Investigating the dynamic relationships between wild shrimp and cultured shrimp prices

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Abstract:

The arrival in the 1980s of cultured shrimp has induced large changes on the World shrimp market. In this new environment, the French Guyana fishery production now competes strongly with some of these farm-produced shrimp. Consequently, this fishery has experienced serious export problems during the last decade. This paper is an attempt to analyse the specific impact of the arrival of cultured shrimp to the World markets on the French Guyana domestic market price and on the fleet activity. More particularly, we analyse the relationship between this market and one cultured shrimp market usually considered as benchmark, the Thai market. We use econometric techniques to test whether these two markets are co-integrated and whether causalities exist between them. For these tests, we rely on the two monthly price series recorded over the period 1986-93. The tests indicate that the two series are non-stationary and co-integrated. Therefore, a long-term relationship exists between the two markets, and the identification of the valid dynamic model requires a two-dimension error correcting model. The analysis of the impulse responses derived from the model indicates that the French Guyana market responds significantly to shocks on the Thai market. In other respects, a significant causality from the Thai market to the French Guyana market has been identified. Finally, we discuss these results by briefly analysis the main consequences they induced for the fishery’s dynamics.

Introduction

The French Guyana shrimp fishery is a trawl fishery that has been exploiting the stock of two Penaeid shrimp species (*Penaeus brasiliensis* and *Penaeus subtilis*) off the coast of the French Guyana[1] since the early 1960s (see Dintheer et al. (1991) for the fishery's history review). The fishery is organised like an industrial fishery. The fleet which is made up of about 70 vessels owned by several companies, lands annually around 4,000 tons of shrimp to the port of Cayenne (the main town of the French Guyana). The fishing season covers the whole year with no seasonal closure and is characterised by a six-month period of high catch rate (from January to June) and a six-month period of lower catch rate (from July to

[1] The French Guyana is a French over-seas department located at the north-eastern border of Brazil.
December). The fishery can be considered as a single-species fishery since *Penaeus subtilis* represents now about 95% of the catches. Since the mid 80s, the global resource access and the fleet effort are limited by a licence system combined to an annual TAC. The production that used to be exported to the Japanese and US markets is now exported toward the European market and mainly purchased by French processors. This market re-orientation is an indirect result of the «francisation» program through which, from 1985 onwards, the French government substituted the American and Japanese companies present in the fishery by French companies (Béné 1996). The French Guyana shrimp used to enjoy very high commercial values that compensated the small catches volumes and induced that Cayenne was the 5th French fishing port (in value terms) at the end of the eighties. However, important financial and commercial difficulties have entailed a 28% decrease (in value terms) between 1990 and 1993 (IEDOM 1993).

During the same period, the world shrimp market has been experiencing very important qualitative and quantitative modifications due to the massive arrival of farm-produced shrimp, essentially from Asia and South-America. One major change was the new influence that the price of these cultured shrimp impose on the world market. It is indeed now widely-admitted that large part of the price fluctuations on Tokyo and New-York shrimp marketplaces is closely related to some cultured species considered since few years as market indicators. Amongst these cultured species that largely influence the world shrimp marketplaces, there are the «Black Tiger» *Penaeus monodon* from Thailand and the Ecuadorian *Penaeus vannamei*.

These species are precisely the two species against which the French Guyana’s species *Penaeus subtilis* competes on the French market (Anon 1991).

It appears therefore interesting to try to analyse the effect of the recent expansion of the cultured shrimp on the world shrimp markets. More precisely, and with specific regard to the French Guyana fishery, one may wonder whether the prices of this small shrimp fishery is actually influenced by the dynamics of one of the cultured shrimp considered as world-market price indicator. The plan of the paper is therefore organised to investigate this question. Firstly, a description of the cultured shrimp arrival on the world market is proposed, and we review the main commonly-admitted reasons of their commercial predominance on the wild shrimp species. We then rely on the co-integration approach to test whether there exists a relation between the price of the French Guyana shrimp and that of the Thai Black Tiger. The econometric procedures of the time series analyses are then presented and the results are detailed. Finally, we discuss these results by briefly analysis the main consequences they induced for the fishery's dynamics.

### Impact of the cultured shrimp on the markets

#### Arrival of the cultured shrimp on the world markets

Between 1980 and 1990, the world total shrimp production increased from about 1.6 millions tons to more than 2.6 millions tons. Shrimp represents now more than 30% (in value terms) of the total sea products traded over the world (Anon 1993). This significant growth of shrimp on the sea products markets has been resulting from a simultaneous increase of the demand and the supply. However, the volume of wild captured shrimp remained constant during this period (Fig.1).
This suggests that the increased demand was thus met with the addition of the cultured shrimp (Csavas 1992). It is for instance estimated that the shares of cultured shrimp on the two main markets (the Japanese and US markets) represent now more than 46% and 50% of their respective shrimp imports (Ferdouse 1992).

The European market, which is the third world market, was reached by cultured shrimp somewhat later than the US and Japanese markets. The European demand however seems to accelerate and E.U. imports expanded from about 132,000 tons in 1981 to 365,000 tons in 1991 (Anon 1993). Historically, the cold species has dominated in most of the European countries before 1980. Nevertheless, the tropical shrimp species have made significant inroads into some of the E.U. countries during the last decade, essentially through the arrival of new cultured shrimp species. It is admitted that the most significant change in E.U. shrimp imports and consumption is precisely the emergence of cultured shrimp (Hottlet 1992).

With respect to shrimp consumption, one can 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).

**The production and commercial advantages of the cultured shrimp**

The predominance of the cultured shrimp on the wild shrimp may be explained by the following reasons (Csavas 1992):

- **The year-round production of cultured shrimp.**

Shrimp catch is usually seasonal and especially unreliable, even in tropical waters. This high variability is due to the fact that for most species, the larvae and post-larvae stages growth in inshore nurseries which appear to be under large influence of continental climatic factors. Consequently, the catches volumes may fluctuate considerably from one year to the other and entails idle processing capacities over extended periods. Conversely, cultured shrimp production can be planned, and may even be continuous in the tropics. This improves the efficiency of the processors and reduces global storage costs.

- **The flexibility of the farm production.**

Farm offers a better opportunity for the industry to adapt to the consumers' demand regarding species and sizes. Shrimp catch is usually determined in this respect by geography and season. Shrimp farmers can select more flexibly from a wide range of cultivable species and produce the sizes required.

- **The autolysis problem.**

Because of the rapid autolysis that starts in crustaceans immediately after their death, prompt icing is pursued on shrimp trawlers. However, during each trip, in addition to gear manipulating, the main task for the crew consists of separating the shrimp from the by-catch, that usually represent more than 90% of the total trawler catch, and of sorting the shrimp into different commercial categories before the icing procedure. As a consequence, it is not rare to observe wild shrimp catches affected by commercial devaluation following autolysis problems.
All these reasons may explain why, even if cultured shrimp is known to offer relatively poorer culinary qualities than wild shrimp, it is systematically preferred by processors. This preference is particularly obvious in the French market. Asia expanded its cultured shrimp imports from 17% in 1981 to 38% in 1993, and Latin America from 8 to 25.6% (Anon 1993). In the same time the West-African captured shrimp, that used to be the main supplying region to the French market with more than 35% of the French imports in 1981, saw its share declined to 10% in 1992. The most impressive growth in exports of shrimp to the French market was enjoyed by Ecuador (from nil in 1986 to about 6,000 t in 1993) but Thailand’s exports also grew significantly (1,700 t in 1986 to 6,800 t in 1991) (Josupeit 1992).

Econometrics method

The co-integration techniques are an econometric «package» which is now classically used to analyse simultaneously the long-run and short-term interrelations that are thought to exist between several time series. We therefore apply these techniques to analyse and quantify the dynamic relationships that might appear to link the French Guyana shrimp market to the Thai shrimp market. The estimation method comprises four successive steps that are (1) the stationary analysis, (2) the co-integration analysis, (3) the measure of causalities and estimation of the impulse response functions, and (4) the estimation of the Error Correction Model (ECM).

Stationary analysis

The first step consists in determining the integration orders of the series through unit root tests. One most commonly used test is the Dickey-Fuller (DF) test and its extension, the Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1981). The auto-regressive order is determined by minimising the Akaike (MC) and Hannan Quinn (HQ) criteria. Three situations have to be considered, regarding to the possible presence of an intercept and a time trend in the process. This amounts of considering successively the three following cases (where \( \varepsilon_t \) is assumed to be a white noise for the DF test):

**case (1): True model** \((H_0): y_t = \alpha + \beta y_{t-1} + \varepsilon_t\) **Estimated model** \((H_1): y_t = \alpha + \rho y_{t-1} + \beta t + \varepsilon_t\)

\(H_0: \) joint hypotheses: \(\rho = 1, \beta = 0\). If \(H_0\) is not rejected then we consider the case (2)

**case (2): True model** \((H_0): Y_t = y_{t-1} + \varepsilon_t\) **Estimated model** \((H_1): y_t = \alpha + \rho y_{t-1} + \varepsilon_t\)

\(H_0: \) joint hypotheses: \(\rho = 1, \alpha = 0\). If \(H_0\) is not rejected then we consider the case (3)

**case (3): True model** \((H_0): Y_t = y_{t-1} + \varepsilon_t\) **Estimated model** \((H_1): y_t = \rho y_{t-1} + \varepsilon_t\)

\(H_0: \) joint hypotheses: \(\rho = 1\). If \(H_0\) is not rejected, the series is non-stationary.

If a series is found to be non-stationary (occurrence of a unit root), the test is then re-applied to the series in first-difference. The differencing procedure is then pursued as long as the series is non-stationary. The number of differencing procedures necessary to stationarise the series indicates the integration order of
Testing for cointegration between integrated series can be performed through the Johansen's approach (Johansen 1988, 1989). This technique infers the statistical properties of the co-integration vector by linking this vectors to the canonical correlation between the levels and first differences of the process. This process which is adjusted with lagged differentials may furthermore include deterministic components (a drift $\mu$ and several possible exogenous variables $D_t$). The Johansen’s procedure, which is based on Maximum Likelihood (ML) approach, is now widely adopted. We thus merely recall its theoretical background:

Consider the $K$-dimensional VAR($p$) in levels with a drift $\mu$ in the process:

(1) $Y_t = \sum \Pi_i Y_{t-1} + \mu + \epsilon_t$

where $Y_t = (Y_{1t}, Y_{2t}, ..., Y_{Kt})$’ is assumed to possess no more than one unit root in each of its component series $Y_{kt}$. (1) can be rewritten in first differences:

(2) $\Delta Y_t = \sum \Gamma_i \Delta Y_{t-1} - \Pi Y_{t-p} + \mu + \epsilon_t$

where $\Gamma_i = -\Pi_i + \Pi_i + ... + \Pi_i$ for $i = 1, ..., p-1$ and $\Pi = \Pi_1 - \Pi_2 - ... - \Pi_p$. Since $\Delta Y_t$ is stationary, the stationarity of the right-hand side of (2) will depend on $\Pi Y_{t-p}$ and on the rank of $\Pi$. If $\Pi$ is a full-rank matrix, then $k$ linear independent combinations of $\Pi Y_{t-p}$ must be stationary. This implies that the components of $Y_t$ cannot be co-integrated since no more than $K-k$ independent co-integrated relations have to be observed in the process. On the other hand, if $\Pi Y_{t-p}$ is stationary, each component of $Y_t$ must be stationary around a trend. The VAR model can therefore be written in levels but must include a trend. In turn, if the rank of $\Pi$ is zero, no linear combination of $Y_t$ is stationary. A VAR model in differences has thus to be estimated.

The last case corresponds to the situation when the rank of $\Pi$ is equal to $r$, with $1 < r < K-k$. In this case, the matrix $\Pi$ can be decomposed as $\Pi = \alpha \beta$ where $\alpha$ and $\beta'$ are two $(K \times r)$ matrices. $\alpha$ is termed the loading matrix and $\beta$ is the co-integrated matrix. The basic idea of the Johansen's approach is to determine the rank of $\Pi$ by estimating its $p-r$ smallest eigenvalues and testing whether the latter are significantly different from zero. It is possible to derive the Likelihood Ratio (LR) statistics for testing a specific co-integration rank $r = r_0$ of a VAR($p$)) process against a larger rank, say $r = r_1$. The general test is:

$H_0: r = r_0$ against $H_1: r_0 < r \leq r_1$

More precisely, Johansen (1988) and Johansen and Juselius (1990) propose the two following tests:

- the Trace test: $H_0: r = r_0$ against $H_1: r_0 < r \leq K$
- the $\lambda$-max test: $H_0: r = r_0$ against $H_1: r = r_0 + 1$

Johansen (1988) has shown that the asymptotic distribution of the LR statistic is not a $\chi^2$ distribution and that it depends on the difference $K-r$ between the
dimension of the process and its co-integration rank. This distribution also depends on the value of the intercept μ in the model (2).

The asymptotic percentage points have been tabulated for model with no intercept (μ = 0) for r from 0 to 5 by Johansen (1988) and for r from 0 to 10 by Hall et al. (1989). Johansen and Juselius (1990) tabulated the critical values for intercept in the VAR process (μ ≠ 0). The case where the intercept term is contained in the cointegration relation also requires specify critical values (Johansen and Juselius 1990). An analysis of this case is proposed in Johansen (1991).

**Measure of causalities and identification of the impulse response functions**

The fact that a cointegration relation is shown to exist between two series allows us to study the causality, in the Granger sense, between these two series. The classical Granger causality (GC) test is performed on a VAR model with stationarised series. However, it can be shown (Lütkepohi (1993: 378-379) that CC tests on integrated and cointegrated systems may be conducted like on stable systems, i.e. performed on VAR models in levels.

The impulse functions, which are derived from the MA form of the VAR model, turn out to be useful tools to describe the short-run adjustment dynamics of the system. Indeed, impulse responses show the current and lagged effects on the variables of changes in the innovations. For stable processes, the response coefficients converge asymptotically towards zero. In turn, for integrated or cointegrated VAR 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 and the associated variance decomposition, are still available for the unstable systems.

**Error Correction Model estimation**

If the series have been shown to be co-integrated C(1,1), this means that there exists an effective long-run equilibrium relation from which the variables can not move away for too long. It thus become interesting to try to identify the adjustment process by which these variables come back to this equilibrium relation. This can be done through the estimation of Error Correction Models (ECM). ECM models appear to be closely related to the co-integration concept and to VAR models through the Granger representation theorem (Engle, and Granger 1987). When ECM are represented by VAR models, they are termed VEC (Vector Error Correction) models, the general structure of which reads in first-differences:

\[
\Delta Y_t = \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} - \Pi Y_{t-i} + \mu + \varepsilon_t
\]

**Empirical results**

**The data series**

The French Guyana data were obtained from the Cayenne's port authorities (CCI French Guyana). The price series was derived from the monthly total values (in French Francs FF) declared by the companies, divided by the total exports...
volumes (in kg):

Raw Unit Price = exports in value / exports in volume

for the period during which these series were available, that is, from January 1986 to August 1993. However, the export commercial value depends on the size of the shrimp exported. The size index of the landed shrimp, estimated from the landings, was used as a proxy for the average size of the shrimp exports. The raw series was thus weighted by this size index according to the correcting formula:

Corrected Unit Price = (Raw Unit Price x 100) / size index

We obtain the final French Guyana monthly price series used in the rest of the analyses. This series is noted FGP (for French Guyana Price) from now henceforth and represented on Fig.2.

FIG. 2 here

As far as the cultured shrimp series is concerned, we used the data series of the French shrimp imports recorded by the Food and Agriculture Organisation (FAO-Globefish) in $US, weighted by the FF/$ exchange rate. As indicated in the introduction, both the Ecuadorian and Thai shrimp are now regarded as price indicators on most marketplaces. We might therefore have chosen indifferently one or the other species' prices. However, the Ecuadorian P. vannamei has arrived more recently on the French market and the series of its price is only available since January 1991. In turn, the price series of the Thai shrimp P. monodon covers the whole period 1986-1993. This is the reason why we used the Thai shrimp price H> in the rest of the analysis. Note however that the Ecuadorian and Thai prices display very closed movements, at least for the period where both series were available (c.f. Fig.3).

FIG. 3 here

The TP series is plotted on Fig.2 with the FGP series. The graphical observation is in agreement with the expectation: the two prices show a common long-run trend. In the short-term, however, the two series display separate dynamics that can diverge from one another, but only for short periods.

Stationarity analysis

The visual analysis of the correlograms indicates the probable non-stationary of the two series (details on the data and complementary statistics results are available from the authors). The (MC) and (HQ) criteria were used to determine the autoregressive order of the two series. They both indicate that TP is AR(1) while FGP is AR(1). We therefore performed a simple DF unit root test on H> but an ADF test on FGP to account for the serial correlation. The DF test confirms the non-stationarity of the H> series in levels and the stationary in first difference (Table 1). The P series is therefore 1(1). The test on series in levels also indicates that a constant a has to be included in the process.
Table 1. Unit root test (DF test) on the TP series. If $t_\alpha < t_{\text{tab}} H_0$ rejected

<table>
<thead>
<tr>
<th>Cases</th>
<th>TP in levels</th>
<th>TP in differences</th>
<th>Dickey-Fuller (5%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td>$t_\alpha = -2.944$</td>
<td>$t_\alpha = -10.78$</td>
<td>$t_{\text{tab}} = -3.46$</td>
</tr>
<tr>
<td>(2)</td>
<td>$t_\alpha = -2.812, \alpha \neq 0$</td>
<td>$t_\alpha = -10.84$</td>
<td>$t_{\text{tab}} = -2.89$</td>
</tr>
<tr>
<td>(3)</td>
<td>$t_\alpha = -10.90$</td>
<td>$t_{\text{tab}} = -1.94$</td>
<td></td>
</tr>
</tbody>
</table>

For the FGP series, the ADF test confirms the non-stationarity of the series in levels and the stationary in first difference (Table 2). The series is thus 1(1).

Table 2. ADF test on FGP series. If $t_\alpha < t_{\text{tab}} H_0$ H$_0$ rejected

<table>
<thead>
<tr>
<th>Cases</th>
<th>FGP in levels</th>
<th>FGP in difference</th>
<th>Dickey-Fuller (5%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1)</td>
<td>$t_\alpha = -1.60$</td>
<td>$t_\alpha = -5.26$</td>
<td>$T_{0.05} = -3.46$</td>
</tr>
<tr>
<td>(2)</td>
<td>$t_\alpha = 0.49$</td>
<td>$t_\alpha = -4.97$</td>
<td>$T_{0.05} = -2.89$</td>
</tr>
<tr>
<td>(3)</td>
<td>$t_\alpha = -1.34$</td>
<td>$t_\alpha = -4.75$</td>
<td>$T_{0.05} = -1.94$</td>
</tr>
</tbody>
</table>

Both series are integrated 1(1). However, the shrimp consumption on French market being strongly seasonal, we might wonder whether the seasonal demand pattern does not induce a seasonal component in the price dynamics. From Fig.3, it seems that if a price seasonality exists, it is relatively weak. We applied a 12-terms moving average procedure to decompose the two series in trend and seasonal components and test for seasonality in the de-trended series. It emerges that the two price series display slight seasonalities. Having established that the series are 1(1), we thus had to test for the possible presence of seasonal unit roots. For this purpose, we applied the test proposed by Dickey, Hasza and Fuller (1984). The results indicate that the data does not exhibit non-stationary seasonality. So in the rest of the paper, we use seasonal-unadjusted data, as suggested by Wallis (1974), in order to avoid distorting the dynamics of the estimated models.

Cointegration analysis

The monthly series were analysed after log-transformation of the two raw series. They are noted LH$>$ and LFGP. The (HQ) criterion indicates an auto-regressive order $p = 9$. As the data have been shown not to exhibit non-stationary seasonalities, we can ignore the seasonal integration problem and apply the standard Johansen’s procedure on $L'P$ and $LFGP$. In this case, the basic model accounts for deterministic seasonality through seasonal dummies in the VAR(9) model. The seasonal coefficients were centred with respect to the mean in order to maintain the asymptotic distribution in the rank co-integration tests. The model in differences then reads as follows:
where $D_t$ is the matrix which embodies the deterministic dummy (seasonal and outlier) variables. A preliminary test indicates that the model must include an intercept $\mu$ in the co-integration relation. The Trace and $\lambda$-max tests were performed on the model (4). They both indicate the existence of one co-integration relation (c.f. Table 3).

Table 3. Results of the LR Trace and $\lambda$ - max tests for the model (4).

<table>
<thead>
<tr>
<th>Eigenvalues</th>
<th>$H_0$: r</th>
<th>Trace</th>
<th>Trace_90</th>
<th>$\lambda$ - max</th>
<th>$\lambda$ - max_90</th>
</tr>
</thead>
<tbody>
<tr>
<td>0.2273</td>
<td>0</td>
<td>25.37</td>
<td>17.79</td>
<td>21.41</td>
<td>10.29</td>
</tr>
<tr>
<td>0.1297</td>
<td>1</td>
<td>3.96</td>
<td>7.50</td>
<td>3.96</td>
<td>7.50</td>
</tr>
</tbody>
</table>

Critical values: Johansen and Juselius (1990)

The statistics of the matrix $\Pi = a\beta$ are given in Table 4. The estimated co-integration relation is:

(5) $LFGP_t = -4.86 + 1.809 LTP_t$

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

Table 4. Estimation of the long run equilibrium relationship.

(normalisation with respect to $LFGP$)

<table>
<thead>
<tr>
<th>Matrix $\Pi$</th>
<th>$LTP$</th>
<th>$LFGP$</th>
<th>Constant</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\Delta LTP$</td>
<td>0.271 (-4.248)</td>
<td>0.150 (4.248)</td>
<td>0.729 (4.248)</td>
</tr>
<tr>
<td>$\Delta LFGP$</td>
<td>0.177 (2.236)</td>
<td>-0.098 (-2.236)</td>
<td>0.477 (-2.236)</td>
</tr>
</tbody>
</table>

Loading matrix $\alpha$

| $\Delta LTP$ | 0.150 (4.248) |
| $\Delta LFGP$ | 0.098 (-2.236) |

Co-integration vector $\beta$

<table>
<thead>
<tr>
<th>$LTP$</th>
<th>$LFGP$</th>
<th>Constant</th>
</tr>
</thead>
<tbody>
<tr>
<td>-1.809</td>
<td>4.861</td>
<td></td>
</tr>
</tbody>
</table>

Measure of Causality and impulse response functions identification

Recall that GC tests can be conducted on VAR in levels even for co-integrated systems. Given that the auto-regressive order $p = 9$, the final estimated VAR(9) model is:
Note that the constant is not significant in (6.1). CC tests and instantaneous causalities tests were then performed based on (6.1 - 6.2). The Thai market appears to Granger cause the French Guyana market (Table 5) while the reverse is not verified. This result indicates that the Thai market precedes the French Guyana market. In turn, no instantaneous causality is significant, whatever the sense of the causality.

Table 5. Results of causality tests.

<table>
<thead>
<tr>
<th>Granger causality</th>
<th>Instantaneous causality</th>
</tr>
</thead>
<tbody>
<tr>
<td>H0</td>
<td>F statistics</td>
</tr>
<tr>
<td>LTP → LFGP</td>
<td>2.39</td>
</tr>
<tr>
<td>LFGP → LTP</td>
<td>1.12</td>
</tr>
</tbody>
</table>

The responses of LTP and LFGP series to one-time unit impulse (forecast error) in variable were estimated from the VAR model (6.1 - 6.2). The response functions and their associated two-standard error bounds are displayed on Fig.4.

The top right diagram depicts the response of LFGP produced by a shock to L'P. The diagram shows that the deviation on LFGP occurs 1 month after the shock to LTP. The impact then continues to increase during the 3 following months (i.e. 4 lags after the initial impulse). We furthermore observe that the 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, 3, and 4 appear to be significantly different from zero. As far as the response function of L'P to a one-time unit LFGP impulse is concerned (bottom left diagram), none of the response coefficients is significant. These results are in line with the CC analysis. This was also confirmed by analysing the forecast error variance components, which were also estimated from the VAR(9). The analysis indicates that after 9 lags, more than 95% of the forecast error of LH> is still explained by the series itself while at the same lag 25% of the error of LFGP is already explained by the LH> series (c.f. Appendix). All these elements allow us to conclude that the Thai market is a leader market with respect to the French Guyana market.

Error correction model estimation

Recall that the Hannan-Quinn criterion indicates a VAR(9) in levels. Under the constraint of one co-integration relation, the ECM, in its Johansen's formulation, is therefore a VEC(8) in differences where the outliers and the dummy variables
accounting for the seasonality of the series are also introduced. The estimated model is:

\[
\begin{align*}
\begin{pmatrix}
\Delta LTP_t \\
\Delta LFGP_t
\end{pmatrix} &= \sum_{i=1}^{l} \mathbf{\Gamma}_i \begin{pmatrix}
\Delta LTP_{t-i} \\
\Delta LFGP_{t-i}
\end{pmatrix} + \begin{pmatrix}
-0.271^* & 0.150^* & 0.729^* \\
0.177^* & -0.098^* & -0.477
\end{pmatrix} \begin{pmatrix}
LTP_{t-1} \\
LFGP_{t-1}
\end{pmatrix} + \\
(0.143^*)D_{9103} + (0.151^*)D_{9008} + (-0.104^*)D_{9104} + (0.042)D_{9009} + (-0.122^*)D_{9006} + (0.018)D_{9005} + (-0.031^*)D_{9111} + \psi D^e
\end{align*}
\]

where (*) denotes significant coefficients with respect to t-statistics (5%) and the matrix \(D^#\) accounts for the seasonality of the series (c.f. Table 7 below). The coefficients of the matrix \(\mathbf{\Gamma}_1\) are detailed in appendix. The analysis of these \(\mathbf{\Gamma}_i\) coefficients indicates that the monthly change in the cultured shrimp price only depends on its own past changes (considered at lags 2, 3 and 5). In turn, the monthly change in the French Guyana’s shrimp price depends on its own past changes (at lags 1, 4 and 5) but also on past changes of the cultured shrimp price (at lags 5 and 6).

The five deterministic outliers \(\{D_{9103}, \ldots, D_{9111}\}\) included in the VEC model (7) permit to obtain the normality. Table 6 indicates the events associated to these outliers.

**Table 6.** The outliers introduced in the VEC model.

<table>
<thead>
<tr>
<th>Dummy variable</th>
<th>period</th>
<th>Series</th>
<th>Events</th>
<th>Reference</th>
</tr>
</thead>
<tbody>
<tr>
<td>(D_{9009})</td>
<td>Sept. 1989</td>
<td>(\Delta LTP)</td>
<td>Black Tiger price bottoms out.</td>
<td>Globefish (1990)</td>
</tr>
<tr>
<td>(D_{9008})</td>
<td>Aug. 1990</td>
<td>(\Delta LFGP)</td>
<td>First signs of overproduction in the French Guyana fishery</td>
<td>-</td>
</tr>
<tr>
<td>(D_{9103})</td>
<td>March 1991</td>
<td>(\Delta LTP)</td>
<td>Drug residue scandal in the Thai production</td>
<td>Csavas (1992)</td>
</tr>
<tr>
<td>(D_{9111})</td>
<td>Nov. 1991</td>
<td>(\Delta LFGP)</td>
<td>Social movements (strike) of the French Guyana fishermen</td>
<td>Béné and Moguedet (1993)</td>
</tr>
<tr>
<td>(D_{9206})</td>
<td>June 1992</td>
<td>(\Delta LTP &amp; \Delta LFHP)</td>
<td>Virus epidemic affects the Ecuadorian production</td>
<td>Globefish (1992)</td>
</tr>
</tbody>
</table>
The LTP and LFGP series display slight significant seasonalities (Table 7). The positive seasonal coefficient observed in December, April and June for the LTP series correspond to the consumption peaks that are known to occur for Christmas and New-Year's Eves, Easter and summer start periods (Anon 1991, Jospuit 1992). On the contrary, the negative seasonal coefficients observed in the LFGP series in March and April correspond to the storage problems seasonably encountered by the French Guyana fishery during this period. Indeed, recall that these months coincide to its highest landings volumes period.

Table 7. The seasonal coefficients. Only 11 dummy variables were included into the model because of algorithm procedure limitation (Sept. coefficient not included). Significance tested by t-statistics.

<table>
<thead>
<tr>
<th>Month</th>
<th>Jan</th>
<th>Feb</th>
<th>March</th>
<th>Apr</th>
<th>May</th>
<th>June</th>
</tr>
</thead>
<tbody>
<tr>
<td>ALTP</td>
<td>0.062*</td>
<td>0.036*</td>
<td>0.031</td>
<td>0.056*</td>
<td>-0.011</td>
<td>0.044*</td>
</tr>
<tr>
<td>∆LFGP</td>
<td>-0.020</td>
<td>-0.013</td>
<td>-0.054*</td>
<td>-0.041**</td>
<td>-0.010</td>
<td>-0.002</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Month</th>
<th>July</th>
<th>Aug</th>
<th>Oct</th>
<th>Nov</th>
<th>Dec</th>
</tr>
</thead>
<tbody>
<tr>
<td>∆LTP</td>
<td>0.047*</td>
<td>0.030**</td>
<td>-0.028</td>
<td>0.016</td>
<td>0.032*</td>
</tr>
<tr>
<td>∆LFGP</td>
<td>0.017</td>
<td>-0.001</td>
<td>-0.019</td>
<td>-0.052*</td>
<td>0.003</td>
</tr>
</tbody>
</table>

* significative at 5% ** : significative at 10%

Table 8. Results of Ljung-Box (L-B) and LaGrange Multiplier (LM) tests on the VEC model (7).

<table>
<thead>
<tr>
<th>Tests</th>
<th>d.f</th>
<th>χ² estimated</th>
<th>p-values</th>
</tr>
</thead>
<tbody>
<tr>
<td>χ² normality</td>
<td>2.841</td>
<td>0.58</td>
<td></td>
</tr>
<tr>
<td>L-B</td>
<td>44</td>
<td>79.316</td>
<td>0.00</td>
</tr>
<tr>
<td>LM(1)</td>
<td>4</td>
<td>9.328</td>
<td>0.05</td>
</tr>
<tr>
<td>LM(4)</td>
<td>4</td>
<td>1.855</td>
<td>0.76</td>
</tr>
</tbody>
</table>
The goodness of fit of the estimated VEC model (7) was tested through the Ljung-Box and LaGrange Multiplier tests (Table 8). The $\chi^2$ value of the Ljung-Box test is superior to the $\chi^2$ tabulated, indicating an auto-correlation of the residuals at 5%. In turn, the LaGrange multiplier (LM) tests indicate that this auto-correlation is neither of order 1 nor 4. The 20 first auto-correlation coefficients confirm the occurrence of some significant auto-correlations of higher order.

Discussion

The development in the 1980s of farm-produced shrimp, essentially in Asia and South America, has induced dramatic changes on the world shrimp market. The commercial success that follows this development was important that the price of some of these cultured shrimp has rapidly become reference on the different shrimp world marketplaces.

In this context, the analysis of the auto-regressive processes that characterise the price time series of the French Guyana and Thai shrimp puts forward some worth-noticing points. As the cultured shrimp price series is integrated of order one without any remaining auto-regressive process, we deduce that the monthly variations of this price follows a random walk. We can also conclude that the price determination comes from the instantaneous market clearing. This result is in accordance with a very competitive world-wide market hypothesis. On the other hand, the French Guyana's price is also integrated of order one but with a eleven-months lag for the auto-regressive process. The French Guyana's price however, does not display any seasonal pattern. Thus, this may be interpreted as a small and specific market behaviour where the few buyers and sellers involved keep in mind the price value over approximately one year.

The existence of the co-integration relation corroborates the hypothesis of a common long-run equilibrium relationship between the two shrimp markets. Furthermore, the rejection of a trend coefficient in this co-integration regression signifies that the two markets can be considered as fully integrated without any drift. However, the fact that the coefficient of the co-integration vector is greater than one clearly indicates that the French Guyana's price is characterised by a more important variability along this long-run equilibrium. This result can be explained by the respective sizes of the two markets and by the difference between the production procedures involved (year-planned production for the cultured shrimp versus fluctuating and uncontrolled catches for the wild shrimp).

The analysis of the error correction model offers a good insight on the short-run dynamics that links the two prices within their common relation. The monthly changes in the cultured shrimp price appears to only depend on its own past variations, whereas the monthly changes in the French Guyana's shrimp price turns out to depend on its own past deviations but also on past deviations of the Thai price.

Finally, the analyses of causalities, impulse response functions and forecast error variance decomposition all indicate that the Thai price is a leader for the French Guyana price on French markets. More precisely, the causality tests show that the Thai market price Granger causes the French Guyana market price, or, in usual term, that the Thai market precedes the French Guyana market in their joint dynamics. The impulse responses functions show that a shock on TP induces a deviation in $FGP$ that remains significant for the three following months, while a shock on $FGP$ has no impact on $H>$. The error variance decomposition indicates that globally one fourth of the French Guyana price variation can be attributed to
the Thai market variations.

These results induce important consequences for the fishery dynamics. This means that the local price fluctuations are largely dependent to events exogenous to the fishery, and that the companies have not the entire control of the price of the product they supply to the markets. The fluctuations of this price turn out to be essentially induced by cultured shrimp prices variations. Nevertheless, the high culinary qualities and taste specificity of this species involves that its price behaviour is mainly depending on its own market conditions.

References


Bené Ch., 1996. Effects of market constraints, the remuneration system, and resource dynamics on the spatial distribution of fishing effort, Can. j Fish. Aquat. Sci. (in press).


Captions of the figures

FIG. 1. Volumes of the world total cultured and captured shrimp production, over the period 1980-1990 (source: Chauvin 1992, Traditional shrimp harvest: supply overview. In H de Saram and T. Singh (op. cit.): 5-72.).


FIG.3. Simultaneous variations in the prices (in French Francs) of the Thai shrimp Penaeus monodon.

(TP series) and the Ecuadorian cultured shrimp Penaeus vannamei, on the French market, over the period Jan91 - Jul.93 (source: FAO-Globefish).

FIG.4. Impulse response functions estimated from the VAR(9) model using the Klock and Van Dijk's procedure (1978).
Appendix

Forecast error variance decomposition of LTP and LFGP series (in percent)
(S.E.: standard error coefficients).

<table>
<thead>
<tr>
<th>Lags</th>
<th>S.E.</th>
<th>LTP</th>
<th>LFGP</th>
<th>S.E.</th>
<th>LTP</th>
<th>LFGP</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.04095</td>
<td>100.00</td>
<td>0.00</td>
<td>1</td>
<td>0.09862</td>
<td>75.50</td>
</tr>
<tr>
<td>2</td>
<td>0.05375</td>
<td>99.00</td>
<td>1.00</td>
<td>2</td>
<td>0.10098</td>
<td>73.10</td>
</tr>
<tr>
<td>3</td>
<td>0.06522</td>
<td>98.00</td>
<td>2.00</td>
<td>3</td>
<td>0.10529</td>
<td>71.20</td>
</tr>
<tr>
<td>4</td>
<td>0.07481</td>
<td>98.40</td>
<td>1.60</td>
<td>4</td>
<td>0.32829</td>
<td>96.80</td>
</tr>
<tr>
<td>5</td>
<td>0.08034</td>
<td>98.00</td>
<td>2.00</td>
<td>5</td>
<td>0.06549</td>
<td>24.20</td>
</tr>
<tr>
<td>6</td>
<td>0.08307</td>
<td>97.70</td>
<td>2.30</td>
<td>6</td>
<td>0.65721</td>
<td>24.20</td>
</tr>
<tr>
<td>7</td>
<td>0.08469</td>
<td>97.50</td>
<td>2.50</td>
<td>7</td>
<td>0.65762</td>
<td>24.20</td>
</tr>
<tr>
<td>8</td>
<td>0.08617</td>
<td>97.40</td>
<td>2.60</td>
<td>8</td>
<td>0.06578</td>
<td>24.30</td>
</tr>
<tr>
<td>9</td>
<td>0.08725</td>
<td>95.70</td>
<td>4.30</td>
<td>9</td>
<td>0.65793</td>
<td>24.30</td>
</tr>
<tr>
<td>10</td>
<td>0.08735</td>
<td>95.60</td>
<td>4.40</td>
<td>10</td>
<td>0.06040</td>
<td>24.30</td>
</tr>
<tr>
<td>11</td>
<td>0.08898</td>
<td>94.10</td>
<td>5.90</td>
<td>11</td>
<td>0.06187</td>
<td>24.30</td>
</tr>
<tr>
<td>12</td>
<td>0.08972</td>
<td>91.00</td>
<td>9.00</td>
<td>12</td>
<td>0.65846</td>
<td>24.30</td>
</tr>
</tbody>
</table>

Coefficients of the $\Gamma_i$ matrix of the VEC model (7) where (*) denotes significant coefficients with respect to t-statistics (5%).
FIGURE 3