ELSEVIER

Agricultural Economics 28 (2003) 139–150

AGRICULTURAL ECONOMICS

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Expenditure on different categories of meat in Greece: the influence of changing tastes

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Received 5 March 2001; received in revised form 25 January 2002; accepted 22 July 2002

Abstract

This paper employs a latent variable approach to isolate the effects of changing tastes on the share of total meat expenditure on different categories of meat products in Greece during the period 1965–1995. We find that changes in the relative expenditure on different categories of meat cannot be explained by changes in the relative prices of the different meat products and increased expenditure alone. For pork products in particular, the increase in the share of expenditure has been greater than would be expected as a result of the relative fall in their price. The increase can therefore be associated with changes in taste. This finding is of general interest to those conducting empirical research into consumer behaviour both in economies where there have been significant changes in patterns of food consumption, and where, as in the case of many less industrialised economies, rapid structural changes in food consumption patterns are still to come. It is also of importance to policy makers in assessing the effectiveness of advertising or promotional campaigns in influencing longer term changes in consumer preferences for different products.

JEL classification: C32; D12

Keywords: Latent variable; AIDS; Tastes; Meat; Greece

1. Introduction

The application of demand analysis to deriving price and income/expenditure elasticities associated with food products has generated a wide-ranging literature. However, the isolation of factors other than price and expenditure on levels of demand is less widespread. In this paper, we apply a modified almost ideal demand system (AIDS) in which the latent variable component is analysed in order to determine how these other factors, characterised here as 'tastes' have influenced the demand for different categories of meat products in Greece between 1965 and 1995.

Leybourne (1993) argued that AIDS models with random walk coefficients had advantages over the conventional estimation of AIDS, particularly where these models could not be characterised as cointegrating relationships. However, the lack of cointegration does not necessitate the treatment of all coefficients in the model as random walks. Such an approach requires fairly strict assumptions about, inter alia, the exogeneity of prices and expenditures, and complicates

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0169-5150/02/$ – see front matter
PII: S0169-5150(02)00103-2

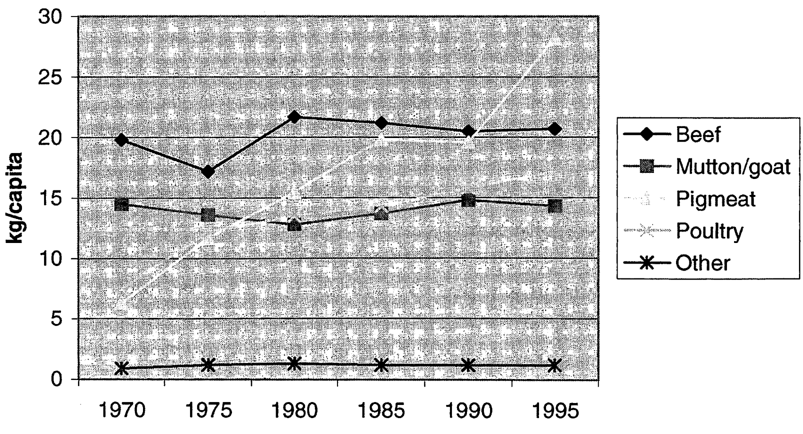
the subsequent interpretation of the estimates. The approach taken in this paper allows only for random walk coefficients in each of the intercepts. This greatly simplifies both estimation and interpretation, and is fully consistent with a non-cointegrated system.

In Section 2 we briefly review the pattern of meat consumption in Greece over the post-World War II period. Section 3 introduces the model, and Section 4 discusses our approach to estimation and inference, paying particular attention to the analysis of the latent variable component. Section 5 describes the data used in the analysis. In Section 6 the results of both the estimated system and the latent variable component are presented and discussed in the context of observed influences on meat consumption in Greece. Concluding remarks are made in Section 7.

2. Food consumption in Greece

During the post-war period, significant changes have occurred in Greek consumption patterns. Up until the early 1970s, total consumption expenditure increased rapidly (by almost four times in constant 1970 prices) due to increases in real income and price stability. Even though expenditure on all commodity categories increased in real terms, the rate of increase was not uniform, resulting in significant changes in the distribution of total expenditures among commodity categories. Although post-war food expenditure increased, it decreased as a share of total consumption expenditure. However, this decrease was associated with an internal redistribution in favour of meats and fish and against bread and cereals and oils and fats (Demoussis, 1985; Andrikopoulos et al., 1987; Mergos and Donatos, 1989; Klonaris, 1999). Two of the main causes of these changes have been contended to be the initiation of the custom connection between Greece and the European Economic Community in 1962, and the rural exodus associated with increased levels of urbanisation which took place in the 1960s.

Nevertheless, the increase in meat consumption was not uniform for all meat items. Between 1965 and 1995, the budget shares of beef, and of lamb and mutton, have remained more or less unchanged. However, the shares of pork and chicken have exhibited a more complicated pattern over time, whilst the expenditure share for frozen meats decreased by almost three times. Although it is not possible to calculate trends in physical per capita consumption from the expenditure data used in this study, data from FAOSTAT (2002) for similar categories of meat products demonstrate that beef, sheep and goat meat consumption during the period 1970–1995 remained relatively static. By contrast, the per capita consumption of pig meat increased significantly from about 6 kg per capita in 1970 to 27 kg per capita in 1995. Similar growth is observed in the consumption of chicken in the 1970s, but the growth is slower for the remainder of the period (Fig. 1).



Source: FAOSTAT (2002)

Fig. 1. Per capita consumption of main categories of meat 1970–1995.

A recent paper by Karagiannis et al. (2000), which employs an error correction model within an AIDS specification, discusses the influence of changing habits on meat consumption behaviour in Greece. The authors suggest that changing habits have exerted an upward influence on the consumption of beef, lamb and mutton, and pork, but have not influenced the demand for chicken over the period. In discussing the results of the estimation reported in this paper, particular attention is paid to comparing them with those given in Karagiannis et al. (2000).

3. Model specification and estimation

The almost ideal demand system of Deaton and Muellbauer (1980) remains one of the most popular specifications of demand systems in the applied literature. The linearised AIDS can be expressed as:

$$s_t = \alpha + \theta p_t + \beta q_t, \quad \text{where} \quad q_t = \ln(E_t) - p_t' s_t \quad (1)$$

where s_t is a $(m+1) \times 1$ vector of expenditure shares, p_t a $(m+1) \times 1$ vector of logged prices and E_t the total expenditure. Denoting a $1 \times (m+1)$ vector of ones as $j'_{m+1} = (1, 1, \ldots, 1)$ the restrictions required by theory are:

$$j'_{m+1}\alpha = 1, \qquad j'_{m+1}\beta = 0 \quad (2)$$

$$j'_{m+1}\theta = 0, \qquad \theta' = \theta \quad (3)$$

The first set of restrictions (2) are imposed automatically since one of the equations is redundant and has errors which are the linear sum of the others. The remaining two (3) are homogeneity and symmetry restrictions, respectively.

In order to ascertain the influence of tastes on the expenditure shares of the different categories of meat, we employ an AIDS specification in which *each* cost share equation is appended with a stochastic error term such that:

$$s_{it} = \sum_j \theta_{ij} p_{jt} + \beta_i q_t + \mu_{it} + e_{it}, \quad \text{where}$$
$$\mu_{it} = \lambda_i + \mu_{it-1} + v_{it} \quad (4)$$

where s_{it} are the expenditure shares on the ith commodity, p_{it} are the logged prices of those commodities, and q_t is the level of expenditure. This is the standard type of share equation with the exception that there is now an unobserved random walk component included. Separating out the drift from the random walk gives:

$$s_{it} = \sum_j \theta_{ij} p_{jt} + \beta_{i1} q_t + \lambda_{it} + \tau_{it} + e_{it}, \quad \text{where}$$
$$\tau_{it} = \tau_{it-1} + v_{it} \quad (5)$$

where τ_{it} is the pure random walk component. We need only assume that v_{it} and e_{it} are stationary, but it involves no great loss in generality to assume that v_{it} is serially uncorrelated. Note that under $\mathrm{Var}(v_{it}) = 0$ then $\tau_{it} = \tau$ (a constant intercept), and (5) is a cointegrated system if the variables are non-stationary, with changes in expenditure determined in part by changing tastes if $\lambda_i \neq 0$. Therefore, this general framework allows for both cointegrated and non-cointegrated models. The taste effect could be measured by estimating the model and constructing:

$$\hat{\mu}_{it} = \hat{\lambda}_i + \hat{\mu}_{it-1} + \hat{v}_{it} \quad (6)$$

which is the path of changes in tastes for the ith equation.

The approach set out above is related to those used in studies by Moschini and Meilke (1989), Burton and Young (1992), Eales and Unnevehr (1993), and Mangen and Burrell (2001), which employ a stochastic set of share equations to allow for structural change in the system parameters by using time trends. Moschini and Meilke (1989) suggest that structural change is likely to affect all equations simultaneously and therefore assert that a common time path for all parameters can be assumed. The direction of any bias in structural change is determined by calculating the difference in the share of a good in expenditure before and after the structural change. Burton and Young (1992) criticise this approach as only allowing for changes in taste in one direction. They allow the taste changes to take a quadratic form, thus allowing for changes in the direction of taste formation. The approach used in the current paper also allows for periods when the influence of changing tastes is moving in favour of a given commodity, and periods when it is moving against and is, we contend, more sensitive to these changes than the approach set out in Burton and Young (1992) as the model specification is less restrictive.

4. Estimation and inference

There are several avenues that could be pursued in estimating the system. The model is simply a special case of a random walk parameter model in which all parameters, except the intercept, have been constrained to have a zero variance. The estimation of these models can be approached via maximum likelihood using the Kalman filter (or related filters) (e.g. Leybourne, 1993) or via spectral approaches (see Harvey, 1989). Alternatively a regression based approach (see Maddala and Kim, 1998, p. 470) can be used. The Kalman filter has been the most popular method for the estimation of these models. The regression based method is less computationally expensive and is analytically simpler. This also facilitates the use of other procedures such as bootstrapping which would be difficult with the Kalman filter since using maximum likelihood procedures requires the use of iterations.

Inference in the regression approach should not differ substantially from a case where the parameters of the model are estimated using maximum likelihood. However, the test of the zero variance of the random walk component (which is of particular interest) has a null on the edge of the parameter space. If the random walk component is a latent regressor in a regression then, in the case where all variables are $I(1)$,[1] H_0: $\sigma_v^2 = 0$ (the variance of the changes in tastes in the random walk component) can also be viewed as the cointegration hypothesis. Since H_0 is on the edge of the parameter space (since σ_v^2 must be non-negative) the regularity conditions required to justify inference within a maximum likelihood framework break down (as noted in Harvey, 1989, condition (iv), p. 210). The tests outlined by McCabe and Leybourne (Maddala and Kim, 1998, p. 206) provide an alternative route for testing this hypothesis, however, critical values for these tests in a multivariate framework are unavailable. Recent work on the bootstrapping of non-stationary systems (Li and Maddala, 1997) seems applicable to the problem at hand and we employ these techniques to determine the critical value of the McCabe–Leybourne test for no cointegration. Leybourne (1993) also performs a test for fixed coefficients, however, this differs from the one presented in this paper.

[1] The data has been tested for unit roots and was found to be broadly consistent with being $I(1)$. Results of these tests are available from the authors on request.

5. Data

The data consists of an annual time series (1965–1995) of consumption expenditure on seven categories of meat: beef; lamb and mutton; pork; chicken; frozen meats; sausages, bacon and ham; and all 'other meats', collected from the national accounts published by the National Statistical Service of Greece. Since the data is aggregated over households, prices have been obtained from the implicit price indices formed by dividing current expenditures by real expenditures (Paasche indices are used in the allocation). The unit endogenous price was used in order to be consistent with Deaton and Muellbauer (1980).

Fig. 2 depicts the evolution of shares in expenditure (in current prices) on the different categories of meat. It can be observed that the share of expenditure on both beef and pork has increased over the period, the share of the later increasing significantly until the mid 1980s since when it has declined somewhat. The share of expenditure on lamb and mutton, the other significant category, remained relatively stable throughout the period but declined marginally in the 1990s. Of the other categories, the share of expenditure on chicken and frozen products has declined and the share of expenditure on sausages, bacon and ham has risen.

Fig. 3 shows how the relative prices of the different meats have varied over the period in comparison with each other and with the general consumer price index. Of particular interest is the growing divergence between the relative prices of beef, pork, chicken and sausages, bacon and ham which fall relative to the CPI and lamb and mutton, frozen products and other meats which increase more rapidly than the general CPI, during the 1970s and to a lesser degree the 1980s. From the late 1980s to the 1990s, the divergence in the relative prices of all categories, with the exception of pork and chicken had reduced as the price of all meats falls relative to the CPI.

6. Results

In this paper, the primary interest is in the structure of the latent variable component and the influence of

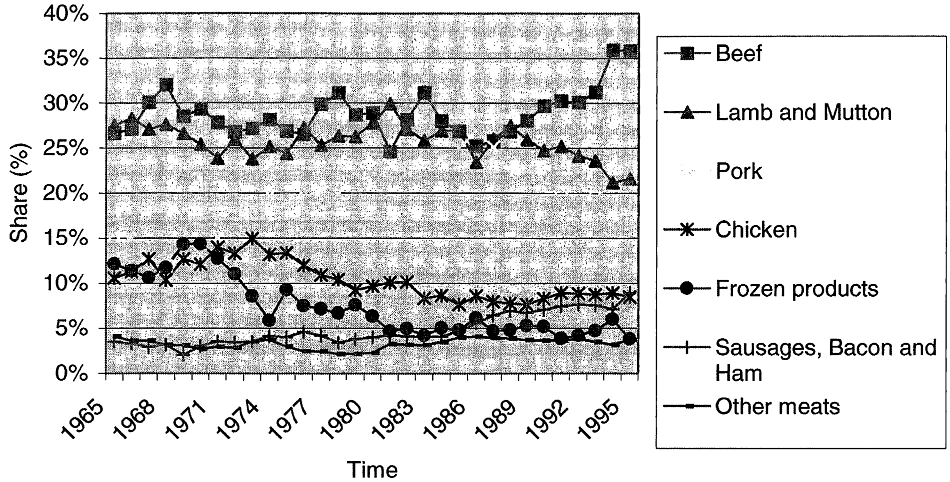


Fig. 2. Evolution of shares of expenditures (in current prices) on Greek meat products.

changing tastes. However, we first report the results of the estimated system and elasticities. The estimated values of the regression coefficients are presented in Table 1. The 'other meats' group was omitted in the estimation. Since the system is fully restricted, the results are invariant to the choice of omitted variable. The estimates of the own price coefficients for all commodity groups except pork and chicken are significantly different from zero. With the exception of those associated with pork and chicken and other meat, the majority of the cross price parameter estimates are also significantly different from zero. The parameter estimates for the effect of changes in real expenditure on the expenditure shares are significantly different from zero for all groups with the exception of pork and frozen meats. Of particular note in the

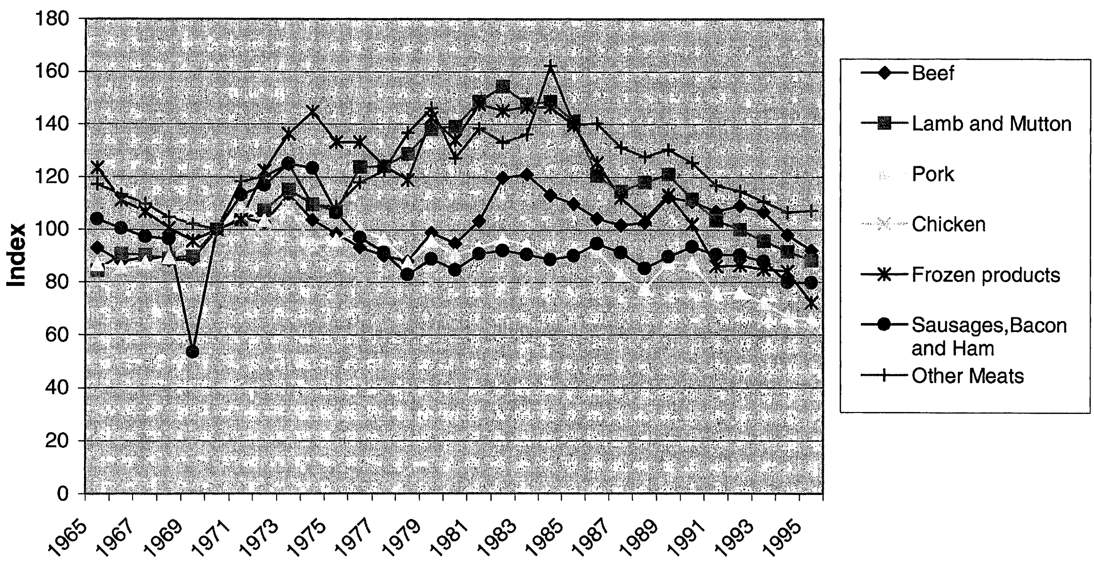


Fig. 3. Price indices of selected meats relative to the general price index.

Table 1
Regression coefficients and standard errors

	Beef	Lamb and mutton	Pork	Chicken	Frozen products	Sausages, bacon and ham
Beef price	0.1166[a] (0.0418)	−0.0546[a] (0.0188)	−0.0938[a] (0.0269)	0.0740[a] (0.0213)	−0.0287 (0.0199)	−0.0164 (0.0093)
Lamb price	−0.05463[a] (0.0188)	0.2175[a] (0.0185)	−0.06310[a] (0.0165)	−0.04905[a] (0.0130)	−0.03655[a] (0.0128)	−0.001664 (0.0063)
Pork price	−0.09385[a] (0.0269)	−0.06310[a] (0.0165)	0.04335 (0.0283)	−0.02697 (0.0172)	0.1294[a] (0.0162)	0.01453[a] (0.0073)
Chicken price	0.07402[a] (0.0213)	−0.04905[a] (0.0130)	−0.02697 (0.0172)	0.01162 (0.0195)	−0.01056 (0.0134)	−0.001424 (0.0062)
Frozen price	−0.02873 (0.0199)	−0.03655[a] (0.0128)	0.1294[a] (0.0162)	−0.01056 (0.0134)	−0.04888[a] (0.0182)	−0.02019[a] (0.0063)
Sausage price	−0.01643 (0.0093)	−0.001664 (0.0063)	0.01453[a] (0.0073)	−0.001424 (0.0062)	−0.02019[a] (0.0063)	0.02512[a] (0.0038)
Other meat price	0.00304 (0.0107)	−0.01251 (0.0087)	−0.003412 (0.0093)	0.002358 (0.0081)	0.01547[a] (0.0069)	5.746E−005 (0.0033)
Expenditure	0.2715[a] (0.0416)	−0.1993[a] (0.0222)	−0.04771 (0.0327)	0.05218[a] (0.027)	−0.01208 (0.0283)	−0.02879[a] (0.0117)
Time trend	−0.006957[a] (0.0021)	0.002826[a] (0.0010)	0.004783[a] (0.0016)	−0.00297[a] (0.0013)	−0.001898 (0.0013)	0.002784[a] (0.00057)

The values in parentheses are standard errors.
[a] 95% level of confidence.

Table 2
Tests for symmetry and homogeneity

Wald for symmetry and homogeneity	30.84
Prob (Wald)	0.07635

context of this paper is that the estimates of the trend parameter are significantly different from zero in all but the case of frozen meats.

Table 2 presents the results of the joint test for symmetry and homogeneity. The Wald test statistic suggests that the null hypothesis is not rejected at the 95% level of confidence.

In Tables 3 and 4 we present the Marshallian and Hicksian own and cross price elasticities of demand, respectively, for the seven commodity groups. In both the Marshallian and Hicksian cases, the own price elasticities, with the exception of lamb and mutton, which is insignificantly different from zero, are negative.[2] The uncompensated own price elasticity values for beef, pork, chicken and sausages, bacon and ham fall within the range −0.45 to −0.93, while those for frozen meat and for 'other meats' exceed −1. A similar pattern is seen in the compensated elasticities. The elasticities differ from those in Karagiannis et al. (2000) in several respects. A key difference discussed below is that the own price elasticity for lamb reported in Table 3 is insignificantly different from zero, whilst in Karagiannis et al. (2000) the short run elasticity is reported as being −0.46 (although no indication of significance is given). Similarly, the own price elasticity of beef, pork and sausages, bacon and ham are of a greater magnitude in Karagiannis et al. (2000) than in the current paper. However, they are more in line with those reported in Eales and Unnevehr (1993), with the exception of pork, which is found to be a luxury in that study, perhaps reflecting different consumption patterns between Greece and the US.[3]

In investigating the cross price elasticities, a number of observations warrant discussion. All cross price elasticities with respect to the price of lamb reported in Table 4 are insignificantly different from zero. It is also observed that both frozen meats and sausages, bacon and ham are substitutes for pork (the former a strong substitute). This may be explained by the fact that the category pork represents fresh meat portions and that there is little consumer resistance to substituting frozen cuts if the price of fresh pork increases. Consumers will also substitute fresh pork for frozen meats if the price of the latter increases.

The expenditure elasticities recorded in Table 5 are all positive, with the exception of 'other meat'. For beef and chicken, the expenditure elasticity is greater than one, a result which supports that reported in Karagiannis et al. (2000). Again, the estimate for lamb and mutton, and sausages, bacon and ham are of a lesser magnitude. To a certain extent the discrepancies can be explained by differences in the number of commodity groups used (seven in this paper and five in Karagiannis et al. (2000)), and in the model specification. However, the contrast in the estimates relating to lamb and mutton are more striking. We contend that the small, and insignificant elasticities[4] are consistent with the fact that lamb is a staple good in Greece and changes in its price are not expected to have significant impacts on its consumption relative to that of other meats. By contrast, in relating the expenditure and price elasticities, it can be observed that whilst beef is highly expenditure elastic, it is price inelastic and thus changes in the consumption of beef can be explained largely by changes in the level of income.

Having briefly discussed the results of the estimated system itself, we now examine in more detail the latent variable taste components and discuss the implications of their structure. Table 6 presents the results of the joint test for the removal of the estimated random walk variables from our system of estimated equations. In Table 6 the McCabe–Leybourne statistic is reported. This test rejects the null hypothesis that the random walk components provide no additional information to the estimated system at a high level of confidence, since the null of cointegration is rejected at the 95% level of significance. The critical value is approximate and has been produced by using a stationary bootstrap using 1000 trials (these procedures are rather lengthy

[2] All of the uncompensated own price elasticities with the exception of lamb and mutton are significantly different from zero. In the case of the compensated own price elasticities both lamb and mutton and beef are not significantly different from zero.

[3] It is acknowledged, following Eales and Unnevehr (1993), that the magnitude of the elasticities may differ (in general be of a greater magnitude) if supply side instruments are incorporated. Their incorporation is, however, beyond the scope of this paper.

[4] The authors appreciate the assistance from an anonymous referee in drawing their attention to this interesting point.

Table 3
Uncompensated elasticities

	Beef	Lamb and mutton	Pork	Chicken	Frozen products	Sausages, bacon and ham	Other meat
Beef price	−0.8670[a] (0.2748)	−0.4323[a] (0.1538)	−0.5099[a] (0.2060)	0.1597 (0.09530)	−0.1695[a] (0.07657)	−0.1024[a] (0.04814)	−0.02044 (0.06572)
Lamb price	0.01092 (0.1358)	0.04318 (0.1628)	−0.09349 (0.1345)	−0.1106 (0.08734)	−0.08447 (0.07256)	0.03083 (0.03744)	−0.02308 (0.06040)
Pork price	−0.4094 (0.2775)	−0.2596 (0.2064)	−0.7308[a] (0.2545)	−0.1127 (0.09774)	0.6796[a] (0.09168)	0.08604[a] (0.04293)	−0.009416 (0.07218)
Chicken price	0.5722[a] (0.2718)	−0.6063[a] (0.2378)	−0.3607 (0.2105)	−0.9394[a] (0.2226)	−0.1400 (0.1738)	−0.03822 (0.1122)	0.006221 (0.1993)
Frozen price	−0.3404 (0.2761)	−0.4509 (0.2483)	1.777[a] (0.2689)	−0.1256 (0.2467)	−1.647[a] (0.2981)	−0.2645[a] (0.1168)	0.2140 (0.1248)
Sausage price	−0.1686 (0.2429)	0.1194 (0.2495)	0.4182[a] (0.1980)	0.03201 (0.2312)	−0.3745[a] (0.1729)	−0.4503[a] (0.1326)	0.02084 (0.1034)
Other meat price	0.4055 (0.6116)	−0.1001 (0.5760)	0.1088 (0.4884)	0.1836 (0.6366)	0.5507 (0.2897)	0.05412 (0.1556)	−1.116[a] (0.5103)

The values in parentheses are standard errors.
[a] 95% level of confidence.

Table 4
Compensated elasticities

	Beef	Lamb and mutton	Pork	Chicken	Frozen products	Sausages, bacon and ham	Other meat
Beef price	−0.3073 (0.2967)	0.06820 (0.1208)	−0.1300 (0.1876)	0.3599[a] (0.09019)	−0.02551 (0.06961)	−0.008781 (0.04481)	0.04346 (0.06737)
Lamb price	0.07627 (0.1351)	0.1016 (0.1554)	−0.04913 (0.1327)	−0.08723 (0.08708)	−0.06766 (0.07184)	0.04176 (0.03708)	−0.01562 (0.06025)
Pork price	−0.1914 (0.2763)	−0.06471 (0.1748)	−0.5828[a] (0.2435)	−0.03475 (0.1043)	0.7357[a] (0.08676)	0.1225[a] (0.04240)	0.01547 (0.07339)
Chicken price	1.006[a] (0.2522)	−0.2181 (0.2177)	−0.06597 (0.1979)	−0.7842[a] (0.2320)	−0.02828 (0.1681)	0.03441 (0.1079)	0.05578 (0.2010)
Frozen price	−0.09916 (0.2705)	−0.2351 (0.2497)	1.941[a] (0.2289)	−0.03931 (0.2336)	−1.585[a] (0.2837)	−0.2241 (0.1194)	0.2415 (0.1293)
Sausage price	−0.05248 (0.2678)	0.2232 (0.1982)	0.4971[a] (0.1721)	0.07355 (0.2307)	−0.3447[a] (0.1836)	−0.4309[a] (0.1270)	0.03410 (0.1026)
Other meat price	0.3806 (0.5900)	−0.1224 (0.4719)	0.09197 (0.4364)	0.1747 (0.6295)	0.5443 (0.2915)	0.04997 (0.1503)	−1.119 (0.5210)

The values in parentheses are standard errors.
[a] 95% level of confidence.

Table 5
Expenditure elasticities

Beef	1.942[a] (0.2265)
Lamb and mutton	0.2267 (0.1236)
Pork	0.7562[a] (0.2676)
Chicken	1.506[a] (0.2857)
Frozen products	0.8372 (0.4855)
Sausages, bacon and ham	0.4030 (0.2831)
Other meat	−0.08626 (0.5289)

The values in parentheses are standard errors.
[a] 95% level of confidence.

Table 6
Tests for cointegration

McCabe–Leybourne test for no cointegration	0.227
Approximate critical value (95%)	0.206
System R^2	0.99

to describe and readers are referred to Li and Maddala (1997) for a full outline of these methods).

This result has significant implications for our interpretation of the estimated system and for our understanding of the nature of change in demand itself. The rejection of the removal of the random walk components for the system of estimated equations suggests that there is significant support for rejecting the assumption of cointegration in this data. Thus, the data cannot be represented as a cointegrated system.[5] The implication for our interpretation of the nature of changing demand over the period under examination is that it has been influenced by changing tastes and that these changes have occurred in a stochastic manner.

Fig. 4 presents each of the random walk latent variable components as indices of their initial period values. These estimated series can be interpreted as the change in expenditure shares not explained by changes in relative prices or in real expenditure. For pork, 'other meat' and sausages, bacon and ham, changing tastes have resulted in positive changes in the share of expenditure on these products over and above those resulting from changes in relative prices and/or total expenditure on meat. For lamb and mutton, the effect of changing tastes has been slightly positive over the period, whilst for chicken, an initial stimulus to demand has been followed by a sharp negative impact on the share of expenditure as a result of changing tastes. The shares of expenditure of both beef and frozen products have been negatively affected by changes in tastes throughout the period.

Although the results reported above cannot be used to explain the causes of the changes in tastes, they can be rationalised in a number of ways. It is likely that a key reason for the changes in tastes in favour of pork was a significant government campaign promoting the consumption of pork during the late 1960s and the 1970s. This contributed to a significant increase in per capita pork consumption from 6 kg per capita in the 1960s to 25 kg per capita during the 1990s (Bank of Agriculture, unpublished data). The changes in tastes for the commodity group sausages, bacon and ham follows a similar path to that of pork. This is unsurprising given that the origin of most of these products is pork meat.

By contrast, a campaign attempting to promote chicken consumption did not appear to have as great an influence on consumer habits. This may, in part, have been due to the dampening effect of increased market prices resulting from restrictions in the supply of chicken, a finding which concurs with that reported by Karagiannis et al. (2000). The initial increase in the latent variable index does, however, suggest that the campaign partially offset the decline in consumption that would be expected from an increase in the relative price of chicken vis-à-vis pork. However, the first symptoms of salmonella during the 1990s may have resulted in a further decrease in chicken consumption, evidenced by the sharp decline in the latent variable component and an associated increase of pork consumption. It is possible that this offset any expected increases in chicken consumption due to the relative health attributes of white meat.

In the case of red meat, where health concerns may be expected to have had a greater downwards influence, there appears to have been no significant change in the tastes for lamb and mutton during the period under examination. By contrast, beef consumption does appear to have been detrimentally affected during the 1980s. It is possible that health concerns influenced meat consumption patterns against beef more than against lamb and mutton since these are staple goods,

[5] This conclusion appears to contradict that in Karagiannis et al. (2000). However, it is now accepted that tests under the unit root as a null hypothesis and the null hypothesis of stationarity commonly contradict each other within any given sample, and as yet there is no resolution to this problem.

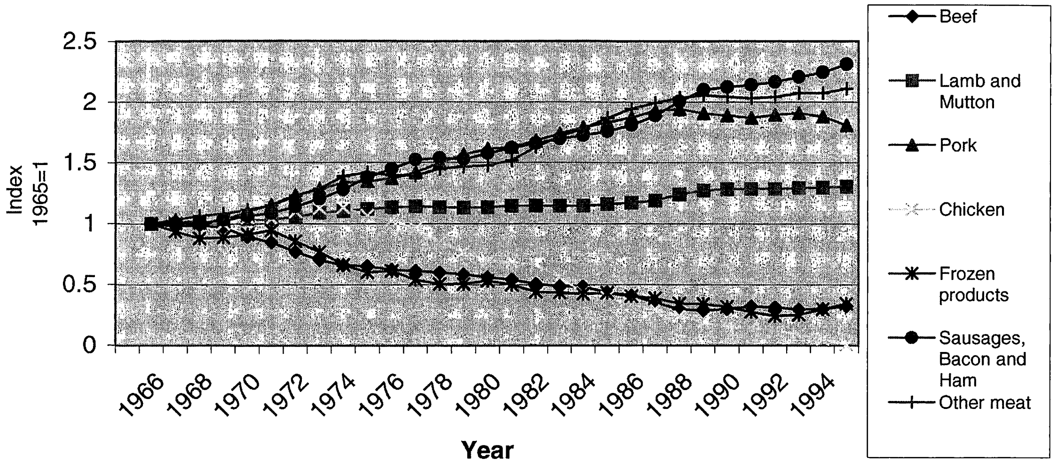


Fig. 4. Trends in the taste (latent variable) components as indices of their initial values.

the consumption of which are strongly associated with traditional Greek cuisine.

The category of 'other meats' includes meat by-products and game (deer, wild boar, etc.) which are mainly consumed in taverns and restaurants. Due to increases in the real income during this period, Greek consumers increased the number of meals taken away from home which in turn created a change in tastes towards meat consumed in restaurants, such as pork, meat by-products, game, sausages, ham and bacon.

The importance of the traditional cuisine in influencing the formation of tastes is also demonstrated by comparing the results of this paper to those reported in Burton and Young (1992) for the UK. Whilst in this paper the effect of taste has been to favour pork, and to a lesser extent lamb and mutton, as shares of expenditure have had a downwards influence on the budget share of beef and chicken, the conclusions reached for the UK were almost diametrically opposed, with a positive effect of taste on chicken (and beef for part of the period) and a negative influence on the budget share for pork and lamb.

7. Conclusions

In this paper we have demonstrated that the share of expenditure on different groups of food is differentially and significantly affected by changes in consumer tastes. In past studies on meat demand, the focus has in the main, been on the magnitude of the elasticities associated with price and expenditure. There have been a smaller number of attempts to explain tastes by investigating structural changes in demand. In this paper, we find that the fact that the latent variable components of the share equations are readily explained by observed changes in consumer behaviour signifies the importance of incorporating analysis of changing tastes in demand systems.

The results of the paper also demonstrate the key role of traditional cuisine in influencing tastes. This factor requires greater attention when comparing the magnitude of price and expenditure elasticities across studies relating to consumption data from different countries or regions.

The methodology used in this paper provides a flexible approach to modelling consumer behaviour. Its strength is that it enables the modelling of changes in tastes as a stochastic trend and in doing so allows for non-cointegrated equations. The estimation of these equations is relatively straight forward, and does not require frequency domain or Kalman filter approaches. More general models can be estimated which allow for evolution in all the parameters in the model. However, such models can easily become overparameterised, and the interpretation of the parameter estimates and calculation of elasticities and can be problematic. A weakness of this approach is that the parameter estimates must be interpreted as long-run parameters, and the simultaneous modelling of dynamics is likely to be over-ambitious even in a medium-sized system with a large sample size. The properties of the standard errors and subsequent test statistics generated analytically or by bootstrapping also remain open to conjecture. We believe that in applied circumstances, some of these problems might be mitigated by complementing the data with further non-sample information, perhaps by enforcing curvature restrictions, and even more generally, by relying on Bayesian or entropy-based approaches which introduce priors on the parameters of the equations and on the degree of taste shifts.

Acknowledgements

The authors would like to thank two anonymous referees for their useful comments and assistance in clarifying the interpretation of the results.

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