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ABSTRACT

Evidence of market integration between farmed pangasius and wild-caught whitefish is provided in the literature, pointing towards pangasius prices being determined on the large international whitefish market. In the presence of price transmission in the value chain, global growth of pangasius farming does therefore not in itself reduce the farm-gate prices in Vietnam. In this paper, price transmission in the pangasius value chain from farmers in Vietnam, via export to final consumption in Germany, is tested using the Johansen cointegration framework. Price transmission is identified both between farm-gate prices and export prices in Vietnam and between export prices and retail prices in Germany. The Law of One Price was rejected in both cases, indicating imperfect price transmission. Weak exogeneity tests of market leadership identify a value chain with downstream market leadership consisting of German retailers leading exporters, which themselves lead farmers. The implication is that growth of Vietnamese pangasius farming can continue, ceteris paribus, without reducing prices substantially. Vietnamese farmers can invest in expansions without fearing self-inflicted price falls, but farmers and local communities remain prone to fluctuations following from supply and demand changes at the international whitefish market outside of their control.

1. Introduction

Traditionally, the international whitefish market was supplied by wild species such as cod, haddock, saithe, hake and Alaska Pollack. However, in recent decades, farmed species such as pangasius and tilapia have become an important part of the international whitefish market. In 2016, 15.4 million tonne of whitefish were supplied globally, from which half were farmed; farmed pangasius makes up 14% of the total supply. While the supply of wild caught species has fallen slightly during the last two decades (1997–2016), the annual global supply of pangasius has grown by 21%, and the global supply of tilapia has grown by 10% (FAO, 2019a; 2019b). Existing literature shows close market integration between European countries, as well as market integration between the different whitefish species (Asche et al., 2002); Nielsen, 2005; Nielsen et al., 2009, 2011; Norman-Lopez, 2009; Bronnmann, 2016; Bronmann et al., 2016). These findings reveal the existence of a large international whitefish market. With the total whitefish supply being seven times larger than the pangasius supply (see Appendix B), growth of the pangasius supply might induce only minor price reductions for the Vietnamese farmers, assuming an effect of price transmission.

In this study, we investigate price transmission and to test for market leadership in the pangasius value chain. For this purpose, we look the entire the supply chain from farmers in Vietnam, via export, to final consumers in Germany and assess implications for Vietnamese farmers. The Johansen cointegration framework is applied for monthly price series from 2007 to 2012 to reveal knowledge about price transmission by testing for cointegration, the Law of One Price and weak exogeneity between different nodes of the value chain. The knowledge on price transmission in the Vietnamese-German pangasius value chain

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is quite important for various reasons.

First, when price transmission is detected, farmers can invest in production expansions without fearing supply-induced price reductions. Pangasius is an integrated part of the international whitefish market (Bronnmann, 2016; Bronmann et al., 2016) and even large growth thus induces only negligible price reduction in the international market. Moreover, with price transmission, only a minimal price reduction is transmitted back to Vietnamese farmers. Without price transmission, theory predicts that growth of the pangasius supply, all other things equal, leads to substantial reductions in farm-gate prices, which may undermine investments.

Second, knowledge of price transmission is also interesting, since it makes pangasius farmers and local communities in Vietnam prone to price fluctuations induced by changing supply and demand within the whitefish market. Whitefish supply is determined by quotas for the main stocks such as cod in the Barents Sea and Alaska pollack in the North Pacific Ocean. Moreover, as shown by Jensen et al. (2014), improved fisheries management causes a wild supply growth in the long run, leading to a downward pressure on prices of farmed substitutes. Changing whitefish demand may also affect prices through growing income, food safety initiatives and increasing preferences; for example, for sustainable fisheries and farm practices. This occurs in particular when the value chain is controlled by retailers, as indicated by weak exogeneity tests. No matter the reason, increasing prices enhance investments of farmers and the livelihood of local communities in Vietnam, while price reductions do the opposite. Without price transmission, pangasius farmers and local communities are decoupled from the international whitefish market.

The implications of integration of markets supplied by both farmed and wild-caught fish were studied theoretically by Anderson (1985), who showed that aquaculture growth reduces prices of both farmed and wild fish, and increases the stock sizes of wild fish in the long run when the stocks initially are exploited above the maximum sustainable yield. Also departing from market integration, Jensen et al. (2014) identified theoretically how improved fisheries management reduces growth in aquaculture.

A broad empirical literature on market integration for whitefish exists. This includes Asche et al. (2002), who identify a European whitefish import market, integrated across countries with close integration of cod markets. Nielsen (2005) finds market integration between European countries for cod, saithe and hake at first-hand sale, whereas Nielsen et al. (2009) identify a closer market integration of total whitefish supply in Europe, consisting of landings and import, on the fresh rather than the frozen market. Tilapia imported to the US is found integrated with red snapper, dab, and blackback flounder in the US fresh fillet market by Norman-Lopez (2009). Nielsen et al. (2011) find farmed trout and wild-caught redfish markets integrated in Germany, while Bronnmann, 2016 identify close market integration between Alaska pollack, cod, saithe, tilapia and pangasius in Germany. Finally, some demand studies identify substitution of whitefish; for example, Thong et al. (2017), who estimate worldwide pangasius demand, and Bronnmann (2016), who finds that pangasius is part of the German whitefish market and indicates substantial substitution between pangasius and wild whitefish species. Hence, the literature shows the existence of an international whitefish market including pangasius integrated across countries and species.

Price transmission for whitefish is studied by Jaffry (2004), who identifies asymmetric price transmission between retail and auction prices for hake in France, with retailers responding slower to auction price reductions than to increases. Simioni et al. (2013) also identify asymmetric price transmission in the domestic value chain in France, but for cod. Asche et al. (2007) find price transmission in the cod value chain from Norway to Portugal, while Fernández-Polanco and Llorente (2015) find evidence of price transmission for hake imported to Spain, but not for hake in the domestic value chain. Furthermore, Ankamah-Yeboah and Bronnmann (2017) find price transmission and market leadership by retailers in the German parts of the Alaska pollack and cod value chains. Finally, Pham et al. (2018) study the pangasius value chain from Vietnam to Poland and find symmetric price transmission of the whole chain, except between retail and wholesale prices in Poland.

To the best of our knowledge, this paper is the first to identify price transmission and downstream market leadership of the full pangasius value chain and to show that farm-gate prices are largely unaffected by their own growth. It is also the first to point out the dependency of the whitefish market for local communities in Vietnam.

The paper is structures as follows: After this introduction, the model is specified in Section 2. The pangasius industry is presented in Section 3 together with data, while the results are provided and discussed in Section 4. Section 5 concludes the paper.

2. Model specification

In this article, we apply the standard market integration modelling based on cointegration as a complementary measure of detecting market governance through product prices along the value chain. Traditionally, a market integration test of the relationship between two market prices is expressed in logarithm as:

$$lnp_t^1 = \alpha + \beta lnp_t^2 + e_t \tag{1}$$

where p_t^i is the price in market i at time t and the parameter α is a constant term that reflects transportation costs and quality differences between markets 1 and 2. The error term e_t is assumed to be white noise. The parameter β determines the long-run relationship between the prices. If $\beta = 0$, there is no relationship between the price series, while $\beta = 1$ indicates that the LOP holds and there is complete market integration. In case of $\beta \neq 1$ or $\beta \neq 0$, the price changes transferred from one level of the chain to the other are incomplete in the long. Lags can be introduced into Eq. (1) to account for dynamic adjustment patterns (Ravallion, 1986). How Eq. (1) is handled depends on the nature of the stochastic processes of the vector of prices. In the presence of non-stationary price series, Eq. (1) produces spurious results. Cointegration analysis presents an alternative solution (Engle and Granger, 1987). The standard unit root test based on the Augmented Dickey-Fuller model (Dickey and Fuller, 1979, 1981) is used in this study.

A cointegration relationship exists when linear combinations of non-stationary variables integrated of the same order are stationary. In identifying cointegration, the Johansen (1988) approach is adopted over the Engle-Granger two step approach due to the interest of hypothesis testing on the cointegrating space. The vector error correction (VEC) representation of the Johansen (1988) cointegration test is given by

$$\Delta p_t = \sum_{i=1}^{k-1} \delta_i \Delta p_{t-i} + \Pi p_{t-1} + \mu + \varepsilon_t \tag{2}$$

where Δp_t contains a vector I(0) of n price series. The parameter μ is the deterministic component composed of the constant, trend, structural break, and seasonality; δ_i is the short run parameter. The long-run relationship (i.e., the cointegration vector) is captured by the matrix Π, defined as , where $\beta' p_t$ is stationary if a cointegration relationship exists and the matrix Π has a reduced rank (r): $0 < r < n$. To conclude, regarding market integration, a rank of $n - 1$ should be identified, where n is the number of price vectors.

The rank of the matrix Π is tested with modified chi-square tests called the trace and maximum eigenvalue statistics (Johansen, 1988). If $rank(\Pi) = 0$, then no cointegration relationship exists and if $rank(\Pi) = n$, then there is full rank – indicating that all the price series are stationary. As we conduct trivariate and bivariate tests, the ranks of

Π are expected to be 2 and 1, respectively. The number of k lags selected into Eq. (2) are based on one that addresses the first moment dependence in the data. Eq. (2) can be written out as trivariate and bivariate representations in Eqs. (3) and (4), respectively, assuming a lag of order 1 for simplicity:

$$\begin{pmatrix} \Delta p_t^1 \\ \Delta p_t^2 \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \end{pmatrix} + \begin{pmatrix} \delta_{11} & \delta_{12} \\ \delta_{21} & \delta_{22} \end{pmatrix} \begin{pmatrix} \Delta p_{t-1}^1 \\ \Delta p_{t-1}^2 \end{pmatrix} + \begin{pmatrix} \alpha_1 \\ \alpha_2 \end{pmatrix} (p_{t-1}^1 - \beta p_{t-1}^2) + \begin{matrix} \varepsilon_t^1 \\ \varepsilon_t^2 \end{matrix} \tag{3}$$

$$\begin{pmatrix} \Delta p_t^1 \\ \Delta p_t^2 \\ \Delta p_t^3 \end{pmatrix} = \begin{pmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \end{pmatrix} + \begin{pmatrix} \delta_{11} & \delta_{12} \\ \delta_{21} & \delta_{22} \\ \delta_{31} & \delta_{32} \end{pmatrix} \begin{pmatrix} \Delta p_{t-1}^1 \\ \Delta p_{t-1}^2 \\ \Delta p_{t-1}^3 \end{pmatrix} + \begin{pmatrix} \alpha_{11} & \alpha_{12} \\ \alpha_{21} & \alpha_{22} \\ \alpha_{31} & \alpha_{32} \end{pmatrix} \begin{pmatrix} p_{t-1}^1 - \beta_{13} p_{t-1}^3 \\ p_{t-1}^2 - \beta_{23} p_{t-1}^3 \end{pmatrix} + \begin{matrix} \varepsilon_t^1 \\ \varepsilon_t^2 \\ \varepsilon_t^3 \end{matrix} \tag{4}$$

where all parameters are as defined before. The β matrix contains the cointegrating vectors that define the long-run relationships of the vector of price series, p_t. The LOP is tested in this setting with the restrictions $\beta' = [1\ \ -1]$ for the bivariate and $\beta' = \begin{bmatrix} 1 & 0 & -1 \\ 1 & 0 & -1 \end{bmatrix}$. To strengthen the conclusion of cointegration, further restrictions following Hjalmarsson and Österholm (2010) are imposed on the bivariate ($\beta' = [1\ \ 0]$ and $\beta' = [0\ \ 1]$) and trivariate ($\beta' = \begin{bmatrix} 0 & 0 & 1 \\ 0 & 1 & 0 \end{bmatrix}$, $\beta' = \begin{bmatrix} 0 & 0 & 1 \\ 1 & 0 & 0 \end{bmatrix}$ and $\beta' = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \end{bmatrix}$) systems. This is a post-estimation test of stationarity among the variables, and it checks that the cointegration system is not driven by any stationary or near unit root processes. Rejection of these restrictions implies that stationarity is rejected.

The parameter α represents the speed with which prices adjust to the equilibrium. According to Engle and Granger (1987), the existence of a cointegrating relationship between variables implies that causality exists in at least one direction, hence one of the α' s should be different from zero. This is used to test for weak exogeneity ($\alpha = 0$) of the respective price vectors. A weakly exogenous variable indicates that the variable does not adjust and points towards market leadership. In the current setting of the value chain analysis, it is used to indicate which actors in the chain have power. In terms of global value chain interpretation, we use it to signal market governance and who is the leader in the chain. It is acknowledged that, similar to the market power literature, this is a weak measure since it is based on only price data; hence, it can be only used as an indication or signal.

3. The pangasius industry and data

Pangasius catfish (Pangasius hypophthalmus) is one of the fastest growing aquaculture species globally (FAO, 2010; 2019a), with annual production of over 1.1 million tonne (FishStatJ, 2019). Vietnam is the major producer, representing more than 75% of the global production. Pangasius catfish is raised popularly in freshwater bodies in the Mekong Delta, along the Mekong River. Previously, the fish were raised in cages and pens. Since Vietnam began to experience globalization, Vietnamese farmers have farmed pangasius in ponds with pellet feed, yielding a very high productivity of up to 500 tonnes/ha. In 2012, the total pangasius farming area of Vietnam was 3.586 ha, of which households account for 48.7%, farming companies for 49.1% and farmer collectives for only 2.1% (Tung et al., 2014). Recently, the farming area has increased significantly, peaking in 2015 at 5.900 ha (Mard, 2016). However, the farming area owned by households has diminished dramatically, and farming companies currently account for approximately 80% of the total farming area in the delta.

In 2018, Vietnam produced nearly 1.42 million tonnes of pangasius and over 97% of the production was used as raw material for more than 140 processing plants (EPA, 2014; VASEP, 2019a). Approximately 71% of the 5.400 ha farming areas in 2018 have received Global G.A.P or Aquaculture Stewardship Council (ASC) certification. Most processed products are used for export, and Vietnam represents over 90% of the global export value of pangasius, with an export value in 2018 of $2.26 billion (VASEP, 2019a). The US, China and the EU are the largest export markets of pangasius, accounting for approximately 60% of the total export volume in 2018 (VASEP, 2019b).

Pangasius products are exported mostly as frozen fillets, accounting for 98% of the total export value (EPA, 2014). Belgium, the Netherlands, the UK and Germany are the major retail markets in the Europe. In the period of 2007–2012, Germany imported approximately 2,355 tonnes of pangasius fillets per month (Table 1). Germany is the third largest importer of pangasius, with import value in 2017 of $23 million (VASEP, 2018). Pangasius frozen fillets are sold by the most popular retailers in Germany such as Aldi, Schwarz Group (Lidl), Rewe and Edeka. Pangasius has the advantage of low price, and the taste is comparable to other whitefish such as cod or haddock.

For this study, four different sources of data were used. The farm gate data were collected in Vietnam from the Vietnamese Association of Seafood Exporters and Producers (VASEP, 2015). VASEP obtains quantity and prices of pangasius products supplied by farmers to processors per month in different locations. The farm gate price used in this study is the monthly average farm-gate price of pangasius in the Mekong Delta over the period of 2007–2012.

Monthly export data for Vietnamese pangasius are available from the website of the International Trade Centre (www.trademap.org), WTO. Recently, pangasius products, i.e., frozen fillets, have been classified by two separate codes: HS030429 (frozen fish fillets) for the period 2007–2012 and HS030462 (frozen fillets, pangasius species) after 2012. Although the product code HS030429 applied to all types of frozen fish fillets, cross checking the data prior to 2012 revealed that more than 95% of frozen fish fillets exported from Vietnam were pangasius (the remainder were other fishes such as tilapia and carp). The monthly data on exported value ($) and quantity (kg) distributed to import countries were obtained, and the average export price was calculated for the period 2007–2012. All prices were converted to USD, and the analysis was conducted using nominal values (USD) per kilogram.

The retail data consist of two proprietary sets of data. In both datasets, the European Article Number (EAN) Code identifies each item. The first dataset was obtained from the Consumer Scan of the GfK Panel Service, Germany for the year 2007. In this dataset, households record the total expenditure (EUR) on, and volume purchased, of a specific EAN in one purchase trip. From this, the price per unit is calculated. The second proprietary dataset consists of scanned sales data (Symphony IRI Group 2012) containing sales data from the 1st week of 2008 to the 52nd week of 2012. In this dataset, each observation consists of the scanned EAN, the paid price (EUR) and the number of bought units. For our analyses, we focus on price series for frozen natural pangasius fillets.

All prices are converted to USD/kg and are aggregated to monthly averages. Fig. 1 shows the price movements of the three different price series. The Vietnamese producer and export prices are much lower than the retail prices in the German market. On average, retail prices are about ten times as high as the producer prices of pangasius and four times that of export prices. Farm, export and retail prices follow each

Table 1
Summary statistics of prices and quantity exports to Germany, monthly average 2007–2012.

	FARM (US $/kg)	EXPORT (US $/kg)	RETAIL (US $/kg)	Export to Germany (tonne/month)
Mean	1.010	2.909	12.676	2355
Maximum	1.631	3.749	17.795	4,268
Minimum	0.812	2.288	9.516	1,030
Std. Dev.	0.177	0.290	2.132	764

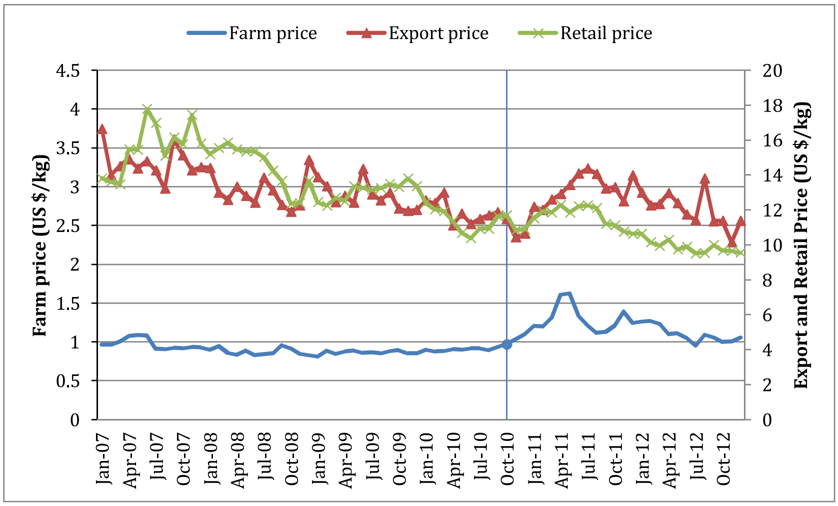

Fig. 1. Pangasius price development in the value chain (2007–2012).

Table 2
Augmented Dickey-Fuller unit root test.

Model Specification	Farm	Export	Retail
Level Model			
None	−1.771(0)	−0.824(1)	−0.920(0)
Intercept	−1.755(0)	−4.104(0)***|-1.801(3)	−0.830(0)
Trend	−2.163(0)	−4.790(0)***|-2.218(3)	−3.583(0)**
First Difference			
None	−7.526(0)***	12.046(0)***	−9.360(0)***
Intercept	−7.473(0)***	NA | -11.998(0)***	−9.400(0)***
Trend	−7.418(0)***	NA | -11.910(0)***	NA
Breakpoint (*Level*)			
Intercept (2010m09)	−3.759(0)		
Intercept and Trend (2010m12)	−4.572(0)		
Trend (2007m08)	−2.560(0)		
Breakpoint (*First Difference*)			
Intercept (2011m04)	−8.875(0)***		
Intercept and Trend (2011m04)	−8.741(0)***		
Trend (2011m01)	−7.563(0)***		

Lags selected using the Schwarz Information Criterion and modified Hannan-Quinn (|##), NA indicates no test conducted at the specification due to rejection in levels. *** and ** indicate statistical significance at 1% and 5% levels.

other with mainly a decreasing trend until April 2009 and follow each other again after December 2010. First by increasing, and then, by decreasing afterwards. Between these periods, some temporary deviations appear that may indicate a structural break. The periods prior to and after the break show a stable relative price[1] . Price levels in the farmgate prices were generally higher than before the break while that of the export and retail prices were lower on the average.

4. Results and discussion

As the test for cointegration depends on the time series properties of the data, we first conduct the Augmented Dickey-Fuller (ADF) test to ensure that all prices are integrated of the same order. The ADF test results are shown in Table 2. A unit root is tested with all the ADF model specifications, including restricted and unrestricted deterministic terms. A unit root is confirmed first with failing to reject the null in levels and subsequently rejecting in first difference.

As shown in Table 2, the producer price series shows unit root process for all specifications using the Schwarz information criterion to select the optimal lags. A unit root is identified in an export price series when both the intercept and trend terms are restricted. For the retail price series, a unit root process is identified by intercept and none specifications. Since the producer price appears to show a structural break, which may be caused by the variation of exchange rate in 2010 and beginning 2011, the breakpoint unit root test using the Dickey-Fuller min-t statistic as the breakpoint selection is tested. Specifications with intercept and trend terms are used. The results are presented in the lower part of Table 2. Different break dates are identified with different specifications both in levels and first difference. The evidence still confirms the presence of a unit root process. In matching the break dates with the development of the price series in Fig. 1, September 2010 seems to be the most plausible breakpoint (see vertical line in Fig. 1). As a result, a break dummy variable is created for this period and included in the cointegration test and vector error correction models.

Given the presence of unit root in the price series, or the lack thereof, which would be made evident subsequently, the cointegration-based test of Johansen (1988) market integration is used to uncover causality and relationships among prices in the pangasius value chain. We begin with a trivariate model to uncover relationships in a common system. Thereafter, bivariate relationships are examined in isolation. Johansen (1988) suggests two statistics for the number of ranks of the cointegration vectors in the system; these are the maximum eigenvalue and the trace value. The evidence of market integration should present a rank (r) of one less the number of endogenous variables (K) included in the system (i.e., r = K-1).

All models presented are specified with seasonal monthly dummies (M), without trend and constant terms (1) and a lag length indicated in parenthesis in Table 3, thus 1 M(#). The estimation of the model was based on a search process of one that finds evidence of cointegration, fulfils residual normality and homoscedasticity, and has no serial correlation. The Ljung-Box test statistic was used to test for serial correlation. These misspecification tests are presented in the bottom of Table 3. Where cointegration is identified in multiple specifications, information criteria are used as a guide.

Considering the trivariate cointegration models, the null hypothesis of no cointegration (r = 0) and the presence of cointegration with rank of one (r = 1) are significantly rejected at the 5% level for both the trace and max-Eigen value statistics. However, the presence of cointegration with rank of two (r = 2) fails to be rejected. This implies that the three nodes of the value chain are cointegrated and with a rank of

[1] Due to the relatively smaller values of the producer price, the scale of the axis conceals the variation in the prices before the break. See Appendix A for a chart of only the producer price.

Table 3
Johansen cointegration test.

Statistic Name	Trivariate Farm-Export-Retail	Bivariate Farm-Export	Export-Retail	Farm-Retail
Model	1M(4)	1M(1)	1M(2)	1M(2)
λ_{trace}				
$r = 0$	48.854***	13.878**	12.493**	13.939**
$r = 1$	23.712**	0.837	0.463	0.363
$r = 2$	4.915			
λ_{max}				
$r = 0$	25.142**	13.041**	12.030**	13.575**
$r = 1$	18.797**	0.837	0.463	0.363
$r = 2$	4.915			
$\beta_{FR}\beta_{ER}$				
LOP: $P(x^2)$	11.521***	11.803***	11.234***	13.073***
Post Stationarity Test				
$\beta_F \sim I(0)$	16.792***	12.103***		13.154***
$\beta_E \sim I(0)$	19.691***	4.373**	11.042***	
$\beta_R \sim I(0)$	14.149***		10.698***	5.793**
P(Normality)	0.078	0.065	0.361	0.271
P(Homoscedasticity)	0.421	0.051	0.352	0.107
P(Q-Stat) |P(LM-Stat)	0.118	0.464	0.513	0.412

*** and ** indicate statistical significance at 1% and 5% levels.

two, we can conclude market integration. A test of the Law of One Price (LOP[2]) is rejected at the 1% significance level. This is an indication that the degree of market integration is partial.

As shown previously, the test on the stochastic processes of the series is sensitive to model specification and lags used. Hence, in other to ensure that the evidence of cointegration is not driven by stationary time series processes, the Hjalmarsson and Österholm (2010) post-estimation stationarity test is conducted. The likelihood ratio test of the restrictions, respectively indicated in Table 3 as β_F, β_E and β_R, are all rejected at the 1% significance level. This is an indication that the model is not driven by stationary variables.

For the bivariate models, we investigate relationships between producers and exporters, exporters and retailers and then producers and retailers. The rank zero tests in all bivariate models are significantly rejected at the 5% level but not so at the rank of one. Hence, we conclude the evidence of market integration between actors in the chain. The bivariate analysis is conducted and confirms the findings of Asche et al. (2002), which indicate that, if three prices share a common stochastic trend, then it must be that all prices in the system are pairwise cointegrated. The LOP test is rejected for all market pairs, thereby confirming the case of the trivariate models. The Hjalmarsson and Österholm (2010) restrictions are also rejected; hence, we are able to indicate that the markets pairs are integrated. As a result, price changes passing through in the value chain are incomplete in the long run.

Components of the underlying vector error correction model of the cointegration test are reported in Table 4. The speed of adjustment parameters/loading matrix indicated by the α's shows how quickly actors respond to changes or shocks towards the long run equilibrium. Considering the trivariate model, it can be observed that the upstream actors (producers and exporters) adjust to changes in the system but the downstream retailers do not. From the magnitudes and significance of these parameters, exporters are the quickest to react and they react faster in response to their own price deviations from the retailers (α_{i1}) than from producer price deviations from retailers (α_{i2}).

Weak exogeneity in the trivariate model is implemented by the joint test of the α's using the LR test. The producer and exporter models are significantly rejected at the 1% level while the retailer model fails to be rejected. This signals that retailers have the leading market role in the value chain and hence exert a certain amount of control/power or governance over the flow of information, which is reflected in the product prices.

In the examination of the relationship between the market pairs, the speed of adjustment parameter shows that the equilibrium adjustment process is fastest for exporters when reacting to deviations from the retail market changes. However, they are not responsive to producer prices. The unidirectional pattern observed with the weak exogeneity test is that, for any market pair, the downstream actors exert power over the upstream actors in the value chain. The block exogeneity test presents cross price causality in the short run and the p-values are shown in the lower part of Table 3. Generally, no significant relationships exist in the short run except for the retail model in the trivariate model. In this model, the producer and exporter prices are joint predictors of the retail prices at the 5% significance level, but not individually.

The coefficient of the structural dummy break shows significance in model types that include farmgate prices. As can be seen, the structural break has a positive and significant effect on farmgate equations and shows that the break resulted in general increase in farmgate price levels while it resulted in a decrease in export prices. These effects are visible and confirmed in Fig. 1. The implication of the finding of imperfect price transmission and downstream market leadership is that price differentiation initiatives, such as certification for sustainable farm practices in the Aquaculture Stewardship Council scheme, benefit not only downstream companies but also Vietnamese pangasius farmers.

The implication of imperfect price transmission together with the existence of the international whitefish market is that global growth of pangasius farming, ceteris paribus, can be expected only to have a negligible effect on the farm-gate price in Vietnam. The reason is that the global whitefish supply to Germany is 7 times larger than for pangasius alone (Appendix B). The small effect is also confirmed by Thong et al. (2017) who estimate an inverse Almost Ideal Demand system and identify inflexible prices with own price flexibilities on average on −0.42. Thus, farmers, all other thing equal, can invest in production expansion without fearing undermining self-inflicted price reductions. This is important since pangasius is a cheap whitefish competing primarily on price.

Finally, the implication of the simultaneous presence of price transmission and market integration is that Vietnamese pangasius farmers, as well as the livelihood of local communities depend on the whitefish market. The income of Vietnamese farmers is determined by whitefish supply and demand. While farming continues to grow, wild catches are fluctuating. Total catches from the largest whitefish stock worldwide, Alaska pollack in the North Pacific Ocean, declined 20% over the last two decades (1997–2016), but are increasing again after a low in 2009 (FAO, 2019b.). Catches from another large stock, cod in the Barents Sea, are fluctuating along the same trend (ICES, 2018). For Alaska pollack none of the three main stocks at the Aleutian Islands, the Eastern Bering Sea and the Gulf of Alaska are overfished (North Pacific Fisheries Management Council, 2018). The Barents Sea cod stock also has full reproductive capacity and fishing mortality is within the precautionary level (ICES, 2018). Hence, while it remains a matter of speculation, the current healthy stock status indicates high supplies also in the coming years, which may induce a continued downward pressure on prices of whitefish including pangasius.

Whitefish demand is determined by income and preferences with a large share of whitefish being consumed in high-income countries. While income in these countries is increasing, the income growth is often limited. Preferences for healthy food and for sustainable caught and farmed fish may, however, be increasing and induce an upward pressure on prices of pangasius.

5. Conclusions

Price transmission of the Vietnam-Germany pangasius value chain has been studied by testing cointegration of price series, the Law of One Price and weak exogeneity. Market integration is found by both

[2] Imposed with the following restriction on the cointegrating space: $\beta' = \begin{bmatrix} 1 & 0 & -1 \\ 0 & 1 & -1 \end{bmatrix}$

Table 4
Vector error correction (VEC) estimates and tests.

	Trivariate VEC Model			Bivariate VEC Models					
	ΔFarm	ΔExport	ΔRetail	ΔFarm	ΔExport	ΔExport	ΔRetail	ΔFarm	ΔRetail
Loading Matrix									
α_{i1}	−0.291**	0.435***	0.062	−0.260***	0.079	−0.386***	0.038	−0.278***	0.082
α_{i2}	−0.213	−0.810***	0.094						
Weak Exogeneity Test									
$\alpha_{i1} = \alpha_{i2} = 0$	12.923***	16.966***	2.253	11.167***	0.882	8.961**	0.144	11.063***	1.438
Break Dummy	0.094***	−0.077**	−0.028	0.067***	−0.022	0.020	−0.011	0.073***	−0.030
P-value (Block Exogeneity)									
ΔFarm	–	0.120	0.068		0.579				0.856
ΔExport	0.667	–	0.063	0.865			0.790		
ΔRetail	0.720	0.735				0.439		0.418	
Joint = 0	0.810	0.378	0.025						

*** and ** indicate statistical significance at 1% and 5% levels.

bivariate and trivariate testing in the whole supply chain from Vietnamese farmers to German consumers. The Law of One Price is rejected in all cases, indicating imperfect price transmission. Thus, prices in different nodes can fluctuate, but they cannot continue to deviate too much from each other. Weak exogeneity tests fail to reject downstream market leadership of the value chain.

The implication of price transmission is that Vietnamese farmers may benefit from achieving a part of the downstream price premiums of, for example, ASC or Global G.A.P certification (Tran et al., 2013). According to the Centre for the Promotion of Import from developing countries, in The Netherlands, the consumer price of frozen fillet with ASC certification at Europe in 2018 was double the price of the product without the ASC (CBI, 2018). The implication of simultaneous price transmission and a whitefish market is that Vietnamese pangasius farmers can expand without fearing that self-inflicted price reductions undermine investments. A further implication is that incomes of Vietnamese pangasius farmers and the livelihood of their local communities are determined by the whitefish market.

That self-inflicted price reduction is not a barrier for growth indicates that investment in the sector may be economically viable. However, there are usually several factors affecting success or failure of growth strategies, so that individual actors have to carefully consider strengths and weaknesses. The method can be applied elsewhere to identify how aquaculture growth affects prices and income of the farmers. It can also be applied elsewhere to provide knowledge to farmers and policy makers on how fish market developments affect farmers' income and livelihood of local communities.

The paper has a limitation in use of the data in short period from 2007 only to 2012. While farm gate and export price are cheap and easy to obtain, the retail scanner data are very expensive so that we did not have data up to date. However, the data show a very clear trend in price and good enough for testing our hypothesis and therefore the data used did not impaired our conclusion. The results are based on a relatively small dataset that can be extended in the future. Moreover, the aquaculture sector develops dynamically, implying that there is a need to take the latest market data into account before companies decide on expansion. Nonetheless, the findings represent important background knowledge when considering initiatives to expand production.

Declaration of interests

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

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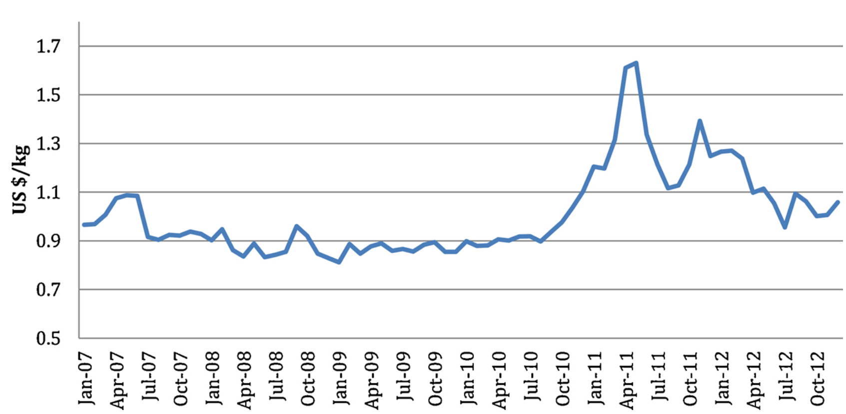
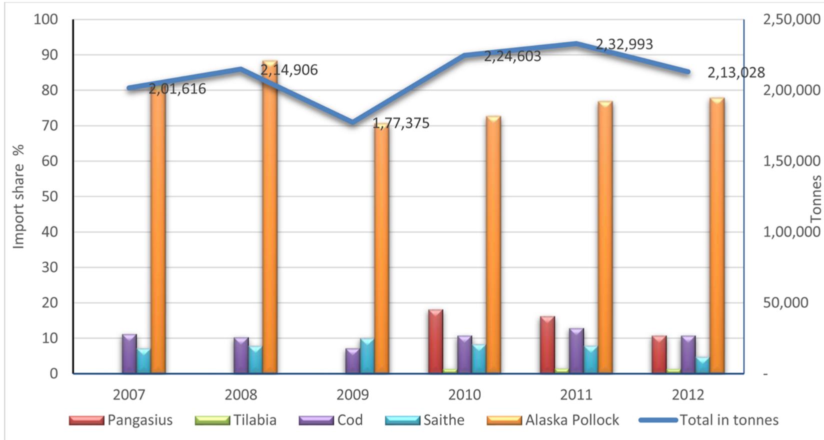
Appendix A. Average farm gate price US $/kg 2007–2012



Appendix A. German import of frozen fillet of whitefish in 2007–2012



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