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Demand Equations with some Nonstationary Variables: The Demand for Farmed and Wild Salmon in Japan

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Frank Asche
Cathy Roheim Wessells

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Frank Asche*
Stavanger University College and
Centre for Fisheries Economics, Norwegian School of Economics and Business Administration
Box 2557 Ullandhaug
N-4091 Stavanger, Norway
E-mail: Frank.Asche@tn.his.no

and

Cathy Roheim Wessells
Department of Environmental and Natural Resource Economics,
Office of the Dean, 111 Woodward Hall, University of Rhode Island, Kingston, RI 02881
E-mail: crw@uri.edu

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Abstract

There seems to be substantial evidence in favor of treating price series as nonstationary. There are also strong arguments for treating expenditure shares as stationary. With this structure in the data, it is generally not valid to estimate share equations, which most commonly-used demand systems are formulated in, as there will not be a long-run relationship between data series that are integrated of different orders. We show that one can estimate a demand system using traditional tools if the shares are stationary and prices are nonstationary but cointegrated. Futhermore, with nonstationary prices, one can use the Generalized Composite Commodity Theorem to test the validity of grouping goods into an aggregate, while using only the data on the disaggregated goods. This is in contrast to the normal assumptions of weak separability, which is typically not tested because one needs additional data. We use these tools to investigate the extent of competition in the Japanese import market between wild-caught North American salmon and farmed salmon and trout.

Keywords: Demand system, nonstationary data series, cointegration, aggregation, salmon

JEL Classification: D12, C32

1. INTRODUCTION

During the last decade, attention to time series properties of economic data has become an important element in the process of modeling long-run relationships. Modeling systems of demand equations are largely an exception. In general, multiple data series must be stationary or integrated of the same order to form a long-run relationship with each other.[1] Most economic data series are found to be nonstationary, and they will then form long-run relationships with each other only when they are cointegrated (Engle and Granger, 1987). Several studies have shown that it is reasonable to treat the set of prices used in demand analyses as nonstationary data series. This includes market delineation studies (Ardeni, 1989; Goodwin and Schroeder, 1991; Doane and Spulber, 1994), aggregation studies (Lewbel, 1996; Davis, Lin and Shumway, 2000) and demand studies (Ng, 1995; Attfield, 1997; Karagiannis and Mergos, 2002).[2] On the other hand, expenditure share variables are normally thought of as stationary since they are bounded between zero and one.[3] Since most common demand systems have shares as left hand side variables, this indicates that one would not expect to find long-run relationships between the dependent and independent variables as one then has equations with data series integrated of different orders. This implies that there cannot exist demand equations with such data structures.[4]

In this paper we will show that, when the price series are cointegrated, this data structure can provide the long-run relationships that are necessary for demand relationships to exist. Moreover, traditional econometric tools can be used to estimate systems of demand equations as normal inference theory will apply. This is in contrast to the case when all data series (including the budget shares) are nonstationary, when traditional econometric tools and normal inference theory will not apply.

Another question when specifying demand systems is whether it is reasonable to investigate a group of variables independently of all other goods in a consumers' bundle. There are two fundamentally different approaches to validate aggregation – a) different forms of separability and b) relationships between prices (Deaton and Muellbauer, 1980a). Weak separability has been the preferred criterion used to justify investigation of only a limited group of goods. Since this is a difficult criterion to test, weak separability is primarily assumed. Another aggregation criterion suggested by Hicks (1936) and Leontief (1936) is the Composite Commodity Theorem (CCT), where one investigates relationships between prices. However, this criterion has not been used in empirical analyses because it must hold as an exact relationship. Lewbel (1996) provides the Generalized Composite Commodity Theorem (GCCT) that is empirically testable. In contrast to the case for weak separability, a test of whether the GCCT holds only needs the data of the demand system. In this paper we will use the GCCT to test if a group of goods can be analyzed independently of the other goods in the consumer's bundle.

The empirical application will estimate import demand for high-valued frozen salmon in Japan. Until the late 1980s, this flow consisted almost exclusively of wild sockeye salmon from North America, primarily Alaska. However, during the 1980s salmon farming was a growing industry. In the early 1990s, there was a considerable growth in production of farmed salmon trout and coho in Chile and Norway, which was largely exported to Japan. By the late 1990s, the Japanese imports of both farmed salmon trout and farmed coho were larger than imports of wild-caught North American salmon. Throughout this period, Alaskan fishermen have seen their prices for salmon decreasing. There have been a number of industry explanations for this. After record high prices in 1988, the Exxon Valdez oil spill of 1989 was blamed for causing low market prices. During the early 1990s, further Alaskan salmon price declines resulted in legal actions involving allegations of collusive behavior among Japanese buyers. During the remainder of the 1990s, attention turned to the dramatic increase in production of farmed salmon as a further cause of loss to market share in the Japanese market for wild-caught Alaskan salmon. It is therefore of interest to investigate to what extent farmed salmon and trout have become substitutes for wild North American salmon in its principal market, Japan.

This paper will be organized as follows. First, a discussion of estimation of a demand system with some non-stationary data series is provided. Then the approach suggested by Lewbel using the GCCT to determine product groups and appropriate aggregation is examined. The paper continues with the empirical example and discussion.

2. THE ALMOST IDEAL DEMAND SYSTEM AND NONSTATIONARY DATA

We will use the Almost Ideal Demand System (AIDS) of Deaton and Muellbauer (1980b), as this seems to be the most popular demand system in empirical work. Each equation in the AIDS can be written as

$$(1) \qquad w_{it} = \alpha_i + \sum_j \gamma_{ij} \ln p_{jt} + \beta_i \ln\left(\frac{X_t}{P_t}\right) + e_{it},$$

thus for observation t, w_{it} is the budget share of good i, p_{jt} is the price of commodity j, X_t is total expenditure on the goods in the system and $\ln P_t$ is a translog price index defined by

$$(2) \qquad \ln P_t = \alpha_0 + \sum_i \alpha_i \ln p_{it} + \frac{1}{2}\sum_i \sum_j \gamma_{ij} \ln p_{it} \ln p_{jt} .$$

Economic theory implies the following restrictions:

Homogeneity: $\sum_j \gamma_{ij} = 0$;

Symmetry: $\gamma_{ij} = \gamma_{ji}$;

and the fact that all budget shares must sum to one requires that $\sum_i \alpha_i = 1$, $\sum_i \gamma_{ij} = 0$, and $\sum_i \beta_i = 0$.

The AIDS is linear except for the translog price index lnP_t. This problem has traditionally been circumvented in most applied work as suggested by Deaton and Muellbauer, by using a scaled Stone's price index, AP_t^*, where A is a constant scaling factor and $\ln P_t^* = \sum_i w_{it} \ln p_{it}$, which makes the system linear. Deaton and Muellbauer argue that if prices are highly collinear, Stone's index should be a good approximation of the translog index.

The use of Stone's price index has been questioned by several authors (Green and Alston, 1990; Pashardes, 1993; Buse, 1994; Moschini, 1995; Asche and Wessells, 1997). There are several explanations for this, including simultaneity bias (measurement errors), omitted variables and that Stone's index is not invariant to units of measurement. To avoid these problems, Moschini (1995) suggested three alternative Stone indices that correspond to loglinear versions of the classical Laspeyre, Paasche and Tornquist indices. We will use the loglinear analogue to a Laspeyre index, which can be written as

(3) $$\ln P_t^L = \sum_{i=1}^{n} w_i^0 \ln p_{it} .$$

Each w_i^0 is the budget share at some fixed point, and constant. This index has the additional advantage that it cannot be endogenous due to variation in the expenditure shares within the index.

If all variables are stationary, the demand system can be estimated using traditional econometric techniques. When all data series are nonstationary but integrated of the same order, the demand equation represents a long-run relationship between prices and shares only if the prices are cointegrated. Attfield (1997) and Karagiannis and Mergos (2002) show that this is the case in their systems.

If the expenditure shares are stationary, but the prices nonstationary, a different approach must be taken. It is well known that an equation must be balanced in the sense that all variables or linear combinations of variables must be integrated of the same order for there to be long-run relationships (Hendry, 1995). This does not mean that all variables must be integrated of the same order. For instance, a classical Error Correction Model contains both $I(0)$ and $I(1)$ data series. This is possible because when the $I(1)$ data series are cointegrated, their linear combination will be $I(0)$, so that all components in the equation are $I(0)$. Moreover, Sims, Stock and Watson (1993) show that ordinary inference theory applies for any parameter that can be represented as the parameter on an $I(0)$ variable.

If we know that the prices in equation (1) are nonstationary, but the budget shares are stationary, we know that, in general, there will not be a long run relationship between these variables (Engel and Granger, 1987). If so, the data cannot be used to estimate demand equations for these goods. However, if the nonstationary prices are cointegrated, they will form a linear combination that is stationary. This will make the equation balanced and we can continue to estimate demand equations. Moreover, the results of Sims, Stock and Watson (1990) indicate that ordinary inference theory will apply to the parameters on all the prices. This is because the cointegration vector is identified only up to a nonsingular transformation, and can be represented normalized on all the prices. That is, with n prices in the system, the cointegration vector can be given n different representations depending on which price one normalizes upon:

$$(4) \qquad \gamma_{ik}\left(\ln p_{kt} + \sum_{j,j\neq k}\frac{\gamma_{ij}}{\gamma_{ik}}\ln p_{jt}\right)$$

Hence, each γ_{ik} parameter, for all i and k, can be represented as the parameter on the cointegration vector, which is stationary, and not only on the individual nonstationary price.

When there is more than one cointegration vector present, the different cointegration vectors represent independent linear relationships between the variables in question. Stock and Watson (1988) show that in a system with n data series integrated of the same order, there will be n-r cointegration vectors and r stochastic trends. With n nonstationary prices, one can find at most n-1 cointegration vectors. This corresponds to the situation when all prices follow the same stochastic trend.

As noted above, because of the adding-up condition, n-1 is also the number of independent demand equations in a system with n prices. One will therefore expect to find n-1 cointegration vectors when investigating the relationship between the prices in such a demand system. Finding fewer cointegration vectors implies further restrictions on the demand equations that are specific to the data in question, as one or more of the price vectors will be linear combinations of the others.[5] This should not influence the economic interpretation of the elasticities derived from the estimated parameters. However, one must be sure that all the prices belong to the cointegration vectors, as normal inference theory will break down for the parameters of prices that are not part of the cointegration relationships.

A question of some interest is whether to expect that the relative prices are stationary if the nominal prices are nonstationary.[6] Lewbel (1996) found nonstationary nominal prices, and to his surprise, nonstationary relative prices as well, in the empirical analysis. He did not discuss the theoretical reasons why this might be the case, but it is of some interest to look closer at this issue as relative prices will be stationary only under fairly strong assumptions about market structure. Given nonstationary nominal prices, stationary relative prices, $\ln p_j - \ln p_k$, imply that each pair of prices are cointegrated with a parameter of unity.[7] If this is the case, the nominal prices are proportional and the Law of One Price holds. This special case will only occur when the markets for the goods in question are fully integrated. As a result, one can generally expect that with nonstationary nominal prices, the relative prices are nonstationary as well.

It is also of interest to discuss the parameter on the real expenditure term in this context. If the real expenditure series is stationary, ordinary inference theory will apply. However, if it is nonstationary, there can be two outcomes. First, if expenditure is cointegrated with the prices, ordinary inference theory will apply since it can be added to the expression in equation (4). If expenditure is not cointegrated with the prices, it will be an $I(1)$ term that does not form a relationship with any other variable in the system. The true β parameters will then be zero, and the demand system will be homothetic.

3. AGGREGATION

As noted in the introduction, there are two fundamentally different approaches to validate aggregation – different forms of separability and relationships between prices (Deaton and Muellbauer, 1980a). Creating groups of goods that are investigated in isolation from the rest of the consumer's bundle is normally justified with a weak separability assumption. Weak separability gives conditions for the structure of consumers' preferences so that it is valid to investigate the demand for a limited number of goods. However, whether a weak separability assumption is valid depends on the relationship between the goods in question and all other goods in the consumer's bundle. To test weak separability one needs data on all goods, and it involves estimation of a much larger demand system. Tests are in general difficult to conduct and have low power (Lewbel, 1996). This leads most researchers to assume weak separability without any testing. However, this also makes the results questionable since one can raise doubts with respect to the validity of the separability assumption, as for example, Winters (1984).

Aggregation based on relationships between prices is formulated in the CCT (Hicks, 1936; Leontief, 1936). The CCT basically states that a group of goods with proportional prices can be aggregated and be represented as a single good with a single price. However, the CCT must hold as an identity. Alternatively, the GCCT of Lewbel (1996) gives an empirically operational version of the CCT. Furthemore, Lewbel shows that an AIDS aggregates consistently within each group.

The GCCT can be represented as follows. Define ρ_i as the ratio of the price of good i to the price index of the group of interest, P_I.

(5) $\rho_i = p_i / P_I$

If the distribution of the relative price ρ_i is independent of the group index P_I, the GCCT will hold. Let $r_i = \ln \rho_i$ and $R_I = \log P_I$. Lewbel shows that for nonstationary prices, this is equivalent to finding that u_t in the relationship

(6) $r_i - R_I = u_t$

is nonstationary, or that the relative price ρ_i is not cointegrated with P_I.[8]

When testing whether a group of goods can be aggregated with the GCCT, one needs only price and quantity data for the goods of interest (the quantity data is necessary to construct the price index). Hence, in contrast to weak separability, the GCCT can be easily be used to verify that the group of goods considered is a valid group using only the data required to estimate the demand system of interest. If this condition fails, there may of course be other criteria that provide support for the group. However, if these are not testable, they provide much poorer support for investigating the group of interest separate from other goods.

4. BACKGROUND AND DATA

The Japanese market for salmon is the largest in the world (with the European Union and U.S. as the two other large markets). Japanese fishermen land substantial quantities of low-valued chum salmon, but virtually all high-valued salmon is imported. Traditionally, wild salmon in substantial quantities has been available only in the northern Pacific, with the U.S. (Alaska) as the largest harvester. The U.S. and Canada were traditionally the only large exporters of salmon in the world, and most of their high-valued salmon was sent to Japan until the late 1980s. In 1988, Japanese imports were 85% from the U.S. and 9% from Canada. Sockeye was the primary species imported with a share of about 85%, followed by coho.

After several technological breakthroughs, salmon farming became a viable commercial sector during the 1980s. As the pioneers were European, the preferred species was Atlantic salmon, although operators quickly started farming salmon trout and (in the Pacific) coho, targeting the Japanese market.[9] These species are suitable for Japanese tradition because of their deeply red flesh.[10] In the early 1990s, Japanese imports for these species increased rapidly. Of the imported high-valued species, the market shares in 2000 were 35% for salmon trout, 34% for farmed coho, and 31% for sockeye. The market share for imported farmed coho and salmon trout were close to zero as late as 1990.

World salmon prices have decreased substantially over the period. While this is primarily due to productivity growth and increased production of farmed salmon (Tveteraas, 2000), there have been a number of hypotheses for the causes of price declines for wild salmon. In Alaska the oil spill of the Exxon Valdez in 1989 and alleged collusive behavior among Japanese buyers in the early 1990s have been blamed for adverse effects on Alaskan salmon prices. How much of the price decline for Alaskan salmon can be blamed on the oil spill, collusion among Japanese buyers, and increased farmed salmon production is unknown. It is possible that the decrease in prices of farmed salmon, due to productivity gains, has influenced the development in Alaskan salmon prices. However, there was no significant market share of farmed salmon in the Japanese market until the mid-1990s. The objective of this analysis is to ask, given this newly-structured market for salmon in Japan, does the presence of farmed salmon in the Japanese market influence prices for wild salmon from Alaska?

In our empirical analysis, we use Japanese import data on a monthly basis from 1994 to 2000. The data contains import values and quantites for salmon trout, Chilean coho, and North

American sockeye.[11] Since all coho in Chile is farmed and virtually all production of farmed coho is done in Chile, this variable can also be labeled farmed coho. All production of salmon trout is farmed, while all North American sockeye are wild-caught.

5. EMPIRICAL APPLICATION

The first step in the empirical analysis is to investigate the time series properties of the data. In Table 1, the results from Augmented Dickey-Fuller (ADF) tests are reported.[12] As shown, all prices and the price index are nonstationary in levels, but stationary in first differences, while the shares and the real expenditure variable are stationary in their levels.[13] For these data to justify estimation of demand equations, we must first confirm that the prices are cointegrated. In Table 2 we show the results from bivariate cointegration tests using Johansen's test (Johansen, 1991). The tests indicate that all series follow the same stochastic trend. As noted above, Stock and Watson (1988) show that in a system with n data series integrated of the same order, there will be n-r cointegration vectors and r stochastic trends. When estimating this multivariate system we expect to find n-1 or two cointegration vectors.[14] However, as one can see from Table 3, the multivariate test indicates only one cointegration vector. We estimate relatively many parameters given the length of our data set. The power of the tests may therefore be relatively low as we are exposed to what Hendry (1995) labels the "curse of dimensionality." Hence, although the results are not entirely clear with respect to how many stochastic trends there are in this system of prices, it is clear that there is at least one cointegration vector. For our purpose this is sufficient, since the prices then will make up a stationary vector, and hence we can estimate the demand system.

To test for the GCCT, we must test for cointegration between the relative prices of the different species with respect to the price index. The first step is then to investigate the time series properties of the relative prices. These are conducted with an ADF test without a trend and the results are reported in Table 4. All relative prices seem to be stationary. Accordingly, they cannot be cointegrated with the price index as this was shown to be nonstationary in Table 1. Hence, we must conclude that the GCCT holds for this group of goods, and we can justify looking at salmon separately from other goods in the consumers' budget.

We proceed to estimate the demand system with the equation for sockeye deleted to avoid singularity of the covariance matrix. The estimated parameters are reported in Table 5 together with goodness-of-fit information and autocorrelation tests for the different equations. The explanatory power of all equations is good, and there is no evidence of dynamic misspecification. A trend parameter in the demand equations can be interpreted as a systematic shift in tastes (Deaton and Muellbauer, 1980b). Our results indicate that there is a positive trend for farmed salmon and salmon trout, while the trend is negative for the wild-caught salmon. There also exist monthly differences in demand, with the largest budget shares of sockeye occurring during July through November, and for farmed coho in January through June.

In Table 6, the elasticities are reported. The first section of the table reports the compensated elasticities. These are all statistically significant, and indicate that all types of salmon are strong substitutes. The strong substitution effects seem to be reduced when the expenditure effects are taken into account, as several of the uncompensated cross-price elasticities then become statistically insignificant and basically zero.[15] As one can see the expenditure effects are also relatively strong, with salmon trout relatively inelastic, and the two remaining products expenditure elastic. Still the elasticities give substantial evidence that salmon trout competes with sockeye.

6. CONCLUDING REMARKS

There is substantial evidence in favor of treating price series that are commonly employed as right hand side variables in demand systems as nonstationary. There are also strong arguments for treating expenditure shares as stationary. With this structure in the data, it is generally not valid to estimate share equations, which is the most common specification used in demand systems, as there will not be a long-run relationship between data series that are integrated of different orders. In this paper we show that one can estimate a demand system if the shares are stationary and prices are nonstationary but cointegrated. In this case, ordinary inference theory will also apply. Hence, the traditional approach where one estimates the system as a seemingly unrelated regression is also a valid approach with this structure on the data.

Another common problem in applied demand analysis with disaggregated data is to justify why one treats goods in question as a separate group. Normally a weak separability assumption is made. However, one needs additional data to validate this assumption. In this paper we argue that an alternative is to use the Generalized Composite Commodity Theorem of Lewbel (1996). To test whether this theorem holds for a group of goods, one only needs the data that are used when estimates a demand system. Hence, the GCCT can be used to easily validate that treatment of the goods in question as a separate group, provided that the theorem holds. The results from the demand analysis will then be more reliable since we cannot argue that the aggregation assumption is invalid.

We use these tools to investigate the extent of competition in the Japanese import market between wild-caught North American salmon and farmed salmon and trout. The budget shares are stationary, while the prices are not. However, the prices are cointegrated so that the share equations can be estimated with traditional tools. We also find that the GCCT holds, and accordingly one can treat the three goods as a separate group.

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Table 1. Dickey-Fuller tests

Variable	Test statistic, levels	Test statistic, first differences
Trout price	-1.951	-3.820*
Coho price	-1.878	-4.538*
Sockeye price	-1.841	-4.520*
Trout share	-4.575*	
Coho share	-7.405*	
Sockeye share	-8.155*	
Expenditure	-5.516*	

* indicates significant at a 5% level. Critical value at a 5% level is –3.467.

Table 2. Bivariate Johansen tests on prices

Variables in the test	H_0:Rank=p	Max	Trace	Autocorrelation[a]
Coho and trout	p=0	30.73*	37.21*	1.618 (0.108)
	p<=1	6.473	6.473	0.992 (0.465)
Coho and sockeye	p=0	17.89*	25.95*	0.955 (0.499)
	p<=1	8.068	8.068	1.048 (0.417)

* indicates significant at a 5% level and ** indicates significant at a 10% level.
[a] *p*-values in parenthesis

Table 3. Multivariate Johansen test on prices

H_0:Rank=p	**Max**		**Trace**	
	Test statistic	Critical value at a 5% level	Test statistic	Critical value at a 5% level
p=0	31.790*	22.0	45.590*	34.9
p<=1	8.955*	15.7	13.800	20.0
p<=2	4.842	9.2	4.842	9.2

* indicates significant at a 5% level.

Table 4. Dickey-Fuller tests, relative prices

Variable	Test statistic
Trout vs. Index	-3.047*
Coho vs. Index	-5.194*
Sockeye vs. Index	-3.438*

* indicates significant at a 5% level. Critical value at a 5% level is –2.899.

Table 5. Parameter estimates

Equation:	Trout		Sockeye	
Parameter	Estimate	St. Error	Estimate	St. Error
Trout price	-0.099	0.070	0.105	0.043
Coho price	-0.006	0.052	-0.036	0.034
Sockeye price	0.105	0.043	-0.069	0.048
Expenditure	-0.181	0.027	0.091	0.029
D1	-0.132	0.048	-0.081	0.052
D2	-0.299	0.048	-0.093	0.053
D3	-0.271	0.049	-0.148	0.055
D4	-0.222	0.048	-0.164	0.053
D5	-0.123	0.045	-0.131	0.049
D6	-0.144	0.044	-0.008	0.048
D7	-0.339	0.051	0.542	0.056
D8	-0.308	0.060	0.603	0.066
D9	-0.357	0.049	0.579	0.054
D10	-0.191	0.047	0.398	0.051
D11	-0.028	0.045	0.164	0.048
Trend	0.002	0.000	-0.003	0.000
Constant	3.307	0.397	-1.078	0.438
R^2	0.852		0.927	
L-B(12)[a]	14.76		19.14	
p-value of L-B	0.255		0.085	

[a] L-B(12) is a Ljung-Box test against autocorrelation up to the 12th order.

Table 6. Elasticities[a]

	Equation:		
	Salmon trout	Farmed coho	Sockeye
Compensated elasticities			
Trout	-0.913	0.342	0.664
	(0.194)	(0.179)	(0.126)
Farmed coho	0.276	-0.566	0.188
	(0.145)	(0.178)	(0.099)
Sockeye	0.637	0.224	-0.852
	(0.121)	(0.117)	(0.139)
Uncompensated elasticities			
Trout	-1.094	-0.132	0.208
	(0.193)	(0.183)	(0.124)
Farmed coho	0.131	-0.948	-0.179
	(0.145)	(0.174)	(0.063)
Sockeye	0.464	-0.230	-1.289
	(0.130)	(0.128)	(0.150)
Expenditure elasticities			
	0.498	1.310	1.261
	(0.074)	(0.079)	(0.085)

[a] Standard errors in the parentheses

Endnotes

[1] A data series is said to be integrated of order p, denoted $I(p)$, if it becomes stationary after being differenced p times. A stationary data series has constant mean, variance and covariances depending only on the lag length not on time. If any of these moments is dependent on time the data series will be nonstationary.

[2] We will show that one can generally expect that with nonstationary nominal prices, the relative prices are nonstationary as well.

[3] However, the mean and variance of the shares need not be stable. Ng (1995), Attfield (1997) and Karagiannis and Mergos (2002) find shares in their AIDS model to be nonstationary.

[4] If the shares are treated as nonstationary, a cointegration framework must be used to confirm long-run relationships in the data and to provide valid statistical inference. For instance, Attfield (1997) use the Fully Modified Least Squares Estimator of Phillips and Hansen (1990) and Karagiannis and Mergos use an Error Correction Model. Ng (1995) investigates an AIDS model with nonstationary data equation by equation.

[5] This is true even though there are relationships between the demand equations through the symmetry restriction.

[6] The AIDS model and most other demand systems can be specified in terms of the relative prices $\ln p_j$-$\ln p_k$ when homogeneity is imposed. The system will then contain only stationary variables if the relative prices are stationary (possibly with exception of the expenditure term).

[7] It follows from Johansen and Juselius (1994) that only in the case of n prices and n-1 cointegration vectors, can the cointegration vectors can be normalized so that pairs of prices can be given stationary representations. However, for these pairs to be the relative prices, appropriate restrictions on the cointegration parameters must also hold.

[8] Most common functional forms for demand systems, including the AIDS, will aggregate consistently under the GCCT.

[9] The main producers of salmon trout are Norway and Chile, while Chile is virtually the only producer of farmed coho.

[10] Sockeye is the salmon species with the deepest red color, favored by Japanese consumers. However, sockeye are not as biologically feasible to farm on a commercial basis.

[11] North American sockeye is an aggregate of Alaskan and Canadian sockeye.

[12] The tests are reported with six lags. However, the results are, with a few exceptions, insensitive to the choice of lag length. For the exceptions, the deviations are on few lags, and when lag coefficients are statistically significant at higher lags.

[13] For completeness it is worthwhile to note that the relative prices of trout and sockeye with respect to coho are nonstationary as the ADF test statistics respectively are –2.632 and –2.134.

[14] The multivariate Johansen test was estimated with three lags. This seems sufficient to model the dynamics in the system as LM tests against up to 12th order autocorrelation in each equation gave the following test statistics with *p*-values in parentheses: Salmon trout, 0.561 (0.863); Coho, 1.305 (0.242) and Sockeye, 0.601 (0.832).

[15] This is not uncommon in applied work in particular when one is looking at disaggregated goods. See e.g. Chalfant (1987).