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An Empirical Investigation of Interproduct Relationships Between Domestic and Imported Seafood in the U.S.

Youngjae Lee and P. Lynn Kennedy

This study seeks to identify interproduct relationships between domestic catfish and a representative selection of imported seafood. In doing so, this study uses multivariate cointegration and structural analyses. Multivariate cointegration analysis suggests that six imported seafood product groupings form a common market with domestic catfish. Structural analysis reveals that 1) domestic and imported catfish are net and gross quantity substitutes; 2) domestic catfish and imported seafood are normal goods; 3) six imported seafood products are identified as gross quantity substitutes for domestic catfish; and 4) according to the derived Allais coefficients, interaction intensities of imported seafood for domestic catfish (from greatest to least) are as follows: tuna, shrimp, salmon, tilapia, catfish, and trout.

Key Words: catfish, multivariate cointegration, quantity substitutability, seafood imports, structural analysis

JEL Classifications: D12, F10, F11, F13

This study begins with two basic questions about the U.S. seafood market. The first question is "Do domestic and imported seafood belong to a common market?" The second question is "and if these two groups do belong to the same market, how do they compete against each other in that market?"[1] In this study, we seek to answer both questions using domestic catfish and a representative sampling of imported seafood by using suitable econometric techniques as discussed subsequently.

The underlying economic concept of cointegration is that market forces will prohibit persistent deviation from interproducts' long-run behavioral path in the market (Bose and McIlgorm, 1996; Dolado, Jenkinson, and Sosvilla-Rivero, 1990; Harris, 1995). For example, if the price of domestic catfish is considerably higher than the price of imported catfish, it would then be a reasonable assumption to suppose that U.S. seafood consumers would shift away from domestic catfish to imported catfish

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[1] Stigler's arbitrage-based definition of a market assumes that prices of close substitutes move together because arbitrage ensures that the law of one price holds for close substitutes. However, although it is implicitly assumed that substitution and arbitrage are the main determinants in delineating a market, none of the previous studies actually investigate the degree of substitutability. Hence, the relationship between market delineation and demand structure is still not properly addressed in the literature (Asche, Salvanes, and Steen, 1997).

subject to their budget constraint. Therefore, as a result of this shift in consumer demand, the price of domestic catfish will decline. This process should prohibit persistent long-run deviations in the equilibrium, although significant short-run deviations may occur. Consequently, cointegration analysis provides a suitable framework for analyzing the long-run price relationships between domestic catfish and various imported seafood products. Figures 1 and 2 show the short-run deviation between domestic catfish and six representative imported seafood groups used in this study from January 1989 to December 2007.

The quantity elasticity provides essential information needed in answering the second question (i.e., interproduct competition). In particular, cross-quantity elasticity will provide information that serves to identify interproduct relationships in the market. Because such elasticities can be obtained through a system of inverse demand equations rather than a single equation model, structural analyses would be required. Previous studies have developed the various specifications of inverse demand systems (Lee and Kennedy, 2008). These inverse demand systems have often been used to estimate quantity elasticities so as to assist in the identification of substitutability among seafood (Eales, Durham, and Wessells, 1997; Park, Thurman, and Easley, 2004).

This study was conducted as follows. In the next section, two analytical methodologies, cointegration and structural analyses, are reviewed from an empirical perspective. This study then discusses trends in U.S. seafood imports and reviews how these trends relate to market behavior. Section four discusses the empirical results obtained from implementation of the two underlying analytical methodologies. In the final section, the article concludes with a discussion of limitations and potential future research opportunities.

Analytical Methodology

Cointegration Analysis

The concept of cointegration expects time series variables to be nonstationary in behavior. It also expects that monthly time series might contain seasonal components. Therefore, it is a prerequisite to examine seasonal unit roots of the monthly price series for their inclusion in the cointegration analysis. For a seasonal unit root test, this study adopts the Hylleberg-Engle-Granger-Yoo test procedure developed by Hylleberg et al. (1990) as follows:

$$\begin{aligned}\phi(B)^{*}p_{13t} = &\sum_{k=1}^{12}\pi_k p_{k,t-1} + m_0 t + m_1 \\ &+ \sum_{k=2}^{12} m_k S_{kt} + \varepsilon_t,\end{aligned} \tag{1}$$

where p_{kt} is a polynomial of monthly price series in the backshift operator, t is a time

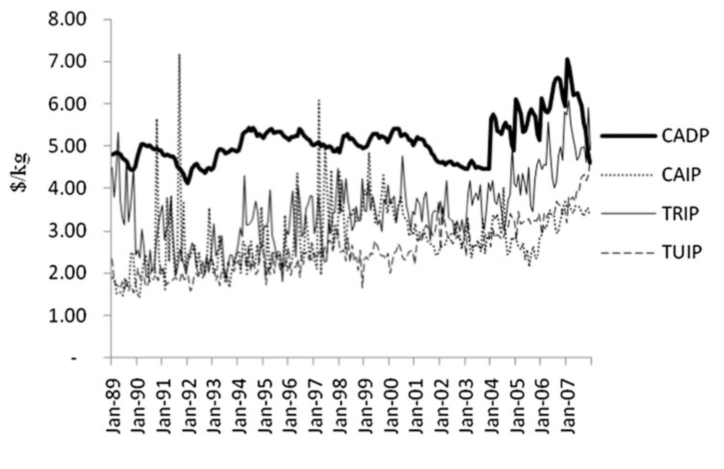

Figure 1. Long-Run Equilibrium and Short-Run Deviation Between the Prices of Domestic Catfish (CADP) and Imported Catfish (CAIP), Trout (TRIP), and Tuna (TUIP)

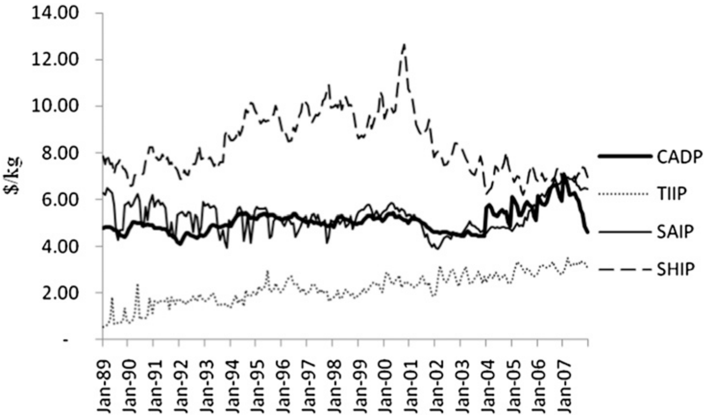

Figure 2. Long-Run Equilibrium and Short-Run Deviation Between the Prices of Domestic Catfish (CADP) and Imported Tilapia (TIIP), Salmon (SAIP), and Shrimp (SHIP)

trend, m_1 is a constant, S_k is a seasonal dummy, and ε_t is the white noise residual. To detect a seasonal unit root at frequency, π_k, we use the *t*-statistics obtained by Equation (1). In using the *t*-statistics, we test the null hypothesis, H_0: $\pi_k = 0$ (seasonal unit root). If the monthly price series shows a seasonal unit root at a specific frequency, then these series are nonstationary. The condition of nonstationarity is necessary for cointegration analysis.

To find the possible stationary linear combinations of nonstationary series, Johansen (1988) and Johansen and Juselius (1990) developed the vector autoregressive (VAR) procedure. The VAR procedure overcomes some problems that are encountered with a single equation procedure, e.g., arbitrary selection of the dependent variable and failure to identify the number of cointegrating vectors for the multivariate case (Bose, Bodmand, and Campbell, 2006; Engle and Granger, 1987). The VAR procedure is based on maximum likelihood within a Gaussian autoregression and not only allows a test of how many cointegrating relations there are in a given system, but also allows hypotheses tests regarding the space generated by the cointegration vectors (Johansen, 1988).

Johansen (1988) and Johansen and Juselius (1990) consider the following model:

$$\Delta p_t = \sum_{i=1}^{k} \Gamma_i \Delta p_{t-i} + \Pi p_{t-k-1} + \Phi S_t + \mu + \varepsilon_t, \tag{2}$$

where Δ is the first difference operator, p_t is a seven-dimensional vector of price variables with lag order $k + 1$, ε_t are the independent, normal innovations of the VAR process with mean zero and nonsingular, covariance matrix Λ, S_t are seasonal dummies, and μ is an intercept. Π is a 7×7 matrix of coefficients. Also $\Gamma_i = -(I - \Pi_1 - \ldots - \Pi_i)$ for $i = 1, \ldots, k - 1$, and $\Pi = -(I - \Pi_1 - \ldots - \Pi_k)$.

Equation (2) contains information on both short- and long-run adjustment to changes in p_t through the estimation of Γ_i and Π, respectively. The coefficient matrix Π contains information about the long-run relationships between the variables in the data vector. The rank of Π_k, r, determines how many linear combinations of p_t are stationary. If $r = N$, the variables in levels are stationary; if $r = 0$ so that $\Pi_k = 0$, none of the linear combinations are stationary. When $0 < r < N$, there exist r cointegration vectors. In this case, one can factorize Π_k; $-\Pi_k = \alpha\beta'$, where both α and β are ($N \times r$) matrices, β contains the cointegration vectors, and α contains the adjustment parameters.

Johansen (1988) and Johansen and Juselius (1990) show that it is possible to determine the number of significant vectors using the likelihood ratio (LR) test. Their proposed LR test for the hypothesis that there are at most "r" cointegrating vectors is given by:

$$LR = -T \sum_{i=r+1}^{N} \ln\left(1 - \hat{\lambda}_i\right), \tag{3}$$

where $\hat{\lambda}_{r+1} \ldots \hat{\lambda}_N$ are the *N-r* smallest squared canonical correlation coefficients between the residuals (u_{it} and v_{it} in Equations [4] and [5], respectively) obtained by first regressing $\Delta p_t (t = 1.2, \ldots, T)$ on its lagged differences as follows:

$$\Delta p_{it} = \alpha_{i0} + \beta_{i1}\Delta p_{it-1} + \beta_{i2}\Delta p_{it-2} + \cdots + \beta_{ik}\Delta p_{it-k} + u_{it}, \tag{4}$$

and then by regressing p_{t-k} on the same regressors as follows:

$$p_{it-k} = \alpha'_{i0} + \beta'_{i1}\Delta p_{it-1} + \beta'_{i2}\Delta p_{it-2} + \cdots + \beta'_{ik}\Delta p_{it-k} + v_{it}. \tag{5}$$

Here "i" represents a vector of N price variables. The asymptotic distribution of the LR test statistic is given by a multivariate version of the Dickey-Fuller distribution (Johansen and Juselius, 1990). Full details for Johansen's test as to its theoretical background and application are provided in Dickey and Rossana (1994).

Structural Analyses

Although multivariate cointegration analysis indicates the existence of an integrated market, long-run relation parameters, given by the cointegration regression, are difficult to interpret directly for the degree of substitutability among products because they are not based on a structural model (Asche, Salvanes, and Steen, 1997). This disadvantage of cointegration analysis will be supplemented by structural analysis. Based on utility maximization of a given budget constraint, a structural model provides the degree of substitutability among products by which consumers maximize their utility given income and can be measured when market conditions change. Previous studies have developed well-defined inverse demand systems such as the Differential Inverse Rotterdam Demand System (DIRDS), the Differential Inverse Central Bureau of Statistics (DICBS) demand system, the Differential Inverse Almost Ideal Demand System (DIAIDS), and the Differential Inverse National Bureau of Research (DINBR) demand system to estimate quantity elasticities. In this study, however, we emphasize the concept of how inverse demand systems can be used to obtain reliable estimates. In an inverse demand system, monthly data are more appropriate than annual or quarterly data because, as Eales, Durham, and Wessells (1997) discussed, when we model consumer demand with high-frequency time series data, it is possible that quantities consumed are predetermined.[2]

Now, let us summarize a representative seafood consumer's behavior as follows:

$$\max_{q} U(\boldsymbol{q}) \text{ s.t. } \textstyle\sum_i p_i q_i = 1. \tag{6}$$

As shown by Brown, Lee, and Seale (1995), the solution of Equation (6) leads to the Differential Inverse Generalized Demand System (DIGDS) as follows:

$$w_i d \ln p_i = (h_i - \theta_1 w_i) d \ln Q + \sum_j \left(h_{ij} - \delta_{ij}\theta_2 w_i + \theta_2 w_i w_j\right) d \ln q_j, \tag{7}$$

where θ_1 and θ_2 are nesting parameters, δ_{ij} is the Kronecker delta, and $d \ln Q = \sum_i w_i d \ln q_i$ is the Divisia volume index. The following restriction on parameters of Equation (7) hold:

$$\sum_i \left(h_{ij} - \delta_{ij}\theta_2 w_i + \theta_2 w_i w_j\right) = 0 \text{: adding up,} \tag{8}$$

$$\sum_i (h_i - \theta_1 w_i) = -1 \text{: adding up,} \tag{9}$$

$$\sum_j \left(h_{ij} - \delta_{ij}\theta_2 w_i + \theta_2 w_i w_j\right) = 0 \text{: homogeneity, and} \tag{10}$$

$$\left(h_{ij} - \delta_{ij}\theta_2 w_i + \theta_2 w_i w_j\right) = \left(h_{ji} - \delta_{ji}\theta_2 w_j + \theta_2 w_j w_i\right) \text{: symmetry.} \tag{11}$$

Scale and quantity elasticities at means, $q_i = \bar{q}$, can be derived from Equation (7). These elasticities are as follows:

$$\mu_i = h_i / w_i - \theta_1 \text{: scale elasticity,} \tag{12}$$

$$\eta^c_{ij} = h_{ij}/w_i - \delta_{ij}\theta_2 + \theta_2 w_j \text{: compensated quantity elasticity, and} \tag{13}$$

$$\eta_{ij} = \eta^c_{ij} + w_j \mu_i \text{: uncompensated quantity elasticity.} \tag{14}$$

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[2] We investigated the predeterminedness of monthly quantities supplied with a pair of Wu-Hausman tests; see Hausman (1978), and Thurman (1986), and Wu (1973) for discussions of the tests. The test statistics could support the null hypothesis of predetermined quantities.

The other nested models can be obtained by restricting θ_1 and θ_2 appropriately in Equation (7) as follows:

$$w_i d \ln p_i = h_i d \ln Q + \sum_j h_{ij} d \ln q_j \tag{15}$$
DIRDS for $\theta_1 = 0$ and $\theta_2 = 0$,

$$w_i d \ln \frac{p_i^*}{P} = c_i d \ln Q + \sum_j h_{ij} d \ln q_j \tag{16}$$
DICBS for $\theta_1 = 1$ and $\theta_2 = 0$,

$$dw_i = c_i d \ln Q + \sum_j c_{ij} d \ln q_j \tag{17}$$
DIAIDS for $\theta_1 = 1$ and $\theta_2 = 1$,

$$dw_i - w_i d \ln Q = h_i d \ln Q + \sum_j c_{ij} d \ln q_j \tag{18}$$
DINBR for $\theta_1 = 0$ and $\theta_2 = 1$,

where $c_i = h_i + w_i$, $c_{ij} = h_{ij} + \delta_{ij} w_i - w_i w_j$, $d \ln P = \sum_i w_i d \ln p_i$ is the Divisia price index, and p_i^* is the non-normalized price.

The conditions of adding up, homogeneity, and symmetry for these four nested models can be directly redefined by restricting parameters θ_1 and θ_2 in Equations (8), (9), (10), and (11). It is also straightforward to obtain scale elasticities and compensated/uncompensated quantity elasticities for DIRDS, DICBS, DIAIDS, and DINBR models by restricting θ_1 and θ_2 in Equations (12), (13), and (14).

U.S. Seafood Imports

In 2006, U.S. imports of seafood were valued at \$13.4 billion, amounting to \$6.7 billion more than 1996 seafood imports, implying that a 100% increase in the value of seafood imports took place over the past decade. The volume of seafood imports was 2.45 million metric tons, representing an increase of 1.01 million metric tons from 1996. This increase amounts to a 70% increase in import volume over the same 10-year span. Because the rate of increase for import value is greater than import volume, imported price has increased from \$4.65/kg in 1996 to \$5.47/kg in 2006. In 2006, seafood imports consisted of 2 billion kilograms of fresh and frozen products valued at \$11.7 billion, 328 million kilograms of canned products valued at \$1.3 billion, 40 million kilograms of cured products valued at \$206.5 million, 3.3 million kilograms of caviar and roe products valued at \$32.4 million, and 24 million kilograms of other products valued at \$119.4 million, respectively.

From 1996 to 2006, the amount of U.S. seafood imports has increased continuously with relatively little fluctuation in total volume and/or value. Shrimp imports were \$4.1 billion and 0.590 million metric tons in 2006, representing increases of 67% for value and 123% for quantity from 1996. The unit price of imported shrimp decreased from \$9.30/kg to \$6.97/kg during this period of time (1996–2006). Shrimp imports accounted for 31% of the value and 24% of the quantity of total seafood imports in 2006. Salmon imports were \$1.5 billion and 0.242 million metric tons in 2006, representing increases of 278% for value and 190% for quantity from 1996. Unlike that of shrimp, the unit price of imported salmon increased from \$4.93/kg to \$6.43/kg during this 10-year period of time. Salmon imports accounted for 11% of the value and 10% of the quantity of total seafood imports in 2006. Tuna imports were \$0.9 billion and 0.275 million metric tons in 2006. Although the value of tuna imports increased by 48% during this 10-year period, the quantity of tuna imports was steady or had slightly decreased. Consequently, the unit price of imported tuna increased from \$2.28/kg to \$3.39/kg during this period. Tuna imports accounted for 7% of value and 11% of quantity of total seafood imports into the U.S. in 2006.

Although the import quantity and value of catfish, trout, and tilapia are relatively small compared with shrimp, salmon, and tuna, the U.S import growth rates of catfish, trout, and tilapia were much larger than those of shrimp, salmon, and tuna during this sample period of time. The U.S. import trends of catfish, trout, and tilapia are shown to be similar to that of salmon imports, which represent both import value and quantity increase but the increase in value is greater than the increase in quantity so that the unit price of imports have increased during this time period. Catfish imports were \$111 million and 34,000 metric tons in 2006, representing increases of 35-fold for value and

30-fold for quantity from 1996. Tilapia imports were $483 million and 158,000 metric tons in 2006, representing increases of 11-fold for value and eightfold for quantity from 1996. Trout imports were $22 million and 6.9 thousand metric tons in 2006, representing increases of 313% for value and 225% for quantity from 1996.

Empirical Results

Data Description

Our analysis includes five fin fish species (catfish, trout, tuna, tilapia, and salmon) and one crustacean species (shrimp). We obtained monthly price and quantity data for each of these products from January 1989 to December 2007 from different sources. Price and quantity data for domestic catfish (round weight processed) come from the National Agricultural Statistics Service. Quantity and value data for imported seafood are obtained from the National Marine Fisheries Service. The unit prices of imported seafood are obtained by dividing the total value by volume of imports. The obtained quantity and price data represent an actual quantity amount (i.e., kilograms) and actual price (i.e., $/kg). Before using these actual data in both cointegration and structural analysis, we normalize price and quantity data by following the method as suggested by Lee and Kennedy (2009). The descriptive statistics for both the normalized budget share and quantity are summarized in Table 1.

Cointegration Analysis

Table 2 presents the results of the seasonal unit root tests for the individual price variable. We use log normalized price series, $\ln p_i$, rather than normalized price series, p_i, to test for seasonal unit roots of the seven individual monthly price series in Equation (1). The regression equations include an intercept, time trend, and 11 seasonal dummy variables. To test individual seafood products for seasonal price series unit roots at frequency, π_k, we used t-statistics that were obtained from Equation (1). Based on the test results, we reject the null hypothesis of seasonal unit roots for nonzero frequency, $\pi_{k\neq1}$, at the 1% level, but we fail to reject the null hypothesis in the case of zero frequency, π_1, because all π_1 statistics are greater than the critical value. We also use the F-test suggested by Beaulieu and Miron (1993). The test results strongly reject the null hypothesis because all the calculated F-values are higher than the critical values along with the fact that π_2 and at least one member of each of the subsets of test statistics $\{\pi_{odd},\pi_{even}\}$ are significantly different from zero. Thus, overall test results indicate that the price series do not contain any seasonal unit roots at any seasonal frequency other than zero. Therefore, these series are nonstationary (a necessary condition for cointegration analysis).

This study uses the LR test to determine the number of cointegration vectors among seven fish price series. Test results are reported in Table 3. Test results indicate that there are six

Table 1. Shares and Variation in Budget and Quantity of Seven Seafoods: January 1989 to December 2007

	Budget Share			Quantity Share		
Type	Mean	Minimum	Maximum	Mean	Minimum	Maximum
Catfish (D)[a]	0.21848	0.11464	0.30966	0.22546	0.12854	0.31868
Catfish	0.00270	0.00000	0.02381	0.00497	0.00001	0.03965
Trout	0.00170	0.00039	0.00447	0.00267	0.00051	0.00868
Tuna	0.13138	0.06189	0.33388	0.28160	0.11767	0.59017
Tilapia	0.01947	0.00011	0.07559	0.04033	0.00047	0.13569
Salmon	0.11324	0.04980	0.23292	0.11531	0.03774	0.21596
Shrimp	0.51304	0.35346	0.65401	0.32966	0.17373	0.48617

[a] (D) represents domestic product.

Table 2. Results of Tests for Seasonal Unit Roots in Monthly Aggregate Series[a]

	0	π	π/2		2π/3		π/3		5π/6		π/6		π/2	2π/3	π/3	5π/6	π/6
	π_1	π_2	π_3	π_4	π_5	π_6	π_7	π_8	π_9	π_{10}	π_{11}	π_{12}	$F_{3,4}$	$F_{5,6}$	$F_{7,8}$	$F_{9,10}$	$F_{11,12}$
p_{cd}	3.37	−5.54	−2.65	−5.48	−4.85	3.54	−0.44	−5.45	−5.80	1.75	0.34	−10.55	20.05	19.12	15.74	17.90	57.51
p_{ca}	−1.65	−4.32	−7.13	−0.08	−6.67	−0.99	−4.54	0.76	−6.29	0.67	−4.58	−2.63	25.42	23.73	10.63	19.80	16.89
p_{tr}	−1.81	−5.08	−5.93	−3.65	−6.79	2.16	−6.22	−1.24	−6.89	1.15	−2.45	−4.74	26.45	25.64	20.33	24.00	17.84
p_{tu}	−1.82	−3.20	−4.74	−1.67	−4.67	−0.23	−4.49	−2.38	−4.11	−0.70	−3.55	−4.22	13.03	11.07	14.00	9.23	18.95
p_{ti}	−0.65	−3.81	−5.96	−4.73	−5.27	1.06	−3.23	−2.68	−6.01	4.11	−3.91	−4.55	33.26	13.96	10.76	27.68	23.55
p_{sa}	4.41	−5.33	−4.22	−4.66	−5.29	3.64	−3.56	−5.98	−6.77	2.35	0.91	−7.62	21.72	21.67	28.88	25.37	29.28
p_{sh}	3.68	−5.01	−4.69	−5.84	−5.36	4.81	−1.88	−6.49	−4.90	4.27	0.68	−7.95	31.91	29.06	25.06	21.92	32.44

[a] 1 = $p_{cd}, p_{ca}, p_{tr}, p_{tu}, p_{ti}, p_{sa}$, and p_{sh} represent the prices of domestic catfish, and imported catfish, trout, tuna, tilapia, salmon, and shrimp, respectively; 2 = seasonal unit roots are tested in log levels of prices; 3 = the estimation equations include a constant, 11 seasonal dummies, and a time trend; 4 = standard errors are OLS standard errors; 5 = critical values of the test statistics are given in Table A.1 of Beaulieu and Miron (1993, pp. 325–26).

Table 3. Results of the Likelihood Ratio Test

Eigenvalue	Likelihood Ratio	Ho:r	F-Value	$Pr > F$
1.0077	0.1639	1	8.46	0.0001
0.5474	0.3291	2	6.73	0.0001
0.4427	0.5093	3	5.69	0.0001
0.1641	0.7348	4	3.92	0.0001
0.0997	0.8553	5	3.48	0.0004
0.0624	0.9406	6	3.03	0.0175

cointegrating vectors because the null hypothesis, $r = 6$, cannot be rejected at the 1% level of significance. This finding implies that all imported seafood products considered within this study comprise a common market with domestic catfish.

Quantity Substitutability

To measure the degree of quantity substitutability among these cointegrated seafood products, this study conducts structural analysis. We first seek to identify what specification of the inverse demand models best fit the data used in this study. To do this, we test the nesting parameters, θ_1 and θ_2, estimated using the DIGDS model. In doing so, Equation (7) is modified as an empirical form as follows:

$$
\begin{aligned}
\bar{w}_{it}\Delta \ln p_{it} &= h_i \Delta \ln Q_t + \sum_j h_{ij}\Delta \ln q_{jt} \\
&\quad - \theta_1 \bar{w}_{it}\Delta \ln Q_t \\
&\quad - \theta_2 \bar{w}_{it}\Delta \ln(q_{it}/Q_t) + \alpha_{i+\varepsilon^{it}},
\end{aligned} \tag{19}
$$

where $\Delta \ln p_{it} = \ln p_{it} - \ln p_{it-12}$ represents seasonally adjusted series, $\bar{w}_{it} = \frac{w_{it} + w_{it-12}}{2}$ represents 2-year moving average of monthly budget share, and α_i is constant. The estimated nesting parameters are tested according to the restrictions in Equations (15) through (18). The test results, reported in Table 4, show that the DIGDS model could not be reduced to one of the nested models. These results are similar to those found by other previous studies in quite different empirical applications (Brown, Lee, and Seale, 1995; Eales, Durham, and Wessells, 1997; Park, Thurman, Easley, 2004).

Following Lee and Kennedy (2008), we test the statistical validity of the models directly

Table 4. Restrictions on the Generalized Model That Yields Alternative Functional Forms

	Restrictions		Test Results	
Models	θ_1	θ_2	F-Value	$Pr > F$
DIRDS	0	0	65.23	0.0001
DIAIDS	1	1	3144.22	0.0001
DICBS	1	0	304.66	0.0001
DINBR	0	1	2698.34	0.0001

DIRDS = Differential Inverse Rotterdam Demand System; DIAIDS = Differential Inverse Almost Ideal Demand System; DICBS = Differential Inverse Central Bureau of Statistics; DINBR = Differential Inverse National Bureau of Research.

using Equations (15) to (18) and did not detect a superior model among the four different models. However, this study finds that the negativity condition is satisfied in the DIRDS, DICBS, and DIGDS models.

In the second step, we seek to estimate scale and quantity elasticities. In doing so, we used the previous test results. For example, because the negativity condition is satisfied in the DIRDS, DICBS, and DIGDS models (Barten and Geyskens, 1975), we determine whether these three models satisfy the negativity of scale effect. Among these three models, DIRDS is perfectly satisfactory in this regard.

Based on this result, this study uses the DIRDS model to estimate elasticity coefficients. In the estimation procedure, we impose homogeneity and symmetry restrictions on the econometric regressions of the model. Although the DIRDS model using these data did not show problems related to singularity in the residuals matrix, we delete one equation in the system to ensure freedom from singularity in the residuals matrix obtained from the SUR (Seeming Unrelated Regression) econometric model. Equation (15) is slightly modified for estimation as follows:

$$(20) \quad \bar{w}_{it}\Delta \ln p_{it} = h_i \Delta \ln Q_t + \sum_j h_{ij}\Delta \ln q_{jt} + \alpha_{i+\varepsilon^{it}}.$$

After estimating elasticity coefficients, h_i and h_{ij}, we calculate scale elasticity, compensated quantity elasticity, and uncompensated quantity elasticity using Equations (12), (13), and (14), respectively.

Finally, we estimate the Allais coefficients to measure the intensity of substitutable interaction among these seafood products by using the following equations:

$$(21) \quad a_{ij} = h_{ij}/w_i w_j - h_{rs}/w_r w_s + (h_i/w_i - h_r/w_r) + (h_j/w_j - h_s/w_s),$$

and

$$(22) \quad \alpha_{ij} = a_{ij}/\sqrt{(a_{ii}a_{ij})},$$

where subscripts r and s refer to some standard pair of goods r and s and are included so we can compare the relative strength of substitutability between the pair i and j and the standard pair r and s (Barten and Bettendorf, 1989).

Table 5 presents quantity and scale elasticity coefficients estimated by the DIRDS model with statistical results of R^2 for the system model's goodness-of-fit and t-statistics. The system R^2 is 0.9545 for the DIRDS model, indicating that this model explains the variation of the price-dependent variable. The t-statistics show that most estimated quantity and scale elasticity parameters are significantly different from zero at the 5% level. The own quantity elasticity parameters are negative and significantly different from zero at $\alpha = 0.05$ except for imported tuna. Of 21 cross-quantity elasticity parameters, nine are significantly different from zero at $\alpha = 0.05$. All seven scale elasticity parameters are significantly different from zero at $\alpha = 0.05$.

The elasticity coefficients have been transformed into scale and quantity elasticities using Equations (12) to (14), in which $\theta_1 = 0$ and $\theta_2 = 0$. Table 6 shows compensated quantity elasticities and scale elasticities. Compensated quantity elasticities represent net effect of quantity on price, whereas scale elasticities represent effect of expenditure on price in this system. The results show three interesting facts related to U.S. seafood imports.

First, results show that the estimated compensated quantity elasticities are very inelastic. Tomek and Robinson (1990) define the relationship between quantity and price elasticities as the inverse relationship, which is defined as follows:

Table 5. Quantity and Scale Elasticity Coefficients Estimated by DIRDS

	Catfish (D)	Catfish	Trout	Tuna	Tilapia	Salmon	Shrimp	dlnQ
Catfish (D)	−0.0174	−0.0007	−0.0023	−0.0002	0.0041	0.0052	0.0114	−0.2179
	(0.04)	(0.04)	(0.22)	(0.02)	(0.00)	(0.24)	(0.21)	(0.00)
Catfish		−0.0001	0.0001	0.0000	−0.0001	0.0001	0.0007	−0.0023
		(0.00)	(0.36)	(0.39)	(0.09)	(0.68)	(0.03)	(0.00)
Trout			−0.0003	0.0001	0.0000	0.0003	0.0006	−0.0018
			(0.00)	(0.58)	(0.37)	(0.09)	(0.00)	(0.00)
Tuna				0.0000	0.0000	0.0001	0.0001	−0.0019
				(0.52)	(0.21)	(0.03)	(0.14)	(0.00)
Tilapia					−0.0018	0.0003	−0.0017	−0.0205
					(0.00)	(0.71)	(0.29)	(0.00)
Salmon						−0.0220	0.0164	−0.1231
						(0.00)	(0.01)	(0.00)
Shrimp							−0.0284	−0.5010
							(0.02)	(0.00)

System $R^2 = 0.9545$. The numbers in parentheses are the *p* values.
DIRDS = Differential Inverse Rotterdam Demand System.

$$(23)\qquad F_{ii} = \frac{1}{E_{ii}} = \frac{\Delta P_i}{\Delta Q_i} \times \frac{Q_i}{P_i}, \text{ and}$$

$$(24)\qquad F_{ij} = \frac{1}{E_{ji}} = \frac{\Delta P_i}{\Delta Q_j} \times \frac{Q_j}{P_i},$$

where F_{ii} and F_{ij} represent own- and cross-quantity elasticities and E_{ii} and E_{ji} represent own- and cross-price elasticities between goods i and j. Thus, if own-quantity elasticity is less than one, in absolute value, own good demand is elastic. If the cross-quantity elasticity, F_{ij}, is less than one in absolute value, demand for good j is greatly influenced by a small change in the price of good i. Therefore, the results of this study indicate that demand for these seafood products is very elastic and the cross-price effect on quantity will be greater than one. This result is consistent with those of Barten and Bettendorf (1989), Lee and Kennedy (2008), and Park, Thurman, and Easley (2004). However, the quantity elasticity in empirical demand models implies that price is a function of the quantity of the particular product as well as the quantities of substitutes. In contrast, the usual demand function makes quantity a function of the price of the product as well as other products' prices. Because different variables are held constant in the equations, the reciprocal of the quantity elasticity is not always a good approximation of the price elasticity (Huang, 1994, 1996; Eales, 1996). As Houck (1965) indicated, the reciprocal of the quantity elasticity equals the price elasticity only if the cross-quantity elasticities are zero.

Second, two goods i and j are net quantity complements if $\eta_{ij}^c > 0$ and net quantity substitutes if $\eta_{ij}^c < 0$. Consistent with expectations, domestic and imported catfish are net substitutes.[3] Trout and tuna are also net quantity substitutes for domestic catfish, whereas tilapia, salmon, and shrimp are net quantity complements for domestic catfish.

Finally, for a normal good, a change in quantity has a negative scale effect, i.e., $\mu_i < 0$. Results reveal that domestic catfish and the other imported seafood products are normal goods.

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[3] Ligeon, Jolly, and Jackson (1996) demonstrated decreasing quantities of imported catfish if domestic prices decrease relative to import prices.

Table 6. Compensated Quantity and Scale Elasticities

	Catfish (D)	Catfish	Trout	Tuna	Tilapia	Salmon	Shrimp	dlnQ
Catfish (D)	–0.080	–0.003	–0.011	–0.001	0.019	0.024	0.052	–0.998
Catfish	–0.252	–0.033	0.026	0.002	–0.048	0.035	0.272	–0.857
Trout	–1.378	0.041	–0.171	0.031	0.025	0.148	0.346	–1.083
Tuna	–0.002	0.000	0.000	0.000	0.000	0.001	0.001	–0.015
Tilapia	0.209	–0.007	0.002	–0.001	–0.091	0.018	–0.088	–1.053
Salmon	0.046	0.001	0.002	0.001	0.003	–0.194	0.145	–1.087
Shrimp	0.022	0.001	0.001	0.000	–0.003	0.032	–0.055	–0.977

Table 7 shows uncompensated quantity elasticities. Uncompensated quantity elasticity represents the gross quantity effect on price, which is the sum of net quantity and scale effects. Therefore, the uncompensated inverse demand of a normal good is more quantity elastic than compensated inverse demand. As seen in Table 7, the seven own uncompensated quantity elasticities are more elastic than those of own compensated quantity elasticities. Not only are imported catfish, trout, and tuna but also imported tilapia, salmon, and shrimp are gross quantity substitutes for domestic catfish, which is a result of the negative scale elasticities of these seafood products.

To calculate the Allais coefficients, this study selected imported catfish and imported trout as the standard pair of goods *r* and *s*, respectively. This selection causes all other Allais interactions to become negative, implying a stronger degree of substitution between the two commodities *i* and *j* as compared with the standard pair of goods *r* and *s* (imported catfish and imported trout). For example, the Allais coefficient between domestic and imported catfish is −0.786. This result implies that the substitutionary relationship between domestic and imported catfish is stronger than that between imported catfish and trout. Therefore, by comparing the magnitude of the coefficients, we can identify the intensity of substitutable interaction between these seafood products. For example, the Allais coefficient between domestic catfish and imported shrimp is –0.978, which is less than the coefficient between domestic and imported catfish (see Table 8). Therefore, we determine that shrimp is a stronger substitute for domestic catfish than is imported catfish. According to the results of this analysis, domestic catfish is the strongest substitute for its own good because the Allais coefficients between domestic catfish and each of the six imported seafood products are greater than –1. For domestic catfish, tuna, shrimp, and salmon display relatively strong interaction intensities, whereas imported catfish, trout, and tilapia show relatively weak interaction intensities.

Conclusions

The purpose of this study was to identify interproduct relationships between domestic catfish and six representative imported seafood products. In doing so, this study uses two different methodologies. At first, this study uses multivariate cointegration analysis to determine

Table 7. Uncompensated Quantity Elasticities

	Catfish (D)	Catfish	Trout	Tuna	Tilapia	Salmon	Shrimp
Catfish (D)	–0.298	–0.006	–0.012	–0.132	–0.001	–0.089	–0.460
Catfish	–0.439	–0.036	0.024	–0.110	–0.065	–0.062	–0.168
Trout	–1.614	0.038	–0.173	–0.111	0.004	0.026	–0.210
Tuna	–0.005	0.000	0.000	–0.002	0.000	–0.001	–0.007
Tilapia	–0.021	–0.010	0.000	–0.139	–0.112	–0.102	–0.628
Salmon	–0.192	–0.002	0.000	–0.142	–0.018	–0.317	–0.413
Shrimp	–0.191	–0.001	–0.001	–0.128	–0.022	–0.079	–0.556

Table 8. Allais Coefficients

	Catfish (D)	Catfish	Trout	Tuna	Tilapia	Salmon	Shrimp
Catfish (D)	–1.000	–0.786	–0.508	–0.991	–0.809	–0.924	–0.978
Catfish		–1.000	0.000	–0.739	–0.752	–0.686	–0.710
Trout			–1.000	–0.358	–0.291	–0.315	–0.346
Tuna				–1.000	–0.877	–0.950	–0.999
Tilapia					–1.000	–0.821	–0.881
Salmon						–1.000	–0.927
Shrimp							–1.000

whether domestic and imported seafood belong to a common market. Multivariate cointegration analysis is preceded by a seasonal unit root test. The multivariate cointegration analysis indicates that six imported seafood products form one common market along with domestic catfish.

Second, this study uses inverse demand systems to identify the degree of quantity substitutability. Consistent with a priori expectations regarding substitutability, domestic and imported catfish are actually net and gross quantity substitutes. The scale elasticities show that all seven seafood products considered in this study are normal goods. For domestic catfish, imported tilapia, salmon, and shrimp are net complements, whereas they are gross quantity substitutes.

Finally, this study calculates the Allais coefficients to determine the intensity of substitutable interaction between domestic catfish and a representative sampling of imported seafood. According to the results, imported salmon, tuna, and shrimp show strong quantity substitutability for domestic catfish. In contrast, imported catfish, trout, and tilapia show relatively weak quantity substitutability for domestic catfish.

One finding that should be noted in this study is that the U.S. domestic catfish industry can be more negatively influenced by imports of major seafood products such as salmon, tuna, and shrimp than from imports of catfish. However, few studies have had the ability to identify consumer behavior with respect to seafood consumption in the U.S. seafood market. Additional studies of this nature would aid policymakers by helping them to better understand the environment in which domestic seafood policy and trade policy interact.

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References

Asche, F., K.G. Salvanes, and F. Steen. "Market Delineation and Demand Structure." *American Journal of Agricultural Economics* 79(1997): 139–50.

Barten, A.P., and L.J. Bettendorf. "Price Formation of Fish: An Application of an Inverse Demand System." *European Economic Review* 33(1989):1509–25.

Barten, A.P., and E. Geyskens. "The Negativity Condition in Consumer Demand." *European Economic Review* 6(1975):227–60.

Beaulieu, J.J., and J.A. Miron. "Seasonal Unit Roots in Aggregate U.S. Data." *Journal of Econometrics* 55(1993):305–28.

Bose, S., M.P. Bodmand, and H.F. Campbell. "Estimation and Significance of Long-Run Relationships Among Fish Prices." *International Journal of Oceans and Oceanography* 1,2(2006):167–82.

Bose, S., and A. McIlgorm. "Substitutability Among Species in the Japanese Tuna Market: A Cointegration Analysis." *Marine Resource Economics* 11(1996):143–55.

Brown, M.G., J.Y. Lee, and J.L. Seale. "A Family of Inverse Demand Systems and Choice of Functional Form." *Empirical Economics* 20(1995):519–30.

Dickey, D.A., and R.J. Rossana. "Cointegrated Time Series: A Guide to Estimation and Hypothesis Testing." *Oxford Bulletin of Economics and Statistics* 56,3(1994):326–53.

Dolado, J.J., T. Jenkinson, and S. Sosvilla-Rivero. "Cointegration and Unit Roots." *Journal of Economic Surveys* 4(1990):249–73.

Eales, J. "A Further Look at flexibilities and Elasticities: Comment." *American Journal of Agricultural Economics* 78(1996):1125–29.

Eales, J.S., C. Durham, and C.R. Wessells. "Generalized Models of Japanese Demand for

Fish." *American Journal of Agricultural Economics* 79(1997):1153–63.

Engle, F.F., and C.W.J. Granger. "Co-integration and Error Correction: Representation, Estimation, and Testing." *Econometrica* 55,2(1987): 251–76.

Harris, R. *Using Cointegration Analysis in Econometric Modeling*. London: Prentice Hall, 1995.

Hausman, J.A. "Specification Tests in Econometrics." Econometrica 46,6(1978):1251–71.

Houck, J.P. "The Relationship of Direct Price Flexibilities to Direct Price Elasticities." *Journal of Farm Economics* 47(1965):789–92.

Huang, K.S. "A Further Look at Flexibilities and Elasticities." *American Journal of Agricultural Economics* 76(1994):313–17.

———. "A Further Look at Flexibilities and Elasticities." *American Journal of Agricultural Economics* 78(1996):1130–31.

Hylleberg, S., R. Engle, C.W.J. Granger, and H.S. Yoo. "Seasonal Integration and Cointegration." *Journal of Econometrics* 44(1990): 215–38.

Johansen, S. "Statistical Analysis of Cointegration Vectors." *Journal of Economic Dynamics & Control* 12(1988):231–54.

Johansen, S., and K. Juselius. "Maximum Likelihood Estimation and Inference on Cointegration with Application to the Demand for Money." *Oxford Bulletin of Economics and Statistics* 52(1990):169–210.

Lee, Y.J., and P.L. Kennedy. "An Examination of Inverse Demand Models: An Application to the U.S. Crawfish Industry." *Agricultural and Resource Economics Review* 37,2(2008): 243–56.

———. "A Demand Analysis of the Korean Wine Market Using an Unrestricted Source Differentiated LA/AIDS Model." *Journal of Wine Economics* 4,2(2009):185–200.

Ligeon, C., C.M. Jolly, and J.D. Jackson. "Evaluation of the Possible Threat of NAFTA on U.S. Catfish Industry Using a Traditional Import Demand Function. *Journal of Food Distribution Research* 27,2(1996):33–41.

Park, H.J., W.N. Thurman, and J.E. Easley. "Modeling Inverse Demand for Fish: Empirical Welfare Measurement in Gulf and South Atlantic Fisheries." *Marine Resource Economics* 19(2004):333–51.

Thurman, W.N. "Endogeneity Testing in a Supply and Demand Framework." Review of Economics and Statistics 68,4(1986):638–46.

Wu, D.M. "Alternative Tests of Independence Between Stochastic Regressors and Disturbances." Econometrica 41,4(1973):733–50.

Tomek, W.G., and K.L. Robinson. *Agricultural Product Prices*. 3rd ed. Ithaca, NY, London: Cornell University Press, 1990.