Testing for Market Integration and the Law of One Price in World Shrimp Markets

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ABSTRACT

Using import price data from Japan, United States, and the European Union, we test if shrimp price movements in these markets indicate an integrated world market for shrimp. We utilize cointegration techniques to investigate if prices in these markets share a common stochastic trend and if the law of one price holds. Results point to strong linkage among Japanese, American, and European markets. The results on the aggregated shrimp markets are checked against the results at a more disaggregated level. Data from wholesale markets in Tokyo, New York, and Europe for specific shrimp products confirm the integrated nature of shrimp markets. Evidence also exists in support of the law of one price in shrimp markets.

Keywords: Market integration; Law of One Price; Shrimp Prices; Cointegration

JEL Classification: Q11; Q22

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Introduction

Shrimp is the most valuable fishery traded in the international market. According to the Food and Agriculture Organization (FAO) it accounted for 18 percent ($10 billion) of total fish products exports in 2002. However, except for a number of studies that concentrate on specific regional markets (Japan and the U.S.) there is not a lot of research attention on the dynamics of shrimp prices in the world markets[1]. Studying the price dynamics of world shrimp markets is important as it comes at a time when U.S. producers of wild caught shrimp are looking for ways to have some control over the prices they receive in the domestic market. If the U.S. market is separate from other main markets in the world, then policies that aim to increase prices in the U.S. would benefit domestic producers, however, if world markets are highly integrated then polices that would increase prices in the U.S. cannot be sustained in the long-run as producers elsewhere in the world adjust their supply patterns.

An integrated and efficient market would not allow price differences to persist in the long-run, price movements in one market would result in adjustments to regain the equilibrium relationships among the prices observed in geographically separated markets. These adjustments may come from two sources: one, suppliers can adjust their behavior based on their profit maximizing goal by shifting supplies towards markets that pay higher prices and second, from the demand side consumers can switch to different sellers and

[1] Ling *et al.* (1998) examined price transmissions in vertically coordinated frozen black tiger shrimp markets, Traesupap *et al.* (1999) estimated Japanese shrimp supply and demand during the 1990s, Bene *et al.* (2000) investigated the impact of cultured shrimp industry on wild shrimp fisheries, and Gillig *et al.* (1998) modeled how volume prices of imported shrimp affected ex-vessel prices in the Gulf of Mexico.

substitute away from the more expensive suppliers. If shrimp markets are integrated, it puts into question the efficacy of polices that try to control prices in one domestic market. The main objective of this paper is to determine the price dynamics and the degree of market integration among the world's important shrimp markets. This paper is organized as follows: we start with a brief background on the main shrimp markets in the world, then a discussion of the empirical approach to modeling market integration, description of the data, discussion of the empirical results, and concluding remarks.

World Shrimp Trade Patterns

It is evident from Figure 1 that Japan and the U.S. are the dominant shrimp import markets in the world. The two countries accounted for 51 percent of total volume during the period 1984-2002. Spain was a distant third with an average share of seven percent followed by Denmark (5.2 %) and France (4.8 %). Taken together, the countries of the European Union (EU) as a region account for a considerable share of imported shrimp in the world. It is interesting to note that the share of Japan has been declining since the mid-1980s reflecting the slowdown in the Japanese economy during the period. In contrast, the U.S. share has been steady during the period and finally overtaking Japan in 1998. The demand for imported shrimp in Spain appears to have been increasing since 1984.

In terms of dollar value, Figure 2 shows Japan and the U.S. accounting for $6 billion of the estimated $10 billion imported shrimp in 2002. The sharp rise in the U.S. demand for imported shrimp is clear from the chart. US shrimp imports grew from $713 million in 1984 to $3.4 billion dollars in 2002, an average annual growth of 7.7 percent. Japan likewise experienced a rapid increase in the value of its imports, peaking in 1995 at $3.8 billion, declining thereafter.

The growth of shrimp imports in the rest of the countries in Figure 1 and 2 are also impressive despite the relatively smaller size of these markets compared to Japan and the U.S. In Spain for instance, imports have been growing at an annual rate of 22.9 percent during the period 1979-2002. Italy and the United Kingdom likewise recorded annual growth rates of 16.4 percent and 12.2 percent, respectively.

Empirical Approach

Cointegration Tests of Market Integration and LOP

In this paper we will apply the cointegration approach to testing market integration and the LOP[2]. Since the introduction of cointegration techniques in Engle and Granger (1987), Johansen (1988; 1991), Johansen and Juselius (1990; 1992) researchers have been conscious about applying the proper statistical procedures on nonstationary data. Cointegration allows for a characterization of stable long-run relationships for nonstationary series which lends itself well to analyzing price series that appear nonstationary but may share a common trend with other price series. The Engle and Granger method was applied first in the strand of literature that initially considered cointegration techniques in price based tests of market integration. Among these early papers are Ardeni (1989), Baffes (1991), and Goodwin and Schroeder (1991), and Diakosavvas (1995).

A shortcoming of the Engle and Granger method is it's basically a bi-variate approach that only accommodates relationships between two price series and thus would not lend itself well to analyzing multivariate systems that characterize for example markets with many sellers and buyers. Results of the Engle and Granger procedure are also sensitive to

[2] Dickey *et al.* (1991) and Hendry and Juselius (2000; 2001) provide detailed primers on how to conduct cointegration analysis and interpret the results.

which price series is used to normalize the other. Hypothesis testing on the estimated cointegration vector is likewise not possible under this approach.

The Johansen method is preferred over the Engle and Granger approach and has proven to be popular in the recent literature on market integration and LOP. Some of the early papers that utilized the Johansen technique include Goodwin (1992) who applied it to international wheat markets. Recently the cointegration based tests proposed in Asche *et al.* (1999) applied to world salmon markets have been applied to other fishery markets as well. Gordon *et al.* (1993) tested for market linkages in Paris fish markets; Gordon and Hannesson (1996) investigated price linkages in European and U.S. markets for cod fish; Asche *et al.* (2004) tested for market integration and LOP in the French whitefish market; and most recently Nielsen (2005) focused on price formation in the first-hand market for whitefish. Numerous other applications of the basic Johansen cointegration techniques as applied to LOP tests exist in the literature covering a diverse set of goods and industries.[3] This paper follows the basic methodology applied in the recent literature.

Econometric Specification

Since the cointegration approach applies to nonstationary series, the first step in the analysis is to test for nonstationarity of the price series. Variables that are nonstationary can be made stationary by differencing, the number of differencing (d) required to make the series stationary identifies the order of integration I(d). The Augmented Dickey-Fuller (ADF) test is employed here to determine the order of integration of each of the price series studied. The ADF test statistic is estimated by running a regression of the form:

[3] Cointegration based LOP and market integration tests are not without its critiques however. Barret (1996), McNew and Fackler (1997), Miljkovic (1999), and Miljkovic and Paul (2001) main critique is that cointegration is not necessary nor sufficient condition for market integration. The claim is otherwise integrated markets may not appear cointegrated if transactions and transportation costs are nonstationary. In addition, when supply and demand shocks are cointegrated across regions prices may exhibit a cointegrated behavior even if underlying markets are not integrated. In the shrimp market studied here there is no reason to believe that this is the case.

$$\Delta P_{it} = \alpha + \beta P_{it-1} + \delta t + \sum_{\gamma=1}^{k} \varphi_{\gamma} \Delta P_{it-\gamma} + \varepsilon_{t} \qquad (3)$$

Depending on the nature of the price series, the constant term α or the time trend δt can be omitted k represents the number of lags of the price variable to be included. The null hypothesis is that the particular price series is nonstationary. The test statistic is computed by dividing β by its standard error. The procedure is first done on the levels and then repeated for the first differences of the price series considered in this paper. Once the price series are confirmed to be integrated of degree one with the rejection of the null hypothesis on the first differences of the price series, they are deemed possible candidates for cointegration.

The Johansen cointegration test is based on a vector auto-regression (VAR) system. Given a price vector P_t it is carried out using the following representation:

$$P_{t} = \sum_{i=1}^{k} \prod_{i} P_{t-i} + \prod_{k} P_{t-k} + \mu + e_{t} \qquad (4)$$

The vector P_t contains the N price series to be tested for cointegration and assumed to be generated by the above unrestricted k_{th} order VAR in levels. Each of the $\prod_i$ is a $(N \times N)$ parameter matrix, μ is a constant term, and e_t is a $(N \times 1)$ vector of disturbances with a mean zero, covariance matrix of Ω, and i.i.d. normal over time. The system of equations represented in (4) can be written in error correction form as:

$$\Delta P_{t} = \sum_{i=1}^{k} \Gamma_{i} \Delta P_{t-i} + \Gamma_{k} P_{t-k} + \mu + e_{t} \qquad (5)$$

where $\Gamma_i = -I + \Pi_1 + \Pi_2 + ... + \Pi_i$ and i =1,…,k-1. The long-run solution to equation (4) is Γ_k. If P_t is a vector of first difference stationary price series, the left hand side of equation

(5) must be stationary, the first k-1 elements are I(0) as well. By assumption the error term in equation (5) are also stationary. Thus, either Γ_k is a matrix of zeroes or P_t contains a number of cointegrating vectors. The rank of Γ_k, r determines the number of stationary linear combinations of P_t. There are three possibilities: one, if $r = N$, the price variables are stationary in levels, second if $r = 0$, there exists no linear combination of P_t that are stationary, and third, when $0 < r < N$, there exists r stationary linear combinations of P_t. A rank of $r = N - 1$ in a multivariate system with N price series would imply that there is only one stochastic trend driving the behavior of prices in the system.

In the third case Γ_k could be factored such that $\Gamma_k = \alpha\beta'$, where α and β are both $N \times r$ matrices of rank r. The cointegrating vectors or the long-run relationships in the system are contained in the β matrix. The adjustment parameters on the other hand are identified in the parameters contained in α. There are two alternative tests that are used to identify the number of significant cointegrating vectors r, the trace test and the maximum eigenvalue test both of which are discussed in detail in Johansen (1995). The two tests have the null hypothesis that there are at most r cointegration vectors. The alternative hypothesis in the trace test is that there exist more than r cointegration vectors while for the maximum eigenvalue test, the alternative hypothesis is that there are exactly $r + 1$ cointegration vectors.

The main advantage of the Johansen approach in testing for market integration and the Law of One Price is that it allows hypothesis testing on the coefficients of both α and β using likelihood ratio tests as outlined in Johansen and Juselius (1990). If we are interested in testing for the Law of One Price, restrictions can be placed and tested on the parameters in the β matrix. In the case of a bivariate system where two price series are examined, the rank of $\Pi = \alpha\beta'$ would be equal to 1 and the dimensions of α and β matrices would be

2×1. LOP is tested by imposing the restriction $\beta' = (1, -1)'$. Since the matrix β contains long-run parameter in the system the test can be considered a test of the validity of LOP as a long-run concept. The LOP test can be easily extended to multivariate cases. Assuming that the rank of the multivariate system is $N-1$, the LOP test becomes a test of whether the columns in the β matrix sum to zero. The β matrix in a multivariate test with four price series can be represented as follows:

$$\beta' = \begin{bmatrix} 1 & -\beta_1 & 0 & 0 \\ 1 & 0 & -\beta_2 & 0 \\ 1 & 0 & 0 & -\beta_3 \end{bmatrix} \tag{6}$$

Also of interest are the adjustment parameters contained in α. The adjustment parameters are related to the concept of weak exogeneity. If all adjustment parameters for one variable are zero, then this variable is said to be weakly exogenous to the long-run parameters in the remaining equations. This implies that the coefficients on the levels of the remaining price series in the system is zero in this particular equation which would mean other price variables are not influencing this variable in the long-run. The implication of testing for weak exogeneity of certain price series in the system is the ability to identify possible market and price leadership from the direction of causality that the adjustment parameters provide.

Data

Two datasets were used in this study. The first dataset covers a more aggregated product group at the country and regional level and second, a more precise product classification in well identified market places. The first set of data covers the aggregate data from the three most important shrimp markets in the world. Price data on imported frozen shrimp (Harmonized Tariff Schedule (HTS) Code: 30613) as a product group for the European Union, Japan, and the United States were assembled[4]. The data were obtained from three main sources: Eurostat, Japan Ministry of Finance, and the U.S. International Trade Commission (ITC) Trade Dataweb.

Shrimp price data from Japan and the U.S. were available for the period July 1997 to June 2005. The data available online from Eurostat for the EU starts from January 1995 and ends in December of 2004. The unit prices which were reported in Yen and Euro for Japan and the E.U. were converted into U.S. dollar equivalents using foreign exchange data obtained from the International Monetary Fund's International Financial Statistics database. Following the usual practice in conducting price based tests of market integration and LOP, the logarithms of the price data were used for the empirical analysis.

A graphical summary of the price data from the three top consumers of imported shrimp is shown in Figure 3. Although there are observable differences in month to month movements, the three price series appear to share a similar pattern over a longer period of time. Prices in Japan appear to dominate both those recorded in the E.U and the United States. However, starting in 1998 prices in the U.S. has moved closer to Japan. The same trend appears to be true for E.U. prices which became closer to the U.S. and Japan prices starting in 2000. The observed price gaps however might not strictly reflect differences in prices paid for the same type of shrimp as the price series are at a highly aggregated level. Most probably the price differentials reflect different composition of frozen shrimp product category imported by the E.U., Japan, and the U.S. The higher prices in Japan could just be reflecting the fact that low count (bigger size shrimps) frozen shrimp accounts for a bigger proportion of its total imports when compared with the E.U. and the United States.

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[4] Specific prices for shrimp in disaggregated product categories up to ten digit HTS were available for the U.S. but the same detailed price data were not readily available for E.U. and Japan.

A second dataset was collected to confirm if similar trends are evident in more specific shrimp product categories. Data on wholesale shrimp prices by count size (i.e. 16-20 shrimps per pound category) in Tokyo, New York, and European wholesale markets were obtained from the Food and Agricultural Organization/Globefish Commodity Update (2001; 2004a) publications for shrimp. The data includes recorded Tokyo wholesale prices of white shrimp from India and Indonesia by different shrimp counts as well as black tiger from Indonesia. New York wholesale prices include those for brown shrimp from the U.S. Gulf of Mexico states, black tiger from Thailand, and white shrimp from Ecuador and Mexico. The price data in Europe include product categories of white shrimp from Ecuador. Price data from the Tokyo wholesale market are the most extensive with data reported from January 1989 to March 2004 while price data for Europe are much shorter covering only January 1989 to January 2000.

Figures 4, 5, and 6 show the general price trends in each product classification. Although there are short-term variations in the prices reported in the three price charts, the price series appear to follow the same long-term trend. It is clear from the charts that lower count shrimp products receive higher prices. Shrimp prices that are in the same market place and are of the same species move closer together indicating strong substitution possibilities within the same species regardless of country of origin. In the New York wholesale markets for shrimp size of 26-30 count per pound Mexican white shrimp receive higher prices when compared with Thailand black tiger and U.S. Gulf of Mexico brown shrimp. The higher price for Mexican shrimp might indicate a premium for quality. In panels where nearby count sizes where included we see that a single long-term price trend prevails across product size categories.

Empirical Results and Analysis

Stationarity of Shrimp Prices

Cointegration analysis requires that all variables in the system are integrated of the same order. Augmented Dickey Fuller (ADF) tests were first conducted in order to ensure that all price series are integrated of the same order. The results of the ADF tests following the general specification in equation (3) are reported in Table 1. Two specifications of the ADF test were used. The first included a constant while the second contained a deterministic trend in addition to the constant. The tests were undertaken in both the levels and first differences of the logarithms of the price series. Optimum lag lengths for each price series were chosen based on the Akaike Information Criteria.

In the three main world markets, we cannot reject the null hypothesis that shrimp prices in levels have unit roots. The results of the ADF tests on the first differences of these price series provide sufficient support for the alternative that the price series are stationary after first differencing. We can conclude that shrimp prices in the European Union, Japan, and the United States are integrated of the first order I(1).

The wholesale prices observed in Tokyo show evidence that the price series in levels are non-stationary and that stationarity is achieved after first differencing. In New York wholesale shrimp prices, ADF tests show nonstationarity at the levels. In three cases however when a trend was included in the basic ADF specification there was evidence at the 5% level that Mexican white shrimp prices are stationary. These findings however were not supported at the 1% level. In the European Union we find evidence of nonstationarity but in two specifications of the ADF test we find evidence for stationarity at the 5% level but again not supported at the 1% level[5]. In all the wholesale price series ADF tests on first differences of the prices indicated the absence of unit roots. As the price series in the two datasets in general appear to be integrated of the same order we can proceed with analyzing if the shrimp prices are cointegrated.

Cointegration Results for Europe, Japan, and the United States

The cointegration based test of market integration confirms the common trend visible in Figure 15. The test was carried out with one lag in the system based on the Akaike Information Criteria[6]. The results of the multivariate test are reported in Table 2 and it indicates that there are two cointegrating vectors or a single stochastic trend for all three prices. We confirm if each country and region contribute significantly to the long-run relationship by undertaking exclusion tests on the long-run parameter estimates of β. The null hypothesis is that parameter estimate for β_i is equal to zero where *i* could be any of the three countries or region. If the null hypothesis is supported it means that particular country or region *i* is not part of the market. Results of the tests confirm the robustness of the cointegration relationship[7]. This result indicates support for market integration in world shrimp markets as prices are confirmed to have a long-run relationship. However, the multivariate test for LOP was rejected for the system with a test statistic of 14.65 which is $\chi^2(2)$ distributed. Table 4 reports the results of weak exogeneity tests and there is evidence

[5] In the case of the Mexican and Indian shrimp prices in New York and Europe, respectively, an alternative stationarity test developed by Philipps and Peron (1988) was used and it indicated support for the presence of unit roots of these prices.
[6] The specification was satisfactory as Jarque-Bera tests indicated support for the null of multivariate normal residuals. The hypothesis of no serial correlation was likewise not rejected after a Lagrange multiplier (LM) test was conducted
[7] The three exclusion tests for the U.S., Japan, and Europe returned the respective test statistics of 23.87, 18.53, and 21.34 that are distributed $\chi^2(2)$ all indicating rejection of the null hypothesis at the 1% level.

that each of the three price series were endogenous confirming the linkages of these three market places.

Since the multivariate Johansen test indicated that the three prices share a common stochastic trend it must be that all prices in the system are pairwise cointegrated. Table 3 reports the results of pairwise cointegration tests. It confirms the findings in the multivariate test as each pair was shown to be cointegrated based on either the maximum eigenvalue test or the trace test. The results of the test for the LOP also gives us an indication why it fails in the system of three prices. LOP is supported for Japan and the United States but fails whenever Europe is in the pair. These results could be an artifact of how the price trends in each country or region was affected by aggregation across product types and imports source. In the next section we investigate if these results hold up when using a more disaggregated data set.

Cointegration Results for Wholesale Markets in the E.U., New York, and Tokyo

We use a second dataset covering wholesale markets to check the consistency of the results in the aggregated dataset. The wholesale market for the 41-50 count shrimp was first examined. The European Union data available for this product segment is brown shrimp from India. The E.U. price was matched with the wholesale prices recorded in Tokyo and New York. Table 5 reports the results of the cointegration and LOP tests conducted using the Johansen procedure.

There is evidence of cointegration in all possible combinations of the price series in the three wholesale markets. Furthermore, both maximum eigenvalue and trace tests confirm that there are two cointegrating vectors. The two cointegrating vectors imply a common stochastic trend and strong market integration among the wholesale prices observed in the three areas. The hypothesis that the Law of One Price prevails in the three markets was likewise supported. This outcome imply that LOP tests using aggregated price series might lead to incorrectly rejecting the hypothesis.

Exclusion tests on the long-run parameter estimates of β were conducted and indicate the robustness of the estimated cointegration relationships. The likelihood ratio tests for weak exogeneity of each of the price series show that Tokyo prices are weakly exogenous with respect to the other wholesale markets (Table 6). Weak exogeneity of Tokyo prices imply that Japan could be the primary marketplace in the world and that long-run prices elsewhere follow the lead of the Japanese market. In cases where shrimp from Mexico and Ecuador were included, weak exoegenity tests indicate that wholesale prices in New York for shrimp from these two countries might be insulated from price fluctuations from other wholesale markets. On the other hand it is clear that the European price series is not weakly exogenous in all combinations of wholesale prices in the three marketplaces.

Bivariate Cointegration and LOP Tests

Bivariate cointegration tests were conducted on the remaining price series. The results are summarized in Table 7. The market integration hypothesis is supported across all combinations of price series in each market segment. Long-run relationships in the price pairs are indicated by the presence of cointegrating vectors. An exception was the price pair N.Y.-Mexico white and Tokyo Indonesia white in the 16-20 count category where the price pair appear not to share a common stochastic trend. In a number of price pairs the weak exogeneity of wholesale prices in Tokyo was supported. The Law of One Price was also strongly supported in six cases, had weaker support in three out of the thirteen price pairs investigated. Price pairs that are composed of nearby product segments like 16-20 and 13-15 count shrimps also indicate that there is integration in both geographic and product space.

Concluding Remarks

In this paper we set out to investigate the price dynamics and the degree of market integration in world shrimp markets. The answers to these research questions are important in light of the expectations of U.S. domestic producers that trade sanctions against shrimp exporters will lead to higher prices in the U.S. market. The belief is justified if price determination in the U.S. shrimp market were based solely on demand and supply conditions in the U.S. However in integrated markets, arbitrage forces prices in different locations to follow a similar trend as each location adjusts to changing demand and supply conditions elsewhere.

The empirical results provide consistent evidence that the world's three main markets for shrimp are integrated. The long-run relationship among the prices holds despite the differences in the level of tariffs imposed by Europe (12 %), Japan (4 %), and the United States (0 %). Furthermore, despite the trade barriers, the Law of One Price holds in a variety of product segments. In fact in product segments where the Law of One Price is supported overwhelmingly, price information from one marketplace will reflect market conditions for the world market.

The implication of these findings is since the U.S. is part of a global shrimp market antidumping tariffs that will increase shrimp prices domestically will only have a short-term impact. Exporters will realign their supplies to flow into marketplaces where higher prices are offered. In the demand side, consumers also adjust their consumption and substitute away from higher priced imports to exporters that are willing to offer lower prices. The combination of these effects dampens the impact of targeted tariffs.

Finally, the observed strong linkage of U.S. domestic shrimp prices to prices around the world at present suggest a possible method to evaluate the success of future marketing efforts by domestic producers. The goal of marketing U.S. domestic shrimp as a premium product is to insulate it from the effects of price movements in other marketplaces. The methodology employed in this study can be used in future research to check if indeed marketing efforts has resulted in segmenting the domestic shrimp industry from the rest of the world.

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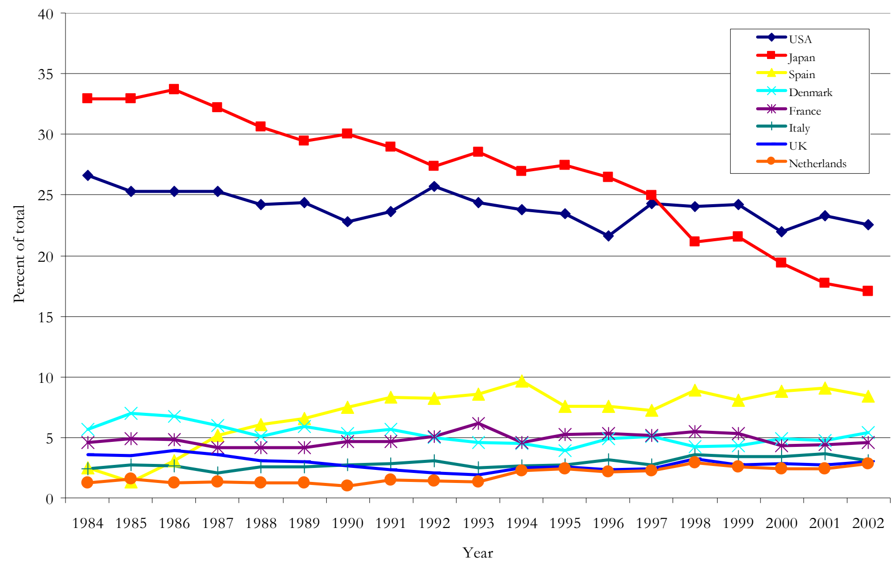

Figure 1. Annual Volume Shares Fresh, Chilled or Frozen Shrimp Imports by Major Importing Countries, 1984-2002

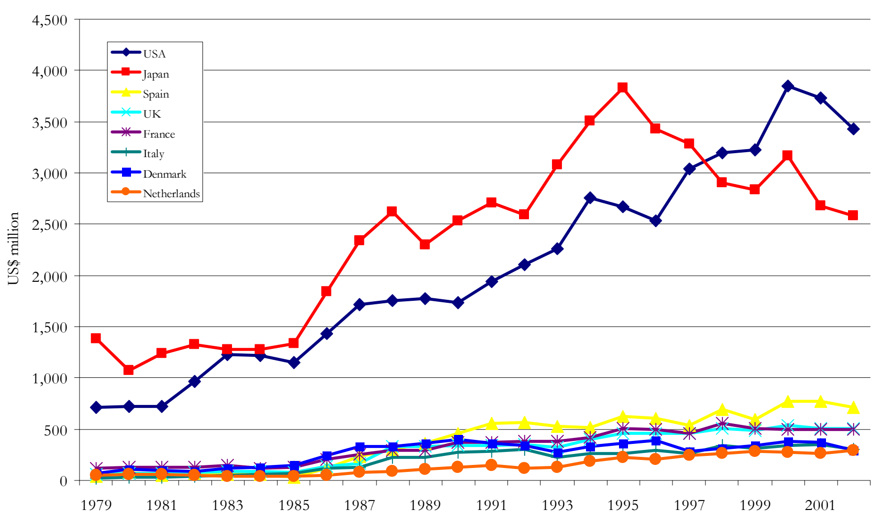

Figure 2. Annual Shrimp Imports by Major Importing Countries in US$ million, 1979-2002

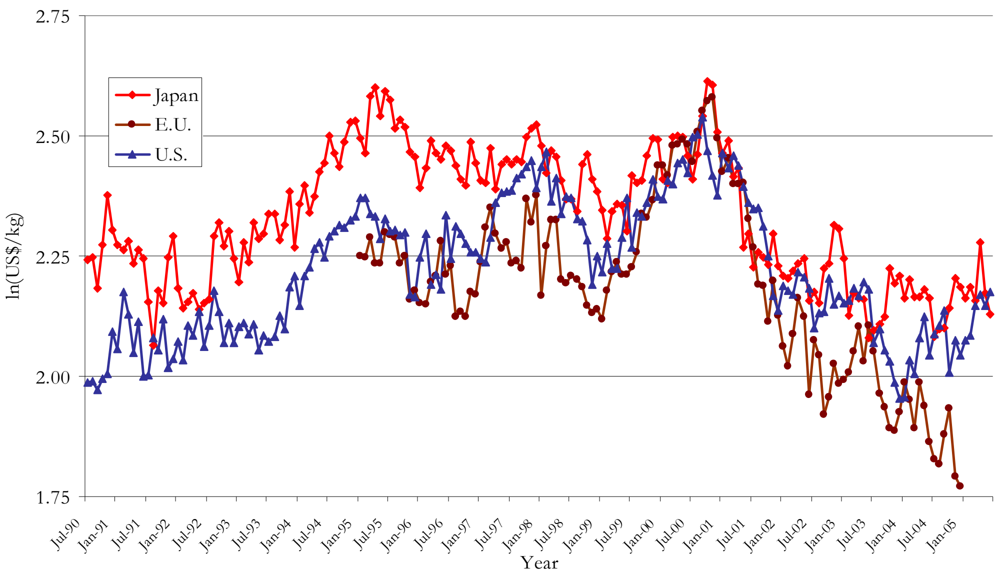

Figure 3. Shrimp Prices in the European Union, Japan and the United States, 1990-2005

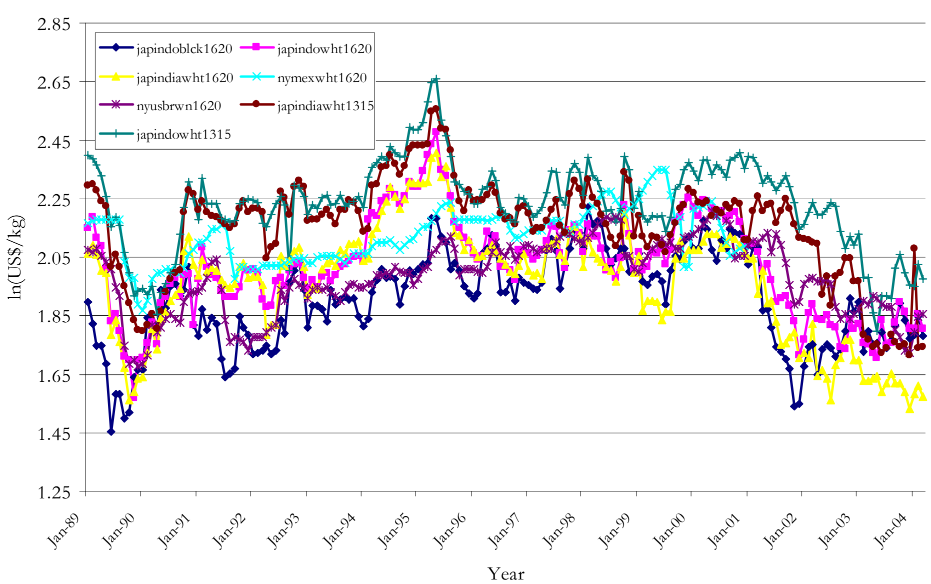

Figure 4. N.Y. and Tokyo Wholesale Markets Prices: 16-20 and 13-15 count size

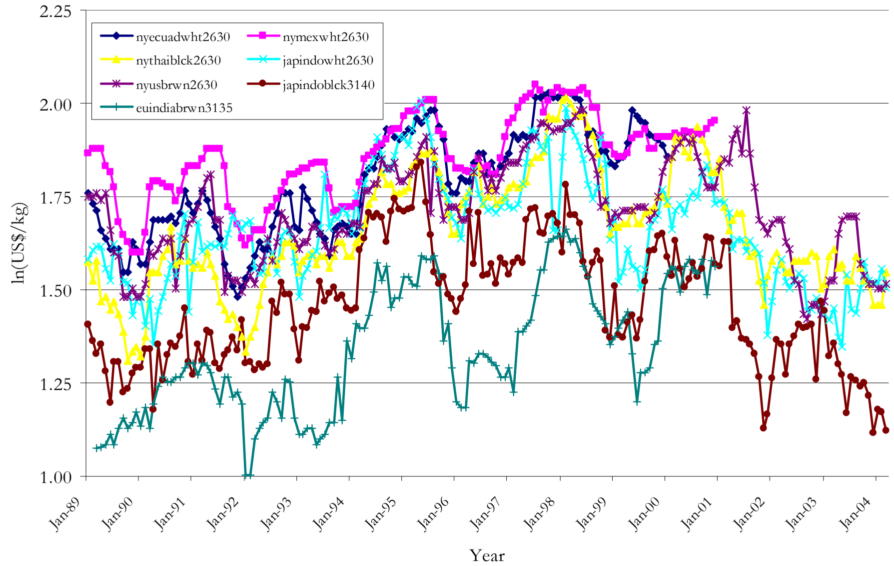

Figure 5. N.Y., Tokyo, and E.U. Wholesale Markets Prices: 26-30, 31-35, and 31-40 count size

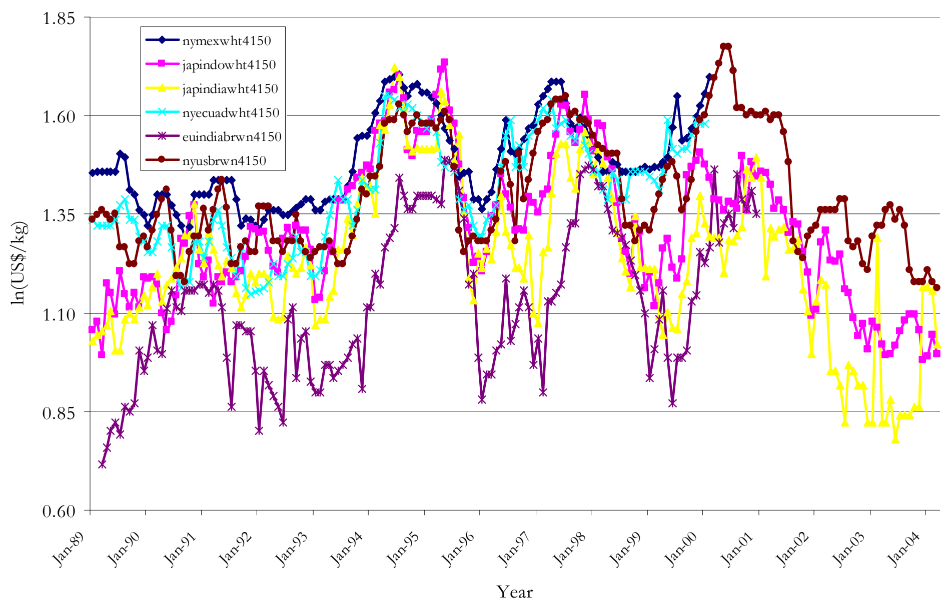

Figure 6. N.Y., Tokyo, and Europe Wholesale Markets Prices: 41-50 count size

Table 1
Augmented Dickey-Fuller Tests Results

Variable	Levels with constant	 with constant and trend	First Differences with constant	 with constant and trend
World Markets				
EU	-1.051	-1.804	-12.484**	-7.818**
Japan	-1.862	-2.151	-8.696**	-8.720**
United States	-2.153	-2.016	-18.452**	-18.485**
Wholesale Markets				
New York				
U.S. Brown 16-20	-2.520	-2.623	-10.521**	-10.493**
Mexico White 16-20	-1.852	-3.626*	-7.001**	-6.982**
U.S. Brown 26-30	-2.609	-2.582	-8.264**	-8.250**
Mexico White 26-30	-2.411	-3.658*	-7.687**	-7.676**
Thailand Black Tiger 26-30	-2.328	-2.162	-7.028**	-7.072**
Ecuador White 26-30	-1.828	-3.140	-9.039**	-9.001**
U.S. Brown 41-50	-2.343	-3.309	-11.419**	-11.415**
Mexico White 41-50	-1.683	-3.853*	-8.618**	-8.633**
Ecuador White 41-50	-1.629	-2.539	-10.613**	-10.581**
Tokyo				
India White 13-15	-1.887	-2.079	-15.566**	-15.547**
Indonesia White 13-15	-2.626	-2.631	-10.720**	-10.697**
India White 16-20	-1.729	-2.017	-14.714**	-14.701**
Indonesia White 16-20	-2.244	-2.272	-10.949**	-10.923**
Indonesia Black Tiger 16-20	-2.962	-2.965	-11.366**	-11.357**
Indonesia White 26-30	-2.202	-2.224	-12.603**	-12.590**
Indonesia Black Tiger 31-40	-2.055	-1.992	-16.786**	-16.806**
India White 41-50	-2.678	-2.889	-16.179**	-16.166**
Indonesia White 41-50	-2.333	-2.392	-13.043**	-13.098**
Europe				
India Brown 31-35	-2.110	-3.914*	-13.369**	-13.321**
India Brown 41-50	-2.987*	-3.263	-13.813**	-13.771**

**Indicates significance at the 1% level. *Indicates significance at the 5% level.

Critical value at 1% level is -3.48 with constant and -4.03 with a trend.

Critical value at 5% level is -2.88 with a constant and -3.45 with a constant and trend (MacKinnon 1996).

Table 2

Multivariate Johansen Test for Cointegration

H_0: rank = r	Max test	Critical value 5%	Trace test	Critical value 5%
r==0	28.24**	25.82	52.03**	42.92
r<=1	20.99**	19.39	23.79*	25.87
r<=2	2.80	12.52	2.80	12.52

**denotes rejection of the hypothesis at 5% level

*denotes rejection of the hypothesis at 10% level

Table 3

Bivariate Johansen Tests for Cointegration and the Law of One Price

Variables	H_0: rank = r	Max test	Trace test	Law of One Price
U.S. and Japan	r==0	26.13***	31.33***	0.129
	r<=1	5.21	5.21	
U.S. and E.U.	r==0	26.34***	29.07***	7.54***
	r<=1	2.72	2.72	
Japan and E.U.	r==0	21.39***	25.26***	14.91***
	r<=1	3.87	3.87	

***denotes rejection of the hypothesis at 1% level

Table 4

Weak Exogeneity Tests

Variable	Test statistic	*p*-value
European Union	9.467	0.008
Japan	17.704	0.000
United States	15.653	0.000

Table 5

Multivariate Johansen Test for Cointegration and LOP – Prices of 41-50 Count Shrimp

Price Series	Max Test			Trace Test			LOP Tests	
	$r = 0$	$r = 1$	$r = 2$	$r = 0$	$r <= 1$	$r <= 2$	L.R.	*p*-value
NY-USBr/Tok-IndWht/EU-IndBr	35.5***	21.80**	9.76	67.10***	31.57***	9.76	5.60	0.06
NY-USBr/Tok-IndoWht/EU-IndBr	28.20**	20.22**	11.33	59.74***	31.54***	12.52	5.06	0.08
NY-MexWht/Tok-IndWht/EU-IndBr	37.36***	17.39*	8.58	63.32***	25.96**	8.58	2.20	0.33
NY-MexWht/Tok-IndoWht/EU-IndBr	27.14**	19.72**	9.72	56.58***	29.44**	9.72	1.37	0.51
NY-EcuWht/Tok-IndWht/EU-IndBr	36.15***	19.28*	10.91*	66.34***	30.19**	10.91	3.11	0.21

Note: NY= New York, Tok = Tokyo, EU = European Union, USBr = U.S. Brown, IndWht = India White, IndoWht = Indonesia White, IndBr = India Brown, MexWht = Mexico White, and EcuWht = Ecuador White
***, **,* deontes rejection of the hypothesis at 1%, 5%, and 10% level, respectively.

Table 6

Weak Exogeneity Tests – Prices of 41-50 Count Shrimp

Price Series	Test Statistic	*p*-value
New York - US Brown	8.714	0.013
Tokyo - India White	2.819	0.244
European Union - India Brown	25.492	0.000
New York - US Brown	8.611	0.013
Tokyo - Indonesia White	0.386	0.825
European Union - India Brown	16.364	0.000
New York - Mexico White	2.040	0.361
Tokyo - India White	7.226	0.027
European Union - India Brown	26.878	0.000
New York - Mexico White	2.84	0.242
Tokyo - Indonesia White	7.443	0.024
European Union - India Brown	16.589	0.000
New York - Ecuador White	3.790	0.150
Tokyo - India White	4.412	0.110
European Union - India Brown	22.982	0.000

Table 7

Bivariate Johansen Tests for Cointegration and LOP-Wholesale Shrimp Prices

Price Series	H_0: rank=r	Max test	Trace test	Weak Exogeneity	Law of One Price
31-40/31-35 Count					
Tokyo-Indonesia Black	r==0	19.25*	26.32**	1.708	0.587
E.U.-India Brown	r<=1	12.52	7.07	9.133***	
26-30 Count					
N.Y.-U.S. Brown	r==0	26.67***	30.03**	4.672**	0.033
Tokyo-Indonesia White	r<=1	3.36	3.36	16.001***	
N.Y.-Thailand Black	r==0	19.00*	23.80*	2.009	14.194***
Tokyo-Indonesia White	r<=1	4.80	4.80	9.720***	
N.Y.-Mexico White	r==0	25.36***	34.25***	4.480**	0.311
Tokyo-Indonesia White	r<=1	8.89	8.89	13.60	
N.Y.-Ecuador White	r==0	17.95*	25.28*	3.643*	0.511
Tokyo-Indonesia White	r<=1	7.33	7.33	5.549*	
16-20 Count					
N.Y.-U.S. Brown	r==0	21.83**	27.83**	15.827***	6.960***
Tokyo-Indonesia White	r<=1	5.99	5.99	0.252	
N.Y.-U.S. Brown	r==0	20.16**	24.18*	14.90***	5.780**
Tokyo-India White	r<=1	4.03	4.03	0.238	
N.Y.-U.S. Brown	r==0	38.67***	45.90***	29.93***	3.592
Tokyo-Indonesia Black	r<=1	7.24	7.34	0.005	
N.Y.-Mexico White	r==0	22.32**	30.16**	4.315**	14.390***
Tokyo-Indonesia White	r<=1	7.85	7.85	9.917***	
N.Y.-Mexico White	r==0	26.03***	36.51***	7.167***	15.156***
Tokyo-India White	r<=1	10.48	10.48	10.370***	
N.Y.-Mexico White	r==0	25.43***	45.33***	3.150	4.068**
Tokyo-Indonesia Black	r<=1	19.90***	19.90***	2.318	
16-20/13-15 Count					
N.Y.-U.S. Brown	r==0	19.93**	24.42*	10.269***	2.541
Tokyo-Indonesia White	r<=1	4.49	4.49	4.283**	
N.Y.-U.S. Brown	r==0	21.86**	25.77*	13.629***	5.322**
Tokyo-Indonesia White	r<=1	3.91	3.91	2.424	

Note: ***, **, * denote rejection of the hypothesis at 1%, 5%, and 10% levels, respectively.