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Impact of cultured shrimp industry on wild shrimp fisheries: analysis of price determination mechanisms and market dynamics

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Abstract

Cultured shrimp production has been growing dramatically on the world market over the last 15 years and some of the farm-raised species are now considered as price-indicators on the main market places. One may, therefore, expect the price of these cultured shrimp to have an impact on the price formation of other species and especially on wild shrimp with which they compete. In this paper, the authors address this question in the case of the wild shrimp *Penaeus subtilis* exploited by the French Guyana fishery (South America) and competing on the French market with the cultured Thai shrimp 'Back Tiger'. A series of econometric tests issued from the co-integration theory is performed between the price series of the two products. These tests indicate that the two series are co-integrated and that the black tiger market acts as a market leader for the French Guyana shrimp product. The authors then discuss the reasons of the current predominance of farm-raised shrimp on wild-caught product (and in particular the French Guyana shrimp) and identify the constraints that the French market demand induces on both producers and importers. In the light of this analysis, a commercial strategy that would mitigate the impact of the Thai shrimp on the French Guyana product is suggested. © 2000 Elsevier Science B.V. All rights reserved.

Keywords: Wild shrimp; Cultured shrimp; Co-integration; Price determination; Market analysis; French Guyana

1. Introduction

Commodity prices are known to be volatile (Deaton and Laroque, 1992). Understanding the price determination mechanisms of these commodities is therefore of great practical importance, particularly for low and middle incomes countries where price variability can have detrimental impacts on both consumers and producers of agricultural commodities. In addition to the micro-economic aspect of the problem, the volatility of primary commodities also poses major

macro-economic difficulties for these less-developed countries because their economies often depend on exports of a small number of primary commodities.

This is the case of the French Guyana, a French overseas territory located at the north-eastern border of Brazil, the economy of which almost exclusively depends on the exploitation of the wild 'brown' shrimp, *Penaeus subtilis*, living off the north part of South America between Venezuela and Brazil. In French Guyana (noted FG henceforth), about 70 semi-industrial trawlers exploit the portion of the stock found in the FG Exclusive Economic Zone (Béné, 1997a). The export of the production represents 40% (in value terms) of the total export of the FG Territory (IEDOM, 1990, 1993). Any shock on

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the price of this product has, therefore, tremendous effects on the fishing sector and subsequently on the whole socio-economy of the Territory.

The fleet lands annually around 4000 tonnes of shrimp to the port of Cayenne (the main town of FG) and the production, which used to be exported to the Japanese and US markets during the seventies, is now exported towards the European market (Béné, 1996). 95% are purchased by French processors at the Rungis Fish market in Paris, France. The very high commercial value of the FG shrimp compensated the small catch volumes and explains that Cayenne managed to become the fifth French fishing port (in value terms) at the end of the 1980s (IEDOM, 1990).

During the same period, the world shrimp market has been experiencing drastic changes due to the ever increasing supplies of tropical farm-produced shrimp, exported essentially from Asia (China, Indonesia, Thailand, Vietnam) and Latin America (Ecuador). Due to their strong competitiveness because of low production costs, the cultured shrimp of these countries have rapidly made significant inroads into most world shrimp markets. World aquaculture production progressed from about 59 000 to 719 000 tonnes (a 1118% increase) between 1981 and 1991, making the cultured shrimp share grow from 3 to 27% of the total shrimp world production (Rosenberry, 1991).

This substantial increase of the cultured shrimp share induced major changes in the dynamics and structure of the world shrimp markets. While the tropical wild shrimp have dominated the markets until the 1970s, it is now widely admitted that a large part of the movements observed on the two largest market places (Tokyo and New York) are closely related to the price and supply of cultured shrimp (Globefish, 1990). Some of the cultured species are now considered world market indicators. Among these are the 'black tiger' (*Penaeus monodon*) from Thailand and the 'Ecuadorian white' (*Penaeus vannamei*). Black tiger has been the main species produced over the 1990s. In 1997 it represented 17% of the total world production (Josupeit, 1999). In the same year the Ecuadorian white shrimp was the most imported species to the US market with more than 61 000 tonnes.

The European market, third world market after Japan and the USA, was touched by this cultured shrimp 'overflow' somewhat later than the two other main markets. Historically, the cold species has dom-

inated in most of the European countries before 1980. Nevertheless, the warm-water shrimp species have made significant intrusions into the EU countries during the 1980s, and it is acknowledged that the most notable changes in EU shrimp markets over the last decade is precisely the emergence of these tropical farmed-raised shrimp (Hottlet, 1992). With respect to shrimp consumption, one can now divide Europe into two parts: Mediterranean countries and Northern countries. Mediterranean countries prefer larger-sized warm-water shrimp which are generally cooked or grilled, shell-on and head-on. In contrary, cold-water species have always and continue to be the preferred species in Northern European countries (Josupeit, 1992).

In France, the import of tropical farmed-raised shrimp have also displayed major increases over the last 15 years, especially the black tiger and the Ecuadorian white. The Ecuadorian white enjoyed the most important import expansion in the French markets between 1986 and 1993 (from 0 to about 6000 tonnes), while the Thai black tiger grew also significantly, from 1700 tonnes in 1986 to 6800 tonnes in 1991 (Josupeit, 1992).

Although these species are acknowledged by French traders and processors to offer relatively poorer organoleptic (i.e. culinary) qualities than the wild species exported by the FG's companies, market analyses show that the black tiger and Ecuadorian white shrimp are the two species against which *P. subtilis* primarily competes on the French market (Anon., 1991). It, therefore, seems particularly relevant to try to determine the exact impact that these two farm-raised species induce on the price of *P. subtilis*. Indeed given their growing predominance on the international and French markets, they might have a significant influence on the trade of the FG brown shrimp and control to some extent the determination of its price. If this is the case, one can expect the price of *P. subtilis* to follow (possibly with some time delay) the movements of the cultured species and, therefore, to observe co-movements between the price series. The present paper is organised to investigate this hypothesis. Our objective is to verify the reality of this price relationship, and, if it does exist, to detect and quantify the effects that the cultured shrimp's price movements/variability may induce on the price of the wild FG shrimp.

One possible way to analyse such dynamic linkages is to use dynamic simultaneous equations systems (DSES). The basic premise of the DSES is that the series are assumed to be embodied in a system of simultaneous equations where the classifications into endogenous and exogenous and the causal structure of the model are both given a priori. One of the main purposes of the DSES is to forecast the effects of changes in the exogenous variables on the endogenous variables. Keithly et al. (1993), for instance, use this kind of approach to estimate the impact of the world (exogenous) cultured shrimp prices on the US (endogenous) shrimp price dynamics. The DSES approach has however been criticised in recent years on several grounds. First, the classification of variables into endogenous and exogenous is arbitrary and the causal structure cannot then be tested. Secondly, the coefficients in the simultaneous equations are constant and cannot reflect the adaptive behaviour of the economic agents. An alternative approach is to use the co-integration approach where dynamic linkages between variables can be tested without imposing any a priori structure to the model, and thereby avoid the delicate question of exogeneity–endogeneity of the variables. The present paper is the first attempt to address the question of the impact of farm-raised shrimp

markets on the wild-caught shrimp industry using this co-integration framework.

Further details about the principles of the co-integration technique and how it can be used to interpret the data are given in the next section where we also introduce the data series. The econometric analyses and the test results are then presented. We conclude by discussing these results and their relation to market issues, intervention policy and commercial strategies for the FG shrimp industry.

2. Data and methods

2.1. The data series

The series (French Guyana Price, FGP) of the price of the wild shrimp *P. subtilis* exported by the FG fishing companies to the Rungis Fish Market was obtained from the Cayenne's port authorities (CCI, several years). It covers the period January 1986–August 1993 on a monthly basis. Since shrimp are usually sold within different size categories at different prices, the FGP series is in fact the price of the aggregated panel made up by the seven size-categories commercialised by the FG companies. This FGP series is illustrated in Fig. 1.

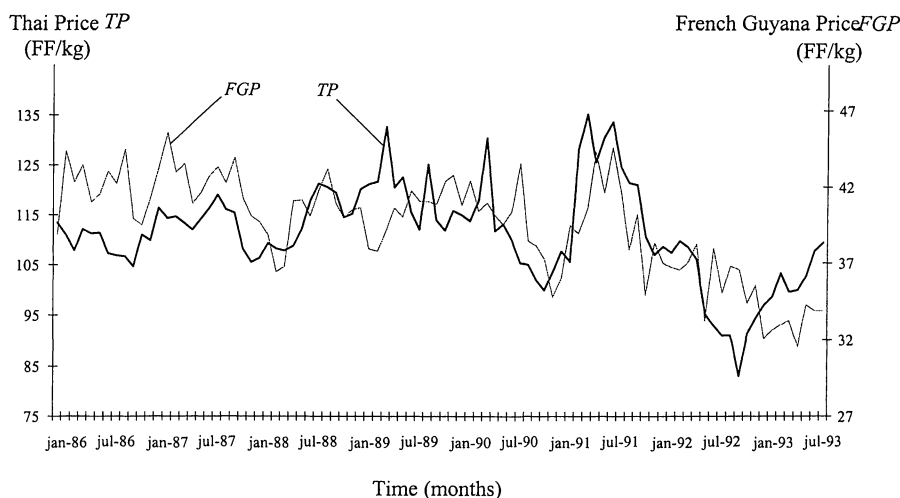


Fig. 1. French Guyana (FGP) and Thai (TP) price time-series over the period January 1986–August 1993. The two series diverge from one another on the short-term (due to their respective dynamics) but display a long-run common trend. Note that the mean of the Thai shrimp price is greater than that of the FG shrimp. This is because the FG aggregated panel corresponds to an 'averaged' category which is smaller (and therefore associated with a lower average price) than the Thai category used for the analysis (source: FGP series: CCI French Guyana, TP series: FAO-Globefish).

As far as the cultured shrimp species are concerned, the data of both the Ecuadorian shrimp *P. vannamei* and Thai shrimp *P. monodon* could theoretically be used to test our hypothesis. However, the price of *P. vannamei* has been recorded on regular bases by the Rungis market authorities only since January 1991. On the other hand, the price of *P. monodon* imported to the Rungis market has been recorded over a longer period. In particular the data of the largest commercialised category of this species is available over the years 1986–1993 (FAO-Globefish, 1990, 1992). We, therefore, used the Thai series (noted TP for Thai price) in the rest of the analysis. The TP data is plotted in Fig. 1 along with the FGP series. When normalised, the two prices display a long-run common trend, confirming graphically the hypothesis of price dependence.

2.2. Co-integration concept, associated tests and expected interpretation

From an econometric viewpoint, this common trend can be associated with the concept of co-integration. The co-integration concept was developed by Engle and Granger (1987) to account for the stochastic property displayed by time-series when they exhibit co-movements and verify long-run equilibrium relationships. In practice, this co-integration concept is associated to a series of statistical concepts which provide an econometric self-contained framework based on four successive steps: (1) unit root test, (2) co-integration analysis, (3) measure of causalities and (4) estimation of the error correction model (ECM) and identification of the impulse response functions. These different steps will be used in the present article to analyse and quantify the dynamic relationships that seems to link the FG shrimp market to the Thai shrimp market. The readers interested in a more technical presentation of these econometric techniques are invited to refer to the original articles indicated in the respective sections.

2.2.1. Unit root tests

For two (or more) series to be co-integrated, they must first be individually integrated. The first step of the analysis, therefore, consists in determining

whether the two series considered FGP and TP are integrated (of order 1 or greater), that is, whether they have unit roots. There are several ways to test for the presence of unit roots. In this study, we use the Dickey–Fuller (DF) and augmented Dickey–Fuller (ADF) tests (Dickey and Fuller, 1981) with the null hypothesis H_0 of no unit root (i.e. the data generating process is non-stationary), against the alternative of presence of unit roots (stationarity of the process). The ADF accounts for possible autocorrelations in the dependent variable.

2.2.2. Co-integration analysis

If both series are integrated of order 1, i.e. $I(1)$, they may also be co-integrated $C(1, 1)$. Several tests have been developed to test for co-integration. The most widely used tests are the Engle and Granger's 'two steps' procedure (Engle and Granger, 1987) and the Johansen's maximum likelihood test (Johansen, 1988, 1991). For our analysis, we use the latter procedure which presents several advantages. In particular it allows the estimation of co-integration vectors using a full information maximum likelihood (FIML) procedure that can be applied to fully dynamic unrestricted models. Furthermore, this FIML process, which is adjusted with lagged differentials, can include deterministic components (a drift μ and several possible exogenous variables). The result, which depends on the number of lags accounted for in the Vector Auto-Regressive (VAR), implies to work out the relevant lag length. This can be done through the calculation of Schwarz (SC) and Hannan–Quinn (HQ) criteria.

2.2.3. Granger causalities tests

If the two price series are $C(1, 1)$, a causality relationship (in the Granger sense, 1988) may exist between them. The Granger causality (GC) test allows to determine whether one series (let us assume in our case the cultured species price TP), precedes the other series (the wild species price FGP), and thus plays the role of 'leader'. In that case, TP is said to 'Granger causes' FGP (noted $C_{TP \rightarrow FGP}$). The common movements may, however, be perfectly synchronous (instantaneous causality¹ noted $C_{TP \Rightarrow FGP}$) which would mean that both species influence each other's price. Granger Causality tests are normally performed on

¹ Note that in practical the two types of causalities may coexist.

VAR models with stationarised series, i.e. in levels if the series are $I(0)$ and in first-difference if they are $I(1)$. Lütkepohl (1993, pp. 378–379), however, showed that GC tests on integrated and co-integrated systems can be conducted on VAR in level.

2.2.4. Impulse response functions

The impulse functions (or dynamic multipliers) of a system are computed from the moving average (MA) form (i.e. the innovations) of the VAR model of that system (see Lütkepohl, 1993, pp. 97–99 and 379–382 for a more precise description). The impulse responses show the current and lagged effects on the variables of changes in the innovations. They are useful tools to identify the short-run adjustment dynamics of a multivariable system². In our case (if a real link exists between the two price series), the impulse response functions can be used to estimate the degree and duration of a shock affecting one price series to the other series. They will, therefore, be useful to determine to what extent the volatility of one price may actually affect the other price.

2.2.5. Error correction model estimation

The existence of a long-run equilibrium relation between two series implies the presence of a short-run dynamic adjustment process. The existence of this short-run adjustment is related to the fact that the two series cannot diverge too far from the other before market force operates to restore the equilibrium. This adjustment mechanism refers to the concept of ECMs initially developed by Davidson et al. (1978). When an ECM is represented by a VAR model, it is termed vector error correction (VEC) model. A matrix embodying the deterministic dummy (seasonal and outliers) can be added in the VEC estimation procedure. In our case, assuming that the two price series FGP and TP are linked by a long-run equilibrium, the estimation of the VEC model will permit to identify and quantify the short run dynamic adjustment process that takes place when one series diverges from the long run

common trend. Furthermore, the identification of the matrix including the dummy variables will allow us to identify the deterministic seasonal and/or isolated events that play(ed) a major role in the price auto and cross-determination.

3. Empirical results

3.1. Stationarity analysis

The visual analysis of the correlograms indicates the probable non-stationarity of the two series. Both the Akaike (AIC) and HQ criteria indicate that TP is AR(1) while FGP is AR(11). We, therefore, performed a simple DF unit root test on TP but an ADF test on FGP to account for the serial correlation. The DF test confirms the non-stationarity of the TP series in levels and the stationarity in first difference (Table 1). The test on series in levels also indicates that a constant has to be included in the process. For the FGP series, the ADF test confirms the non-stationarity of the series in levels and the stationarity in first difference (Table 2). Both series are thus integrated $I(1)$.

Shrimp consumption on the French market appears to be strongly seasonal (Josupeit, 1992). This seasonal demand pattern may, therefore, induce a seasonal component in the price dynamics. From Fig. 1, it seems that if a price seasonality exists, it is relatively weak. The two series were, therefore, decomposed in trend and seasonal components and seasonality was tested in the detrended series. The two price series display slight seasonalities. The Dickey et al. (1984) test was thus performed to test for the possible presence of seasonal unit roots. The result (not shown) indicates that the data do not exhibit non-stationary seasonality. Following Wallis (1974), we used seasonal-unadjusted data in the rest of the analysis in order to avoid distorting the dynamics of the estimated models.

3.2. Co-integration analysis

The monthly series TP and FGP were log-transformed (noted LTP and LFGP from now henceforth). As the data do not exhibit non-stationary seasonalities, the standard Johansen's procedure could be applied. In this case, the basic model accounts for deterministic seasonality through seasonal dummies in the VAR

² For stable systems (i.e. with no unit root), the impulse response coefficients converge asymptotically towards zero. In turn, for integrated or co-integrated processes, the existence of one or more unit root entails that the responses do not converge systematically to zero. Their use as descriptive tool of the short-run adjustment is, in this case, somewhat more puzzling. Nevertheless, the impulse response functions are still computable for unstable systems.

Table 1
Unit root test (DF test) on the TP series. Test procedure: if $t_\alpha < t_{tab}$, H_0 rejected

Case	H_1 estimated model	H_0 joint hypothesis	TP in levels	TP in differences	Dickey–Fuller (5%)
1	$y_t = \alpha + \rho y_{t-1} + \beta t + \varepsilon_t$	$\rho = 1, \beta = 0$	$t_\alpha = -2.944$	$t_\alpha = -10.78^a$	$t_{tab} = -3.46$
2	$y_t = \alpha + \rho y_{t-1} + \varepsilon_t$	$\rho = 1, \alpha = 0$	$t_\alpha = -2.812^a$	$t_\alpha = -10.84^a$	$t_{tab} = -2.89$
3	$y_t = \rho y_{t-1} + \varepsilon_t$	$\rho = 1$		$t_\alpha = -10.90^a$	$t_{tab} = -1.94$

^a H_0 rejected.

Table 2
ADF test on FGP series. Test procedure: if $t_\alpha < t_{tab}$, H_0 rejected

Case	H_1 estimated model	H_0 joint hypothesis	FGP in levels	FGP in difference	Dickey–Fuller (5%)
1	$y_t = \alpha + \rho y_{t-1} + \beta t + \varepsilon_t$	$\rho = 1, \beta = 0$	$t_\alpha = -1.60$	$t_\alpha = -5.26^a$	$t_{0.05} = -3.46$
2	$y_t = \alpha + \rho y_{t-1} + \varepsilon_t$	$\rho = 1, \alpha = 0$	$t_\alpha = 0.49$	$t_\alpha = -4.97^a$	$t_{0.05} = -2.89$
3	$y_t = \rho y_{t-1} + \varepsilon_t$	$\rho = 1$	$t_\alpha = -1.34$	$t_\alpha = -4.75^a$	$t_{0.05} = -1.94$

^a H_0 rejected.

model. The seasonal coefficients were centred with respect to the mean in order to maintain the asymptotic distribution in the rank co-integration tests. When restricting the model to include only seasonal dummies and an intercept, the null hypothesis of no autocorrelation in the residuals is accepted. Both the HQ and SC criteria indicate an autoregressive order $p=9$. The joint hypothesis of the rank order and the form of the model for deterministic components was then tested using the trace test and the Pantula’s principle (Pantula, 1989). The test procedure and the results are reported in Table 3. The test indicates that there is one co-integration relation and that the model must include an intercept μ in the co-integration relation. The statistics of the co-integration relation are given in Table 4. The estimated co-integration relation³ is

$$LFPG_t = -4.86 + 1.809LTP_t \tag{1}$$

3.3. Granger causality test

Given the autoregressive order $p=9$, the final estimated VAR(9) model is

$$LFPG_t = \sum_{i=1}^9 \alpha_i LTP_{t-i} + \sum_{j=1}^9 \alpha_j LFPG_{t-j} + \Psi D_t \tag{2.1}$$

³ The ADF test conducted on the residuals of the co-integration relation (1) indicates that the residuals are stationary. The DF statistic $t_\alpha = -3.88$ is smaller than the critical value $t_{tab} = -3.37$ at 5% tabulated by Engle and Yoo (1987).

Table 3
Trace test: determination of the rank order and identification of the deterministic components of the model^{a,b}

Joint hypothesis H_0	r	Models		
		1	2	3
Eigenvalues	0	0.1960	0.1834	0.2044
	1	0.0216	0.0008	0.0813
Trace test	0	20.399	17.295	26.640
	1	1.852 ^c	0.070	7.205
Critical values	0	17.794	13.308	22.946
	1	7.503	2.706	10.558

^a Three models are tested. Model (1): intercept in the co-integration relation but no deterministic trend in the levels. Model (2): deterministic trends in the levels. Model (3): trend in the co-integration relation. The test procedure relies on the Pantula’s principle: starting with the joint hypothesis H_0 : model (1) and $r=0$, one compares the trace with the critical values. If trace > critical values, H_0 is rejected and one carries on the test with the new joint hypothesis H_0 : model (2) and $r=0$, etc.

^b Critical values in Johansen and Juselius (1990).

^c No rejection of H_0 .

$$LTP_t = 0.624 \sum_{i=1}^9 \alpha_i LTP_{t-i} + \sum_{j=1}^9 \alpha_j LFPG_{t-j} + \Psi D_t \tag{2.2}$$

with t -values in brackets. The constant is not significant in Eq. (2.1). The GC and instantaneous⁴ causal-

⁴ For the instantaneous causality test, $j=0 \dots 9$ in Eqs. (2.1) and (2.2) and we tested for $\alpha_j=0$ with $j=0$.

Table 4
Estimation of the long run equilibrium relationship (normalisation with respect to LFGP)^a

	LTP	LFGP	Constant
ΔLTP	-0.271 (-4.248)	0.150 (4.248)	0.729 (4.248)
ΔLFGP	0.177 (2.236)	-0.098 (-2.236)	-0.477 (-2.236)
Cointegration vector β	-1.809	1	4.861
Loading matrix α	ΔLTP 0.150 (4.248)	ΔLFGP -0.098 (-2.236)	

^a (): *t*-statistics; matrix Π=αβ.

ity tests were then performed on the VAR models (2.1) and (2.2). The Thai market appears to Granger cause the FG market (Table 5) but the reverse is not supported. This result indicates that the Thai market precedes the FG market. No instantaneous causality was found significant, whatever the sense of the causality.

3.4. Impulse response functions

The impulses responses functions were estimated using the Klock and van Dijk’s (1978) procedure. The responses of LTP and LFGP series to 1-time unit impulses in variable were estimated from the VAR models (2.1) and (2.2). These response functions and their associated two-standard error bounds are displayed in Fig. 2. Of primary interest for the present analysis are the top-right and bottom-left diagrams. The top-right diagram depicts the response of LFGP following a shock imposed on LTP. The deviation on LFGP occurs 1 month after the shock to LTP. The impact then continues to increase during the following 3 months (i.e. four lags after the initial impulse). The response function does not return to zero due to the presence of the unit root. However, under the two-standard error criterion, only the coefficients of the lags 2–4 appear to be significantly different from zero. For the response of LTP to a 1-time unit impulse on LFGP (bottom-left

diagram) none of the response coefficients is significant.

3.5. Error correction model estimation

Under the constraint of one co-integration relation, the ECM in its Johansen’s formulation, is a VEC(8)⁵ in differences where the outliers and the dummy variables accounting for the seasonality of the series are also introduced. The estimated model is

$$\begin{aligned}
 & \begin{pmatrix} \Delta LTP_t \\ \Delta LFGP_t \end{pmatrix} \\
 &= \sum_{i=1}^8 \Gamma_i \begin{pmatrix} \Delta LTP_{t-i} \\ \Delta LFGP_{t-i} \end{pmatrix} \\
 &+ \begin{pmatrix} -0.271^* & 0.150^* & 0.729^* \\ 0.177^* & -0.098^* & -0.477^* \end{pmatrix} \begin{pmatrix} LTP_{t-1} \\ LFGP_{t-1} \\ 1 \end{pmatrix} \\
 &+ \begin{pmatrix} 0.143^* \\ 0.039 \end{pmatrix} D_{9103} + \begin{pmatrix} 0.151^* \\ 0.042 \end{pmatrix} D_{8909} \\
 &+ \begin{pmatrix} -0.104^* \\ -0.122^* \end{pmatrix} D_{9206} + \begin{pmatrix} -0.018 \\ -0.093^* \end{pmatrix} D_{9008} \\
 &+ \begin{pmatrix} -0.031 \\ -0.111^* \end{pmatrix} D_{9111} + \Psi D^\# \tag{3}
 \end{aligned}$$

where (*) denotes the significant coefficients with respect to *t*-statistics (5%) and the matrix *D*[#] accounts for the seasonality of the series (Table 6). The coefficients of the matrix Γ_{*i*} of the VEC model (3) are detailed in Eq. (4)

Table 5
Granger causality tests. Test procedure: if *F*<*F*_α=2.04, H₀ rejected; if *t*<1.96, H₀ rejected

Granger causality		Instantaneous causality	
H ₀	<i>F</i> -statistics	H ₀	<i>t</i> -statistics
LTP→LFGP	2.39	LTP⇒LFGP	-0.27
LFGP→LTP	1.12	LFGP⇒LTP	-0.27

⁵ Recall that the HQ criterion indicates a VAR(9) in levels.

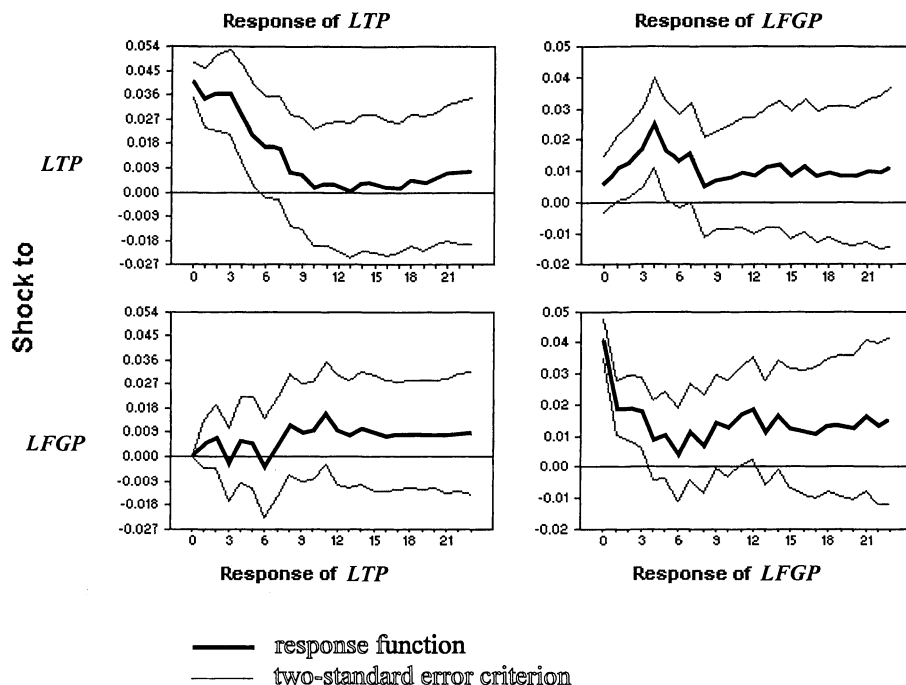


Fig. 2. Impulse response functions estimated from the VAR(9) model. Dark lines: impulse response coefficients, dashed lines: associated two-standard error coefficients. The top-right diagram shows the response of LFGP series to a 1-time impulse on LTP. The deviation on LFGP occurs 1 month after the shock on LTP. The impact increases during the following 3 months and does not return to zero (due to the presence of the unit root). Under the two-standard error criterion, only the coefficients of the lags 2–4 are significantly different from zero. Bottom-left diagram: response of LTP to a shock on LFGP: none of the response coefficients is significant.

Table 6
The seasonal coefficients^a

Month	January	February	March	April	May	June	July	August	October	November	December
ΔLTP	0.062 ^b	0.036 ^b	0.031	0.056 ^b	-0.011	0.044 ^b	0.047 ^b	0.030 ^c	-0.028	0.016	0.032 ^b
$\Delta LFGP$	-0.020	-0.013	-0.054 ^b	-0.041 ^c	-0.010	-0.002	0.017	-0.001	-0.019	-0.052 ^b	0.03

^a Only 11 dummy variables were included into the model because of algorithm procedure limitation (September coefficient not included). Significance tested by *t*-statistics.

^b Significant at 5%.

^c Significant at 10%.

Table 7
The outliers introduced in the VEC model

Dummy variable	Period	Series	Events	Reference
D_{8909}	September 1989	ΔLTP	Black tiger price bottoms out	Globefish (1990)
D_{9008}	August 1990	$\Delta LFGP$	First signs of overproduction in the French Guyana fishery	Béné and Moguedet (1996)
D_{9103}	March 1991	ΔLTP	Drug residue scandal in the Thai production	Csavas (1992)
D_{9111}	November 1991	$\Delta LFGP$	Social movements (strike) of the French Guyana fishermen	Béné (1997b)
D_{9206}	June 1992	ΔLTP and $\Delta LFGP$	Virus epidemic affects the Ecuadorian production	Globefish (1992)

$$\begin{aligned}
& \sum_{i=1}^8 \Gamma_i \begin{pmatrix} \Delta LTP_{t-i} \\ \Delta LFGP_{t-i} \end{pmatrix} \\
&= \begin{pmatrix} 0.069 & -0.040 \\ 0.020 & -0.443^* \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-1} \\
&+ \begin{pmatrix} 0.349^* & 0.072 \\ 0.122 & -0.191 \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-2} \\
&+ \begin{pmatrix} 0.214^* & -0.099 \\ 0.190 & -0.077 \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-3} \\
&+ \begin{pmatrix} 0.051 & 0.025 \\ 0.106 & -0.214^* \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-4} \\
&+ \begin{pmatrix} -0.199^* & 0.081 \\ -0.211^* & -0.248^* \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-5} \\
&+ \begin{pmatrix} 0.098 & -0.154 \\ -0.351^* & -0.121 \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-6} \\
&+ \begin{pmatrix} 0.296 & -0.083 \\ 0.096 & -0.051 \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-7} \\
&+ \begin{pmatrix} 0.031 & 0.145 \\ -0.015 & -0.062 \end{pmatrix} \begin{pmatrix} \Delta LTP \\ \Delta LFGP \end{pmatrix}_{t-8} \quad (4)
\end{aligned}$$

The expression (4) indicates that monthly change in the cultured shrimp price ΔLTP depends only on its own past changes (significant at lags 2, 3 and 5), while monthly change in the FG's shrimp price $\Delta LFGP$ depends on its own past changes (significant at lags 1, 4 and 5) but also on past changes of the cultured shrimp price (significant at lags 5 and 6). The LTP and LFGP series display slight significant seasonalities. These are displayed in Table 6. Seasonal coefficients are significant in December, April and June for LTP and in March and April for LFGP. Finally, five deterministic outliers $\{D_{9103}, \dots, D_{9111}\}$ were included in the VEC model (3) to obtain the normality⁶. Table 7 indicates the events associated to these outliers.

⁶ The goodness of fit of the estimated VEC model (3) was tested through the Ljung–Box and Lagrange multiplier (LM) tests. The χ^2 value of the Ljung–Box test is superior to the χ^2 tabulated, indicating an autocorrelation of the residuals at 5%. In turn, the LM tests indicate that this autocorrelation is neither of order 1 nor 4. The 20 first autocorrelation coefficients confirm the occurrence of some significant autocorrelations of higher order.

4. Interpretation and conclusion

Shrimp aquaculture has become a substantial component of the international shrimp industry. Comprising only 3% of the total world shrimp production in 1981, cultured shrimp composed over 30% in 1994. Total world production of farm-raised shrimp increased from about 1140% over the same period, growing from 59 000 million tonnes in 1981 to 733 000 million tonnes in 1994 (Anderson and Fong, 1997). The shrimp species that have proven especially suitable for culture are located mainly in tropical and semi-tropical regions. With the exception of China which produces a temperate species (*Penaeus chinensis*), shrimp farming occurs in tropical South-east Asian and Latin American countries and produce warm-water species. Thailand, Indonesia and India are currently the major cultured shrimp producing countries in Asia and Ecuador has been the major producer of cultured shrimp in Latin America for more than a decade. Most of this farmed production enters the export market. The main importers are Japan, the United States and Western Europe, accounting for about 85% of the world trade. The pattern of global trade for cultured shrimp thus presents a one way-flow, from developing tropical producers countries to the three main markets in the developed world.

A few papers have been released that describe the major changes which took place in shrimp markets and international trade following the expansion of farm-raised production (see for instance Csavas, 1992; Yang, 1994; Ling et al., 1996; Anderson and Fong, 1997) but these are mainly descriptive. Surprisingly, given the magnitude of the phenomenon, very little attention has been devoted to analysing the impact(s) of the expansion of shrimp farm-based production on the wild shrimp industry. Keithly et al. (1993) investigated the impact of shrimp aquaculture on US wild shrimp prices. Using a DSES model incorporating US import demand, US supply, US dockside demand and Japanese demand and supply relationships, they concluded that if there were no shrimp aquaculture, USA import prices in 1988 and 1989 would have been about 70% higher. Apart from this analysis, which unfortunately relied on a disputable mod-

elling approach, no other empirical analysis⁷ has been performed to evaluate and analyse the impact(s) that the growth in farm-raised shrimp industry has induced on the fishing industry exploiting wild species.

This paper attempted to address this question in the case of the wild brown shrimp *Penaeus subtilis* exploited by the FG fishery. This product is exported at 95% to the Rungis Fish market in Paris, France, where it struggles against several farmed raised species imported from Asia and Latin America (Béné and Moguedet, 1996). It competes in particular against the cultured black tiger shrimp arriving from Thailand. Given the fact that black tiger is now widely recognised as a price indicator on international markets (Globefish, 1990) one could expect the price of FG shrimp to be affected by the evolution of the black tiger price. To test that hypothesis we used the co-integration approach applied to the price time-series of the two species over the period 1986–1993.

The analysis pointed out several interesting conclusions regarding the price behaviour of the FG product. First, considering the FG series individually, the fact that FG is integrated of order 1 with a 11-months lag in the autoregressive process suggests the case of a small and specific market where the few buyers and sellers keep in mind the price value over approximately 1 year and take it into account for the current transaction.

Secondly, when considered along with TP, the analysis indicates that the two series are co-integrated. The existence of the co-integration relation corroborates the hypothesis of a common long-run equilibrium relationship between the two markets.

In econometrical terms, since we used the log-transformed series, the co-integration vector value 1.809 (see Eq. (1)) represents the cross-elasticity of demand for *P. subtilis* with respect to *P. monodon*. The positive sign of the cross elasticity implies that the two species are substitutes, which is in accordance with expectations. However, if we accept the hypothesis that this value is significantly different from unity, this indicates that the relation between the two raw (i.e. not log-transformed) prices is not linear. In particular, it means that when prices are

high, the relation is in favour of the FG species. This may be explained by the combination of two mechanisms. First, as already mentioned, FG shrimp, like most of the wild species, is known to provide higher organoleptic qualities than cultured species. At the same time, the large commercial categories, characterised by relatively higher prices, are usually purchased by consumers who look for high quality products. These consumers are thus willing to pay a higher price for this taste difference and therefore preferentially purchase wild shrimp. In turn, less expensive shrimp (which correspond to smaller commercial categories) are purchased by consumers who are interested in 'more affordable' quality-price ratios. Consequently, for these smaller categories, the low production cost cultured shrimp are favoured.

Even if this consumer behaviour is described here with respect to categories preference, it may also be valid for time preference: when shrimp prices are globally high for all categories, this sea-product (and especially the farm-raised species), lose their attractiveness for the consumers who usually purchase them for their attractive quality/price ratio. Consequently, during these high price periods, only consumers who look for the luxurious component of the market continue to purchase shrimp, and they turn for this purpose to the wild shrimp, renowned for their higher organoleptic qualities.

This consumer behaviour, which induces the non-linearity observed in the equilibrium relation between the wild and the cultured shrimp prices, may represent an opportunity for the FG shrimp companies to alleviate the dominance of the black tiger on their product. We shall return to this point later in this discussion.

The analysis of the others tests (ECM estimation, impulse function calculation, and causality tests) provide a more thorough insight into the nature of the long-run equilibrium relationship that exists between the two markets. They indicate in particular that the relationship is not 'symmetrical' and that, as presumed, the black tiger price does play a significant role in the price determination of the FG brown shrimp both at the long-run and short-run scales. First, the Granger causality analysis reveals that the black tiger price precedes the FG price in their common trend and therefore acts as a leader for the FG shrimp price. Secondly, the impulse response functions derived from

⁷ For theoretical analyses, see Anderson (1985) or Berk and Perloff (1985).

the VEC model indicate that the FG shrimp price responds significantly to shocks occurring on the Thai market. This was confirmed by the ECM which indicates that monthly changes in the FG price depends on the FGP series' own past deviations but also to past deviations in the black tiger price series.

All these results indicate that, on the Rungis market, the price of the FG product is largely determined by the black tiger price. In particular, fluctuation in the brown shrimp price appears to be largely induced by shocks occurring on the black tiger shrimp market.

Unfortunately, black tiger, like farmed shrimp in general, although benefiting from year-round production schemes, are known to display some degree of instability, both in price and production levels. Anderson and Fong (1997, p. 32), for instance, refers to the 'Boom-or-bust' nature of farmed shrimp industry and there are many examples of production collapses in the literature, that occurred even in the group of top exporter countries (e.g. China passed from 199 000 million tonnes in 1992 to less than 50 000 million tonnes in 1993, and Taiwan passed from 114 500 million tonnes in 1987 to 32 000 in 1989). Several factors can be identified as causes of this variability in production. These factors range from the micro to global scale. At the micro-level, self-pollution generated by overfeeding, raised-temperature farming, overstocking, misuse of antibiotics, lack of sanitation control, and ultimately disease, are some of the common factors that can affect the production (Csavas, 1994). At the meso-level, urban and industrial pollution, over-development and general degradation of the coastal habitat are usually identified as major causes of production fluctuation (Flaherty and Karnjanakesorn, 1995). Finally, at the macro level, the possible spreading out of viral epidemic through the whole countries, or climatic events such as El Niño also influence the production level. For instance, in Ecuador, during El Niño years, the shrimp larvae used for natural stocking are usually more abundant, due to the warmer waters brought by the El Niño phenomenon. This generally leads to an increase in the production. On the other hand, in Asia, El Niño events have been shown to modify the monsoon pattern and induce severe droughts, which negatively affect the shrimp production.

This volatility, which affects market prices on a random basis, is further exacerbated by periodic 'tensions' induced by the highly seasonal nature of

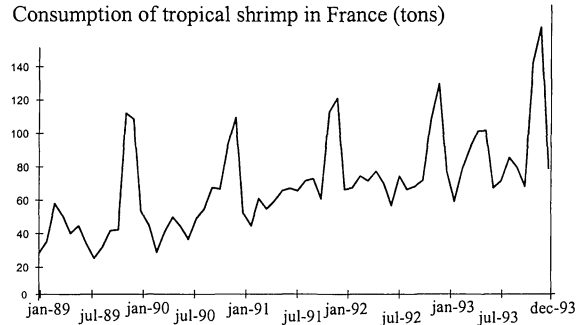


Fig. 3. Consumption of tropical shrimp in France (monthly data in tonnes) over the period 1989–1993. The series shows the general increasing trend that has characterised the French market over that period but also the high seasonality of the demand. Tropical shrimp is essentially consumed in France during Christmas and New year eve's periods (source: FIOM Paris).

the demand on the importer markets. In Japan for instance, consumption rates are particularly high during the golden week (29 April–9 May) and Christmas holiday period (Josupeit, 1999). European markets are also characterised by a high seasonality in the demand. In France, analyses indicate that the tropical shrimp is essentially consumed during Christmas and New year eve's periods. For illustration, Fig. 3 represents the consumption of tropical shrimp on the French market over the period 1989–1993. The figure does illustrate the high seasonality of the demand with a significant peak in December.

In this context, one of the main constraints faced by importers is the adjustment of the supply to the market dynamics and their objective is to ensure an adequate supply of shrimp that will permit to follow, or even anticipate, the demand dynamics⁸. In this respect, it is widely admitted that due to the possibility of control of their inputs, shrimp farms can plan (to a certain extent) their production scheme and therefore are more reliable and flexible than fisheries, where landings is usually seasonal and especially unreliable,

⁸ In Japan, in order to minimise uncertainties resulting from shrimp trading, vertical co-ordination mechanisms such as contractual prices are now commonly used between Japanese shrimp importers and Asian shrimp packers. The use of these contractual price shifts trade from a spot market to a situation of bilateral contracting, which creates a substantial degree of dependence between buyers and sellers (see Ling et al., 1998 for a more detailed analysis).

even in tropical waters. This high variability is due to the fact that for most wild species, the larvae and post-larvae stages grow in inshore nurseries which appear to be largely under the influence of continental environmental factors (rainfall volume, river discharge, etc.). Consequently, the catch volumes fluctuate considerably on a monthly and seasonal basis, making the supply of the shrimp fisheries particularly unreliable and completely exogenous for the traders.

In the case of the wild brown supplied by the FG companies, the problem of adjustment between local production and French markets demand is further exacerbated by the type of seasonality which affects the fishery landings. As emphasised by Béné and Doyen (2000), the seasonal oscillations of the FG landings are in fact completely dephased with the market dynamics, inducing seasonal storage problems: the period of high landings (February–April) occurs during the period of relatively low demand on the French markets, while the period of high demand (November–December) corresponds to the period of lowest landings in the fishery. The negative seasonal coefficients identified in the ECM for the brown shrimp series in March and April (see Table 6 for recall) are thought to reflect these storage problems.

The dephased dynamics between the FG supply and the French market thus represents a major drawback for the fishery which can not supply shrimp in accordance with the markets dynamics. However, market analyses (Anon., 1991; Josupeit, 1992) also reveal that, along with this supply–demand adjustment issue, farm-raised products provide two other major commercial advantages over wild-caught shrimp for the Mediterranean-type markets⁹: (1) a better flexibility in terms of size-categories with respect to the consumer preference, and (2) a better resistance to processing. First, regarding the consumer's size preference, it is clear that it is easier for farm producers to monitor the growth of shrimp in basins and subsequently to programme the harvest time than it is for trawler skippers to control the size of the wild shrimp caught in their nets. Secondly, with respect to the processing aspect, it seems that farm-raised shrimp better resists (in terms of texture and external aspect) the

different treatments and preparations that shrimp undergoes during the processing stages.

All these reasons drove the French importers to favour farm-raised species over wild species and explain why aquaculture shrimp in general, and the black tiger species in particular, despite their lower culinary quality, were able to 'dethrone' the wild species in only a decade and become the new leaders on the international markets. As demonstrated here in the case of the black tiger and FG wild shrimp, this change in leadership caused the price of the wild-caught species to be now largely determined by the price of the farm-raised product. In particular, it has been shown that a large part of the volatility affecting the FG shrimp results from shocks occurring on the Thai shrimp price.

This result points out an important consequence in term of intervention policies. It indicates that neither the local companies nor the policy-makers at the French national level would be capable of mitigating this exogenous volatility through a price stabilisation policy. The success of such price policy, whatever its theoretical effectiveness and empirical suitability with respect to the specific situation of FG, will always depend on the stability/volatility of the farmed shrimp markets at the international level. These are, however, exogenous, that is to say, out of control for the FG system, and depend on a series of various and combined factors affecting the farm-raised producers at the micro, meso and global levels in their respective countries. Therefore, the only way for local FG exporters to eliminate this exogenous volatility would be to 'cut' the relationship that links the price of the FG product to that of the aquaculture shrimp. This could be attained by creating a 'niche' on the French market where the FG product could be supplied without having to compete against the farm-raised shrimp. The comparative advantage of the brown shrimp in terms of culinary quality and the consumer asymmetrical behaviour evidenced by the non-linearity in the substitute relationship both suggest that a possible commercial strategy would be to target high standard restaurants or even individual consumers seeking high organoleptic quality products and willing to pay higher prices for these products. The development of this niche could be achieved through a marketing programme based on quality labelling and directed to the relevant segment of the market. The niche is likely to be narrow and specific but the modest amount of shrimp exported by

⁹ Recall that Mediterranean consumers prefer large-sized, warm-water, cooked or grilled shrimp.

the French Guyana (4000 tonnes per year) should facilitate the feasibility of the operation. This commercial strategy should allow the FG shrimp sector to alleviate the dominance that farm-raised species impose on its product's price. It will not, however, solve the dephased supply–demand dynamics that characterised the FG case. Within this market niche, the consumption of the FG species, promoted and sold as luxurious product, is indeed very likely to be even more concentrated on Christmas and New year eve's periods than it is now, when both aquaculture and wild-caught sectors are basing their development expectations on the expansion and democratisation of the market. The FG sector will thus have to find a way to reduce the discrepancy that exists between its seasonal production and the French market. This could be achieved through the development of adequate storage capacities and management. Béné and Doyen (2000) provide a theoretical analytical framework regarding this issue.

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