DYNAMIC AND STOCHASTIC STRUCTURE: AN ANALYSIS OF THREE TIME SERIES OF HOUSEHOLD BUDGET STUDIES

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Introduction

THIS paper investigates dynamic and stochastic structure in consumer demand systems. We estimate separate demand systems for three separate time series (one British and two Japanese) of grouped household budget data.

We estimate the usual static specification in which all demand system parameters remain fixed, and two dynamic specifications in which some demand system parameters vary with time or lagged consumption. We find strong evidence that the dynamic specifications are superior to the static model for all three samples.[1]

Our stochastic specification uses an error components structure that allows correlated disturbances across households in a given year. We believe this is the first application of an error components structure in demand system estimation. Unlike the usual stochastic specification, it is consistent with the possibility that, for example, in a particularly cold year most households consume more fuel that usual. We find the time-specific effect very significant in one sample, moderately significant in another, and not significant in the third.

The paper is organized as follows: Section I describes our demand model. Section II describes our data and discusses the estimation problems arising because our data are not for individual households but are grouped by income class. Section III presents our results. Section IV summarizes our findings and discusses the possibility of interpreting the superiority of dynamic over static models as evidence of changing tastes.

I. Model Specification

Model specification involves three major components: (a) functional forms for the demand equations, (b) dynamic structure, and (c) stochastic structure.

A. *Functional Form*

The demand system we estimate is a quadratic expenditure system (QES), a functional form in which the demand equations are quadratic in expenditure. We regard demand theory as a model of expenditure allocation among an exhaustive set of consumption categories, and hence distinguish between "expenditure" (μ) and "income" (i.e., receipts, as reported in our data). In share form, the QES demand equations are given by

$$\begin{aligned} w_i = {} & p_i b_i/\mu + a_i\left[1 - \left(\sum p_k b_k\right)/\mu\right] \\ & + (c_i - a_i)\lambda \prod (p_k/\mu)^{-c_k} \\ & \times\left[1 - \left(\sum p_k b_k\right)/\mu\right]^2, \end{aligned} \tag{1}$$

where $\Sigma a_k = 1$, $\Sigma c_k = 1$, the a's, b's, c's and λ are parameters, w_i is the share of the i^{th} good in total expenditure (μ) on the n included categories, and p_i the price of the i^{th} good; the QES contains $3n - 1$ independent parameters.[2] This demand system was introduced and estimated in Pollak and Wales (1978).[3] The familiar linear expenditure system (LES) is clearly a special case of the QES.

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[1] With the exception of Tsujimura and Sato (1964), ours is the first study that examines dynamic structure using household budget data. Tsujimura and Sato use Japanese household budget data for the period 1951–60 to estimate a dynamic version of the linear expenditure system: they first estimate Engel curves separately for each year, and then use the coefficients to estimate the complete system for the entire period. They conclude that the static model should be replaced by a Duesenberry-type model of habit formation.

[2] $\Sigma(\Pi)$ without indexes or limits of summation means $\Sigma_{k=1}^{n}(\Pi_{k=1}^{n})$.

[3] Howe, Pollak and Wales (1979) give a closed form characterization of the entire class of theoretically plausible demand systems quadratic in expenditure, and estimate an alternative QES form using per capita time series data from the U.S. national product accounts.

B. Dynamic Structure

To estimate a demand system, one must either assume that its parameters remain constant over time or specify how they change. In addition to the usual static specification in which all demand system parameters remain fixed, we estimate two dynamic specifications: (i) a linear time trend specification in which some demand system parameters (the b's) vary linearly with time,

$$b_{it} = b_i^* + \beta_i t \tag{2}$$

and (ii) a lagged consumption specification in which some demand system parameters (again the b's) vary linearly with past consumption

$$b_{it} = b_i^* + \beta_i z_{it-1} \tag{3}$$

where z_{it-1} is a variable representing past consumption. In section IV we discuss the possibility (and the difficulties) of interpreting these dynamic specifications as reflecting taste change. If we accept a taste change interpretation of the lagged consumption specification, its precise meaning depends in part on the variable chosen to represent past consumption. For example, if z_{it-1} represents the household's own consumption of the i^{th} good in period $t-1$, then the model can be interpreted as habit formation; if z_{it-1} represents other households' consumption of the i^{th} good in period $t-1$, then the specification can be interpreted as interdependent preferences (see Pollak (1976)). Both (2) and (3) contain the static model as a special case in which the β's are zero, and both add n independent parameters to the static model.

C. Stochastic Structure

The usual "independent" stochastic specification adds disturbance terms to the share equations and assumes these disturbances are independent across households and over time. In addition to the independent specification we consider an error components structure in which the disturbance is the sum of two terms, one independent across households and over time, and the other a "time specific effect" (TSE) which is the same for all households in a particular year.[4] The independent specification then is a special case of error components structure in which the TSE is absent. The TSE allows a positive correlation between the disturbances of different households in a particular year, so, unlike the independent specification, the error components structure is consistent with the possibility that, e.g., in a particularly cold year most households consume more fuel than usual.

To formalize the error components model, we let u_{rt} denote the $(n-1)\times 1$ vector of disturbances added to the demand system (1), where r denotes the household and t the time period.[5] These disturbances are the sum of two components:

$$u_{rt} = e_t + \epsilon_{rt} \qquad \begin{array}{l} r = 1,\ldots,q_t \\ t = 1,\ldots,T \end{array} \tag{4}$$

where q_t is the number of households in period t and T the number of time periods. We do not assume that household r in period t is the same as household r in period τ, nor even that there are the same number of households in every period. We do assume that the $(n-1)\times 1$ disturbance vectors e_t and ϵ_{rt} are independently normally distributed with zero means and covariance matrices given by Γ and Δ, respectively, where Γ is positive semi-definite and Δ positive definite. The covariance matrix for the u's is thus given by

$$E(u_{rt}u'_{s\tau}) = \begin{bmatrix} \Gamma + \Delta & r = s, & t = \tau \\ \Gamma & r \neq s, & t = \tau \\ 0 & & t \neq \tau. \end{bmatrix} \tag{5}$$

Under these assumptions the log of the likelihood is (aside from an additive constant):

$$L = -\frac{1}{2}\sum_{t=1}^{T}\Bigg\{\log|W_t| + (q_t - 1)\log|\Delta| + \sum_{s=1}^{q_t} u'_{st}\Delta^{-1}u_{st} + \sum_{s=1}^{q_t}\sum_{r=1}^{q_t} u'_{st}\frac{(W_t^{-1} - \Delta^{-1})}{q_t}u_{rt}\Bigg\} \tag{6}$$

where $W_t = \Delta + q_t\Gamma$.[6] Since Γ and Δ are $(n-1)$

[4] Since we do not have data on the same households in successive years, we do not discuss three-component structures involving household-specific disturbance terms.

[5] Since our system of equations involves shares which sum to one for each observation, the disturbances must sum to zero. Hence we may drop one equation from the analysis, and thus are left with an $(n-1)\times 1$ vector of disturbances.

[6] We are indebted to J. Magnus for providing this particularly simple expression for the likelihood function. For a brief outline of its derivation see Magnus (1982, p. 248). It would be convenient for estimation if this expression could be concentrated with respect to the Γ and Δ parameters. Although this has been done for the case when the q_t are equal for all t, as in Magnus (1982), it does not appear to be possible when the q_t differ.

$\times (n-1)$ symmetric matrices, this stochastic specification contains six independent parameters.[7]

II. Data and Estimation

This section describes briefly the British and Japanese samples and discusses the problems of using grouped rather than individual household data in demand analysis.[8]

The British sample involves household budget data for the seven years 1966–72. The data are reported in the *Family Expenditure Survey* (*FES*) series, an annual publication which reports income, expenditure patterns, and some demographic characteristics of a sample of U.K. households. The *FES* cross-classifies households by income and certain demographic characteristics, and reports mean expenditure on various consumption categories for each income-demographic cell. We consider only households consisting of one man, one woman, and two children. The survey reports a total of 37 cells: 4 income classes for 1966; 5 for 1967, 1968 and 1969; and 6 for 1970, 1971 and 1972. To simplify the computations we analyze only three consumption categories: "food," "clothing," and "miscellaneous." Current expenditures on these categories were obtained directly from the surveys, and the corresponding price indexes from *National Income and Expenditure* (1975, table XV, p. 400).

We analyze two separate Japanese samples, one for the years 1953–1964 and the other for 1963–1977. In the earlier sample the data are for households from urban areas only, while in the later sample they are for households from all Japan. In both samples the data report income and expenditure patterns for "typical workers' households" (i.e., wage and clerical worker households consisting of a married couple with two children in which the household head is the only wage earner). The data appear in the *Annual Report on the Family Income and Expenditure Survey* (*FIES*). In each year of each sample the survey groups households by annual income and reports between 15 and 21 income classes. We use reported mean expenditure on three broad consumption categories that correspond closely to those used with the British sample. This yields 215 cells for the earlier sample and 231 cells for the later sample. Price indexes are obtained from the *Annual Report on the C.P.I.*, and aggregated to correspond to our three consumption categories using weights provided in the reports.

Three problems arise because the data reported in the survey are mean expenditure patterns for income-demographic cells rather than expenditure patterns of individual households.

(a) The first is an aggregation problem. If individual household demand equations are quadratic in expenditure, then the mean demand equations for a particular income-demographic cell depend on the variance of expenditure within that cell. Unfortunately, the data do not report these variances, and we "solve" the aggregation problem by treating the reported cell means as if they represent expenditure patterns of individual households.[9] As the number of cells reported in a survey becomes larger, the variance within each cell is likely to be smaller and hence the error resulting from omitting the variance from the grouped demand equations is less serious. In this regard the Japanese samples are better than the British since they contain a much larger number of income classes.

(b) The second is a "weighting" problem which arises because cells contain different numbers of households and thus different amounts of information. To account for these differences we use an estimation procedure which weights each cell by the square root of the number of families it contains. Thus, it is actually the weighted e_t and ϵ_{rt} disturbances which we require to have constant covariance matrices Γ and Δ, and the weighted u_{rt} to have the covariance structure given in (5).[10]

(c) The third problem is choosing an appropriate variable to represent past consumption. We use a weighted average of the consumption of

[7] We impose the restriction that Δ be positive definite by writing it as the product of a lower triangular matrix (ΔL) and its transpose, and estimate the elements of ΔL rather than Δ. The same procedure is followed for Γ, thus giving elements of ΓL. We restrict the diagonal elements of ΔL and of ΓL to be positive to guarantee uniqueness of the factorization.

[8] For a more detailed discussion of the data see Darrough, Pollak and Wales (1981).

[9] We tried to estimate these variances, but since none of our attempts yielded plausible parameter values we do not report these results.

[10] The parameter estimates obtained using unweighted observations are very similar to the weighted ones for the British sample, but differ considerably for the two Japanese samples. This is not surprising since cell sizes vary much less in the British sample than in the two Japanese samples, as discussed in Darrough, Pollak and Wales (1981).

all households in the sample in the previous year, weighting by cell size.[11]

III. Results

This section discusses our results and their implications. For each of our three samples we estimated the QES for six models: the static and two dynamic specifications (time trend and lagged consumption) in conjunction with two stochastic specifications (the usual independent specification and an error components structure allowing a TSE).[12] We first discuss model selection and then the impact of model specification on estimated marginal budget shares and price elasticities.

Dynamic specifications are significant in all three samples. For the two Japanese samples standard likelihood ratio tests show that both dynamic specifications are superior to the static model regardless of the stochastic structure: in all cases the addition of dynamic variation is significant at the 1% level in the Japanese samples. The lagged consumption and time trend specifications are not nested, so formal classical tests are not available; however, since they contain the same number of parameters, a comparison of likelihood values is of interest.[13] On this basis the time trend model is superior: in the Japanese samples it yields a higher likelihood value than lagged consumption in all but one of the four cases, namely, the earlier sample with a TSE. For the British sample dynamic specifications are significant at the 1% level with no TSE and at the 5% level with the TSE. This holds for both the lagged consumption and time trend specifications, and these two dynamic specifications yield virtually identical likelihood values with or without the TSE.

The habit formation or interdependent preferences interpretation of the lagged consumption model requires positive β's.[14] In the earlier Japanese sample β's are positive for all three goods; in the other two samples β's are positive for food and miscellaneous, but negative for clothing. It is tempting to interpret negative β's as reflecting partial stock adjustment for consumer durables rather than habit formation (see Houthakker and Haldi (1960)), although the absence of interest rates suggests that the interpretation is rather loose. The consumer durable interpretation would be more convincing if the β's for clothing were negative in all three samples.

The significance of the TSE depends on the sample considered and the dynamic specification imposed. In the static model the TSE is highly significant in both Japanese samples and marginally significant in the British sample; for both dynamic specifications it is highly significant in the earlier Japanese sample, but insignificant in the later Japanese and British samples.[15] Since the evidence for the dynamic specifications is strong with both stochastic specifications, we consider only the dynamic specifications in evaluating stochastic structure. When we do so, the performance of the TSE is mixed. However, an investigator who confined his attention to the static model would find evidence of a TSE in all three samples.

All six estimated demand models for the British and earlier Japanese samples correspond to "well-behaved" preferences: the implied Slutsky matrix

[11] Although aggregation arguments suggest that higher moments of lagged consumption may also play a role, we ignore them. If one adopts a taste change interpretation of the lagged consumption specification, interdependent preferences imply that the appropriate variable is the average lagged consumption of households assumed to influence those in the cell, while habit formation implies that the appropriate variable is the average lagged consumption of the households in the cell. Since the only data we have are those reported in the previous year's sample, and since the samples are not panels, these are different households. Using average consumption in the previous year's sample is compatible with an interdependent preference interpretation in which tastes depend on the average past consumption of all households with the same demographic profile, regardless of their income or expenditure. The habit formation interpretation is tenuous.

An alternative but substantially more complicated approach would relate the demand system parameters of households in each percentile of the income or expenditure distribution to the consumption of those occupying that position in the previous year.

[12] Log likelihood values for all models and parameter estimates for some are presented in Darrough, Pollak and Wales (1981). We also estimated six LES specifications, but since the QES was significantly superior at the 1% level for all models in all samples, we report only QES results.

[13] New techniques have recently been proposed for testing non-nested hypotheses (e.g., Pesaran and Deaton (1978), Davidson and MacKinnon (1981)), but their application is beyond the scope of this paper.

[14] Interdependent preferences with "snob appeal" are consistent with negative β's.

[15] For the earlier Japanese sample, the TSE is significant at the 1% level for all three dynamic specifications. For the later Japanese sample, it is significant at the 1% level with the static specification; at about the 15% level with lagged consumption; and at about the 50% level with the time trend. For the British sample, the TSE is significant at the 10% level with the static specification but it is not significant with the dynamic specification.

TABLE 1.—MARGINAL BUDGET SHARES AND OWN PRICE ELASTICITIES
LAGGED CONSUMPTION SPECIFICATION, 1953–1964 JAPANESE SAMPLE, 1960 PRICES

Total Expenditure (yen/month)	Marginal Budget Shares			Own Price Elasticities					
	TSE			no TSE			TSE		
	Food	Clothing	Misc.	Food	Clothing	Misc.	Food	Clothing	Misc.
1233	.38	.17	.45	−.61	−0.85	−0.96	−0.74	−1.09	−1.20
1946	.32	.19	.49	−.68	−0.95	−1.04	−0.87	−1.11	−1.22
2834	.24	.21	.54	−.73	−1.01	−1.11	−1.03	−1.15	−1.28
3761	.16	.24	.60	−.77	−1.04	−1.16	−1.21	−1.18	−1.34
3906	.15	.24	.61	−.77	−1.09	−1.23	−1.24	−1.19	−1.35

Notes: All values are calculated at 1960 prices for every third expenditure class reported in the 1953–1964 Japanese sample. Total expenditure refers to the sum of expenditures on food, clothing and miscellaneous.

is negative semi-definite at all sample price-expenditure situations for both of these samples. For the later Japanese sample, the implied Slutsky matrix is negative semi-definite at about 75% of the sample points for each of the six models.

We now examine the impact of model choice on estimated marginal budget shares and price elasticities. We emphasize the results for the earlier Japanese and British samples, confining our discussion of the later Japanese sample to footnotes.

Marginal budget shares are virtually identical for all six models in the earlier Japanese sample; the first three columns of table 1 present estimated marginal budget shares for one model evaluated at a price situation near the sample midpoint. The marginal budget shares vary dramatically with changes in expenditure, a result consistent with our findings that the QES is a significant improvement over the LES, a functional form which implies constant marginal budget shares. The marginal budget share for food drops sharply as expenditure rises, that for miscellaneous rises sharply, while that for clothing increases gradually.[16]

Although marginal budget shares for a particular sample are very similar for all six models, price elasticities are not. For the earlier Japanese sample, the stochastic specification makes a considerable difference, while the static and the two dynamic specifications yield similar results. The second three columns of table 1 report the elasticities corresponding to one of the specifications without a TSE, while the final three columns report those corresponding to one with a TSE. Comparing the two sets of elasticity estimates, we see that those implied by the models with a TSE are uniformly higher (in absolute value) than those corresponding to the usual independent stochastic specification.

For the British sample estimated price elasticities depend on the dynamic specification and, to a lesser extent, on the stochastic specification.[17] Table 2 presents own price elasticities for the static and the two dynamic specifications with no TSE. The lagged consumption specification generally yields higher elasticities (in absolute value) than the static or time trend specification. With all three specifications, the price elasticity for miscellaneous increases with expenditure, although in the static model the increase is small. In both dynamic models, the food elasticity increases with expenditure; in the static model, it decreases with expenditure, although the decrease is small. The clothing elasticities exhibit a more complex pattern: those in the static model decrease with expenditure and exhibit more variation with expenditure than those in the two dynamic specifications. In the lagged consumption model, clothing elasticities also decrease with expenditure, while in the time trend model they increase with expenditure.

The signs of the correlations between the disturbances for the three goods have been calculated

[16] In the other two samples, estimated marginal budget shares are also virtually independent of model specification. For the later Japanese sample marginal budget shares behave in roughly the same way as in the earlier one. For the British sample the marginal budget share for food drops sharply as expenditure rises (from 0.36 in the lowest expenditure cell to 0.20 in the highest) while that for miscellaneous rises sharply (from 0.36 to 0.63); unlike the Japanese, the British marginal budget share for clothing decreases with expenditure (from 0.28 to 0.17).

[17] In the British sample the six models exhibit four distinct elasticity patterns: the time trend models with and without the TSE are virtually identical, as are the lagged consumption models with and without the TSE.

TABLE 2.—OWN PRICE ELASTICITIES
1970 BRITISH SAMPLE

Expenditure (S./week)	Static, no TSE			Lagged Consumption, no TSE			Time Trend, no TSE		
	Food	Clothing	Misc.	Food	Clothing	Misc.	Food	Clothing	Misc.
230.7	−.97	−2.51	−1.59	−0.97	−2.21	−2.80	−.67	−1.77	−1.75
259.0	−.96	−2.23	−1.61	−1.01	−2.08	−2.94	−.70	−1.74	−1.94
280.2	−.96	−2.09	−1.62	−1.05	−2.02	−3.03	−.73	−1.74	−2.07
320.5	−.95	−1.91	−1.64	−1.11	−1.97	−3.18	−.78	−1.77	−2.29
359.3	−.94	−1.79	−1.65	−1.17	−1.95	−3.30	−.82	−1.82	−2.48
476.2	−.91	−1.59	−1.67	−1.37	−2.02	−3.58	−.95	−2.08	−2.91

for the earlier Japanese sample and the British sample based on the lagged consumption specifications. For the correlations derived from our estimates of Δ, the results for the two samples are very similar. All the correlation coefficients are negative, implying that an unusually high expenditure on any one good tends to be associated with unusually low expenditure on the other two.[18] Turning to the correlation coefficients for the TSE (derived from estimates of Γ), the signs are negative for the earlier Japanese sample, except for the correlation between food and clothing.[19]

To compare the magnitudes of the two error components we have calculated the ratio of the variance of the TSE to the total variance of the disturbance for each good (i.e., the appropriate elements of $\Gamma/(\Gamma + \Delta)$). For the earlier Japanese sample these ratios are 0.34, 0.10 and 0.33 for the food, clothing and miscellaneous categories;[20] thus, at least for food and miscellaneous, a substantial part of the variance of the disturbance term is attributable to an effect that is not independent across households, but instead affects all households in a given time period.

IV. Conclusion

In addition to confirming that the QES dominates the LES as a functional form for analyzing household budget data, analysis of our samples suggests three important points. (i) There is some evidence of a time-specific effect. The assumption that disturbances are independent across households or income classes should be imposed with caution. (ii) There is evidence that stochastic and dynamic structures act as substitutes: for example, time-specific effects are more likely to be statistically significant in static than in dynamic specifications.[21] Thus, these structures should not be evaluated separately. (iii) There is strong evidence that dynamic specifications are superior to static models. Although this finding can be interpreted in several ways, it clearly indicates a problem with static models. We now turn to interpretation.

In empirical demand analysis the static model is one in which all demand system parameters remain fixed so demand for market goods is a stable function of goods' prices and expenditure. It is tempting to interpret the statistical superiority of the dynamic specifications (i.e., our finding that allowing some demand system parameters to vary with time or lagged consumption significantly increases the likelihood value) as evidence of "taste change." Four comments are in order.

First, we have not shown that our data are inconsistent with constant tastes. Our results depend on the maintained hypothesis that the demand system is a QES. Suppose the data were in fact generated by some other demand system (e.g., a translog) and that an investigator using the true system could explain all observed variations in consumption patterns as responses to variations in prices and expenditure. Depending on the functional form of the true demand system and the trends in prices and expenditure, we might find the dynamic QES significantly superior to the static QES. The literature contains two distinct ap-

[18] Because of the budget constraint at least one of the three correlation coefficients between the three pairs of goods must be negative.

[19] The same sign pattern holds for the later Japanese sample.

[20] The corresponding figures for the later Japanese samples are 0.02, 0.08 and 0.06.

[21] Our conclusion that the stochastic and dynamic specifications are interrelated is consistent with a similar finding in Howe, Pollak and Wales (1979) using per capita time series data from the U.S. national product accounts.

proaches to taste change which, in this situation, yield different results. The empirical demand analysis literature adopts a specific functional form as a maintained hypothesis. A finding that some demand system parameters depend on time or lagged consumption is then interpreted as evidence of taste change; Pollak (1978) surveys this literature. The revealed preference literature does not impose a specific functional form as a maintained hypothesis: instead, it investigates whether there exist well-behaved preference orderings consistent with the data. The answer depends on whether the data satisfy the strong axiom of revealed preference. Systematic procedures for determining whether a finite set of price-quantity data can be rationalized by well-behaved preferences, and, if they can be, for constructing such preferences, have been investigated by several authors; for a recent treatment and references to the literature, see Varian (1982).[22]

Second, if omitted variables that affect consumption patterns are correlated with lagged consumption or time, then the lagged consumption or time trend specification may be statistically significant even though tastes are constant. For example, with non-separable preferences hours of leisure and stocks of consumer durables affect demand for non-durable goods and services. Similarly, the dynamic variables may capture the effects of omitted demographic variables. The omitted variable view thus interprets the statistical superiority of dynamic specifications over static models as evidence that the static models are misspecified, not as evidence of taste change.

Third, even if taste change is really occurring, neither the time trend nor the lagged consumption specification casts much light on its origin: time trends are inherently uninformative, and lagged consumption is broadly consistent with both interdependent preferences and habit formation. Grouped household budget data are poorly suited to illuminating the underlying taste change mechanism.[23]

Finally, "constant tastes," instead of referring to preferences for "market goods," may refer to preferences for "basic commodities" such as in Becker's household production model.[24] Although the stability of preferences for commodities becomes an interesting issue when commodities are defined in an empirically meaningful way, so long as economists derive demand functions for goods from preferences for goods and estimate them using goods prices, discovering the dynamics of these preferences is crucial for empirical demand analysis. Although constant tastes may be a satisfactory assumption at some level of philosophical abstraction, the static model is not adequate for empirical demand analysis.

[22] The revealed preference tradition is deterministic rather than stochastic: a single "small" violation of the strong axiom is as devastating as many "large" ones.

[23] Per capita time series data from national product accounts are also poorly suited. Individual household budget data would be better than grouped data, and panel data would be best.

[24] See Becker (1965). For a discussion of the distinction between taste change and technical progress in the household production model and the difficulties which arise when commodities are unobservable, see Pollak and Wachter (1975).

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