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A Fish Is a Fish Is a Fish? Testing for Market Linkages on the Paris Fish Market

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Abstract *This paper applies both the Engle–Granger and Johansen cointegration test procedures to determine the existence of market linkage among high-valued (salmon and turbot) and low-valued (cod) fish species using monthly average wholesale price data recorded on the Paris fish market. We find that the price of salmon is determined exogenously to the system of prices examined and that the market for salmon is not linked to the markets for turbot or cod.*

Keywords Cointegration, market linkages, salmon price.

Introduction

A common practice in fisheries economic research is to examine high-valued species (*i.e.*, a high income elasticity) and low-valued species independently. For instance, fresh salmon is considered a high-valued species and is analyzed as a separate fish market by economists where only the relationship among different species of salmon or other high-valued species is assessed (DeVoretz, 1989; Hermann and Lin, 1988; Bjørndal, Salvanes and Andreassen, 1992; DeVoretz and Salvanes, 1993; Bjørndal, Gordon and Singh, 1993). If the markets for high-valued species are linked, say through commodities arbitrage, individual fish prices cannot diverge "too far" from other fish prices before market forces operate to restore equilibrium. In a time-series framework, this implies the existence of a long-run equilibrium relationship among the prices of high-valued fish species and specifically, that these fish prices must be cointegrated (Granger, 1986; Engle and Granger, 1987).[1] On the other hand, if the markets for high- and low-valued fish species are not linked, a cointegrating relationship will not exist among the prices of these different valued fish species.

We would like to thank Trond Bjørndal, Kjell Vaage and two anonymous referees for helpful comments, and Sarah Bibb and Frank Asche for research assistance.

[1] A number of studies have used Engle–Granger procedures to define the spatial characteristics of markets for different commodities (Slade, 1986; Goodwin and Schrolder, 1991; Gordon, Hobbs and Kerr, 1993).

The purpose of this paper is to test for market linkages among high- and low-valued fish species using monthly data for the period 1981–90 recorded at the Marché d'Interel National de Rungis—the largest wholesale fish market in Europe located near Paris. The test for market linkages is based on the existence of cointegrating transformations among different fish prices. Three fish prices are used in testing: salmon, turbot and cod. The former two represent high-valued fish species and the latter represents a low-valued fish species.

Two recent developments in French fish markets emphasize the relevance of the paper. First, the substantial increase in production of farmed atlantic salmon during the 1980s caused a significant decrease in the price of salmon[2] and allowed for increased imports of both fresh and frozen salmon, primarily from Norway and Scotland to France. Bjørndal, Salvanes and Andreassen (1992) report that total imports of salmon to France increased by 208% during the 1980s. In addition, by 1989 Norway held a dominant share of both fresh (62%) and frozen (43%) salmon import supply. Hence, it seems reasonable to enquire as to the market linkages between salmon and other fish species.

Second, Andreassen (1991) argues that retailing of high-valued fresh fish, particularly salmon, in France has changed substantially during the 1980s. In earlier years, high-valued fish was sold mainly through restaurants and specialized fish shops. During the 1980s, however, general supermarkets increased their share of fresh fish sold from 21% in 1981 to 40% in 1989. By the late 1980s these outlets were the largest distributors of fresh fish products in France. One would expect that these changes in the distribution of fresh fish have improved consumer access to both high- and low-valued fish products allowing for the possibility of greater market linkages both among high-valued species and, perhaps, between high- and low-valued species. In this paper, we test for market linkages in both types of fish species.

In testing for cointegration, we employ both the Engle–Granger (1987) and Johansen (1988) procedures. The Engle–Granger test is chosen because of its simple and direct application; whereas, the Johansen test is complicated in application but generates test statistics (*i.e.*, likelihood ratios) with exact limiting distributions and allows for identification of all cointegrating transformations for the set of variables examined (Hall, 1986). This is important because the more cointegrating vectors that exist for the different price series the more stable the long-run relationship (*i.e.*, market linkages) among fish prices.

From the empirical results, a number of implications are possible. If, for the three prices examined, a cointegrating vector exists strategic pricing and marketing of individual fish species should respond to supply variations of the other fish species. This is especially true for the salmon fish-farm industry which has some flexibility in setting the timing and quantity of fish supplied to the market. Moreover, if cointegration is not observed between high- and low-valued fish species, this will provide statistical support for the assumption that different valued fish species can be examined independently. Finally, since we are using spot market data for farmed fish (salmon) and for wild-caught fish (turbot and cod) tests using error-correction models, which are derived from the cointegration models, can be used to define the importance of farmed-fish prices in determining wild-caught fish prices.

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[2] See Bjørndal (1990).

The Rungis Fish Market

Rungis is a wholesale market supplying Paris with fresh fish (*i.e.*, the product form sold on the Rungis market is fresh) and fresh agricultural commodities with annual sales revenue of over 4000 million French Francs (FF). About 33% of total fresh fish in France is distributed via this wholesale market. The volume of fish distributed through the market includes both domestic (70%) and imported (30%) fish (Andreassen, 1991). The Rungis market is not an auction market but rather organized as a spot market operating four days a week. Approximately 80 fish mongers hold permits to sell fish on the market. About 1800 fish buyers frequent the market daily.[3] The market is organized with all supplies in one building thus allowing price information to be quickly passed between buyers and sellers and product quality assessed. Suppliers start with list prices and accept bids from buyers. The market clears on a daily basis with no storage of fish products.

Cointegration tests for different fish species are carried out using wholesale prices. Consequently, market linkages, if observed, reflect links at the wholesale level. In fisheries economic research, it is common to assume separability between high- and low-valued fish at either the retail or wholesale market levels.[4]

On a monthly basis data are available on a variety of fresh fish species for the period 1981 (January) to 1990 (June).[5] The data represent average monthly prices for each species. To facilitate the cointegration testing three fish prices are used; the price of salmon, the price of turbot and the price of cod. In Table 1, the annual quantity and value of these three fish species distributed through the Rungis market for the period 1981–89 are presented.[6] The volume of cod and turbot over this nine year period appears relatively stable compared to salmon, which shows a sevenfold increase in volume traded. Moreover, by 1986 salmon is the single most valuable species on the Rungis market.[7]

Salmon recorded on the Rungis market represent all imported product. Hjelle (1989) states that Norway dominates this fresh fish market with about 90% of salmon traded at Rungis in 1988. The remaining supplies are from United Kingdom and Ireland. Because of the Norwegian position in this market and that for this product monthly prices are available for the period of analysis, the average price of fresh Norwegian salmon is used as the price variable for this fish species.[8] In France, fresh salmon is distributed retail primarily through restaurants, supermarkets and specialized fish shops.

Turbot is a wild-caught fish supplied primarily by the domestic (French) fleet. This is a high-valued species but the quantity sold at Rungis is not large (see Table 1). Turbot is generally considered a substitute for salmon and has a similar retail distribution system.

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[3] The buyers include caterers, restaurants, fish shops and agents for supermarkets and fish-processing industries.

[4] See Bjørndal, Salvanes and Andreassen (1992) for an example of wholesale prices used to estimate market demand functions assuming separability between high- and low-valued fish species.

[5] Price data used in the analyses are collected by Andreassen (1991).

[6] The six months of data available for 1990 are not reported in Table 1.

[7] Bjørndal, Salvanes and Andreassen (1992) investigate the demand conditions for fresh salmon in France using the monthly Rungis data series. In modelling demand, they assume that salmon and an index of other high-valued species represent substitute products.

[8] Scottish salmon is not available on the Rungis market for all months of our analysis.

Table 1
The Value (in ,000 FF) and Quantities (in ,000 metric tonnes) of Important Fish Species Sold at Rungis in the Period 1981–1989

Species	Cod		Salmon		Turbot	
Year	Quantity	Value	Quantity	Value	Quantity	Value
1981	4,904	63,443	874	39,318	991	48,337
1982	4,482	67,518	959	49,485	1,001	54,037
1983	4,200	71,575	1,075	57,340	966	58,856
1984	4,297	79,745	1,488	89,132	1,088	69,508
1985	2,979	85,510	1,485	96,366	1,076	76,130
1986	4,146	96,965	3,049	134,858	993	77,156
1987	3,911	89,744	3,335	158,202	720	62,455
1988	4,454	93,746	4,239	201,138	968	93,770
1989	4,337	108,498	6,577	260,404	613	67,151

Cod is a wild-caught low-valued fish, with the domestic fleet providing about 55% of total cod supply on the Rungis market. Cod is retailed primarily through supermarkets, fish shops and used in the catering industry. The price variable for cod used in the analysis is a divisia price index of domestic and imported cod at Rungis.[9]

Data available on other fresh fish species are combined to form a Divisia price index, and this index is used to deflate the nominal prices of salmon, turbot and cod.[10] Monthly average nominal and real prices of turbot, cod and salmon are shown in Figure 1 for the period 1981–90. In nominal and real terms, all prices show substantial yearly variation, but turbot and cod prices trend upward over the period, whereas salmon prices show a downward trend.

Seasonality in the Rungis fish market is an important factor, with demand for fish increasing in the months of May/June and during the Christmas season (see, Bjørndal, Salvanes and Andreassen, 1992). In order to observe long run market linkages and to facilitate cointegration testing we purge the data of seasonal effects. If the price series are not deseasonalized, test results may lead to false inferences about market linkages. By doing so, the results will provide no information on seasonal relationships among different fish species. Keep in mind, the interesting factor on the Rungis market is the long term decline in salmon prices and the impact of this on market linkages.

In deseasonalizing the data we follow the approach suggested by Jorgenson (1964) and regress each price series separately against a matrix of seasonal (monthly) dummy variables (D_S) and a polynomial in a time trend (T). For each price series i, the regression is written

$$\mathcal{P}_i = T\alpha + D_S\beta + \varepsilon \qquad (1)$$

The estimated vector α will capture trend and cyclical variation in $\mathcal{P}_i$, whereas the

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[9] Trond Bjørndal provided helpful comments on the French fish market.

[10] The species used in building the Divisia price index include dog fish, whiting, skate, sole, brill, monk, hake, sea bream and saithe.

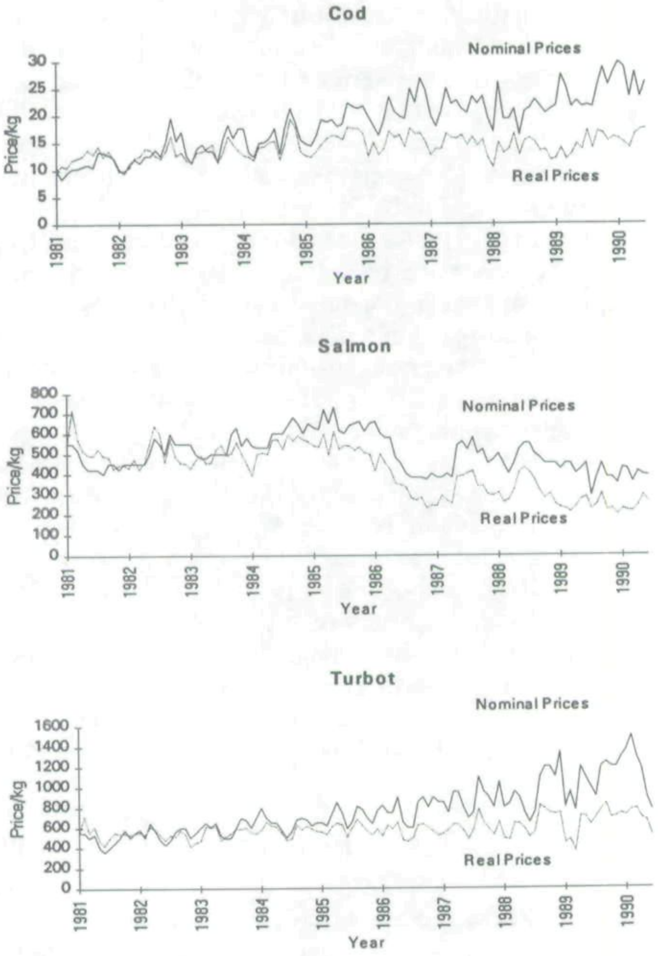

Figure 1.

vector β will measure only seasonal variation. The deseasonalized price is generated by purging the original series of seasonal factors or

$$P_i = \mathcal{P}_i - D_S\hat{\beta} \tag{2}$$

P_i has the property of not summing to zero as would be the case if $T\alpha$ were missing from Equation (1), but rather has the same sum as the original, unadjusted series. Finally, the deseasonalized real prices are used in cointegration testing.

Tests for Cointegration

The concept of cointegration developed from the observation that many economic time series tended to change over time. If the series change in a stable or predictable way it will have a well defined mean and variance. If the series change in an unstable or unpredictable way the mean and variance will not be well defined. However, even if a time series does not have well defined mean and variance in level form (*i.e.*, the actual data values) it may exhibit stable properties if differenced (d) one or more times, and these series are called integrated series of order d. Cointegration extends the single series concept of integration to consideration of two or more series. Suppose that we have two time series each integrated of order one. We know that each series will change in an unstable way, however, it is possible that the difference between the two series will change in a stable way. If so, these two series are said to be cointegrated.

As described by Granger (1986), time series procedures have two advantages over the traditional regression approach (*e.g.*, market demand analysis). First, if the variables of interest are not cointegrated a standard regression analyses of the variables may lead to incorrect inferences because the true value of the coefficients must be zero. Thus cointegration testing can be used as a pre-test procedure to avoid "spurious regressions". Second, if cointegration exists the traditional regression model defines the long run relationship but is not appropriate to define the short-run dynamic behaviour of the system. However, from the cointegration regression, an error-correction model exists and includes variables to measure both short-run disequilibrium and dynamic effects.

The development of cointegration stems largely from work by Granger and Engle (Granger, 1986; Engle and Granger, 1987).[11] The idea is one of testing for the existence of a long-run equilibrium relationship among a series of variables. This equilibrium can be observed by examining the properties of the series investigated. Specifically, a vector X_t with the property that each component (x_t) of X_t having been differenced d times is stationary can be written as $x_t \sim I(\mathrm{d})$.[12] If all components of X_t are integrated of the same order, cointegration exists if a cointegrating vector γ is found such that

$$Z_t = X_t\gamma \sim I(0) \tag{3}$$

That is, Z_t is stationary and represents random disturbances from a (long-run) equilibrium position; with the system again adjusting to the equilibrium. If the components, x_t, are integrated of different orders then the estimated value of $\hat{\gamma}$ in Equation (3) must be zero implying that a long-run equilibrium does not exist. For the case at hand, cointegration implies that fish prices cannot diverge systematically from a long-run equilibrium position with other fish prices. If series are cointegrated short-run deviations are possible but market forces operate to regain the equilibrium. Presumably, the underlying market conditions are such that if the price of one fish species diverges "too far" from the equilibrium level, buyers substitute towards other fish species changing the demand conditions and causing the divergent fish price to return to the long-run equilibrium position.

Engle and Granger (1987) suggest a simple test for cointegration using a two-stage approach. First, test that each variable series is stationary after differencing d times. Second, form the cointegrating vector and test that the errors (Z_t in Equation (3)) are integrated of order zero.

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[11] See Dickey, Jansen and Thornton (1991) for a survey of cointegration techniques and procedures.

[12] Stationarity implies that the probability distribution of the series is invariant with respect to time. In other words, the mean, variance and covariance are constant for all observations.

Although the Engle and Granger procedure is simple and direct, several problems confront the applied researcher. As described by Hall (1986), the procedure assumes that the cointegrating vector is unique. This need not be the case and a number of cointegrating vectors may exist. As well, the Engle–Granger test procedure may not have well defined limiting distributions which may complicate hypothesis testing.

Work by Johansen (1988) and Johansen and Juselius (1990) offer solutions to the problems in Engle–Granger cointegration testing. Briefly described, the method relies upon the concept of canonical correlations from the theory of multivariate analysis. The data are divided into a differenced and a levels part. Under the assumption of $I(1)$ processes the differenced data are stationary. The technique of canonical correlations is used to find linear combinations of the data in levels which are as highly correlated as possible with the differences. It follows that these linear combinations must be stationary, or cointegrated.

Another appealing aspect of the Johansen approach is its completeness in the sense that it provides tests of linear restrictions on the cointegrating vectors as well as estimates of its elements and information about its rank. The assumption is that x_t is generated by an autoregressive form:

$$x_t = \sum_{i=1}^{k} \pi_t x_{t-i} + \varepsilon_t \tag{4}$$

which can be rewritten as

$$\Delta x_t = \sum_{i=1}^{k-1} \Gamma_i \Delta x_{t-i} - \pi x_{t-k} + \varepsilon_t \, , \tag{5}$$

with

$$\Gamma_m = -I + \sum_{i=1}^{m} \pi_i, \; m = 1, \ldots, k-1, \; \pi = I - \sum_{i=1}^{k} \pi_i$$

and ε_t is a random error term.

Since ε_t is stationary, the rank (ρ) of the ''long-run'' matrix π determines how many linear combinations of x_t are stationary. If $\rho = \mathrm{n}$ all x_t are stationary, while if $\rho = 0$ so that $\pi = 0$, Δx_t is stationary and all linear combinations of $x_t \sim I(1)$. For $0 < \pi < \mathrm{n}$, there exist ρ cointegrating vectors, meaning ρ stationary linear combinations of x_t. In that case, π can be factored as $\alpha\beta'$ with both α and β being ($\mathrm{n} \times \rho$) matrices. The cointegrating vectors of β are the error correction mechanisms in the system, while α contains the adjustment parameters.[13]

In the empirical analysis, we first apply the Engle–Granger procedure to the

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[13] This result is known as Granger's Representation Theorem. Engle and Granger (1987) give the original result.

three fish prices of interest and determine if a cointegrating vector exists. Next, we apply the Johansen procedure to test for the existence of more than one cointegrating vector.

Empirical Results

Prior to testing for cointegration, it is necessary to establish the univariate stationarity properties of each of the price variables. A standard procedure is to apply an Augmented Dickey–Fuller (ADF) test of the following form:

$$\Delta P_{it} = \beta_0 + \beta T + \rho P_{i,t-1} + \sum_{\gamma=1}^{k} \alpha_\gamma \, \Delta P_{i,t-\gamma} + \varepsilon_t \tag{6}$$

where Δ is the difference operator and T is a time trend (Dickey and Fuller, 1979). The null hypothesis is that the series is non-stationary [*i.e.*, $\sim I(1)$] against the alternative hypothesis of stationarity [*i.e.*, $\sim I(0)$]. The null hypothesis is tested based on the "t-statistic" value of ρ in Equation (6).

In specifying Equation (6), the lagged differences ($\Delta P_{i,t-\gamma}$) are included to ensure that the residuals (ε_t) are white noise. If too few lags are included, the size of the test changes in an unknown manner, and if too many lags are included the power of the test is reduced. Accordingly, in this paper, the lag length, k, is chosen to minimize the Akaike Information Criterion (AIC).

We first test the null hypothesis that each price series is $\sim I(1)$, and report the results in the first three rows of Table 2. The ADF "t" test statistics are evaluated using a one-tailed test at a 5 percent significance level of -3.45 (Dickey and Fuller, 1979). However, for the three monthly price series none of the test statistics, can reject the null hypothesis of non-stationarity. To allow for the possibility that prices are stationary in first differences, we apply the ADF test to the transformed series. The null hypothesis is that the series is $\sim I(2)$ against the alternative hypothesis that the series is $\sim I(1)$. The results are reported in the last three rows of Table 2. For each price series, the test statistics resoundingly reject the null hypothesis of non-stationarity and accept the alternative hypothesis of stationarity in first differences. We take this as evidence that each price series is $\sim I(1)$, and proceed to test for cointegrating vectors.

In applying the Engle–Granger cointegration test, we assume joint endogeneity for the three fish prices and thus there are three cointegrating regressions to examine. The results are presented in Table 3. Consider first, the regressions with salmon and cod normalized as the dependent variable. The ADF statistic for each regression cannot reject the null hypothesis of no cointegration and leads us to conclude that the two equations are spurious. In contrast to this, the regression with turbot normalized as the dependent variable generates an ADF statistic which is consistent with cointegration. Consequently, the Engle–Granger procedure defines one long-run equilibrium relationship and that this cointegrating vector is associated with turbot price being endogenous to the other two fish prices. Moreover, by conventional statistical testing it would appear that salmon and turbot are related, while cod is not statistically important in any of the equations.

These results, however, should be interpreted with caution. The adjusted R^2

Table 2
Univariate Stationarity Test

	(a) H_0:I(1)^a
Variable	ADF^a
P_S^b	−2.60 (1)^c
P_C	−1.36 (10)
P_T	−2.27 (10)
	(b) H_0:I(2)
ΔP_S^d	−14.03 (0)
ΔP_C	−9.33 (3)
ΔP_T	−8.36 (2)

Notes: [a] The critical values for the ADF test are: 95% −3.45, 99% −4.04.

[b] The subscripts on P define salmon (S), cod (C) and turbot (T).

[c] The optimal lag length, chosen by the Akaike Information Criteria, is in brackets.

[d] Δ represents first differences.

values are extremely low and, as pointed out by Hendry (1986), the ability of the Engle–Granger procedure to detect the presence of cointegration is related to the size of this summary statistic. In addition, although we have presented the estimated coefficients and the "t" statistics in Table 3, the adjusted R_2 values are too low to invoke Stock's (1987) "super consistency" theorem and, consequently, statistical testing may not be valid. For these reasons, we investigate further the possibility of long-run equilibrium relationships among the three fish prices using the Johansen procedure.

Table 3
Engle-Granger Test for Cointegration

Dependent Variable	Independent Variables		$\overline{R}^2$	ADF^a
P_S^b	P_C	P_T		
	12.31	−0.657	0.18	−1.56^d
	(1.83)^c	(−4.45)		
P_C	P_S	P_T		
	−0.002	0.001	0.03	−2.02
	(−1.83)	(0.55)		
P_T	P_S	P_C		
	−0.239	2.27	0.16	−3.91
	(−4.45)	(0.55)		

Notes: [a] ADF is the Augmented Dickey-Fuller Statistic, and the critical value for the ADF statistic at the 95% level is −3.62 (Engle and Yoo 1987).

[b] The subscripts on P define salmon (S), cod (C) and turbot (T) with t-statistics in parentheses.

Table 4
Johansen Test for Cointegration

(a) Trace Test

H_0^a	Test Statistic	Critical Value	
$k = 0$	21.22	21.2^b	23.8^c
$k \leqslant 1$	4.28	10.3	12.0
$k \leqslant 2$	1.71	2.9	4.2

(b) Maximum Eigenvalue

H_0	H_a	Test Statistic	Critical Value	
$k = 0$	$k = 1$	16.94	18.7^b	20.8^c
$k = 1$	$k = 2$	3.57	12.1	14.0

Notes: [a] Tested against a general alternative
[b] 90% significance level
[c] 95% significance level

The Johansen procedure for testing cointegration was implemented using a VAR lag length of 10, which was chosen using the minimum AIC technique. The results of the test are reported in Table 4. The first half of the table shows the trace test results, which test the null hypothesis that the number of cointegrating vectors is less than or equal to k, where k is 0, 1 or 2. The critical test values are given in Johansen (1988). The results show that a null hypothesis of zero cointegrating vectors can be rejected at the 90% level but not at the 95% level;[14] however, the null that $k \leqslant 1$ or $\leqslant 2$ cannot be rejected. These results would appear to indicate the existence of only one cointegrating vector. For completeness we also report the results of a Maximum Eigenvalue test of the null hypothesis $k = 0$ against the alternative $k = 1$. These test statistics are reported in the bottom half of Table 4. Johansen and Juselius (1990) report critical test values for the Maximum Eigenvalue procedure. The results show that a null hypothesis of $k = 0$ cannot be rejected in favour of the alternative $k = 1$ at the 90% level although rejection of the null does occur above the 80% level. On the other hand, a null hypothesis of the existence of only one cointegrating vector ($k = 1$) against the alternative of $k = 2$ cannot be rejected. We conclude that there is some evidence of the existence of one cointegrating vector involving the prices of salmon, cod and turbot and, that a long-run equilibrium relationship may exist for these prices.

Both the Engle–Granger and Johansen procedures provide some evidence of market linkages for the three fish species examined. For our purpose it is important to determine the role of salmon price within this relationship. We do this by estimating the space spanned by the cointegrating vector. The eigenvector matrix is premultipled by the transpose of the inverse of the Choleski decomposition of the S_{kk} matrix (in the notation used in Dickey, Jansen and Thornton, p. 66, 1991). This procedure generates the following vector: $[0.106\ -0.268\ 0.356]'$. Since this matrix is an estimate only of the space spanned by the cointegrating vector, we test the hypothesis that the restricted vector $[0\ -0.26\ 0.35]'$ is the cointegrating vector, where the order of the variables in the vector is the price of salmon, cod

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[14] Johansen and Juselius (1990) argue that a 90% confidence interval is appropriate for this test.

TABLE 5
Error Correction Models

	Dependent Variable		
Independent Variables	$P_S{}^a$	P_C	P_T
$EC_{t\text{-}1}{}^b$		-0.33 (-3.83)	-0.46 (-6.07)
$P_{S\ t\text{-}1}$	-0.34 $(-0.316)^c$	—	—
$P_{C\ t\text{-}1}$		-0.20 (-2.12)	

Notes: [a] The subscripts on P define salmon (S), cod (C) and turbot (T).
[b] error correction term lagged one period.
[c] t-statistics in parenthesis.

and turbot respectively. A likelihood ratio statistic is used in testing the restriction.[15] The estimated ratio is 0.044 and, thus, we cannot reject the null hypothesis that the restricted vector is a valid representation of the cointegrating vector.

This is an interesting result. The estimated Johansen cointegrating vector implies that cod and turbot prices bear some statistical relation to each other while the price of salmon is exogenous to these two prices.

We investigate the possibility for the exogeneity of the salmon price further by constructing error-correction models of the three fish prices (Engle and Granger, 1987). The general error-correction model is estimated as:

$$\Delta y_{it} = -\rho_i Z_{it-1} + lagged(\Delta X_{it}, \Delta y_{it}) + e_{it} \tag{8}$$

where y and the vector X represent the different fish prices and Z_{it-1} is the lagged value of the estimated error term from Equation (3). The Z_{it-1} is included to allow for long-run equilibrium adjustments and ΔX_{it} and Δy_{it} are included to allow for short-run dynamics.

Again assuming joint endogeneity for the three fish prices, we estimate various specifications of the error-correction models and test various lag structures. Model specification is based on "t" statistics and a minimum AIC. The estimated parameters for the final equations chosen are reported in Table 5. With salmon prices as the dependent variable, lagged values of cod and turbot prices do not appear in the specification, as well, the error correction term is statistically insignificant. This is consistent with the price of salmon being exogenous to the other two prices. Notice that in the other two specifications, in Table 5, the error correction term is significant, implying that these two prices may be endogenous

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[15] The test statistic is:

$$-2lnQ = \sum_{i=1}^{3} ln[(1 - \lambda_i^k)/(1 - \lambda_i)]$$

where λ^k and λ are the eigenvalues for the restricted and unrestricted models respectively.

to the system. These results support a conclusion that salmon price is not a part of the long-run cointegrating system and, in the short-run, is determined exogenously from the other price series.

Conclusion

The purpose of this paper is to test for market linkages between high- and low-valued fish species recorded on the Rungis fish market near Paris. Rungis is a wholesale spot market for the distribution of domestic and imported fresh fish. Both Engle–Granger and Johansen cointegration procedures are used to test for long-run equilibrium relationships among the prices of salmon, cod and turbot. Results from testing, using both procedures, show weak evidence for the existence of a cointegrating vector for the three-price system. However, testing for the values of the cointegrating vector shows salmon price to be independent of the other two price series. This result is supported by an error-correction model, which shows salmon prices to be determined exogenously to the system and not linked to the markets for cod or turbot. The hypotheses that cod and turbot prices are endogenous to the system cannot be rejected. At the wholesale level cod and turbot prices cannot be represented as substitute products for salmon. For cod, this result is consistent with the common assumption that markets for high- and low-valued species are independent. But, the results also indicate that salmon price is independent of the high-valued species turbot. We conclude that the practise of including high-valued species as substitute commodities in salmon market demand studies may lead to a misspecified model. A better procedure would be to pre-test the variables to ensure a cointegrating relationship exists prior to market demand studies.

The results show that the price of salmon is determined exogenously to the price of turbot and cod. Because farmed atlantic salmon represents by far the greater quantity of fresh salmon on the Rungis market, we read this result to indicate that price variations for farmed fish do not determine or influence price variations in wild-caught fish, such as cod or turbot. In more general terms, this implies that strategic pricing and marketing of farmed salmon need not be restricted by variation in supply of either a high-valued species such as turbot or a low-valued species such as cod.

Finally, for fresh salmon the notion that "A Fish is a Fish is a Fish" is apparently not supported by the data.

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