison purposes all six error correction models are estimated, and insignificant variables are dropped from the specification. The estimation results are based on the Engle–Granger two-step procedure and presented in Table V. The error correction models provide a reasonable fit to the data, although equations (4) and (5) have problems of heteroscedasticity and ARCH. Apart from (3ii) all the overall own-price elasticities are positive, a result similar to that for the long-run elasticities from the cointegration regressions in Table III. All the overall cross-price elasticities are negative, as in Table III. In equations (4) and (5) the family size variable has an overall positive effect on food demand, although in the cointegration regressions in Table III the sign of the coefficient is negative. All the overall income elasticities are positive, as in Table III. The short-run income elasticities based on equations (4) and (5) in Table V are about 0·30, although this value is only about 0·15 for equation (3). These short-run elasticities are lower than the long-run elasticities given in Table III, as economic theory suggests.

4. DEMAND FOR FOOD MODELS IN THE NETHERLANDS: THE TIME-SERIES ESTIMATES

This section estimates time-series models of demand for food in the Netherlands, and compares the results with those of the USA. The Netherlands data contains information not only on overall food expenditure and the price index of food but also on consumers' expenditure and price indexes for sub-categories of foods and other commodities. This permits the use of demand-system as well as single-equation time-series modelling. Although the focus of this paper is on the

Table V. Results of error correction models

Variable	$\Delta \log PCFC_t$		$\Delta \log S_t^*$		$\Delta \log TS_t^*$	
	(3i)	(3ii)	(4i)	(4ii)	(5i)	(5ii)
Sample	(1916–76)		(1931–89)		(1931–89)	
Constant	0·003	0·002	0·006	0·005	0·009	0·007
	(0·002)	(0·002)	(0·005)	(0·004)	(0·005)	(0·005)
$\Delta \log PCFC_{t-1}$	0·517	—	—	—	—	—
	(0·143)					
$\Delta \log PCFC_{t-2}$	0·289	—	—	—	—	—
	(0·108)					
$\Delta \log S_{t-1}^*$	—	—	0·512	0·432	—	—
			(0·117)	(0·107)		
$\Delta \log TS_{t-1}^*$	—	—	—	—	0·507	0·422
					(0·116)	(0·107)
$\Delta \log FP_t$	0·224	−0·105	0·488	0·416	0·499	0·433
	(0·061)	(0·034)	(0·089)	(0·080)	(0·085)	(0·078)
$\Delta \log FP_{t-1}$	—	0·071	−0·227	−0·163	−0·220	−0·160
		(0·045)	(0·095)	(0·077)	(0·093)	(0·077)
$\Delta \log PCY_t$	0·077	0·152	—	—	—	—
	(0·026)	(0·035)				
$\Delta \log PCY_{t-1}$	0·138	−0·067	—	—	—	—
	(0·043)	(0·036)				
$\Delta \log PCY_{t-2}$	—	0·057	—	—	—	—
		(0·028)				
$\Delta \log Y_t^*$	—	—	0·485	0·551	—	—
			(0·074)	(0·072)		
$\Delta \log Y_{t-1}^*$	—	—	−0·174	−0·147	—	—
			(0·096)	(0·095)		
$\Delta \log TY_t^*$	—	—	—	—	0·491	0·555
					(0·070)	(0·072)
$\Delta \log TY_{t-1}^*$	—	—	—	—	−0·189	−0·148
					(0·094)	(0·090)
$\Delta \log NFP_t$	−0·209	—	−0·638	−0·425	−0·650	−0·441
	(0·064)		(0·217)	(0·164)	(0·207)	(0·162)
$\Delta \log NFP_{t-1}$	−0·309	−0·255	0·251	—	0·230	—
	(0·078)	(0·076)	(0·188)		(0·179)	
$\Delta \log NFP_{t-2}$	0·339	0·142	—	—	—	—
	(0·062)	(0·058)				
$\Delta \log \bar{N}_t$	—	—	−0·447	—	−0·539	—
			(0·462)		(0·448)	
$\Delta \log \bar{N}_{t-1}$	—	—	0·807	—	0·859	—
			(0·468)		(0·449)	
EC_{t-1}	−0·614	−0·259	−0·210	−0·133	−0·259	−0·158
	(0·091)	(0·081)	(0·063)	(0·054)	(0·068)	(0·059)
R^2	0·613	0·572	0·859	0·841	0·866	0·843
S.E.	0·012	0·012	0·022	0·022	0·209	0·022
χ^2_{auto} (2)	0·880	0·371	0·539	0·581	2·713	1·159
χ^2_{ARCH} (1)	0·221	0·002	6·380[b]	7·901[c]	8·200[c]	7·534[c]
χ^2_{hetro} (d.f.)	9·50	16·21	23·90[b]	24·98[b]	30·10[b]	28·08[b]
F_{Chow}[a]	1·503	1·126	0·328	0·305	0·333	0·270

Notes:
Figures in parentheses underneath the coefficient values are standard errors.
χ^2_{auto} (2), χ^2_{ARCH} (1), χ^2_{hetro} (d.f.) and F_{Chow} are Breusch–Godfrey's serial correlation, Engle's ARCH, White's heteroscedasticity and Chow's predictive failure tests respectively.
[a] The starting date for forecasting failure is 1972 for the first two models and 1982 for the last four models.
[b] Significance at the 10% level.
[c] Significance at the 5% level.

use of single-equation demand analysis, the almost ideal demand system (AIDS) of Deaton and Muellbauer (1980b) is used in Section 4.1 to derive estimates of income, price and family size elasticities and to test the homogeneity postulate.

4.1. The AIDS Model

Developed by Deaton and Muellbauer (1980b), the AIDS model is based on the Engel function:

$$w_{it} = \alpha_i' + \beta_i' X_t \tag{6}$$

where w_{it} is the budget share of the ith commodity consumed at time t and X_t is the aggregate household expenditure at time t. Using the PIGLOG cost function, the AIDS model is derived as:

$$w_{it} = \alpha_i + \beta_i \log(E/P)_t + \Sigma_j \gamma_{ij} \log p_{jt} \tag{7}$$

where E is representative expenditure on commodity i, α_i, β_i, γ_{ij} $(i,\ j = 1, 2, \ldots, n)$ are constant parameters, and P_t is an overall price index derived from:

$$\log P_t = \alpha_0 + \Sigma_i \alpha_i \log p_{it} + 1/2(\Sigma_i \Sigma_j \gamma_{ij} \log p_{it} \log p_{jt}) \tag{8}$$

Demand theory requires that the following conditions are satisfied:

$$\Sigma_i \alpha_i = 1, \Sigma_i \beta_i = 0, \Sigma_i \gamma_{ij} = 0 \qquad \text{(adding-up condition)} \tag{9}$$

$$\Sigma_j \gamma_{ij} = 0 \qquad \text{(homogeneity condition)} \tag{10}$$

$$\gamma_{ij} = \gamma_{ji} \qquad \text{(symmetry condition)} \tag{11}$$

Following Ray (1980) equation (7) can be further developed by including a family size variable which allows for the effect of economies of household size:

$$w_{it} = \alpha_i + \beta_i \log(E/P)_t + \Sigma_j \gamma_{ij} \log p_{jt} + \theta_i \log \bar{n}_t \tag{12}$$

The Netherlands data set includes all the series necessary to estimate the demand system equations denoted by equation (12). The system equations include total household consumer expenditure (VI) and expenditure on the following commodities: Food (V11), Housing (V22), Clothing (V33), Hygiene and medical care (V44), Education (V55) and Other consumption (V66). The appropriate price indexes, P1, P11, P22, P33, P44, P55, and P66, are also provided. In estimating equation (12), we use the general price index provided by MM rather than that derived from equation (8). A time trend is also added to the specification to capture the effect of changes in consumers' tastes. Table VI presents the estimation results for the unrestricted equations.

In Table VI the Wald test implies that the homogeneity restriction is accepted for the demand equations for Food, Education, and Other commodities, but rejected for Housing, Clothing, and Hygiene and medical care. The family size variable is significant in the Food, Housing and Other consumption equations, although the sign is negative in this last case. The coefficients of own-price for Housing, Clothing, Education, and Other consumption are not significant and, apart from Hygiene and medical care, most of the cross-price coefficients are not significantly different from zero. The expenditure coefficients are all significant except for that of Clothing.

Table VI. The unrestricted parameter estimates for the Netherlands (AIDS). Estimation Sample: 1960–88

	Commodity					
Parameter	V11	V22	V33	V44	V55	V66
α_i	0·873[b]	−0·019	0·408[a]	−0·288[b]	0·100	−0·073
	(0·141)	(0·129)	(0·160)	(0·107)	(0·165)	(0·101)
β_i	−0·148[b]	−0·023[a]	−0·011	0·067[b]	0·036[b]	0·078[b]
	(0·015)	(0·014)	(0·016)	(0·011)	(0·017)	(0·011)
γ_{i1}	0·172[b]	−0·053	−0·054	−0·100[b]	0·069	−0·033
	(0·039)	(0·035)	(0·044)	(0·029)	(0·045)	(0·027)
γ_{i2}	−0·029	0·046[b]	−0·025	0·082[b]	−0·099[b]	0·026
	(0·023)	(0·021)	(0·027)	(0·018)	(0·027)	(0·017)
γ_{i3}	−0·037	0·002	0·039	−0·085[b]	0·109[b]	−0·028
	(0·031)	(0·029)	(0·035)	(0·024)	(0·037)	(0·023)
γ_{i4}	0·010	−0·004	−0·082[b]	0·102[b]	−0·042	0·018
	(0·025)	(0·023)	(0·029)	(0·019)	(0·029)	(0·018)
γ_{i5}	−0·091[b]	0·010	−0·033	0·099[b]	−0·037	0·052[b]
	(0·032)	(0·030)	(0·037)	(0·025)	(0·038)	(0·024)
γ_{i6}	−0·043	0·027	0·101[b]	−0·072[b]	0·017	0·030
	(0·031)	(0·013)	(0·034)	(0·023)	(0·035)	(0·021)
θ_i	0·220[b]	0·136[b]	0·053	−0·072	−0·126	−0·210[b]
	(0·078)	(0·071)	(0·089)	(0·059)	(0·091)	(0·056)
R^2	0·998	0·994	0·993	0·997	0·956	0·992
DW	2·287	2·102	2·901	2·434	2·190	2·149
χ^2_{Wald} (d.f.)	1·743	4·579[b]	12·13[b]	5·00[b]	1·234	0·178

Notes:
Figures in parentheses are the standard errors. In reporting the results, the estimated coefficients of the time trends are omitted.
[a] Significance at the 10% level.
[b] Significance at the 5% level.

Based on equation (12) the expenditure, own-price, cross-price, and size elasticities (ex_i, ep_{ii}, ep_{ij} and $e\bar{n}_i$) can be calculated from:

$$ex_i = 1 + \beta_i / w_i \tag{13}$$

$$ep_{ii} = -1 + \gamma_{ii} / w_i - \beta i \tag{14}$$

$$ep_{ij} = \gamma_{ij} / w_i - \beta_i w_j / w_i \tag{15}$$

$$e\bar{n}_i = \theta_i / w_i \tag{16}$$

The resulting own-price, expenditure and size elasticities are presented in Table VII.

Table VII. Unrestricted elasticity estimates for the Netherlands (AIDS)

Commodity	Own-price	Expenditure	Size
Food	−0·290	0·516	0·719
Housing	−0·769	0·896	0·615
Clothing	−0·706	0·920	0·130
Hygiene and medical care	−0·131	1·614	−0·661
Education	−1·189	1·219	−0·767
Other consumption	−0·595	2·254	−3·376

The own-price and expenditure elasticities for all commodities have the expected signs and plausible magnitudes. The results for the family size elasticity are rather interesting in that they can be given an intuitively appealing interpretation. The positive coefficients for the first three commodities, and the negative coefficients for the remaining three, are consistent with the interpretation that, in order to feed extra mouths from a limited budget, spending patterns must be changed. As family size increases for a given level of expenditure and prices, families are forced to adjust their pattern of demand towards "essential" commodities such as Food, Housing, and Clothing, and away from "less essential" commodities such as Hygiene and medical care, Education, and Other consumption.

4.2. A Single-equation Demand for Food Model

In this sub-section the demand elasticities for the Netherlands are estimated using a single equation approach. The purpose of this analysis is to compare the results with those of the US demand function. The proposed demand model is based on: $\log TS_t^* = f(\log TY_t^*, \log P11_t, \log NFP_t^*, \log \bar{N}_t)$, in which $\log TS_t^* = \log(V11_t/P1_t)$ and $\log TY_t^* = \log(AGGY_t/P1_t)$ are real total food expenditure and income, i.e. the nominal food expenditure and income variables ($AGGY$) deflated by the price index of total consumers' expenditure in their log forms. Total rather than per capita data is used since the estimation results from the Netherlands per capita data are not satisfactory. The price index of non-food items is not given in the information pack, so the non-food price index is constructed as: $\log NFP_t^* = \Sigma_i w_{it} \log PI_t$, where w_{it} is the share of the ith non-food commodity in relation to total non-food expenditure at time t ($i = 2, 3, 4, 5, 6$ representing Housing Clothing, Hygiene and medical care, Education, and Other consumption, respectively) and PI_t is the price index for the ith non-food commodity ($I = 22, 33, 44, 55$, and 66).

All variables are tested for their order of integration. The non-food price variable is a trend stationary series, real food expenditure, real income, and household size are I(1), and the food price variable is I(2). The long-run equilibrium relationship is tested using the Engle–Granger two-step cointegration test as well as the Johansen cointegration approach, and both tests support the assumption of cointegration. The Johansen approach detects one cointegration vector. The results are presented in Table VIII.

Table VIII. The Netherlands cointegration test results. Estimation Sample: 1960–88

Engle–Granger two-step cointegration estimates:

$$\begin{aligned}\log TS_t^* = {} & \underset{(0{\cdot}567)}{2{\cdot}460} + \underset{(0{\cdot}046)}{0{\cdot}792} \log TY_t^* + \underset{(0{\cdot}165)}{0{\cdot}571}\, \Delta \log P11_t \\ & - \underset{(0{\cdot}050)}{0{\cdot}329} \log NFP_t^* - \underset{(0{\cdot}244)}{0{\cdot}703} \log \bar{N}_t\end{aligned} \tag{17}$$

$R^2 = 0{\cdot}988$ S.E. $= 0{\cdot}018$ $CRDW = 1{\cdot}28$ ADF(4) $= 3{\cdot}94^{a}$ χ^2_{wald} (1) $= 57{\cdot}8$

Johansen cointegration estimates (normalized based on the coefficient of $\log TS_t^*$):

$$\begin{aligned}\log TS_t^* = {} & 1{\cdot}879 + \underset{(0{\cdot}023)}{0{\cdot}574} \log TY_t^* + \underset{(0{\cdot}182)}{0{\cdot}870}\, \Delta \log P11_t \\ & - \underset{(0{\cdot}028)}{0{\cdot}085} \log NFP_t^* - \underset{(0{\cdot}109)}{0{\cdot}149} \log \bar{N}_t\end{aligned} \tag{18}$$

Notes:
The statistics in Table VIII are the same as those in Table III. Figures in parentheses are the standard errors.
[a] Significance at the 10% level.

Comparing the Netherlands cointegration results from Table VIII with those for the equivalent US model (5i) in Tables III and IV, it is apparent that the signs on the income, cross-price, and family size variables are the same, i.e. positive, negative, and negative, respectively. The long-run income elasticity for the Netherlands is 0·792 and that of the USA is 0·755, so that the two income elasticity estimates are very similar. The coefficient on the Netherlands price variable in Table VIII is the food demand elasticity with respect to the change in the price level, and thus not directly comparable with the price coefficient for the US price variable in Tables III and IV.

Given the existence of cointegration, the corresponding error correction model is estimated and insignificant variables are deleted from the general specification. This procedure yields the following model:

$$\begin{aligned}\Delta \log TS^*_t = {} & \underset{(0\cdot008)}{0\cdot011} + \underset{(0\cdot198)}{0\cdot336}\ \Delta \log TS^*_{t-1} + \underset{(0\cdot182)}{0\cdot389}\ \Delta \log TS^*_{t-2} \\ & + \underset{(0\cdot182)}{0\cdot478}\ \Delta \log TY^*_t - \underset{(0\cdot177)}{0\cdot324}\ \Delta \log TY^*_{t-1} \\ & + \underset{(0\cdot194)}{0\cdot752}\ \Delta\Delta \log P11_t - \underset{(0\cdot252)}{0\cdot697}\ \Delta \log NFP^*_t \\ & + \underset{(0\cdot286)}{0\cdot509}\ \Delta \log NFP^*_{t-1} - \underset{(0\cdot214)}{0\cdot751}\ EC_{t-1} \end{aligned} \tag{19}$$

$$R^2 = 0\cdot768 \qquad \text{S.E.} = 0\cdot015 \qquad \chi^2_{\text{auto}}(2) = 0\cdot583 \qquad \chi^2_{\text{ARCH}}(1) = 3\cdot017$$
$$\chi^2_{\text{Hetro}}(16) = 21\cdot29 \qquad F_{\text{Chow}} = 1\cdot497$$

where the statistics are the same as those in Table IV. The break point for the Chow (1960) forecasting failure test is 1982.

Once again, the income and cross-price elasticities are positive and negative, respectively. The overall income elasticity (= 0·163) is very close to the average of the income elasticities (= 0·15) for (3i) and (3ii) in Table V. Similarly, the overall cross-price elasticity (= −0·188) is very close to the average of the cross-price elasticities (= −0·146) for the same models. The family size variable is also insignificant.

5. TIME-VARYING PARAMETER MODELS FOR THE USA AND THE NETHERLANDS

This section deals with our own chosen task (i.e. Task 5 of MM) of constructing time-varying parameter (TVP) models for the USA and the Netherlands. The rationale underlying our use of TVP modelling is that, particularly where long periods of time are involved, it seems reasonable to assume that the parameters in the food demand models may not be stable. Engle and Watson (1987) suggest that a TVP model can be applied in a number of circumstances including behavioural changes, unobserved variables and model misspecification. TVP modelling in the context of cointegration analysis is discussed in Granger (1991) and Granger and Lee (1991). Song, Liu, and Romilly (1996) provide an application of these ideas to the relationship between aggregate consumption and income in China.

Our emphasis is on the estimation of TVP models for the USA, since there are consistent annual observations available for the whole of the period 1929–89. This period covers the Second World War, during which one would expect significant changes in consumer behaviour and

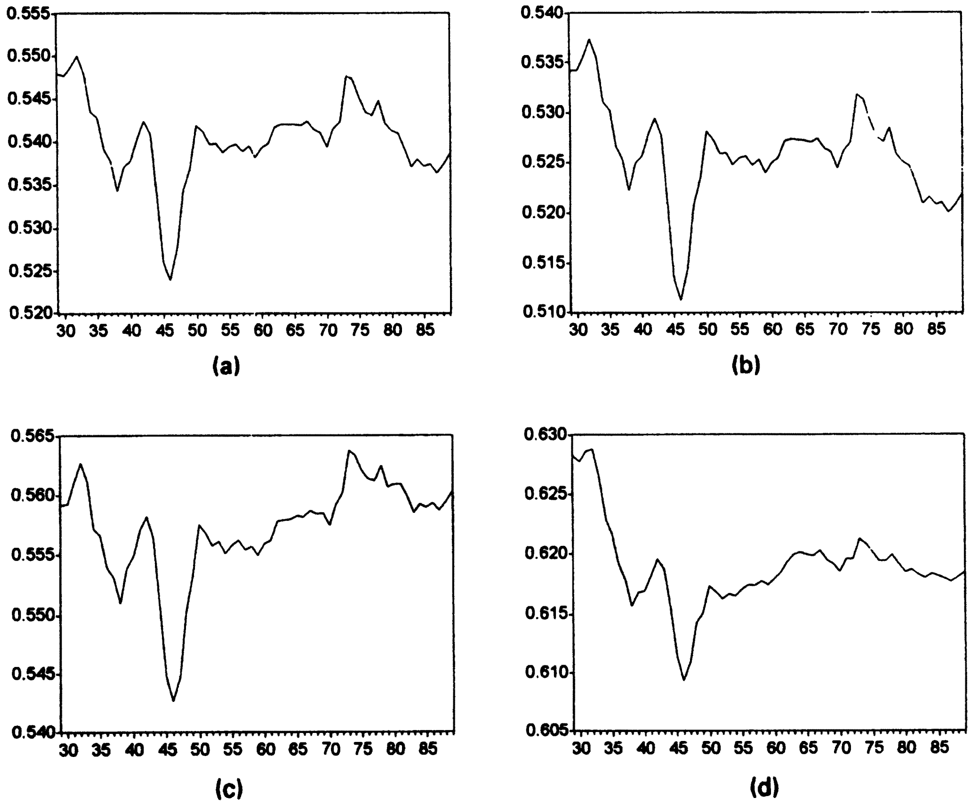

Figure 1. Kalman filter estimates of the income elasticities (USA)

consequently in parameter values such as that of the income elasticity of demand for food. The equivalent data set for the Netherlands is smaller, running from 1960 to 1988, and consequently less appropriate to the task in hand. The US model is based on equation (4i) from Table III, and uses real per capita food expenditure and income. The Netherlands model is based on equation (17) from Table VIII and uses real total food expenditure and income. Following Engle and Watson (1987), the time variation in the model parameters is specified as a random walk and the Kalman (1960) filter algorithm is used for estimation. Details of the estimation procedures and results for both countries are given in the full logbook.

Figure 1 shows the Kalman filter estimates of the time paths of the US income elasticities of food demand. Figure 1(a) is derived by allowing the income elasticity to vary over time, *ceteris paribus*. Figure 1(d) shows the time path when all parameters (including the constant) are allowed to vary. The intermediate cases are given in Figures 1(b) and 1(c).

The pattern of variation across the four panels is similar, showing a strong downward trend in the period 1929 to 1946, and another downward trend from the mid-1970s. Our tentative hypothesis is that the 1930s recession and the Second World War created a 'feel-bad' factor which increased precautionary savings from a given income and reduced spending on food, particularly on luxury items. Food rationing is also likely to have depressed food expenditures. The overall effect is to reduce the income elasticity of demand for food. After the Second World War, increasing prosperity creates a 'feel-good' factor which reduces precautionary saving and increases the income elasticity. A similar explanation would apply to the effects of the oil-price shocks of 1973/74.

Time variations in the US own-price, cross-price, and family size elasticities are smaller and less amenable to a straightforward explanation. In the case of the Netherlands results, the time variation in the income, cross-price, and family size elasticities is very small, although the time variation for the own-price elasticity is relatively large. Our regrettable ignorance of the history of the Netherlands prevents us (perhaps fortunately) from suggesting explanations for these findings. For both the USA and the Netherlands the TVP income, own-price, cross-price, and family size elasticities are always positive, positive, negative and negative, respectively. The Netherlands income elasticity is always somewhat higher than that of the USA. It is possible that alternative specifications of the way in which the parameters vary over time could produce significantly different results. This is a lengthy undertaking, however, and time constraints did not permit the exploration of alternative specifications to the random walk process used above.

6. FORECASTS FOR FOOD DEMAND IN THE NETHERLANDS

This section deals with Task 3 set by MM. Our approach is to provide forecasts for the demand for food in the Netherlands for the period 1989–2000, using both fixed and time-varying parameter models under four sets of assumptions. The forecasts for the Netherlands' real total food expenditure (in log form) are presented in Table IX. Series 1, 3, 5, and 7 are the forecasting results from the fixed parameter model (17). Series 2, 4, 6, and 8 are for the corresponding time-varying parameter model. The estimation sample is 1960–88 and the forecasting starts from 1989. Series 1 and 2 (Scenario 1) show the forecasting results assuming that income increases by an annual rate of 2%, *ceteris paribus*. Series 3 and 4 (Scenario 2) are derived on the assumption that both income and food price increase by 2% annually, *ceteris paribus*. Series 5 and 6 (Scenario 3) assumes that income and non-food price increase by 2% annually, *ceteris paribus*. Series 7 and 8 (Scenario 4) are based on the assumption that income, food and non-food prices increase by 2% annually, *ceteris paribus*. These assumptions are of course somewhat arbitrary.

Table IX shows that in Scenarios 1 and 2 (series 1, 2, 3, 4) the forecasts generated from the fixed parameter model are higher than the TVP model over the whole forecasting period. In Scenarios

Table IX. Food demand forecasts for the Netherlands 1989–2000[a]

Year	Series 1	Series 2	Series 3	Series 4	Series 5	Series 6	Series 7	Series 8
1989	5·00387	4·98775	5·07235	5·06088	4·96128	4·95602	5·02975	5·02915
1990	5·11019	5·08692	5·18004	5·16152	5·02417	5·02282	5·09402	5·09742
1991	5·21864	5·18808	5·28098	5·26414	5·08830	5·09095	5·15953	5·16702
1992	5·32925	5·29126	5·38133	5·34687	5·15375	5·16047	5·20582	5·21608
1993	5·44209	5·39651	5·55791	5·52020	5·22050	5·23137	5·33631	5·35506
1994	5·55717	5·50385	5·61170	5·56208	5·28854	5·30366	5·34306	5·36189
1995	5·67458	5·61337	5·75172	5·69574	5·35802	5·37746	5·43515	5·45983
1996	5·79431	5·72504	5·87296	5·80903	5·42882	5·45267	5·50747	5·53667
1997	5·91644	5·83895	5·99667	5·92464	5·50105	5·52941	5·58128	5·61509
1998	6·04101	5·95515	6·12285	6·04255	5·57474	5·60768	5·65658	5·69508
1999	6·16808	6·07367	6·25155	6·16282	5·64987	5·68750	5·73334	5·77664
2000	6·29770	6·19457	6·38284	6·28550	5·71887	5·76732	5·86334	5·87077

[a] Real total food expenditure (in logs).

3 and 4 (series 5, 6, 7, 8), the pattern is generally reversed, with the TVP model forecasts higher apart from 1989 and 1990 (series 5 and 6) and 1989 (series 7 and 8).

7. CONCLUSIONS

Our results indicate that the income elasticities of food demand for the USA and the Netherlands are of the same order of magnitude for both cross-section and time-series data, where the time-series elasticity is derived from long-run models. This order of magnitude is 0·5 to 0·8. The short-run time-series income elasticities are around 0·15 to 0·30. These results are reasonably encouraging, although the range of the values suggests that parameter instability may be a problem.

The results for the own-price, cross-price, and family size elasticities are less encouraging. Their values are subject to considerable variation, and the sign on the own-price food elasticity is invariably positive for the aggregate time-series models (4) and (5). This contrasts with the negative value of the own-price food elasticity derived using the Tobin specification (3). In the case of the AIDS model for the disaggregate time-series Netherlands data the sign on the own-price food elasticity is negative. The overall results from our static AIDS model seem satisfactory, although estimation of a dynamic AIDS model would have provided an interesting point of comparison.

The results of the TVP modelling of the aggregate data do not add much to these conclusions. Similar income elasticities are produced and the sign on the own-price elasticity remains positive in every time period. The forecasts for Netherlands food demand produced by the fixed and time-varying parameter specifications show some degree of variation between specifications and over time.

APPENDIX: A SUMMARY OF THE LOGBOOK

This appendix attempts to condense the full logbook of this experiment, numbering over 40 pages and 122 equations split into six chapters, into a very short summary which concentrates on the problems encountered in the modelling process.

First, we duplicated Tobin's results from the 1941 US budget survey data and tested for heteroscedasticity using White's test. The results differed according to the form of the test. In one case the null hypothesis of homoscedasticity was rejected, in the other it was accepted. This ambivalent result was a warning of further difficulties to arise when all four US budget survey data sets were considered. The (incorrect) use of Izan weights to remedy heteroscedasticity suggested that the cross-sectional income elasticities were not significantly different from each other, while the use of the $\sqrt{m}$ treatment indicated that the 1972 income elasticity was significantly different from the others. But the way in which we applied this latter treatment did not eliminate heteroscedasticity from the 1941 and 1960 budget survey data, so it was decided to rely on the results from the Izan weightings in writing Section 2 of this paper. Details of these procedures are contained in Chapters 1 and 2 of the full logbook.

We then turned to an analysis of the revised time-series data for the Tobin variables using the same model specification as Tobin. Our primary concern was to determine the integration and cointegration properties of the data. Considerable effort was expended in establishing the appropriate lag lengths for the ADF test for each variable, and the Perron test was implemented to allow for the possibility of structural breaks. The results from the cointegration tests were not clear-cut: the Engle–Granger two-step test did not find a cointegrating relationship between the variables, but the Johansen test did. We were also concerned with the possible existence of a simultaneity problem in our model specification, and used Hausman and other tests to determine the extent of this problem. As a further check, different model specifications were also examined. Once again, the results were not clear-cut, although the own-price variable in the Tobin specification did exhibit weak exogeneity and so provided a justification for the use of the single-equation approach. None of these results are reported in our paper, although details are given in the full logbook in Sections 3.1 to 3.5.

Our next step was to use similar procedures for the alternative US time-series data set provided by MM. After extensive data analysis, a cointegrating relationship was found which also satisfied the homogeneity requirement. But we were puzzled by the fact that the sign on the own-price coefficient, both short- and long-run, was positive rather than negative. This outcome also occurred when we specified and estimated a similar model for the Netherlands time-series data (this puzzle was later solved at the Tilburg workshop; see our reply to the comments following this paper). The greater disaggregation of the Netherlands data set allowed us to estimate an AIDS model in which all the income and own-price coefficients had the expected sign. This was some consolation for our disappointment at the single-equation results. Details are given in Section 3.6 and Chapter 4 of the full logbook.

At this stage we were not satisfied with the overall results of our modelling procedures. Although the estimated income elasticities were not too dissimilar given the differences in data sets and estimation procedures, the other coefficient estimates showed considerable variation and the persistence of the positive sign for the own-price coefficient was a particular concern. Perhaps our assumption of parameter stability was questionable, especially in view of the long time span covered by the US data? It was decided to use an alternative methodology, that of the time-varying parameter approach, to examine whether a more consistent set of results could be obtained. Details of the estimation procedures and results are given in Chapter 5 of the full logbook.

Finally, we attempted to forecast food demand in the Netherlands using both the fixed and time-varying parameter models developed earlier. This was not a task we approached with any degree of conviction, mainly because of our uncertainty regarding the likely future values of the exogenous variables in our forecasting models. The set of assumptions eventually used was determined more by the impending deadline (given 'holy' status by MM) rather than a detailed assessment of the Netherlands' economic prospects. Details are provided in Chapter 6 of the full logbook.

ACKNOWLEDGEMENTS

The authors would like to express their appreciation of the efforts of Jan Magnus and Mary Morgan in organising 'The Experiment in Applied Econometrics', and also for the referees' comments on this paper. Remaining errors are the responsibility of the authors. The US data are compiled from various US government publications (see organizers' data section for details). We also thank Statistics Netherlands for permission to use the Dutch data. The empirical estimates of the demand models are carried out using the EVIEWS-2.0 and FORECAST MASTER computer packages.

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COMMENTS BY PROFESSOR MICHAEL R. WICKENS (University of York)

In analysing the US data, the Dundee team have focused on the purely econometric aspects without considering the underlying economic theory. (Curiously they only introduce theoretical considerations when they analyse the Netherlands' data.) Their aim has been to see what estimates the new techniques produce, and they perform this analysis very competently. Since I have already commented in general terms on this approach to econometric modelling in my discussion of the CBS team's results, I shall not repeat those remarks here. I will simply add that, even if statistical problems are found with Tobin's estimates, and the new estimates have 'acceptable' statistical properties, it is not clear what the results mean without considering whether the models estimated, and the variable definitions used, are theoretically appropriate.

The main thrust of the Dundee team's paper is a time-series analysis of the United States' and Netherlands' data using cointegration analysis and time-varying parameter models. This follows a brief consideration of the US cross-section data based on Tobin's original model. The cross-section evidence is of a declining income elasticity over time ranging from 0·56 to 0·33 (see Table I), a significant elasticity with respect to family size which is increasing over time, and significant residual heteroscedasticity. Weighted least-squares estimation to take account of outlier residuals has the effect of raising all of the income elasticities approximately to the interval

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