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Research Article

The behavior of hake prices in Chile: is the world market leading?

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ABSTRACT. In this study we analyzed price determination throughout the Chilean hake market chain. To analyze the relationship between different prices participating in this chain, a VECM model was successfully estimated. One cointegration vector was identified. Tests for weak exogenous variables, causality, and significance of different variables were performed, and a parsimonious version of the model was selected. The results obtained in this paper outline a price determination process that, in the end, is governed by world market conditions. Moreover, the diverse links in the hake market chain seem to be well integrated, which implies that there is little room for domestic price determination.

Keywords: hake, fish price determination, time series, market chain, cointegration, market integration, Chile.

El comportamiento del precio de la merluza común en Chile: ¿Es el mercado mundial el líder?

RESUMEN. En este estudio se analiza la determinación de precios a través de la cadena de comercialización de merluza chilena. Para analizar la relación entre los diferentes precios que participan en esta cadena, se estimó con éxito un modelo VECM. Se identificó un vector de cointegración. Se realizaron pruebas para las variables exógenas débiles, causalidad, y significancia de diferentes variables, y se seleccionó una versión parsimoniosa del modelo. Los resultados obtenidos describen un proceso de determinación de precios que, al final, se rige por las condiciones del mercado mundial. Por otra parte, los diversos eslabones de la cadena de comercialización de merluza parecen estar bien integrados, lo que implica que hay poco espacio para la determinación de los precios internos.

Palabras clave: merluza común, determinación de precios, series de tiempo, cadena de comercialización, cointegración, integración de mercado, Chile.

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INTRODUCTION

Prices are important for fishermen, since they determine both their income and production costs. Moreover, the information conveyed by prices is essential for the decisions they must make, such as which species to target, with what fishing equipment use, and who to sell the catch to, among others. Although fishermen usually want higher product prices and lower input prices, most often they have no power to determine or influence the level of these prices. Sometimes they do not understand how prices are set up and think that some special economic agents have the power to determine prices. Therefore, one occasionally hears pleas to the authorities, especially from artisanal fishermen, to control price setting as a way of improving their income. In the public discussion on marine resource price determination, the idea that some agents are setting prices in some link of the marketing chain is recurrent (*e.g.*, that the processing plants managers fix ex-vessel prices) or that it is possible to remove intermediaries that set abusive prices, and in that way improve the life conditions of artisanal fishermen (*e.g.*, some people believe that by eliminating market chain links the prices received by the artisanal fishermen will be higher).

The economic literature includes studies about temporal price behavior as a way of understanding how the market functions and how prices are determined (*e.g.*, Asche *et al.*, 2005; Nielsen, 2005; Tveteras & Asche, 2008; Nielsen *et al.*, 2009; Garcia & Salayo, 2009). A central focus has been to analyze the relation between different prices throughout the market chain, to understand how integrated the market is and how causality runs between prices. These studies have also addressed if buyers or sellers can affect prices, if domestic conditions affect prices set by the market, how do world market prices and the exchange rate intervene in the price setting process, and which prices condition or are conditioned by other prices. Within this framework, one central concern has been to test for market integration and price transmission between different links in the market chain using the variability of prices over time (Asche *et al.*, 2007). While market integration analysis is primarily concerned with determining the market's boundaries, by looking at the behavior between potentially connected prices, price transmission analysis focuses on which prices lead price determination and how demand and supply shocks are transmitted through the market chain. However, the test of relevant issues in this literature has sometimes been blurred by the non-stationary behavior of price series. In this case, traditional econometric approaches are no longer valid. An alternative approach, based on cointegration analysis techniques, has been developed and used to test for causality, market integration and the so called "Law of One Price" (Asche *et al.*, 2004)

In this study, we analyzed price determination throughout the Chilean hake market chain. The basic motivation is to characterize the way prices are determined along this chain. This characterization implies identifying the causality direction of prices between ex-vessel prices, fresh fish consumers prices, and export prices; the integration level of prices along the market chain; and the extent of domestic influence on price determination. Although sometimes in the public discussion of price determination in this market in Chile it is assumed that the "small country" assumption is valid, meaning that prices are directly and fully determined by world market prices, this hypothesis has never been tested rigorously. We present results that test this hypothesis. Moreover, the information obtained from this study is important for the design of management policies toward hake fisheries in Chile, and especially for rejecting the design of policies that attempt to control prices in an administrative way.

Hake fisheries in Chile (*Merluccius gayi gayi)* have been operating for a long time and are considered an emblematic fishery (SUBPESCA, 2010). The geographic distribution of this species is located in maritime zones between the Coquimbo (29º10'S) and the Los Lagos regions (42º00'S), mainly near the coast, from depths of 50 to 500 m (SUBPESCA, 2012). From 2004 to present day, hake has shown a spawning biomass lower than what is considered to be biologically sound; thus, it has been declared in a state of overexploitation and depletion (SUBPESCA, 2012). The maximum extraction level was attained in 2004, when approximately 120,000 ton were landed. However, in recent years the extraction level has gone down to 45,000 tons. Hake fisheries are an important source of employment, concentrated primarily in the Biobío region (36º46'S). On the average, the industrial sector generates 3,500 jobs, fundamentally in the processing industries, while employment in the artisanal sector involves up to 11,000 people. In recent years approximately 50% of the landings have been directed to human consumption goods for exports, generating a value of nearly US$30 million per year (SUBPESCA, 2012). The rest is consumed domestically.

The hake market chain in Chile is composed of several links. This market's productive base is determined by fish harvesting. *Merluccius gayi gayi* is captured by an heterogeneous fleet, where the industrial fleet (vessel length over 18 m) and artisanal fleet (vessel length ≤18 m) are easily distinguished. However, regardless of the vessel's size, all captures have three possible uses: to be sold for direct human consumption on the beach, to be sold as raw material for plant processing, and to be sold at the Metropolitan Fish Market (MFM). This market, located in Santiago de Chile, is the main market for fresh fish in Chile. The MFM, in turn, sells fish to wholesale and retail traders for domestic fresh consumption. The plants process the fish for human consumption and then sell their production to foreign markets and domestic retail traders. Thus, we have a mix of price transmission issues, as a result of the different market links, and of integration issues due to the different market segments, all of which are connected by the sole basic input. The principal links in this market chain are depicted in (Fig. 1).

Thus, we can distinguish between four final destinations for hake captures: fresh consumption (through the MFM); exports (through the processing plants); domestic processed consumption (through the processing plants); and direct human consumption (through sales at the beach). However, due to information availability, in this study, we only analyzed three of these destinations, domestic consumption through the MFM, exports to foreign markets and direct fresh consumption. These are the main components of the Chilean hake market chain. Unfortunately, there was no periodical, reliable price information for the

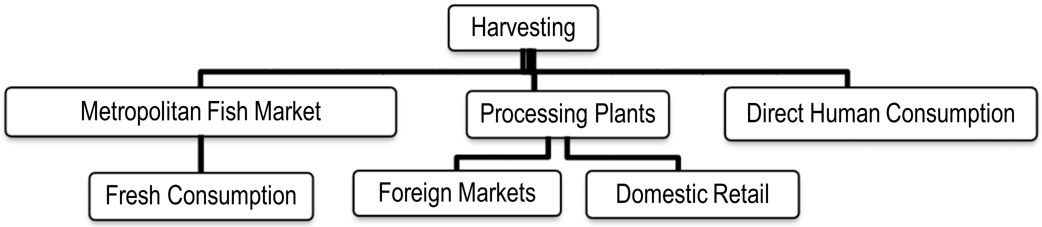

Figure 1. The Chilean hake market chain.

other component (domestic retail). One would suspect that this unobserved price would very closely follow the prices for which we do have information. Moreover, the percentage of hake landed that could potentially be exchanged in this market fluctuated between 8 and 15% of total landings in the 2006-2009 periods, depending on the estimation year. The prices with available information included ex-vessel prices, MFM prices and FOB export prices. For this study, the FOB export price for frozen hake fillet is used as the relevant world market price for Chilean international hake transactions. Actually, one can find different international prices in world markets depending on the product exchanged and the characteristics of the buying and selling countries. So, it should clear that when we talk of the world market price this is a simplification. However, the FOB export price for frozen hake fillet, the main hake export product from Chile, should to a great extent reflect the international demand conditions that Chilean hake meets at the world market relevant for Chile. Nevertheless, since the FOB prices are in foreign currency (US$), we also included the exchange rate of Chilean pesos (CLP) for US dollars (US$) in the analysis. Thus, we analyzed four prices in all. In Fig. 2 the monthly development of these four prices for January 2002 to May 2011 is depicted.

At first sight, it seems that the domestic prices (ex-vessel prices and MFM prices) follow each other relatively closely, while this does not seem to be the case for international prices or exchange rates. However, partial correlation analysis between individual prices does not consider the possibility that more than two prices can interact. Therefore, to disentangle the relationship between all prices a more solid analysis is required.

MATERIALS AND METHODS

Analyzing the relationship between prices to test for market integration is a well-established practice in economics (Asche *et al.,* 2007). When markets are integrated the price development in one market chain is reflected in the prices of other market chains. The so-called "Law of One Price" is a special extreme version of this notion, where the price levels between different market chains are equal, except when a markup reflecting value is added (absolute version), or the rate of change in prices is equal (relative version). This means that different prices in an integrated market follow one other over time. However, if substitution or complementary possibilities are less than perfect, price deviations can be found. Moreover, market power, if used with discretion, can break the temporal relation between prices, as it does when it is used to avoid or increase the price changes generated in a preceding link of the market chain.

From an empirical point of view, the following basic relation helps to clarify the basic ideas:

$$\ln p_{At} = \alpha + \beta \ln p_{Bt} + \varepsilon_t \qquad (1)$$

where p_A and p_B are two different prices in the market chain, ε is a stochastic variable, α and β are parameters, and the subscript t indicates the time period. Parameter α captures transport costs and quality transformations in the product. The relationship between prices is captured by parameter β. If $\beta=0$, there is no relation. If $\beta=1$, the "Law of One Price" is valid, if $0<\beta<1$ then the goods are imperfect substitutes. Thus, the whole discussion of the degree of market integration can be concentrated on the value of β.

One question that arises when considering a price relation like the one depicted in eq. (1) is which direction does price determination occur? Does p_A determine p_B, or is it the other way around, or are both prices simultaneously determined? If one price is exogenously determined with respect to the other, then one can infer that there is a unilateral determination from the first price to the second. Thus, one price leads to the other. In this case, what is required to settle this issue is an exogeneity test applied to both prices.

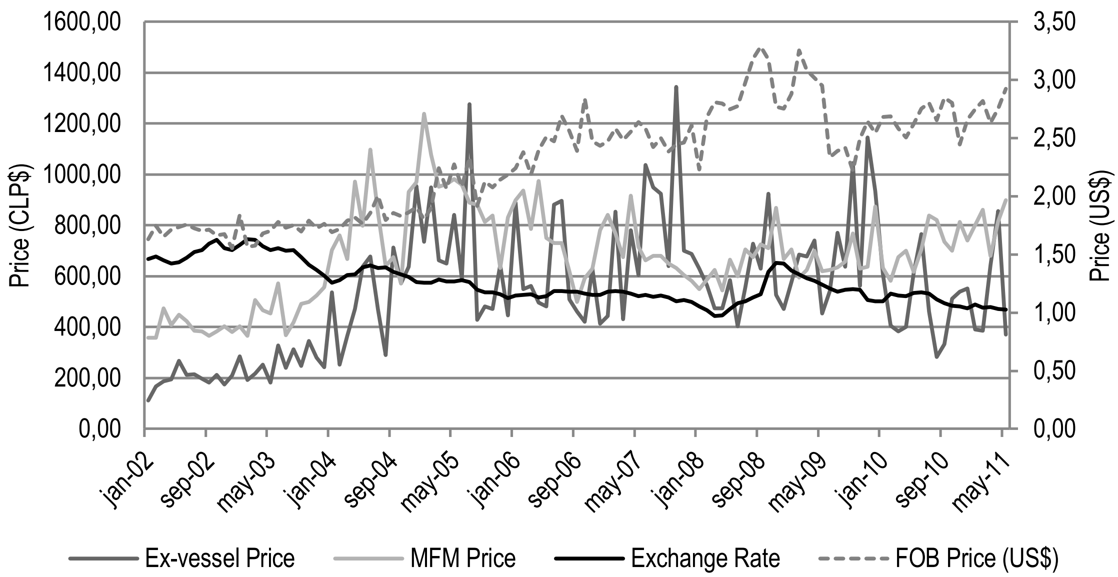

Figure 2. Price development in the Chilean hake market chain (January 2002-May 2011). The left hand axis shows prices for the ex-vessel price per landed kilogram of hake, the Metropolitan Fish Market (MFM) price per kilogram of hake, and the US$ exchange rate, all in CLP. The right hand axis shows the FOB average price per each exported kilogram of frozen hake fillet in US$.

Of course, eq. (1) is a simple and stylized way of looking at price relations. Eq. (1) assumes that there are only two prices that participate in the market chain and that price adjustments occur in the same time period. In practice, more than two prices can participate in the market chain and time adjustment costs might exist. In this case, instead of a one equation system we can have a multiple equation system and the effects between different prices can become apparent throughout different time periods. This has implications for the methodology used to test for market integration and price leadership, but the basic notions embedded in eq. (1) remain valid. That is, the relationship between different prices shall be reflected in the parameters associated with each price and if the "Law of One Price" is valid, in the long run, the sum of the value of these parameters should equal one. This guarantees that relative prices remain constant. Moreover, exogeneity tests in a multivariate framework are necessary to test for price leadership. The refusal to reject weak exogeneity can be interpreted as a leading price in the long run (Asche, 2007). An additional issue that emerges when dealing with temporal price series is that frequently these series are not time stationary, which implies that traditional econometric methods used to empirically analyze price relations, is no longer valid (Engle & Granger, 1987).

A suitable approach to treat non-stationary series in a multivariate context is the vector autoregressive model (VAR) with cointegration analysis (Johansen & Juselius, 1990). This approach can deal with non-stationary price series, allowing for the inclusion of lags in the price interactions, and solving the simultaneity determination of different price variables. It also allows for the analysis of short and long run price responses, permitting the formal testing for price leadership (exogeneity tests) in a multivariate framework.

The model can formally be established as:

$$Z_t = \sum_{t=1}^{k} A_t Z_{t-i} + \Psi D + \mu_t \tag{2}$$

or in its error correction form (VECM) as:

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \Pi Z_{t-k} + \Psi D + \mu_t \tag{3}$$

where Z_t is a (n x 1) vector of (price) variables, t is the subscript that denotes the time period, and n is the number of variables in the model. A_i is a matrix of coefficients of dimension (n x n), D is a matrix of deterministic variables, Ψ is the corresponding parameter vector, μ_t is a column vector of (n x 1) random errors (innovations), and k is the number of lags included in the VAR model. Δ is the difference operator; $\Gamma_i = -I + A_1 + A_2 + \ldots + A_i$, with I as an identity matrix of order n and $\Pi = (I - A_1 - A_2 - \ldots - A_i)$.

In the present case, vector Z_t is the following:

$$Z_t^{'} = \left[p_{E,t}, p_{M,t}, p_{X,t}, e_t \right] \qquad (4)$$

where $p_{E,t}$ is the ex-vessel price in CLP, $p_{M,t}$ is the MFM price in CLP, $p_{X,t}$ is the FOB export price in US$, and e_t is the exchange rate, measured as CLP/US$. For conformity with the empirical section, all of these variables are expressed in natural logs. That is, $x = ln(X)$, where X is the relevant variable in levels.

The model in eq. (3) includes an error correction mechanism (ECM), which allows for the separation of the long and short run relationships. Short run relations are associated with the Γ_i matrices, while the long run relations are associated with the Π matrix. Therefore, the first step is to determine the number of cointegration vectors, r, associated with the long run relations. The procedure used to determine r is known as the method of the reduced rank (Johansen, 1988). Then, it is necessary to test whether the obtained cointegration relations correspond with the expected theoretical values. To accomplish this, the following steps were followed:

- Identify the integration order of the variables.
- Specify the VAR model (lag order, introduction of non-stochastic components).
- Identify the number of cointegration vectors (restricted rank tests).
- Compare the estimated vectors with economic theory.
- Develop tests of weak exogeneity.

When $0 < r < n$, r cointegrating vectors exist. In this case, at most, three cointegration vectors may exist (long run relations). When this happens, one can factor Π such that $\Pi = \alpha\beta'$, where both α and β are $(n \times r)$ matrices. Matrix β contains the cointegrating vectors or the long-run relationships and α encompasses the adjustment parameters.

This model also allows for a short run analysis. This analysis is important to understand how the price variables adjust in the short run. Once the long run relations are identified, eq. (3) can be estimated as follows:

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \alpha \hat{\beta}^{'} Z_{t-k} + \Psi D + \mu_t \qquad (5)$$

where $\hat{\beta}^{'} Z_{t-k}$ corresponds to the estimated cointegration vectors, previously identified.

All variables included in eq. (5) are stationary. Thus, classic econometric tests (t- distribution, F distribution) can be used confidently. That is, once eq. (5) has been estimated, one can test for the significance of the parameters associated with the lagged variables ΔZ_{t-i}. This allows for the presentation of a parsimonious VAR model (PVECM), where all non-significant parameters have been excluded.

Within this framework, testing for Granger causality is possible. Granger non-causality, loosely speaking, implies that the prediction of a variable y_{t+h}, $h=1,\ldots n$, based on the information set Ω_t, does not improve if one includes the behavior of variable x_s, $s \leq t$ in the information set (Granger, 1969). In this context, one says that "x does not Granger cause y". If, on the other hand, when the reverse test between these two variables is carried out and we cannot reject the null that "y does not Granger cause x", then we can conclude that there is causality between y and x, in the described sense. Testing for Granger causality allows us to understand the interactions between the different variables in the model. The test for Granger non-causality within the VECM model framework implies a specific joint test for the parameters in α and Γ related to the potentially non-causal variable in Z. The specific test procedure can be found, *e.g.*, in Lűtkepohl & Krätzig (2004).

Information sources

The price series used in this study were obtained from official sources. The price series for ex-vessel prices was acquired from the National Fisheries Service's database (SERNAPESCA in its Spanish acronym). The MFM price series came from the Agrarian Studies and Policies Office's database (ODEPA in its Spanish acronym). The export prices were obtained from the Superintendent's Office of Customs and the market US$ exchange rate from the Chilean Central Bank. Our information sample was limited in its timespan, since MFM prices (p_M) were available only from the year 2000, and FOB prices (p_X) were available starting in the year 2002. Moreover, our research had an upper time limit of May 2011, due to the availability of collected MFM prices (p_M). This gave us a maximum length in the monthly samples of 113 observations for all variables, between January 2002 and May 2011.

To build the database for estimation purposes, we had to process the original data in different ways and make some assumptions. The ex-vessel price series is based on artisanal prices. The industrial fleet is well-integrated with the processing industry, and thus, their transfer prices for landings are not known, nor are they publicly registered. We assumed that the prices based on artisanal transactions were a good indicator of total ex-vessel prices. Moreover, we had to carefully review the database obtained from SERNAPESCA since several prices were entered incorrectly into the database. In those cases we felt confident we corrected the figures. For instance, we found cases were the reported prices were ten times higher than those in neighborhood fishing coves, and ten times more than prices for the same fishing cove in consecutive months. The MFM prices are collected on a weekly basis as maximum and minimum prices. We took the simple average of the maximum and minimum price for each week and then averaged them over the weeks in the month to build a monthly average price estimate. The export prices are average prices for exports of freeze fillet. This corresponds to the most important commodity exported, in terms of hake. In the last years, this product, on average, makes up 87.3% of total processing plant production. All final prices were expressed in kilograms. Table 1 shows some basic information about the data used.

The average ex-vessel price was lower than the MFM and export price (in CLP), as expected, since the commodities higher up in the market chain include additional production, transportation, marketing, and conservation costs. Moreover, the foreign market variables showed a relatively lower volatility than the domestic ones.

RESULTS

The first step was to analyze stationarity in the time series. To test for stationarity, we used different methodologies and specifications. We used the Augmented Dickey-Fuller test (Dickey & Fuller, 1981), the Phillips Perron test (Phillips & Perron, 1988), and the Dickey Fuller Generalized Least Squares test (Graham *et al.*, 1996). Moreover, we ran these tests in levels, logs, first differences and log first differences. Finally, we tried different specifications for the deterministic components in the test equations, including non-deterministic components, a constant term, a trend term, seasonal dummies, and all the relevant combinations of these components. The results seemed very stable between different specifications and methods. Thus, in the following, we only show results for the more popular Augmented Dickey Fuller (ADF) tests.

In Table 2 the results for the Augmented Dickey Fuller (ADF) tests in log levels for the four variables in the vector for the period January 2002-May 2011 are presented. The optimum lag length for these tests was selected with the Bayesian Minimum Criterion (also called Schwarz criterion).

The ADF test has a null hypothesis of unitary root. The rejection of the null hypothesis indicates stationarity in the variable. The evidence obtained on stationarity for the logs of the ex-vessel and MFM prices was not conclusive. When a constant term, or other deterministic variables, was added to the specification, we could reject the null hypothesis of unitary root. These results were also found for other tests not reported here. They seemed to be solid. If one contemplates the series in Fig. 1, it is apparent that neither series shows a clear trend over time except the first months, in the sample period, where an upward drift is clear. Therefore, the results obtained by these tests simply represent this behavior. In contrast, the logs of the export price and exchange rate did show a clear trend: the first was positive and the second negative over the sample period. This was captured by the ADF test results. In both cases the unitary root null was not rejected. However, to be able to determine the integration level of these variables it was necessary to look at the results for the ADF test in log differences. These results, for the same variables and period, are presented in Table 3.

As can readily be seen in this table, the null hypothesis of unitary root was rejected in all cases at very low significance levels for all of the specifications. Thus, this confirms that the log of the export price and the log of the exchange rate are *I(1)* variables, while the evidence for the log of the ex-vessel price and the log of the MFM price is less compelling. However, the Johansen test for cointegration allows us to deal with variables of different orders of integration (Lűtkepohl & Krätzig, 2004).

To specify the optimal number of lags for the VECM, we used different criteria. According to the Akaike Information criterion, the Final Prediction Error and the Hannan-Quinn criterion, the optimal number of lags in first differences should be one. However, according to the Schwarz criterion this should be zero. Given these results, we made the specification tests for models with lag length one and zero. Finally, based on the specification tests' results, we chose to continue working with the specification with one lag.

We estimated different versions of the model and controlled for their statistical specifications. We used a battery of specification tests for normality, autocorrelation, and autoregressive conditional heteroskedasticity. Finally, we selected a model that included one lag in the VECM, a constant term restricted to the cointegration space, one cointegration vector, centered seasonals, and two impulse dummies in the periods October 2008 and June 2009. All univariate tests suggest that the residuals comply with white noise errors (Table 4).

Table 5 reports the results for the multivariate tests. They show, in general, that the specification was in line with expected results. The system autocorrelation tests show diverging results. While the Lagrange Multiplier-test indicates autocorrelation, the portmanteau test indicates no autocorrelation. Nonetheless, the graphical inspection of the residuals suggests that this potential autocorrelation was not evident.

Table 1. Data description (2000-2011) of different prices (in kg). [1]Free market US$ exchange rate, in Chilean pesos per US$ (CLP/US$); [2]Period January 2000-December 2011; [3]Period January 2000-May 2011; [4]Period January 2002-December 2011.

Variable	N° of observations	Average value	Standard deviation	Minimum	Maximum
Ex-vessel price in CLP (P_E)	144[2]	486.33	250.88	111.26	1276.00
MFM price in CLP (P_M)	137[3]	626.66	205.94	273.13	1239.50
Export price in CLP (P_X)	120[4]	1302.14	197.34	987.70	1972.60
Export price in US$ (P_X)	120[4]	2.34	0.48	1.54	3.29
Exchange rate (e)[1]	144[2]	571.07	75.07	442.94	745.21

Table 2. Augmented Dickey-Fuller tests for variables in log levels (January 2002-May 2011). Numbers in parentheses indicate optimum lag length, selected with Bayesian Minimum Criterion (BIC). *The null hypothesis of unitary root is rejected at 1% significance; **The null hypothesis of unitary root is rejected at 5% significance; ***The null hypothesis of unitary root is rejected at 10% significance; P_E: Ex-vessel price; P_M: MFM price; P_X: export price; e: exchange rate.

Variable	No constant	Constant	Trend	Seasonal dummies + trend	Seasonal dummies
P_E	0.5918 (4)	-3.3404** (1)	-3.5042** (1)	-3.1254 (1)	-3.0175** (1)
P_M	0.6647 (1)	-2.6913*** (1)	-2.7432 (1)	-2.5354 (1)	-2.5532 (1)
P_X	0.6487 (1)	-1.3484 (1)	-2.8598 (1)	-2.5743 (1)	-1.2948 (1)
e	-0.8777 (1)	-1.7620 (1)	-2.8555 (1)	-2.6770 (1)	-1.6547 (1)

Table 3. Augmented Dickey-Fuller tests for variables in log differences (January 2002-May 2011). Numbers in parentheses indicate optimum lag length, selected with Bayesian Minimum Criterion (BIC). *The null hypothesis of unitary root is rejected at 1% significance; **The null hypothesis of unitary root is rejected at 5% significance; ***The null hypothesis of unitary root is rejected at 10% significance; p_E:=ex-vessel price; p_M: MFM price; p_X: export price; e: exchange rate.

Variable	No constant	Constant	Trend	Seasonal dummies + trend	Seasonal dummies
ΔP_E	-9.0518* (3)	-9.0600* (3)	-9.1775* (3)	-8.4911* (3)	-16.7356* (0)
ΔP_M	-15.4790* (0)	-15.4699* (0)	-15.4233* (0)	-14.9167* (0)	-14.9556* (0)
ΔP_X	-15.8533* (0)	-15.8901* (0)	-15.8159* (0)	-14.7825* (0)	-14.8545* (0)
Δ e	-7.0771* (0)	-7.1173* (0)	-7.0839* (0)	-6.4821* (0)	-6.5154* (0)

Table 4. Univariate tests with 12 lags. [1]Non-autocorrelation test; [2]Normality test; [3]Non-autoregressive heterocedasticity test. *Probability of not rejecting the null hypothesis.

Residuals	Portmanteau[1]	Jarque Bera[2]	ARCH-LM[3]
	P-value*	*P*-value*	*P*-value (χ2)*
p_E	0.8159	0.3255	0.9758
p_M	0.8740	0.8272	0.3783
p_X	0.6484	0.5891	0.8587
e	0.7305	0.8910	0.4202

To model the deterministic components in the VECM model, we tested for reduced rank simultaneously with the specification of the deterministic components in the VECM. In Table 6 we present the results for Johansen's cointegration test with the Trace and Max Eigenvalue statistics for three specifications of the deterministic components: constant only in the cointegration relation (Model 2), constant in the cointegration relation and in the first difference model component (Model 3); and constant and trend in the cointegration relation and constant in the first difference component of the error correction model(Model 4). All specifications included centered seasonals, one lag in the VECM, and two impulse dummies in the periods October 2008 and June 2009.

The results were mixed. While the Trace statistic suggests that the proper model could have zero or one

Table 5. Multivariate tests. *Probability of not rejecting the null hypothesis.

Test	Test statistic	*P*-value*
Autocorrelation		
Portmanteau	177.73	0.366
LM	235.81	0.017
Normality tests		
Joint test	3.90	0.866
Skewness only	1.41	0.842
Kurtosis only	2.48	0.647
Autoregressive heteroscedasticity		
ARCH-LM	409.79	0.357

cointegration vector, the Max eigenvalue statistic suggests that it should have zero contegration vectors. Finally, based on the specification tests we chose to work with model 2. This left us with one cointegration vector, as previously discussed in the model specification. The estimated cointegration and adjustment vectors are presented in Tables 7 and 8, respectively.

We carried out different tests on the cointegration space. First, we tested whether the log of the exchange rate (e) could be excluded from the cointegration vector. This amounts to testing the null hypothesis that $\beta_e = 0$. Secondly, we added the hypothesis that the parameter for the log of the metropolitan market price (p_M) could equal unity. Third, we tested instead if the log of the FOB price of frozen hake fillet (p_X) could equal unity. And finally, we tested all these hypotheses simultaneously. The results are presented in (Table 9).

In all cases, the null hypothesis could not be rejected. Thus, the final specifications of the restricted cointegration vector chosen are as presented in (Table 10).

We tested for weak exogeneity. This test is interesting from the perspective of analyzing price determination since the refusal to reject weak exogeneity can be interpreted as a leading price (Asche, 2007). The weak exogeneity test consisted of testing the null of zero coefficient for the alpha coefficients in the adjustment vector for each variable. We tested for each equation and then for several equations simultaneously (Table 11).

We could not reject the null hypothesis that changes in the log of the metropolitan market price (p_M), the log of the export price (p_X) and the log of the exchange rate (e) since they were all simultaneously weak exogenous to the determination of the cointegration vector. Thus, we were able to model the error correction model for the log of the ex-vessel price excluding the determinants of these other variables without any loss of information (Johansen, 1991).

We tested for a parsimonious version of the error correction model. Finally, we ended with a small model, where a likelihood ratio test for 59 excluded variables was not rejected. The likelihood ratio statistic was 54.40, with a *P*-value of 0.6455. The results for the parsimonious model are presented in Table 12. Implicitly, a Granger causality test between the different variables in the cointegration relation was completed when we tested for this reduced model. Basically, the coefficients of the lagged variables Δz_1 in equation (2) should not be significantly different from zero in the equation of the Δz_2 variable, given that the variable is weak exogenous.

The results show that the log of the exchange rate variable was independent of the rest of the variables. It did not Granger cause and it was not caused by any other variable. In contrast, there is some evidence indicating that the log of the export price, p_X, Granger caused the log of the ex-vessel price, p_E. This evidence is not very strong, but the results indicate that p_E did not cause p_X, and that p_X could cause p_E. Finally, the log of the metropolitan market price p_M was independent of p_X, but not independent of p_E. The relation between p_M and p_E was of mutual interactions.

Table 6. Johansen's cointegration test (95% confidence). Model 2: Constant restricted to the cointegration relation, Model 3: Constant unrestricted; Model 4: Trend restricted to the cointegration relation and unrestricted constant. *Significant result.

	Trace			Max eigenvalue		
r	Model 2	Model 3	Model 4	Model 2	Model 3	Model 4
0	57.26916	51.64930	58.15164*	23.92831*	23.82623*	25.46549*
1	33.34085*	27.82307*	32.68615	20.35427	20.34552	20.43868
2	12.98659	7.47754	12.24747	6.53705	6.53588	6.55285
3	6.44954	0.94167	5.69463	6.44954	0.94167	5.69463

Table 7. Cointegration (β) vector. Parentheses indicate *P*-values and square brackets indicate t-values.

$p_{E\ t-1}$	$p_{M\ t-1}$	$p_{X\ t-1}$	e_{t-1}	Constant
1.000	-1.209	-0.950	-0.452	5.321
(0.000)	(0.000)	(0.021)	(0.505)	(0.313)
[0.000]	[-5.268]	[-2.313]	[-0.666]	[1.010]

Table 8. Adjustment (α) vector. Parentheses indicate *P*-values and square brackets indicate t-values.

$\Delta p_{E\ t}$	$\Delta p_{M\ t}$	$\Delta p_{X\ t}$	Δe_t
-0.444	0.081	0.018	-0.002
(0.000)	(0.105)	(0.390)	(0.753)
[-4.139]	[1.623]	[0.859]	[-0.315]

Table 9. Tests on the cointegration space. The lowercase letters indicate free parameter values. The column $\chi 2$ corresponding to the likelihood ratio statistic. [1]Indicates the probability of not rejecting the null hypothesis. [2]The order of the variable in the vector is: p_E, p_M, p_X, e, constant.

Null hypothesis	χ^2	*P* -value[1]
Restriction 1: (1,a,b,0,c)[2]	0.4435	0.5054
Restriction 2: (1,-1,a,0,b)	0.9506	0.6217
Restriction 3: (1,a,-1,0,b)	1.1815	0.5539
Restriction 4: (1,-1,-1,0,a)	1.3248	0.7232

DISCUSSION

A VECM model was estimated to analyze the relation between different prices participating in the hake market chain in Chile. These prices are the log of the ex-vessel price (p_E), the log of the domestic metropolitan market price (p_M), the log of the export price for the main export commodity (p_X) and the log of the exchange rate (e). The model was successfully estimated and the specification tests suggest that it was a good approximation to the data generation process. One cointegration vector was identified and a parsimonious version of the model was selected. Moreover, the test procedure allowed us to identify weak exogenous variables and causality between different variables.

The cointegration vector established a long run statistical relation between different variables. After applying some tests on the cointegration space and on the adjustment vector of coefficients, we concluded that this cointegration vector could be interpreted as a long run relation between p_E, p_M, and p_X. The exchange rate did not participate in this relation. This result indicates

Table 10. Restricted cointegration vector. Parentheses indicate *P*-values and square brackets indicate *t*-values.

$p_{E\ t-1}$	$p_{M\ t-1}$	$p_{X\ t-1}$	$e_{\ t-1}$	Constant
1.000	-1.000	-1.000		1.136
(0.000)[1]	(0.000)	(0.000)	-	(0.000)
[0.000][2]	[0.000]	[0.000]		[22.644]

Table 11. Adjustment vector in the Error Correction Model (α). Parentheses indicate *P*-values and square brackets indicate t-values. LR-test statistic (H_1: unrestricted model) is 2.4658 with *P*-value (χ^2) = 0.4815.

$\Delta p_{E\ t}$	$\Delta p_{M\ t}$	$\Delta p_{X\ t}$	$\Delta e_{\ t}$
-0.509			
(0.000)[1]	-	-	-
[-4.805][2]			

that changes in the log of the exchange rate did not affect these variables relationship in the long run. Moreover, the weak exogeneity tests suggest that both the p_M, and p_X worked as leading prices in this relation. That is, they induced the changes in the p_E in the long run. Finally, the tests for over-identifying restrictions show that the impact of changes both in international as well as domestic prices upstream the market chain were fully transmitted to changes in the ex-vessel price in the long run. This is not the traditional test of the "Law of One Price", which requires a *n -1* cointegration vector to be able to test for this hypothesis in a multivariate framework (Nielsen *et al.*, 2009), but it does show that there was perfect integration in the different markets. Moreover, this result also suggests that there was no evidence of the exertion of market power on price determination for the price received by fishermen in the long run.

It could be argued that this result is simply a reflection of the customary rule used by plant owners to determine ex-vessel prices by considering a fixed percentage of previous months FOB export prices. However, to the best of our knowledge, there is no rigorous evidence that this is the way ex-vessel prices are determined for hake landings in Chile. Even if this were the case, the results presented here show a completely different relation than the one implicit in this customary rule. First of all, the cointegrated vector shows that the relation is between the three principal prices in the hake market chain (p_E, p_M, and p_X) and the p_M price is not even included in the customary rule previously mentioned. Secondly, it points to a long run relation between these three prices that excludes the exchange rate. The customary rule, obviously, includes changes in the exchange rate since the ex-vessel price

Table 12. Parsimonious VECM Model (January 2002-May 2011). *Significance at the 1%; **Significance at the 5%; ***Significance at the 10%; Brackets indicate *t*-values.

Variable	$\Delta\ p_{E\ t}$	$\Delta\ p_{M\ t}$	$\Delta\ p_{X\ t}$	$\Delta\ e_t$
$\Delta\ p_{E\ t-1}$	-0.187*** [-1.954]	0.059 [1.589]	-	-
$\Delta\ p_{M\ t-1}$	-0.375** [-1.994]	-0.395* [-4.513]	-	-
$\Delta\ p_{X\ t-1}$	-0.576 [-1.408]	-	-0.400* [-4.959]	-
$\Delta\ e_{t-1}$	-	-	-	0.330* [4.561]
October 2008	-	-	-	0.149* [6.665]
June 2009	-	-	-0.242* [-3.998]	-
$S2_t$	-	-	0.040*** [1.912]	-
$S5_t$	-	-	-	0.015** [2.021]
$S6_t$	-	-	-	0.021* [2.713]

paid to fishermen is expressed in Chilean pesos (CLP), but the cointegration relation is governed by changes in the relevant prices in the long run. Therefore, this relation does not consider short run changes in the exchange rate. In this sense, we are confronted with a totally different relation than that inherent in the customary rule. Third, both p_M, and p_X changes are leading and fully reflected in the p_E price. This means that changes in domestic conditions also affect and induce changes in ex-vessel prices, which is also not considered in the customary rule. Finally, if the customary rule is not completely stable and changes over time, *e.g.*, by changing the percentage charged on FOB prices, then this should affect the long run relation between prices. Informal conversations with fishermen suggest that this has been the case. On the contrary, our results do not show this, as the leading prices transmit all changes to the ex-vessel prices.

Another way to interpret the stable long run relation between prices is to consider that the different prices are balanced because price arbitrage is at work between the international and domestic markets. This idea is embedded in the single identified cointegration vector where all three prices participate. Basically, the system should be governed by the international price given the “small country” characteristics of the Chilean supply in the world market for fish fillets. The “small country assumption” in international trade means that domestic conditions are not relevant in the world market for price determination (see *e.g.*, Dornbusch, 1980). This price should be transmitted to the domestic market through price arbitrage. If domestic prices are lower than world market prices this should induce a shift in the supply of hake landings from domestic to international buyers. In contrast, if domestic prices are higher than world market prices, this should increase the supply for domestic markets. However, domestic demand restrictions should lead, in due time, domestic prices to converge with the international ones. Thus, the fact that this cointegration exists, and its properties, suggests that the different markets for hake in Chile are integrated and that prices adjust to excess demand conditions, at least in the long run.

If we look at the estimations for the short run component of the VECM (Table 12), it is apparent that the international price did not interact with other prices in the short run. Moreover, the causality tests provided some evidence that the p_X Granger caused the p_E. All of this evidence is consistent with the arbitrage story and the international price determination of the system. Although there was determination from the international market to the ex-vessel price, there were no feedbacks from domestic conditions to the world market.

In contrast, in the short run there was mutual interaction between the p_E and the p_M variables. Thus, causality ran in both directions, although the link from p_E to p_M was not very statistically strong. However, this link could provide us with a clue about how international prices affect domestic metropolitan market prices in the short run through the supply price of hake landings. Higher ex-vessel prices in the short run should signal changes in international prices, which would, in turn, affect domestic markets. As it can be seen, export prices did not directly affect the metropolitan prices in the short run. But to the degree that they did affect ex-vessel prices, they also, in turn, affected metropolitan prices, thus, there is a short run link between the world market price and the domestic metropolitan market price. Of course, this link is not necessary for the long run story to hold. One should presume that most transactions between extractive and domestic marketing firms should take place at prices that already include the world market price.

The results obtained in this paper outline a price determination process that, in the end, is governed by world market conditions. Moreover, the different links in the hake market chain seem to be well integrated. This implies that there is little room for domestic price determination, which suggests that the presence of domestic monopolistic or oligopolistic price setting in this market, as sometimes feared by fishermen, should not be a major concern. To the extent that the market functions well, all attempts to extract monopoly profits should fail in the long run. At the same time, these results preclude the use of non-market mechanisms to artificially increase (or reduce) ex-vessel prices. That is to say that the claim for public price fixing should prove to be equally ineffective at permanently affecting prices. Thus, the only way left for fishermen to increase their income per ton landed would be by increasing the value of landings and final products added.

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