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Agricultural Economics 24 (2001) 289–295

AGRICULTURAL ECONOMICS

www.elsevier.com/locate/agecon

Identification of habit in Japanese food consumption

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Received 18 January 1998; received in revised form 28 February 2000; accepted 10 April 2000

Abstract

Demand models were estimated for five commodities using annual retail level Japanese data. Estimates were made with the Houthakker–Taylor [Houthakker, H.S., Taylor, L.D., 1970. Consumer Demand in the United States: Analysis and Projections. Harvard University Press, New York, MA] state adjustment model. Significant lags were found for meat and cereal, which means that habit is important for these commodities. Some habit affects were also found for seafood.

Keywords: Habit; Dynamic models; Japanese data

1. Introduction

Economists have long considered the time needed to adjust to income and price changes to be an important aspect of food demand. Consumers need more than a single time period to make a full adjustment to changes in prices and incomes. Demand curves become more elastic in the long run than in the short run. These effects have been specified with varying models which include lags.

Any group selling food products to the Japanese market has an interest in the length of time it takes for the Japanese consumer to react to changes in prices and income. Common knowledge suggests that many Japanese consumers are traditional and would not be expected to adjust rapidly to changing market conditions. Groups selling to the Japanese market need to design promotional programs and marketing strategies that take length of adjustment into account.

The purpose of this paper is to estimate the effect of habit by using retail level demand functions for five major food groups. These groups represent the major food groups for at home food consumption in Japan. The theory of consumer demand relates to the retail level. This data set allows demand estimates to be made without the confounding effects of marketing margins encountered with farm or wholesale level demand.

2. Review of literature

There are two ways of accounting for the effect of time in demand analysis. The demand curves can be given different slopes in the long run by using a longer time period as observations, or a shift variable composed of past prices can be included to change the intercept of the short run curve. This manuscript will consider only the latter.

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0169-5150/01/$ – see front matter
PII: S0169-5150(00)00078-5

Pollak (1970) lists three reasons for the difference in long and short run demand curves: (1) the consumer may have contractual commitments; (2) the consumer may be ignorant of consumption possibilities consistent with his/her tastes outside the range of past consumption experiences; (3) goods may be habit forming which implies that past consumption influences current consumption.

Houthakker and Taylor (1970) justify the inclusion of lagged variables in the demand for food because of habit. They emphasize the effect of stocks on the demand for durables, and the effect of habit on non-durables. Phlips (1974) agrees that in the case of non-durables–durables, habit is the dominant effect. Deaton and Muellbauer (1980) also state that habits which develop from past behavior are analogous to stocks of durables which are built up from past purchases.

Manderscheid (1964) argues that the length of run affects demand elasticities. He divides length of run into three periods. In the first, the consumer's response to a price decline may be to purchase heavily in order to build up stocks in addition to increased consumption. In the second, the consumer is less responsive to price changes because a consumer is hesitant to change consumption patterns. In the long run, the consumer may decide to reorganize their consumption patterns.

There have been a number of empirical studies which explore the dynamics of consumer demand for food (Table 1). One of the earliest was by Working (1954). This study was conducted prior to the work of Houthakker and Taylor (1970). Controversy arose in regard to the original work by Working (1954) as O'Regan (1955) and Kuznets (1953) questioned the estimated signs on the coefficients of the lagged variables in the case of the price dependent demand function.

More recent studies have used either the Houthakker and Taylor (1970) state adjustment model or the dynamic AIDS model (Table 1). While the AIDS model is appealing for static analysis, the dynamics are not as well understood or specified as they are with the state adjustment model. In addition, the complexity of the AIDS model makes interpretation more difficult. The empirical findings of these past studies show that the state adjustment model is preferable to using either a static model or a model with a simple lag structure.

The unit of time has varied from annual to weekly data. Generally, the effect of the lagged variables has been larger for the studies using the shorter time periods. As the time period becomes longer, the effect of habit increases while the inventory effect decreases.

The commodities studied have either been a comprehensive group of food and non-food commodities or meat, including various types and cuts. There have been no studies on the groups of food that were used in this study.

3. Conceptual framework

The basic Houthakker–Taylor (Houthakker and Taylor, 1970) state adjustment model can be used to specify the effect of habit for non-durable commodities. The basic hypothesis is that consumption is a function of prices, income and stocks.

$$q_t = \alpha + \beta s_t + \gamma e_t + \eta p_t + u_t \tag{1}$$

where s_t are the stocks at time t and e_t is the income at time t.

Since with non-durables, stocks represent habit, they cannot be measured. Houthakker and Taylor (1970) use the concept of stock depreciation to eliminate stocks from the estimating equation. Specifically, they assume that the rate of stock depreciation is constant and proportional to the amount of stocks at time t, i.e.

$$w_t = \delta s_t \tag{2}$$

where w_t is the 'average using up' of stocks at time t.

From these basic assumptions Houthakker and Taylor (1970) derive the estimating equation as

$$\begin{aligned} q_t = {} & A_0 + A_1 q_{t-1} + A_2 \Delta e_t + A_3 e_{t-1} \\ & + A_4 \Delta p_t + A_5 p_{t-1} + v_t \end{aligned} \tag{3}$$

Estimates for α, β, γ, η, and δ are then derived algebraically from the estimates of A_i.

It should be noted that if δ=2, lagged prices and lagged income are eliminated. The equation is then usually referred to as a Koyck lag.

Over the 29-year time period used in this analysis, Japanese consumers have seen changes in their

Table 1
Review of previous empirical studies

Author	Model	Data	Results
(Working, 1954)	(Price dependent; sum of lagged quantities)	(US annual (meat))	(Significant lagged quantities)
Sexauer (1977)	H–T state adjustment	US annual, biannual, quarterly, monthly (16 commodities)	Effect of inventories declined as time period increased
Wohlgenant and Hahn (1982)	H–T state adjustment	US monthly (beef, pork, chicken)	Inventory important for pork, habit important for chicken, beef inconclusive
Capps and Nayga (1990)	H–T state adjustment	US weekly, biweekly, and monthly scanner (six beef cuts)	Effect of inventories declined as time period increased. Effect of habit increased
Anderson and Blundell (1983)	Static, partial adjustment and autoregressive	Canadian time series (six commodities)	Static, partial adj. models, simple auto rejected in favor of general dynamic
Heien and Durham (1991)	Dynamic quadratic expenditure system	US time series and cross-section (16 commodities)	Habit significant in both cross-section time series and habit stronger in time series
Pope et al. (1980)	H–T state adjustment	US (beef, pork, poultry, fish)	Static and partial adj. rejected in favor of H–T state adjustment
Blanciforti and Green (1983)	Dynamic AIDS	US (11 commodities)	Habit significant for 6 of 11 commodities
Chen and Veeman (1991)	Dynamic AIDS	Canadian quarterly (four meats)	Habit significant for all four commodities
Haden (1990)	Dynamic AIDS	Japanese annual (three types of cigarettes)	Habit significant

culture caused by increased exposure to Western culture and technology. Such changes would be expected to influence food patterns and preferences. Therefore, time was included in the Koyck lag models to measure the effects of changes in tastes and preferences over this period. The lag structure of the state adjustment models should be sufficient to capture the effects of time.

4. Empirical results

The model was estimated using a Japanese data set which included the 29-year period 1963–1991.[2] Prices, quantities and expenditures are at the retail level. Quantities and expenditures were placed on a per capita basis to account for the decrease in household size which took place over this time period. Prices and expenditures were deflated by the consumer price index for all food.

Five commodities were chosen for the analysis; meat, seafood, cereals, vegetables, and fruits. Expenditures on these commodities constitute 56% of Japanese food expenditures. They include only food purchased for at home use.

A simplified version of the state adjustment model was first estimated using seemingly unrelated regression (SUR). This model included lagged values of quantity, but no lagged values of prices or income. The purpose of estimating this model was to delineate important substitutes and complements to use in the state adjustment model. The state adjustment model is subject to problems of multicollinearity making it difficult to identify substitutes and complements. This model included the variable time as well as all complements and substitutes. Symmetry was imposed on all substitute–complement pairs. Pairs were eliminated if the pair had no t-values greater than 1. This left only the complementary relationship between vegetables and three of the other food groups; meat, seafood, and cereal (Table 2). There were no substitutes or complements between fruit and the other food groups.

The quadratic form of the time variable was used to account for possible non-linearities in this variable. The square of time was eliminated if the t-value was less than 1. This resulted in non-linear forms of the time variable for meat and seafood, but linear forms for the other variables.

The lagged value of the dependent variable was included for all food groups, but was eliminated if the t-value was less than 1. The coefficient on lagged quantity was significant for meat and for cereals with t-values of 5.09 and 2.27, respectively (Table 2). t-values were less than 1 for the other three food groups.

One problem with using data sets of this length is the high probability of structural change. Sequential Chow tests were run using OLS with no corrections for serial correlation. There were no significant F-values for seafood and meat at the 0.05 level. There was only one significant value for vegetables at the 0.05 level. However, the test for fruit and cereal yielded several significant values. The highest F-values occurred when the sample was split between 1975 and 1976 for cereal and between 1974 and 1975 for fruit. Examination of the results of splitting the sample between 1975 and 1976 showed the largest change to occur in the constant terms. Therefore, a dummy variable was added to the equations for cereal and fruit which was given the value of 1 for 1963–1975 and 0 otherwise. This addition eliminated the significant structural breaks for these two products.

A matrix of long-run elasticities was computed for the SUR estimates at the mean values of the variables (Table 2). The results generally met with prior expectations. The direct price elasticities ranged from a high of −0.698 for meat to a low of −0.521 for cereal. Total expenditure elasticities ranged from a high of 2.070 for fruit to low of −0.799 for cereal. The cross elasticities showed complementarily between vegetables and all other products but fruit, which is a reasonable result. The lack of important substitution among these products is to be expected since these are aggregate groups.

The state adjustment model was estimated for the five food groups with the same variables that were included in the SUR model except the variable time. Time was eliminated because the state adjustment model already has a built in time adjustment. Following the recommendation of Houthakker and Taylor (1970), the models were estimated with the restriction that the estimated value of δ be equal among own

[2] Statistics Bureau, Annual Report on the Family Income and Expenditure Survey. Prime Minister's Office, Japan (various years).

Table 2
Demand coefficients for the five commodity SUR model[a]

Variable	Seafood	Meat	Cereal	Vegetables	Fruit	Total expenditure	Time
q_{t-1}		0.3986 (5.09)	0.34795 (2.27)				
p_t							
Seafood	−28907 (−11.12)			−2853 (−2.28)			
Meat		−8239 (−6.54)		−2640 (−3.72)			
Cereal			10102 (−5.61)	−2607 (−2.48)			
Vegetables	(−1.91)	−2291 (−2.99)	−2067 (−0.32)	−326 (−10.93)	−16473		
Fruit					−7099 (−11.21)		
e_t	0.04440 (9.79)	0.01956 (7.35)	−0.02320 (−4.26)	0.05422 (13.97)	0.03527 (17.09)		
Time	20.836 (0.11)	174.44 (1.55)	79.39 (1.86)	−720.4 (−13.00)	−512.5 (−15.73)		
Time squared	−13.191 (−2.09)	−9.49 (2.06)					
Dummy			1045 (5.34)		−954.1 (−5.52)		
R^2	0.935	0.998	0.995	0.933	0.976		
DWD	1.797	2.688	2.100	1.665	1.947		
Durbin h		−2.029	−0.55				
Matrix of short run demand elasticities							
Seafood	−0.768			−0.066		0.895	−0.182
Meat		−0.420		−0.093		0.625	−0.091
Cereal		−0.340		−0.010		−0.521	0.041
Veges	−0.090	−0.101	−0.094	−0.557		1.304	−0.400
Fruit					−0.686	2.070	−0.695
Matrix of long run demand elasticities							
Seafood	−0.768			−0.066		0.895	−0.182
Meat		−0.698		−0.155		1.040	−0.151
Cereal			−0.521	−0.015		−0.799	0.063
Veges	−0.090	−0.101	−0.094	−0.557		1.304	−0.400
Fruit					−0.686	2.070	−0.695

[a] Symmetry restrictions imposed (t-values in parentheses).

price, total expenditures, and the price of substitutes and/or complements, i.e.

$$\delta = \left[\frac{A_3}{A_2 - 0.5A_3}\right] = \left[\frac{A_5}{A_4 - 0.5A_5}\right] = \cdots = \left[\frac{A_{k+1}}{A_k - 0.5A_{k+1}}\right] \quad (4)$$

The F-value on the usual test for the compatibility between the data and the above restrictions were all non-significant at the 0.05 level (Table 3).

The linear form of the time variable was added to the model for meat and fruit because infeasible results were obtained without this variable (δ=0). In addition, the dummy variable was dropped from the equation for fruit.

The model results were generally acceptable. R^2 values were all over 0.89 (Table 3). The coefficients on own price change were all negative and highly significant with t-values ranging from −4.48 to −10.66. The coefficients on total expenditure change were also highly significant with t-values ranging from −2.77 to 7.79. The signs were as expected with all products but cereal having positive signs.

Serial correlation was significant for only meat (Table 3). The iterative Cochrane–Orcutt technique was used to obtain estimates for meat.

The estimates for the state adjustment model show a substantial habit effect for meat and cereal (Table 3). There is also evidence of a small habit effect for seafood. The estimates for fruits and vegetables show no habit effect. The values of δ and β are plausible for meat, cereal and seafood.

A matrix of long-run elasticities was computed for the state adjustment estimates at the mean values of the variables (Table 3). The results generally met

Table 3
Demand coefficients for H–T state adjustment model[a]

Variable	Seafood	Meat	Cereal	Vegetables	Fruit	Total expenditure
q_{t-1}	0.95186 (13.04)	0.81922 (14.98)	0.65065 (4.21)	−0.05580 (−0.07)	0.58037 (12.08)	
p_t-p_{t-1} (Seafood)	−32247 (−9.23)			5116.7 (3.17)		
p_{t-1} (Seafood)	−5201.1 (−9.23)			4919.9 (3.17)		
p_t-p_{t-1} (Meat)		−8216.6 (−5.67)		11440 (8.27)		
p_{t-1} (Meat)		−4979.7 (−3.86)		11000 (8.27)		
p_t-p_{t-1} (Cereal)			−7567.5 (−4.48)	−16741 (−6.34)		
p_{t-1} (Cereal)			−6468.0 (−4.48)	−16097 (−6.34)		
p_t-p_{t-1} (Vegetables)	−5383.0 (−3.09)	−2914.4 (−3.58)	1411.2 (1.39)	−17594 (−10.66)		
p_{t-1} (Vegetables)	−868.22 (−3.09)	−1766.3 (−1.86)	1206.2 (1.39)	−16917 (−10.66)		
p_t-p_{t-1} (Fruit)					−7450.0 (−7.27)	
p_{t-1} (Fruit)					−2887.6 (−7.27)	
e_t-e_{t-1}	0.016038 (2.77)	0.024896 (5.85)	−0.012981 (−2.77)	0.016913 (7.79)	0.031108 (5.49)	
e_{t-1}	0.0025868 (2.77)	0.015088 (4.86)	−0.011094 (−2.77)	0.016262 (7.79)	0.012058 (5.49)	
Dummy			920.59 (4.22)			
Time		235.63 (5.16)			−195.49 (−6.33)	
Adj. R^2	0.8943	0.9974	0.9938	0.9609	0.9665	
Durbin h	−1.726	[b]	−0.873	0.298	−0.434	
Test on restr. (F)[c]	0.145	0.281[d]	0.375	0.321	1.368	
Rate (habit use (δ))	0.175	0.870	1.493	1.852	0.481	
Habit (β)	0.126	0.671	1.069	−0.385	−0.050	
LR. Mult.	3.557	4.375	3.526	0.828	0.905	
Matrix of short run demand elasticities						
Seafood	−0.807			−0.145		0.307
Meat		−0.326		−0.102		0.602
Cereal			−0.175	−0.031		−0.201
Vegetables	0.176	0.482	−0.663	−0.654		0.447
Fruit					−0.734	1.863
Matrix of long run demand elasticities						
Seafood	−2.872			−0.5151		0.092
Meat		−1.427		−0.447		2.713
Cereal			−0.618	−0.108		−0.708
Vegetables	0.176	−0.482	−0.663	−0.654		0.447
Fruit					−0.734	1.863

[a] t-values in parentheses.
[b] The model was corrected for serial correlation.
[c] This tests the compatibility between the data and the restriction given in Eq. (4).
[d] This is the value for the model not corrected for serial correlation.

with prior expectations. The direct price elasticities ranged from a high of −2.872 for seafood to a low of −0.618 for cereal. Total expenditure elasticities ranged from a high of 2.713 for meat to a low of −0.708 for cereal. These long run elasticities are generally higher than those estimated with the SUR model (Table 2). Symmetry was not imposed on the state adjustment model. The cross elasticities showed vegetables to be a complement to meat and seafood which is a reasonable result. However, meat and seafood were substitutes for vegetables. Thus, the state adjustment model estimates of the effects of complements and substitutes do not necessarily agree with theory.

5. Summary and conclusions

The state adjustment model shows habit to be important for meat and cereal. There was also an indication of a small habit effect for seafood. Habit was strongest for cereal. This food group is dominated by rice. Rice takes on an almost religious significance by the Japanese.

An important question is: why is habit important for some commodities, but not for others? We have argued that the special view given to rice by the Japanese could result in a strong habit effect. However, this does not explain the strong habit effect for meat. One should note that the habit effect has been found by more than one author for US meat demand (with annual data by Working (1954), and Pope et al. (1980)). Many US consumers base the major evening meal on the meat dish. With this role given to meat, adjustments to prices would be expected to take time. Vegetables, and starches are considered to be a complement to the major dish, and the quantities could be more readily changed in response to income and price changes. This may also hold for the Japanese consumer. Based on this rational, seafood should also show a strong habit effect. In the Japanese diet, the group 'seafood' consists of numerous seafood species, whereas meat consists of mainly beef, pork and poultry. Thus, in response to a price increase for seafood, there is more opportunity to substitute cheaper species than there is for meat. Additional evidence for this argument stems from the lack of a habit effect for fruits and vegetables which consist of numerous alternatives. In addition, the cereal group is limited in type because of the dominance of rice and noodles in the Japanese diet.

The state adjustment model is theoretically appealing. It provides estimates of both the effect of habit and the rate of use of habit. However, with the added variables of lagged prices and total expenditure, it is prone to severe multicollinearity.

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