A Dynamic Approach to Estimate Theoretically Consistent US Meat Demand System

By

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Selected Paper prepared for presentation at the
Southern Agricultural Economics Association Annual Meetings
Orlando, Florida, February 5-8, 2006

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Abstract:

The paper conducts an empirical investigation of the US meat demand system using quarterly data on per capita meat consumption and prices. SUR maximum likelihood is used to estimate a static and dynamic (error correction) linear almost ideal demand systems. Results compare static and dynamic model elasticities.

Keywords: Demand analysis, Dynamic LAIDS, Cointegration

Introduction

The system of equations approach initiated by Stone (1954), including a group of equations (one for each consumer good) in the system and estimating them simultaneously, led to a framework for simultaneously testing some of the restrictions imposed by consumer theory (homogeneity and symmetry). Since then there have been numerous empirical studies of demand systems (Barten, 1969; Christensen et al., 1975; Deaton & Muellbauer, 1980) highlighting the inconsistencies between empirical and theoretical restrictions (i.e., homogeneity and symmetry) in demand analysis. The focus of their research was on the specification of functional form and testing the theoretical restrictions. Barten (1969) rejects homogeneity based on the likelihood ratio statistic obtained from the maximum likelihood estimation of the Rotterdam model. Christensen et al., (1975) conclude the same (reject homogeneity) using transcendental logarithmic utility function to estimate the demand system. Deaton and Muellbauer (1980) who developed the almost ideal demand system (AIDS) reject homogeneity based on F-tests. Deaton and Muellbauer assume that the rejection of homogeneity is a symptom of dynamic misspecification.

In the search for functional form, the AIDS model emerged as most popular functional form in the empirical demand analysis. A recent search on web revealed at least 156 papers either discuss or implement the AIDS model in their studies (Piggott and Marsh, 2004). At the same time, there has been lot of focus on two interrelated problems associated with the specification and the estimation of the AIDS: the first one dealt with the choice between its linear or non-linear specification and the second deals with the choice of an aggregate commodity price deflator (Pashardes, 1993; Buse, 1994; Moschini, 1995). Despite these problems, the AIDS remains one of the better alternatives available for empirical demand analysis. Until recently, the

AIDS model has been estimated with static models, ignoring the statistical properties of the data or the dynamic specification arising from time series analysis. A recurring conclusion in most of these studies was the rejection of homogeneity restriction. Additionally, other unexpected findings have been attributed to poor quality of aggregate data, functional misspecification, and inappropriate use of econometric techniques (Kenzaaenkamp & Barten, 1995).

Recent studies (Ng, 1995; Attfield, 1997; Karagiannis & Mergos, 2002) have suggested that inconsistency between theory and data in demand analysis may arise from inappropriate use of time-series techniques. Ng (1995), using techniques cointegration analysis, concludes that homogeneity holds in many cases. Attfield (1997) finds that homogeneity holds applying the triangular error correction procedure to almost ideal demand systems (AIDS). Balcombe and Davis (1996) proposed the canonical cointegrating regression procedure for estimating the AIDS. This procedure is used in cases where prices follow a distributed lag process, or there is a seasonal pattern. Karagiannis and Velentzas (2000) outlined the potential use of an error correction model (ECM) of the AIDS. Based on the time series properties of the data and existence of a cointegration relationship between the dependent and a linear combination of independent variables, an ECM for the AIDS can be established and econometrically estimated with an iterative seemingly unrelated regression (ISUR) procedure. For time series data, the latter approach seems more appropriate.

Meat demand has been of major interest in applied demand analysis for many countries in recent years. There are a number of studies for US (e.g. Eales and Unnevehr, 1993; Moschini and Meilke, 1989; Nayga and Capps, 1994; Brester and Schroeder, 1995; and Piggott and Marsh, 2004). We further explore the methodology for testing and setting an error correction form of demand systems by presenting a more complete set of alternative tests that can be used to establish long-run demand relationships. The paper provides empirical evidence and measures of elasticity estimates of an ECM-AIDS for meat demand in US over the period 1975(1)–2002(4).

The rest of this paper is organized as follows. The empirical model and the econometric results are presented in the following two sections, respectively. Elasticity analysis results are presented in the Section 4. Finally summary and conclusions are presented.

Empirical Model

We use the most popular AIDS model framework in our study. The AIDS model has many desirable attributes: (a) it is an arbitrary first order approximation to any demand system, (b) it satisfies the axioms of choice, (c) it aggregates over consumers, and (d) it is easy to estimate. The estimated coefficients in a linear approximate almost ideal demand system (LAIDS) model are easy to interpret. It has been extensively used in empirical work (Green and Alston, 1988; Chalfant, 1987). Following past literature, meat is treated as a weakly separable group comprising beef, pork, and poultry (chicken and turkey) in which consumption of an individual meat item depends only on the expenditure of the group, the prices of the goods within the group, and certain introduced demand shifters. The general specification of the AIDS model is given by:

$$w_i = \alpha_i + \sum_{j=1}^{n} \gamma_{ij} \log p_j + \beta_i \log(M / P) \tag{1}$$

where w_i is the share associated with the i^{th} good, α_i is the constant coefficient in the i^{th} share equation, γ_{ij} is the slope coefficient associated with the j^{th} good in the i^{th} share equation, p_j is the price of the j^{th} good, M is the total expenditure on the system of goods given by the following equation: $M = \sum_{i=1}^{n} p_i q_i$ where: q_i is the quantity demanded for the ith good. P is a general price index defined by:

$$\log P = \alpha_0 + \sum_{i=1}^{n} \alpha_i \log p_i + \frac{1}{2} \sum_{i=1}^{n} \sum_{j=1}^{n} \gamma_{ij} \log p_i \log p_j \qquad (2)$$

To comply with the theoretical properties of consumer theory the following restrictions are imposed on the parameters in the AIDS model:

- *Adding up restriction*: $\sum_{i=1}^{n} \alpha_i = 1, \sum_{i=1}^{n} \beta_i = 0, \sum_{i=1}^{n} \gamma_{ij} = 0$, allowing the budget share to sum to unity
- *Homogeneity*: $\sum_{j=1}^{n} \gamma_{ij} = 0$, which is based on the assumption that a proportional change in all prices and expenditure does not affect the quantities purchased. In other words, the consumer does not exhibit money illusion.
- *Symmetry*: $\gamma_{ij} = \gamma_{ji}$, represents consistency of consumer choices.

In empirical studies, to avoid the non-linearity and reduce the multi-colinearity effects in the model, equation (2) is sometimes approximated by a Stone index defined as $\log P = \sum_{i=1}^{n} w_i \log p_i$. We use the simple linear AIDS model in our empirical investigation. Researchers are mostly interested in the demand elasticities; the flexible functional form of the LAIDS model allows us to easily carry out the elasticity analysis. The demand elasticities are calculated as functions of the estimated parameters, and they have standard implications. According to Green and Alston (1990), elasticities in LAIDS can be expressed as: $\eta_i = 1 + \beta_i / w_i$ for income elasticity and $\eta_{ij}^{*} = -\delta_{ij} + w_j + \gamma_{ij} / w_i$ for compensated elasticity. The uncompensated elasticities are computed from $\eta_{ii} = -\delta_{ij} - \beta_i + \gamma_{ij} / w_i$, and $\eta_{ij} = \gamma_{ij} / w_i - \beta_i \; w_j / w_i$.

The LAIDS model estimated ignoring the time series properties of the data has come to be known as static LAIDS, which is also known as the long run LAIDS model. The long run model implicitly assumes that there is no difference between consumers' short run and long run behavior that is, the consumers' behavior is always in "equilibrium." However, in reality, habit persistence, adjustment costs, imperfect information, incorrect expectations, and misinterpreted real price changes often prevent consumers from adjusting their expenditure instantly to price and income changes (Anderson and Blundell, 1983). Therefore, until full adjustment takes place, consumers are "out of equilibrium." This is one of the reasons why most static LAIDS models cannot satisfy the theoretical restrictions (Duffy, 2003). It is therefore necessary to augment the long-run equilibrium relationship with a short-run adjustment mechanism. Moreover, the static LAIDS ignores the statistical properties of the data and the dynamic specification arising from time series analysis. It is well known that most economic data are nonstationary, and the presence of unit roots may invalidate the asymptotic distribution of the estimators. Therefore, traditional statistics such as t, F, and R-square are unreliable, and least squares estimation of the static LAIDS tends to be spurious.

Recent studies (Ng, 1995; Attfield, 1997; Karagiannis & Mergos, 2002) have suggested the use of cointegration and error correction concepts to overcome the spurious regression problem. The concepts of cointegration and the error correction model (ECM) were first proposed by Engle and Granger (1987) and have been widely used by researchers and practitioners in modeling and forecasting macroeconomic activities over the last decade. Engle and Granger (1987) showed that the long-run equilibrium relationship can be conveniently examined using the cointegration technique, and the ECM describes the short-run dynamic characteristics of economic activities. By transforming the cointegration regression into an ECM, both the long-run equilibrium relationship and short-run dynamics can be examined. Secondly, the spurious regression problem will not occur if the variables in the regression are cointegrated.

The variables concerned need to be tested for unit roots, before examining the cointegration relationship. The Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1981), Phillips-Perron (PP) (Phillips, 1987; Perron, 1988) statistics and the unit root testing procedure of Hylleberg et al. (1990) for quarterly data can be employed for this purpose. Once the orders of integration of the variables have been identified, either the Engle and Granger (1987) two-stage approach or the Johansen (1988) maximum likelihood approach can be used to test for the cointegration relationship among the variables in the models.

Once the cointegration relationship between the dependent variables and the linear combination of independent variables in the static LAIDS is confirmed, an ECM of the LAIDS can be established and econometrically estimated with appropriate algorithms. Applications of the ECM-LAIDS can be seen in the studies of demand for food, and meat products (Balcombe and Davis 1996; Attfield 1997; Karagiannis et al., 2000; Karagiannis and Mergos 2002). The ECM of the LAIDS used in this article follows Karagiannis & Mergos (2002) is given by

$$\Delta w_i = \delta_i \Delta w_{t-1} + \sum_{j=1}^{n} \gamma_{ij} \Delta \ln p_j + \beta_i \Delta \ln(M / P) + \lambda_i \mu_{it-1} + \mu_t \quad (3)$$

where Δ refers to the difference operator and μ_{it-1} is the ECM term that measures the feedback effects and is estimated from the corresponding cointegration equation. δ_i and λ_i are the parameters that need to be estimated. The restrictions in the static LAIDS are also applicable here.

Estimation Procedure and Results

Meat data used in the analysis are quarterly observations over the period 1975(1)–2002(4), providing a total of 112 observations. The quantity data are per capita disappearance data from the United States Department of Agriculture (USDA), Economic Research Service (ERS) supply and utilization tables for beef, pork, and poultry (sum of broiler, other-chicken, and turkey) gathered from online sources. The beef price is the average retail choice beef price, the pork price is average retail pork price, and the poultry price was calculated by summing quarterly expenditures on chicken, using the average retail price for whole fryers, and quarterly expenditures on turkey, using the average retail price of whole frozen birds, divided by the sum of quarterly per capita disappearance on chicken and turkey (similar to Piggott and Marsh, 2004). All of the price variables are published in the same source. The total expenditures on meat and budget shares of each meat product are estimated using the price and quantity information discussed above. Table 1 provides descriptive statistics of the variables included in the model.

First we investigate the time series properties (stationarity and cointegration) of the data and test for the appropriateness of the dynamic specification. The data used in the study are seasonally unadjusted quarterly observations. Hence, there exists a need to widen the concept of integration to allow for a mixture of first and fourth differencing being required to attain stationarity. Osborn et al. (1988) use the notation I(a, b) to summarize the required mixture, with the first argument represents the order of non-seasonal (first) differencing and the second argument the order of seasonal differencing necessary for stationarity. Thus, a quarterly series is said to be I(1, 1) if it requires both one quarter and seasonal (four quarter) differencing to become stationary. An I(0, 1) series requires only seasonal differencing, an I(1, 0) series needs only one quarter differencing, and an I(0, 0) series is stationary in levels and does not need differencing.

We use the unit root testing procedure of Hylleberg et al. (1990) to investigate the time series properties of the above data. Test results are presented in Table 2. The null hypotheses for these tests are the series investigated are an I(0, 1). The tests are based on the following regression after augmentation with lagged dependent variables and deterministic components:

$$y_{4t} = \pi_1 y_{t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1} \quad (4)$$

where $y_{1t} = (1 + L + L^2 + L^3)x_t$, $y_{2t} = -(1 - L + L^2 - L^3)x_t$, $y_{3t} = -(1 - L^2)x_t$, and $y_{4t} = (1 - L^4)x_t$

Eq. (4) is estimated initially with all lagged values of the dependent variable up to a maximum lag of eight quarters, plus a constant, trend and three seasonal dummies. A testing down procedure is then followed to eliminate insignificant lagged values of the dependent variable, working from the longest lags towards the shortest, but always subject to the condition that the residuals exhibited no evidence of serial correlation up to the fourth order (Duffy, 2003).

The null hypothesis that X_t is I(0, 1) is not rejected if all $\pi_i = 0$ (i = 1, 2, 3, 4). This is tested by a joint F statistic, denoted as F_{1234} in Table 2. The alternative hypotheses that are worth considering are that each variable is I(1, 0) or I(0, 0). An insignificant t-value for π_1 combined with a significant F_{234} statistic implies that the series is I(1, 0), whereas a significant t-statistic for π_1 and a significant F_{234} statistic indicates that the series is I(0, 0). The F_{1234} statistics in Table 2 indicate that the all of the series used in this study are not I(0, 1). The conjunction of insignificant t-ratios for π_1 (implying non-rejection of $\pi_1 = 0$) and significant values for F_{234} (rejecting the presence of unit roots at the seasonal frequency) leads to the conclusion that the all the series I(1, 0). Fig. 1, and Fig. 2, show the time path of levels and the first differences of budget shares, expenditure and price series respectively.

Having established that the series are I(1,0) (first differencing needed) we continue further to test for cointegration between the variables of Eq. (1) using Engle and Granger (1987) methodology. According to the result reported in Table 3, only the budget-shares of poultry are not cointegrated with prices and expenditure at the 5% significance level. This holds irrespectively of whether or not a time trend is included. These results necessitate the use of a dynamic cointegration test using Eq. (2) and Banerjee et al. (1986) and Kremers et al. (1992)

methodology. The residuals from the earlier cointegration regression are used in this step. Based on the statistical significance of λ_i parameters associated with the residuals, the results in Table 3 indicate the existence of a cointegrated regression equation for all budget shares.

The cointegration relationships in an equation can be modeled using the ECM-LAIDS specification as discussed above. Since the sum of all expenditure shares in the LAIDS model is equal to unity, the residuals variance-covariance matrix is singular. The usual solution is to delete an equation from the system and estimate the remaining equations and then calculate the parameters in the deleted equation in accordance with the adding-up restrictions. In our case we arbitrarily drop the poultry equation from the system. First we estimate the unrestricted static LAIDS models using Eq. (1). We add the deterministic components in the form seasonal dummies and a linear time trend in the model. Estimation is carried out implementing the maximum likelihood (ML) routines for seemingly unrelated regression (SUR). Later we impose the homogeneity and symmetry conditions separately and then combine them to estimate the restricted models. The likelihood ratios estimated from the unrestricted and restricted models are presented in Table 4. Results indicate both homogeneity and symmetry conditions are satisfied by the static model. The estimates from the restricted static LAIDS model are presented in the Table 5 (homogeneity and symmetry constraints imposed).

With regard to the dynamic LAIDS the seasonal dummies and linear time trend are omitted. The Engle and Granger two-step approach is employed for estimating cointegrating regressions. The residuals from these regressions are obtained and incorporated into Equation 3, and then the unrestricted ECM-LAIDS is estimated using the MLE of SUR procedure. The estimates are shown in Tables 8. The estimated parameters δ_i in the ECM-LAIDS are all significantly different from zero which indicates that habit persistence plays an important role in

US meat consumption decision-making process. In other words, the previous distribution of meat expenditure in different products influences US meat consumer's current decision on meat product choice. The coefficients of the error correction terms are all statistically significant at the 1% level and correctly signed, suggesting that any deviations of meat spending from the long-run equilibrium are accounted in the dynamic LAIDS model. With regard to the restriction tests (see Table 4), unfortunately the ECM-LAIDS passes only the symmetry test at the 5% level, while failing the homogeneity test and the joint tests for both homogeneity and symmetry. It indicates that though the dynamic adjustment is likely the correction for misspecification of the functional form and but not the solution for violation of demand theory. The estimates from the restricted dynamic LAIDS model are presented in Table 8 (imposing both symmetry and homogeneity). The parameter estimates from both the restricted models (static and ECM) are used for elasticity analysis.

Elasticity Analysis Results

The estimated Marshallian own-price elasticities from the static model are -0.964, -0.822, and -0.306 for beef, pork, and poultry, respectively (presented in upper half of Table 7). These results mean that per capita beef consumption conditional on meat expenditure is more sensitive to its own price change, while poultry consumption is least sensitive to changes in its own price (consistent with earlier research of Piggott and Marsh, 2004). The Marshallian own-price elasticities from the ECM-LAIDS model are -0.663, -0.985, and -0.661 for beef, pork, and poultry, respectively (presented in lower half of Table 7), suggesting pork consumption is more sensitive to its own price change in short-run, while effect of price change is almost equal for beef and poultry. The Marshallian own price elasiticities are quite different from the static model elasticity estimates.

The compensated cross-price elasticities are positive for beef, pork and poultry indicating they are substitutes (presented in upper half of Table 8). In particular, a one percent increase in pork price causes a 0.28% increase in beef consumption and a one percent increase in poultry price increases beef consumption by 0.06%. Compensated elasticities from the ECM-LAIDS differ in magnitude but are similar in the order i.e. one percent increase in price of pork causes a 0.30 % increase in consumption of beef (presented in lower half of Table 8).

The expenditure elasticity estimates calculated based on the estimates from the static LAIDS model were 1.168 for beef, 0.965 for pork, and 0.557 for poultry (reported in upper half of Table 7). This implies that beef is the most sensitive to changes in total expenditures, followed by pork, and then poultry. This finding means beef is the biggest gainer (loser) of the three competing meats when consumers increase (decrease) per capita expenditures. The order of expenditure elasticities changes completely when we look at estimates derived from the ECM-LAIDS (presented in lower half of Table 7). The estimates 0.49, 1.37, and 1.87 for beef, pork, and poultry respectively, suggest that poultry is the biggest gainer of the three meats in short-run when consumers increase expenditures on meat.

Summary and Conclusions

The objective of this paper was to test theoretical restrictions on a meat demand system using cointegration techniques. Quarterly meat disappearance data spanning from 1975(1) to 2002(4) and average retail prices are used. We investigate the time series properties of the data (stationarity and cointegration) and estimate an ECM-LAIDS model. Elasticities from a static and an ECM-LAIDS models are compared. The static model satisfies all the theoretical restrictions (homogeneity and symmetry) but suffers from the dynamic misspecification. In general, both models give reasonable results of the compensated and uncompensated price elasticities. As for the expenditure elasticities there was a notable difference between the two models with ECM version showing completely opposite results.

The use of time-series techniques has been offered as a potentially promising way for improving the theoretical consistency of demand systems in empirical work by accounting for dynamics in consumer behaviour. Homogeneity, however, is rejected in the ECM-LAIDS. The empirical results in this paper generate elasticity estimates that differ from those generated from static model; little research is available on estimates reliability. Future research may focus on identifying sources of differing results, perhaps through Monte Carlo or other simulations exercises.

References

Anderson, F., and Blundell, R. "Testing restrictions in a flexible demand system: an application to consumer's expenditure in Canada." Review of Economic Studies 50(July, 1983):397–410.

Attfield, C.L.F. "Estimating a cointegrating demand system." European Economic Review 41(Jan, 1997):61–73.

Balcombe, K.G., and Davis, J.R. "An application of cointegration theory in the estimation of the almost ideal demand system for food consumption in Bulgaria." Agricultural Economics 15(Sept, 1996):47–60.

Banks, J., Blundell, R., and Lewbel, A. "Quadratic Engel curves and consumer demand." Review of Economics and Statistics 79(Nov, 1997):527–539.

Banerjee, A., Dolado, J., Hendry, D., and Smith, G. "Exploring equilibrium relationships in econometrics through static models: some Monte Carlo evidence." Oxford Bulletin of Economics and Statistics 48(Aug, 1986):253–278.

Barten, A.P. "Maximum likelihood estimation of a complete system of demand equations." European Economic Review 1(Mar, 1969):7-73.

Brester, G.W., and T.C. Schroeder. "The Impacts of Brand and Generic Advertising on Meat Demand." American Journal of Agricultural Economics 77(Nov, 1995):969–79.

Buse, A. "Evaluating the Linearized Almost Ideal Demand System." American Journal of Agricultural Economics 76(Nov, 1994):781-793.

Chalfant, J.A. "A Globally Flexible, Almost Ideal Demand System." Journal of Business and Economic Statistics 5(April, 1987):233-42.

Christensen, L.R., D.W. Jorgenson, and Lawrence J.L. "Transcendental logarithmic utility functions." American Economic Review 65(June, 1975):367-383.

Coulibaly, N., and B.W. Brorsen. "Explaining the differences between two Previous Meat Generic Advertising Studies." Agribusiness an International Journal 15(Aug, 1999):501-16.

Deaton, A., and J. Muellbauer. "An Almost Ideal Demand System." The American Economic Review 70(June1980):312-326.

Dickey, D.A., and Fuller,W.A., 1981. "Likelihood ratio statistics for autoregressive time series with a unit root." Econometrica 49(June, 1981):1057–1072.

Duffy, M. "Advertising and Food, Drink and Tobacco Consumption in the United Kingdom: A Dynamic Demand System." Agricultural Economics, 1637(2003):1-20.

Eales, J.S., and L.J. Unnevehr. "Simultaneity and Structural Change in U.S. Meat Demand." American Journal of Agricultural Economics 75(May, 1993):259–68.

Engle, R.F., and Granger, C.W.J. "Cointegration and error correction: representation, estimation and testing." Econometrica 55(Mar, 1987):251–276.

Ghysels, E., Lee, H.S., and Noh, J. "Testing for unit roots in time series: some theoretical extensions and a Monte Carlo investigation." Journal of Econometrics. 62(June, 1994):415–442.

Green, R., and J.M. Alston. "Elasticities in AIDS Models." American Journal of Agricultural Economics 72(May, 1990):442–45.

Granger, C.W.J., and P. Newbold. “Spurious Regressions in Econometrics.” Journal of Econometrics 2(July, 1974):111-120.

Hylleberg, S., Engle, R.F., Granger, C.W.J., and Yoo, B.S. “Seasonal integration and cointegration.” Journal of Econometrics 44(April, 1990):215–238.

Johansen, S. “A Statistical Analysis of Cointegrating Vectors.” Journal of Economic Dynamics and Control 12(June, 1988):231-254.

Karagiannis, G., S. Katranidis, and K. Velentzas. “An Error Correction Almost Ideal Demand System for Meat in Greece.” Agricultural Economics 22(Jan, 2000):29-35.

Karagiannis, G., and G. J. Mergos. “Estimating Theoretically Consistent Demand Systems Using CI Techniques with Application to Greek Food Data.” Economics Letters 74(Jan, 2002):37-43.

Keuzenkamp, H.A., and A.P. Barten. “Rejection without falsification: on the history of testing the homogeneity condition in the theory of consumer demand.” Journal of Econometrics 67(May, 1995):103–127.

Kremers, J., Ericsson, N., Dolado, J. “The Power of Cointegration tests.” Oxford Bulletin of Economics and Statistics 54(Aug, 1992):325–348.

Moschini, G. “Units of measurement and the Stone index in demand system estimation.” American Journal of Agricultural Economics 77(Feb, 1995):63–68.

Moschini,G., and K.D. Meilke. “Modeling the Pattern of Structural Change in U.S. Meat Demand.” American Journal of Agricultural Economics 71(May, 1989):253–61.

Ng, S. “Testing for homogeneity in demand systems when the regressors are nonstationary.” Journal of Applied Econometrics 10(March, 1995):147–63.

Nayga, R.M. Jr. and Capps, O. Jr. “Tests of Weak Separability in Disaggregated Meat Products.” American Journal of Agricultural Economics 71(Feb, 1994):800–08.

Pashardes, P. “Bias in estimating the almost ideal demand system with the Stone index approximation.” Economics Journal 103(July, 1993):908– 15.

Perron, P.P. “Trends and random walks in macroeconomic time series: Further evidence and implications.” Journal of Economic Dynamics Control 12(June, 1988):279–332.

Phillips, P.P. “Time series regression with a unit root.” Econometrica 55(March, 1987):277–301.

Piggott, N.E., and T.L. Marsh. “Does Food Safety Information Impact U. S. Meat Demand?” American Journal of Agricultural Economics 86(Feb, 2004):154-74.

Stone, J. R. N. “Linear Expenditure Systems and Demand Analysis: An Application to the Pattern of British Demand.” Economic Journal 64(May, 1954):511-27.

U.S. Department of Agriculture. Economic Research Service. Red Meats Yearbook. Statistical Bulletin No. 921, Washington DC, 1995.

U.S. Department of Agriculture. Economic Research Service. Poultry Yearbook. Statistical Bulletin No. 916, Washington DC, 1995.

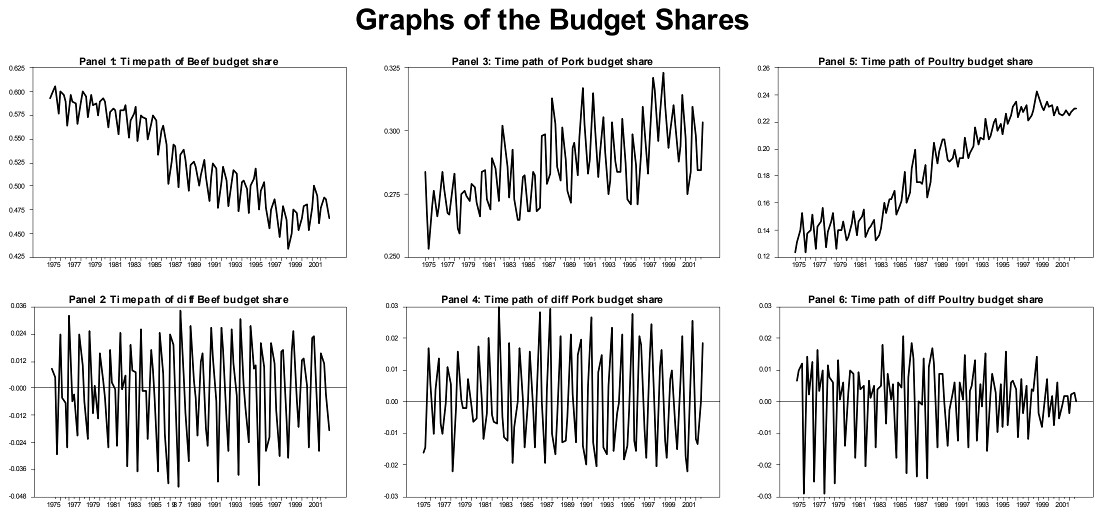

Figure 1. Time Plots in Levels and Differences of Budget Shares

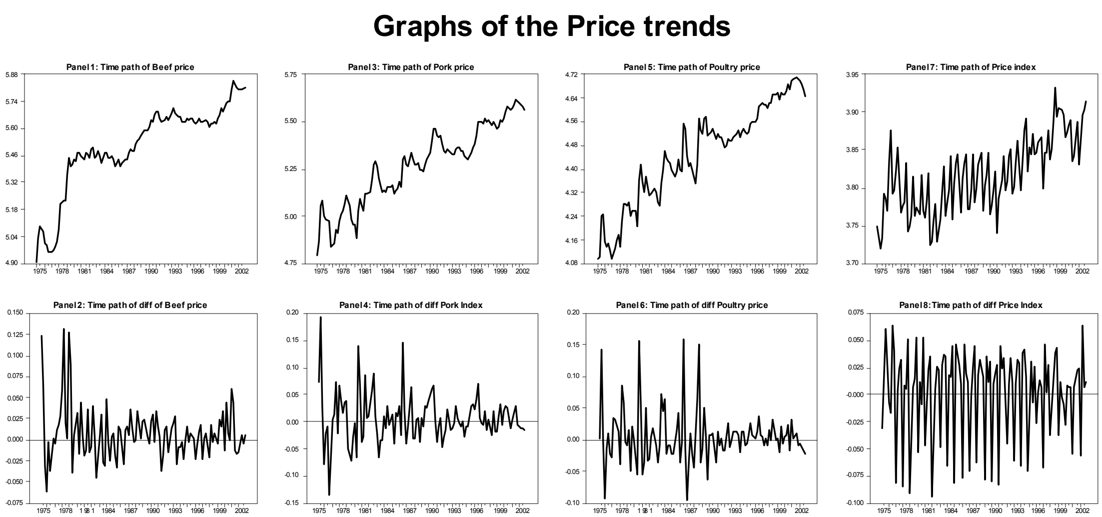

Figure 2. Time Plots in Levels and Differences of Prices and Expenditure

Table 1. Descriptive Statistics of Variables Included in the Model

Variable$_a$	Mean	Std. Dev	Minimum	Maximum
W_1	0.5288	0.0472	0.4345	0.6059
W_2	0.2864	0.0149	0.2534	0.3229
W_3	0.1847	0.0366	0.1232	0.2425
ln p_1	5.5052	0.2235	4.9040	5.8432
ln p_2	5.2688	0.2087	4.7941	5.6166
ln p_3	4.4220	0.1848	4.0096	4.7176
ln (m/P)	3.8170	0.0484	3.7187	3.9319

Note: $_a$ Subscripts refer to (1) Beef, (2) Pork, and (3) Poultry

Table 2. Seasonal Unit Root Test Results (Hylleberg et al., 1990)

Variable	*t*-statistic for Π_1	F_{234}	F_{1234}	Augmentation of Lags	Conclusion
W_1	-1.41	16.09	12.84	1	I(1,0)
W_2	-3.17	10.54	11.53	3	I(1,0)
W_3	-2.83	28.24	29.14	1	I(1,0)
ln p_1	-3.12	42.26	37.08	1	I(1,0)
ln p_2	-2.89	33.92	32.82	2	I(1,0)
ln p_3	-3.04	29.24	29.96	2	I(1,0)
ln (m/P)	-3.24	21.14	21.61	2	I(1,0)
Critical values (5%)	-3.53	5.99	6.47		

Notes: The 5 % critical values are taken from Ghysels et al. (1994); they are appropriate for a test regression that includes, constant, seasonal dummies, and a linear trend and which is estimated from a sample size of 100 observations.

Table 3. Cointegration Test Results

Equation	CI test[b]		Dynamic CI test[c]	
	ADF	PP	λ	*t*-value
W_1	-0.382	-7.294	-0.582	-7.46
W_2	-2.351	-8.315	-1.009	-11.27
W_3	-0.971	-4.595	-0.192	-3.39

Notes: Cointegration tests are based on regression including a constant term and a time trend.
[b]For Engle Granger CI test, the tabulated critical value at 5% is 4.87.
[c]Based on estimation of Eq. (2).

Table 4. Constraints Likelihood Ratio Test Results

	Calculated x^2	p-Value	Degrees of Freedom
	ECM-LAIDS		
Symmetry	0.12	0.7245	1
Homogeneity	21.87	0.0000	2
Homogeneity and Symmetry	23.26	0.0001	3
	Static-LAIDS		
Symmetry	0.72	0.3955	1
Homogeneity	0.06	0.9723	2
Homogeneity and Symmetry	0.84	0.8397	3

Table 5. Estimated Parameters of Static LAIDS for the Meat Demand in US, 1975(1)-2002(4)

Variable	Beef	Pork	Poultry
ln p_1	0.066 (5.08)		
ln p_2	-0.002 (-0.29)	0.048 (5.94)	
ln p_3	-0.064 (6.64)	-0.046 (-0.39)	0.110 (8.56)
ln (m/P)	0.089 (1.84)	-0.009 (-0.31)	-0.079 (-2.19)
q_1	0.026 (6.44)	-0.007 (-2.86)	-0.018 (-6.47)
q_2	0.033 (9.36)	-0.021 (9.10)	-0.012 (-4.87)
q_3	0.030 (9.35)	-0.019 (9.40)	-0.010 (-4.66)
T	-0.001 (-23.39)	0.000 (5.33)	0.001 (4.33)
Constant	-0.271 (1.45)	0.626 (2.23)	-0.544 (-3.81)

Notes: Homogeneity and symmetry constraints imposed. Poultry estimates derived using adding up constraints. The *t*-values are given in the parentheses.

Table 6. Estimated Parameters of ECM-LAIDS for the Meat Demand in US, 1975(1)-2002(4)

Variable	Beef	Pork	Poultry
ΔW_{t-1}	-0.045 (-0.61)	0.217 (2.50)	-0.172 (-1.13)
Δln p_1	0.036 (1.08)		
Δln p_2	0.008 (0.36)	0.035 (1.64)	
Δln p_3	-0.045 (-2.34)	-0.044 (-2.98)	0.089 (4.88)
Δln (m/P)	-0.267 (-6.54)	0.109 (3.76)	0.158 (6.55)
ECM term	-0.322 (-4.65)	-0.612 (-6.76)	0.935 (6.23)

Notes: Homogeneity and symmetry constraints imposed. Poultry estimates derived using adding up constraints. The *t*-values are given in the parentheses. Constant, linear time trend and seasonal dummies omitted.

Table 7. Marshallian and Expenditure Elasticities of the Meat Demand in US, 1975(1)-2002(4)

Product	Beef price	Pork price	Poultry price	Expenditure
		Static LAIDS		
Beef	-0.964	-0.053	-0.151	1.168
Pork	0.010	-0.822	-0.153	0.965
Poultry	-0.121	-0.129	-0.306	0.557
		ECM-LAIDS		
Beef	-0.663	0.162	0.005	0.495
Pork	-0.170	-0.985	-0.220	1.376
Poultry	-0.716	-0.501	-0.661	1.879

Note: Derived from the homogeneity and symmetry imposed estimates.

Table 8. Compensated Price Elasticities of the Meat Demand in US, 1975(1)-2002(4)

Product	Beef price	Pork price	Poultry price
		Static LAIDS	
Beef	-0.344	0.285	0.059
Pork	0.521	-0.542	0.020
Poultry	0.173	0.032	-0.206
		ECM-LAIDS	
Beef	-0.401	0.305	0.095
Pork	0.559	-0.586	0.027
Poultry	0.279	0.043	-0.323

Note: Derived from the homogeneity and symmetry imposed estimates.