The Impact of Imports on Domestic US Shrimp Prices

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ABSTRACT

The US seafood market has fundamentally changed during the last three decades. Stable landings and increasing demand have led to a rapid increase in imports—79% of domestic seafood consumption is estimated to come from imports. Despite several policies supporting vulnerable coastal communities, little attention has been given to the impact of imports on prices obtained by domestic producers. Here we investigate the impact of imports on domestic prices for shrimp, a fishery that in the 1980s was the most valuable in the US, but that has seen real landed values decline by one-half since then. Using cointegration analysis, we show that domestic prices closely track those of "shell-on frozen" imports, indicating that import competition largely drives this trend and that domestic US prices are now set in the global market. A similar market structure is likely to be present for other species facing strong import competition.

Key words: Imports, price determination, shrimp.
JEL codes: Q21, Q22, Q27.

INTRODUCTION

The US seafood market has fundamentally changed during the last three decades. After a surge in domestic production following the expansion of the exclusive economic zone to 200 miles from 1977, the yearly landings of wild fish stabilized at around 4 million metric tons (mt) at the end of the 1980s, and domestic aquaculture production remains limited (Anderson et al. 2023; Garlock, Anderson, et al. 2025). However, seafood demand has been increasing, and this demand has largely been met by increasing imports (Anderson et al. 2019; Shamshak et al. 2019), with the US currently importing an estimated 86% of seafood consumed (NMFS 2024). Stagnant landings of wild fish and a rapidly increasing aquaculture production characterize seafood production globally (Garlock, Asche, et al. 2020; Garlock et al. 2024; Naylor et al. 2023; Asche 2025), and it is not surprising that US seafood imports largely reflect this development. Shamshak et al. (2019) note that four out of the five most consumed seafood species in the US are primarily farmed, and Love et al. (2023) show that imports comprise the main source for most species consumed in the US.

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Authors Atle Oglend and Frank Asche and would like to acknowledge support from the Norwegian Research Council (#352350) and USDA Hatch #7004716.

Received August 12, 2024; Accepted August 26, 2025; Published online February 5, 2026.

Marine Resource Economics, volume 41, number 2, April 2026.
 Published by The University of Chicago Press on behalf of the MRE Foundation. https://doi.org/10.1086/739083

For most major seafood species, there exist global markets with a common price determination process (Smith 2012, 2023; Anderson et al. 2018). Although there is significant room for premiums associated with a number of product attributes, such as fish size, freshness, and other measures of quality (Asche 2025), prices for different products of the same species or species group tend to move together over time (Asche et al. 1999, 2004). As US imports have been increasing for many species, one would accordingly expect that the prices of these species are being set at the global market not only for the imports, but also for the domestic landings being exposed to competition from these imports. Asche et al. (1999) show that this is the case for salmon and provide empirical evidence supporting the model of Anderson (1985). Specifically, in a market where seafood is supplied from both a wild source with a limited resource and aquaculture with the potential to grow, the price will be determined by the cost of production in aquaculture. Valderrama and Anderson (2010) and Asche, Eggert et al. (2022) show that this is true also when the fisheries are regulated with a quota to protect the stock.

There has been surprisingly little attention given to the dynamics of the price determination process for most domestic species in the US and most other countries. The impacts of fisheries management on fishing communities have garnered significant research and policy attention (Kroetz et al. 2015, 2022; NASEM 2021, 2024; Birkenbach et al. 2022; Cojocaru et al. 2022; Asche et al. 2025), and prices and globalization are key drivers of economic outcomes in fishing communities (Smith et al. 2010; Asche et al. 2018; Birkenbach et al. 2019, 2023; NASEM 2021). Despite the importance of the price obtained for fishing communities, in the US salmon is the only species where a formal econometric analysis has been conducted (Asche et al. 1999). Unfortunately, a weak understanding of the price determination process can lead to governance interventions that are damaging. One prominent issue is the strong resistance from fishers to an expansion of US aquaculture production (Knapp and Rubino 2016; Rubino 2023), as a new aquaculture industry is perceived to increase competition and thereby reduce prices. However, if the price is set in a global market, the only impact of an increased domestic supply is a reduction in imports and no change in the competitive position for domestic fishers. Another issue is the stabilization of fishers' income (Kasperski and Holland 2013). The main emphasis in this literature is on species targeted and landed, and there is limited attention given to the sources of price variation.

Here we focus on price determination for US shrimp and directly test whether domestic shrimp prices are exogenously driven by import prices, which can be taken as an estimate of the global price. Shrimp is the most popular seafood in the US, and the domestic shrimp fishery has been one of the largest fisheries in the US for decades (Smith et al. 2017; Asche, Eggert, et al. 2022). Shrimp has been the most consumed seafood product in the US since 2000, and consumption has more than doubled since 1990 (Shamshak et al. 2019; Love et al. 2022). Massive growth in US shrimp consumption is primarily due to the significant increase in imports, mostly from aquaculture. Total world aquaculture shrimp production grew from just over 1 million mt in 2000 to over 4 million tons in 2020 (Garlock et al. 2023; FAO, n.d.). As demand has grown, global shrimp production has experienced bursts of growth punctuated by diseases and trade conflicts (Keithly and Poudel 2008; Asche et al. 2021; Petesch et al. 2021). Over time, innovations and productivity growth have resulted in significantly lower prices, a feature that has been observed for other successful aquaculture species (Asche, Eggert, et al. 2022).

Declining real shrimp prices have had a strong impact on landed value and the economic importance of the fishery. Prior to the 1990s, the shrimp fishery was the most valuable fishery in the US and an important supplier for shrimp domestically. In 1980, the real total landed value of the three main shrimp species (brown, white, and pink shrimp) was $1.283 billion (2024 USD), followed by salmon with $1.172 billion and king crab with $673 million in real landed values (NOAA 2024). While shrimp landings have been relatively stable over time, a steady decline in real price over the last decades has reduced the combined real value of brown, white, and pink shrimp to only $206 million in 2023 (Asche, Oglend, and Smith 2022; NOAA 2024).

The existing literature on US shrimp markets is suggestive of price determination in the global market, but the hypothesis has not been tested directly. Asche et al. (2012) show that US prices are cointegrated with imports but leave open the possibility that prices are jointly determined. Smith et al. (2017) show that domestic US supply shocks due to ecological disturbances can be detected in short-run size-based prices, but in the long run US shrimp obey the law of one price. Petesch et al. (2021) show that the global supply disruption in shrimp aquaculture had lasting effects on US wild-caught shrimp prices, and Asche, Oglend, and Smith (2022) show that the importance of the domestic shrimp landings declined in the price determination process as imports increased. Debaere (2010) shows how trade policies in Europe in the late 1990s and early 2000s diverted global shrimp exports away from Europe towards the US.[1] This caused a reduction in the price of imported shrimp in the US relative to Europe. However, this study does not analyze the interaction of the import market with the market for domestically produced shrimp in the US.

To investigate the price determination process of US wild-caught shrimp, we conduct market integration, weak exogeneity, and Granger causality tests. Together, the analysis allows us to test whether the domestic shrimp price is exogenously driven by the import price, which we assume is representative of the global price. As argued in Debaere (2010), shrimp exporters allocate shrimp to the market that provides the greatest net return, leading to stable relative prices, although trade policies can affect relative prices between markets over time.

Although the shrimp market is heterogenous in that there are many product attributes that have value (Ellis et al. 2016; Hukom et al. 2020), including size, species, and product form, there is strong evidence that these heterogenous products are part of an integrated market where the law of one price (or constant relative prices) holds (Asche et al. 2012; Smith et al. 2017). This allows us to create a domestic price index and an import price index, and we find that in the long run domestic prices are indeed driven by imports.

DATA

Data on the Gulf of Mexico shrimp fishery are from NOAA's ShRCoM database. ShRCoM is the primary source of microdata used for fisheries management in the Gulf of Mexico. The data report daily landing values (USD) and quantities (lb.) for brown, pink, and white shrimp by eight size categories. We aggregate from daily to monthly values.

While our sample does not cover all domestic US shrimp production, previous research has shown that different shrimp markets are integrated geographically and by species. Ankamah-Yeboah et al. (2017) study market integration between the cold- and warm-water shrimp in Europe and find market integration between cold- and warm-water unprocessed shrimp imports. Vinuya (2007) documents that wholesale shrimp markets in Tokyo, New York, and Europe for specific shrimp products are integrated, and provides evidence supporting the law of one price in shrimp markets. Asche et al. (2012) show that brown, pink, and white shrimp prices in the Gulf of

1. Similar effects are reported for other trade shocks for seafood (Santika et al. 2025) as well as more generally (Oglend et al. 2024a).

Mexico are integrated with the law of one price holding. Pincinato and Asche (2016) document that wild and aquaculture shrimp in Brazil are integrated. These studies all point to a highly global and integrated shrimp market. Based on this research, we take the Gulf of Mexico shrimp prices as representative of the price of domestically produced shrimp.

Shrimp import volumes (kg) and values (USD) are from the US Department of Commerce, reported by NOAA Fisheries (https://www.fisheries.noaa.gov/foss/f?p = 215:2:16116515755598). We use the "shell-on frozen" category. This is closest in type to the largest volumes of domestically produced shrimp. In contrast, products like breaded shrimp or shrimp in frozen meals contain significant added value that could distort inferences when comparing with the Gulf of Mexico vessel-level fishery data. Imports by size class are available only from July 1990, and so we start our sample for both imports and domestic shrimp in July 1990. In later years the shell-on frozen category has been further disaggregated into warm-water and cold-water shrimp, with warm-water shrimp disaggregated into farmed and wild origin. We reaggregate all later disaggregation to be consistent with the import classification in July 1990. There are nine size categories reported in the import data. All data series are available through December 2023.

All values are deflated using the US CPI and expressed in December 2023 US dollars. We create real unit values by size category by dividing real monthly landing values (import values) by landing volume (import volume, all expressed in pounds). This approach is valid, as Asche et al. (2012) document that the law of one price holds for white, pink, and brown shrimp prices and all sizes categories for the Gulf of Mexico fishery. Figure 1 shows the time series or all real imported and domestic unit values by size grouping. Both shrimp sources display declining unit values, with a consistent size premia ranking.

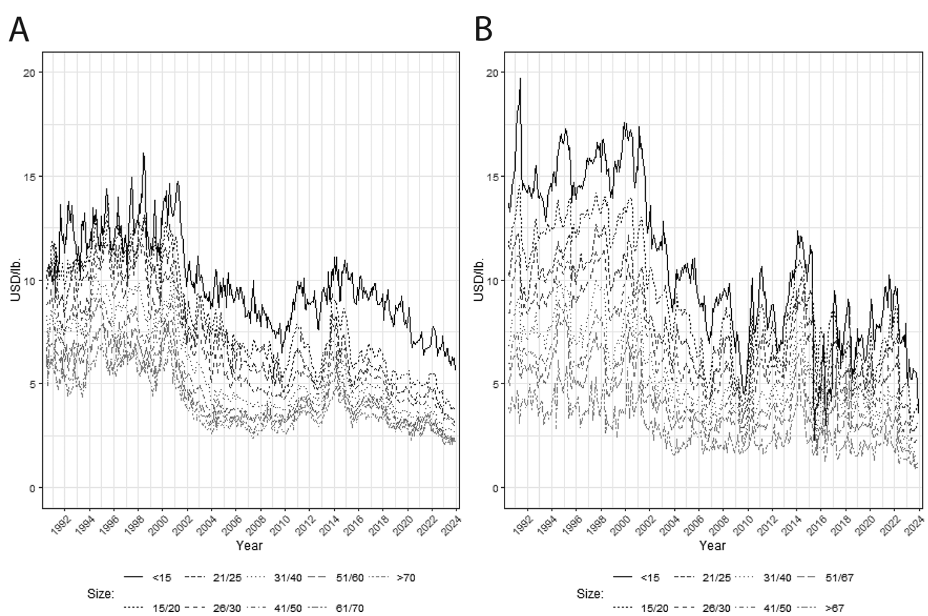

Figure 1. Real Unit Values of Imported and Domestic Shrimp by Size Category, July 1990–December 2023. Values expressed in December 2023 US dollars. (A) Import unit values are imported frozen shell-on shrimp. (B) Domestic unit values are ex-vessel Gulf of Mexico prices of brown, white, and pink shrimp.

To measure the common import and domestic shrimp price component, we create Fisher price indices across all size unit values as shown in figure 1. The Fisher index measures the average price change of a basket of different size categories of shrimp. It effectively measures the changes in the average cost of shrimp, considering both historical and current landing and import patterns. The reference period is set to 2005.

Figure 2 plots the Fisher indices for the domestic and import price measures from July 1990 to December 2023. As expected, the price measures are highly correlated, the volatility is higher for domestic shrimp, and the real price of shrimp declines over the sample period.

Figure 3 shows total Gulf of Mexico catches along with total imports of shell-on frozen shrimp, in million pounds. Domestic shrimp catches have been relatively stable on average over the period, with a slight negative trend. Both series show substantial seasonality, with imports offsetting the domestic supply seasonality—increasing when the main domestic season is winding down. There are some disruptions to the import seasonality towards the end of the sample, potentially because of COVID, but this requires further research. The large and persistent increase in imports in the early 2000s can be ascribed to growth in global aquaculture. We also observe a decline in imports in the early 2010s owing to aquaculture disease issues in this period.

METHODOLOGY

To investigate the dynamic relationship between the domestic and imported shrimp prices, we take a standard approach and use a VAR model.[2] The VAR(p) model is given by the following:

$$\boldsymbol{y}_t = \boldsymbol{\mu}_\mathbf{t} + \sum_{i=1}^{p} \Gamma_i \boldsymbol{y}_{t-1} + \boldsymbol{\epsilon}_t, \tag{1}$$

where $\boldsymbol{y}_t = [p_{IMPt}, p_{DOMt}]'$ is the vector of (log) monthly import and domestic prices measures, $\boldsymbol{\mu}_t$ is a two-dimensional row vector of deterministic terms, Γ_i is the 2×2 matrix of dynamic adjustment parameters for the ith month lagged price vector, and $\boldsymbol{\epsilon}_t$ is a two-dimensional row vector of price components not explained by the price history.

The import price Granger causes the domestic price if including information about the history of the import price improves on the forecast of the domestic price beyond the information contained in the history of the domestic price itself.[3] For instance, disease events or other supply disruptions in aquaculture might cause a higher global price of shrimp. The higher import price leads to demand substitution towards domestic shrimp, with a resulting higher domestic shrimp price. With monthly data the import price provides predictive power on the future domestic price if this effect works itself out over a period longer than a month. Let $\Gamma_{(ij),k}$ be the row i column j element of the Γ_k matrix. Granger non-causality of price j implies $\Gamma_{(ij),1} = \Gamma_{(ij),2} = \cdots = \Gamma_{(ij),p} = 0$.

When prices are cointegrated, we can further disentangle short-run and long-run dynamic causality patterns. The cointegrated VAR can be represented in error correction form as

2. Asche et al. (2004) review different approaches to market integration tests and the theory behind them. While most market integration studies investigate the relationship between the prices, fewer investigate Granger causality or weak exogeneity. Landazuri-Tveteraas et al. (2021), Pincinato et al. (2022), Roll et al. (2022), Polanco et al. (2023), Salazar and Dresdner (2023), and Salazar et al. (2024) provide some recent examples for other species. Oglend et al. (2022) show how the price relationships hold also at the firm level.

3. The econometrics of testing for Granger causality and exogeneity in a cointegrated system is discussed in Johansen and Juselius (1994). Engle et al. (1983) provide a thorough discussion of exogeneity and causality in a time series setting.

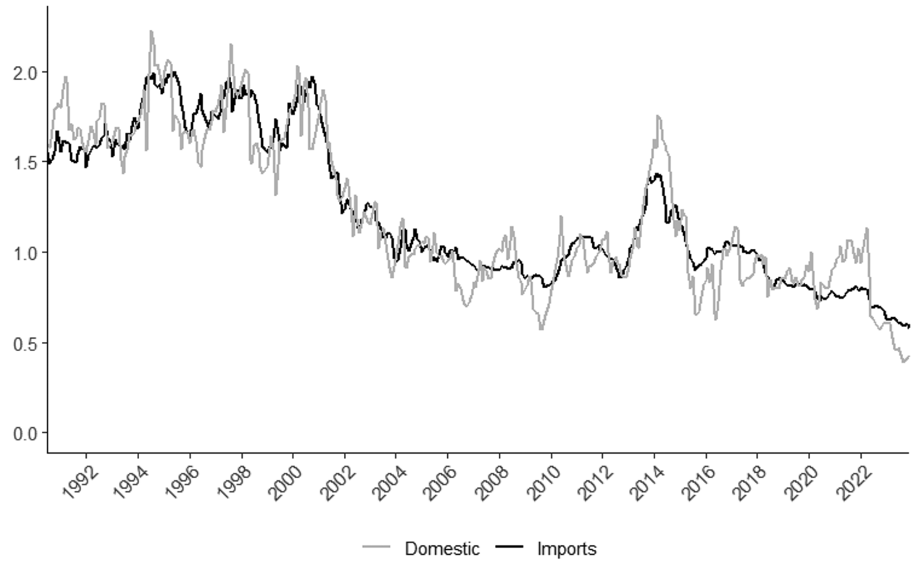

Figure 2. Real Fisher Price Indices, Domestic and Imported Shrimp (2005 = 1)

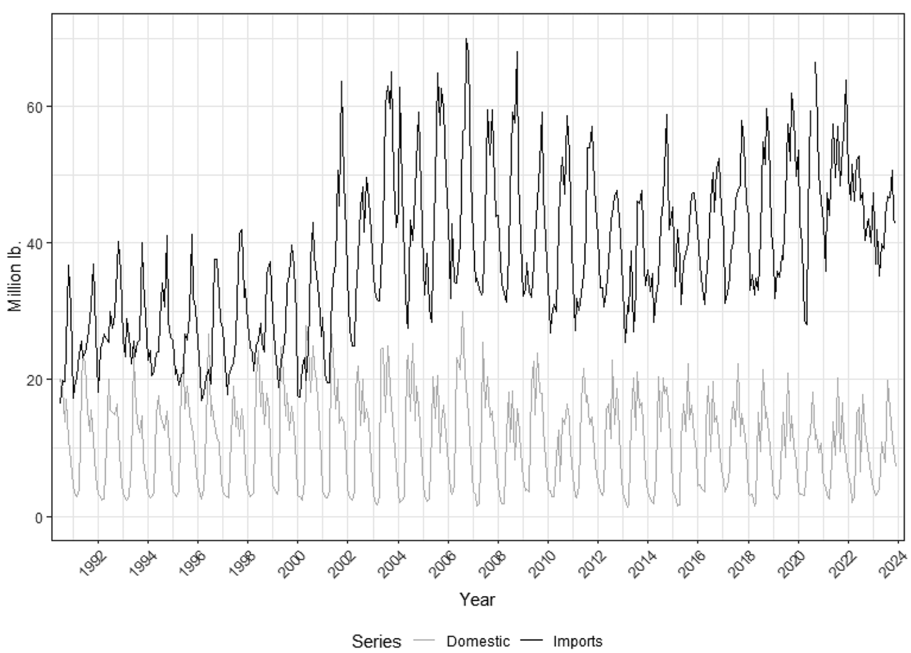

Figure 3. Domestic (Gulf of Mexico) Catches and Shell-On Frozen Shrimp Imports (in million pounds). Domestic catches consist of brown, white, and pink shrimp of size categories 1 (<15) to 8 (>67). Imported shrimp consists of size categories from 1 (<15) to 9 (>70).

$$\Delta \boldsymbol{y}_t = \tilde{\boldsymbol{\mu}}_t + \boldsymbol{\alpha}'\boldsymbol{\beta}\boldsymbol{y}_{t-1} + \sum_{i=1}^{p-1}\tilde{\Gamma}_i\Delta\boldsymbol{y}_{t-1} + \tilde{\boldsymbol{\epsilon}}_t, \tag{2}$$

where $\tilde{\Gamma}_i = -\sum_{j=i+1}^{j=p}\Gamma_j$, and $\boldsymbol{\alpha\beta}' = \sum_{j=1}^{p}\Gamma_j - I$.

When prices are integrated of order one and cointegrated, the 2×2 long-run impact matrix $\boldsymbol{\alpha}'\boldsymbol{\beta}$ has rank one and we can normalize the single cointegrating vector as $\boldsymbol{\beta} = [1, -\beta]$. Cointegration implies the linear combination $\boldsymbol{\beta y}_{t-1} = p_{IMP,t} - \beta p_{DOM,t}$ is a weakly stationary process. If $\beta = 1$ prices move proportionally over time, and the law of one price holds.

The vector $\boldsymbol{\alpha}' = [\alpha_{IMP}, \alpha_{DOM}]$ determines the speed of adjustment of each price series to deviations from long-run equilibrium. Since the long-run impact matrix is $\boldsymbol{\alpha\beta}' = \begin{bmatrix}\alpha_{IMP} & -\alpha_{IMP}\beta \\ \alpha_{DOM} & -\alpha_{DOM}\beta\end{bmatrix}$, it is clear that when $\alpha_{IMP} = 0$ the domestic price has no long-run impact on the import price. In this case, shocks to the domestic price may affect the import price in the short run, but not in the long run, the stochastic trend. If this is the case, the import price is long-run weakly exogenous, and the stochastic trend in the domestic price is determined by the stochastic trend in the import price.

If $\tilde{\Gamma}_{(ij),1} = \tilde{\Gamma}_{(ij),2} = \cdots = \tilde{\Gamma}_{(ij),p-1} = 0$, then price j is Granger non-causal on the short-run movement of price i. The short run is here defined by the choice of lag order p. For instance, if $p = 2$, the short run covers the month-to-month changes in the domestic price. If $p = 3$, the short run refers to price movements over 3 months. See table 1 for an overview of the restrictions for different exogeneity tests in the VAR model.

Weak exogeneity and short-run Granger causality provide all information needed to determine overall Granger causality. In terms of the original VAR coefficients, price i is long-run weakly exogenous if $\sum_{k=1}^{p}\Gamma_{(ij),k} = 0$. If price j is Granger non-causal, then price i is long-run weakly exogenous in the bivariate system. However, the converse is not necessarily true. Price i might be long-run weakly exogenous and price j Granger causal. Price j then has a short-run impact of price i, but no long-run persistent impact.

An example might be transitory seasonal impacts of the domestic price on the import price. Testing both long-run weak exogeneity and Granger non-causality allows us to disentangle short-run transitory and long-run persistent dynamic causality relationships between the prices.

With monthly data we are unable to test for causality relationships that work themselves out within a month. If the import price responds within a month to a domestic shrimp supply or demand shock, this will reveal itself as nonzero instantaneous correlation in the VAR residuals. Causal inference would then require a different research design, such as an instrumental variable approach or theoretical assumptions on the instantaneous correlation structures between the markets.

Table 1. Restrictions for Different Exogeneity Tests in VAR Model

	Restriction	Implication
Granger non-causality	$\Gamma_{(ij),1} = \Gamma_{(ij),2} = \cdots = \Gamma_{(ij),p} = 0$	Price j provides no predictive information on price i.
Short-run Granger non-causality	$\tilde{\Gamma}_{(ij),1} = \tilde{\Gamma}_{(ij),2} = \cdots = \tilde{\Gamma}_{(ij),p-1} = 0$	Price j provides no predictive information on short-run changes in price i.
Long-run weak exogeneity	$\sum_{k=1}^{p}\Gamma_{(ij),k} = 0$ (or $\alpha_i = 0$)	The stochastic trend in price j does not determine the stochastic trend in price i.

Granger non-causality and weak exogeneity are dichotomous concepts. A useful continuous measure of price influence is the forecast error variance decomposition (FEVD). The FEVD gives the proportion of the forecast error variance of price i at some horizon h due to shocks to price j. It provides a continuous measure of which prices "drives" the system at different time horizons. A cost of using the FEVD is that it requires an ordering assumption on the variables in terms of which price is allowed to instantaneously affect the other price. However, both orderings can be applied to check the sensitivity of the ordering assumption.

EMPIRICAL RESULTS

Table 2 shows results for the augmented Dickey-Fuller tests for the null of unit root in the prices. Both domestic and imported prices are integrated of order one, I(1), and contain stochastic trends. Cointegration analysis is then the appropriate methodology.

We start by specifying the unrestricted VAR model. Using an unrestricted VAR model with intercept and linear trend, the Akaike information criterion selects a lag order of 13 months for the VAR model, while the more conservative Hannan-Quinn criterion selects a lag order of 2 months. An alternative specification with monthly seasonal dummies in addition to linear trend suggests a lag order of 2 months for both the Akaike information criterion and Hannan-Quinn criterion.

For robustness we consider all three model specifications in our analysis. In general, the more parsimonious VAR(2) model will provide greater power in testing but might miss longer-run dynamic relationships relevant to understanding Granger causality relationships between the prices. Failure to account for deterministic seasonality might lead to misspecification of the lag order and less reliable dynamic causality relationships.

Likelihood ratio tests for the presence of significant linear trends in the VAR provide no strong evidence for a trend. We choose VAR specifications with only a constant deterministic term (in addition to seasonal dummies for the seasonal specification).

Table 3 reports test results for residual ARCH effects and serial correlation in the estimated VAR(2), VAR(13), and seasonal VAR(2) models. The nonseasonal VAR(2) model shows evidence of heteroskedasticity and unmodeled price dependency at lag orders greater than 10 months. Extending the lag order to 13 reduces the presence of ARCH effects, and to a lesser degree serial correlation effects. The seasonal VAR model eliminates most detectable ARCH effects and serial correlations, suggesting that seasonal dynamics largely account for these characteristics.

We next specify the vector error correction model (VECM). We use an unrestricted constant in the VECM specification. This allows for equilibrium-level differences in the prices while not restricting the linear trend behavior. In either case, cointegration results are highly robust to choice of intercept specification.

Table 4 shows the cointegration rank test results. There is strong evidence for cointegration, even when we allow for very persistent transitory effects in the large 13-month-lag model. The markets are cointegrated; the domestic and import prices share a common stochastic trend.

Table 2. Augmented Dickey-Fuller Unit Root Tests

	Domestic (level)	Import (level)	Domestic (1st diff.)	Import (1st diff.)
ADF statistic	–1.25	–0.41	–9.42**	–6.60**

Note: ADF testing equation includes a constant and a lag order of 5 months. Null hypothesis is unit root. ** $p < 0.01$.

Table 3. VAR Model Fits

Dependent variables: $y_t = [\log(Fisher\_Index\_imp_t), \log(Fisher\_Index\_dom_t)]$			
Deterministic term:	Constant + Trend No Seasonality	Constant + Trend No Seasonality	Constant + Trend Seasonal Dummies
Lag order:	VAR(2)	VAR(13)	VAR(2)
ARCH effects (lag 5)	0.177	0.032	0.221
ARCH effects (lag 10)	0.018	0.033	0.032
ARCH effects (lag 15)	<0.01	0.033	0.270
ARCH effects (lag 20)	<0.01	0.209	0.714
Serial correlation (lag order 5)	<0.01	0.182	0.353
Serial correlation (lag order 10)	<0.01	0.049	0.183
Serial correlation (lag order 15)	<0.01	0.025	0.144
Serial correlation (lag order 20)	<0.01	0.016	0.146
No. of parameters	12	56	34
Log likelihood	1,303.91	1,340.26	1,413.11

Note: The p-values are for the null of no ARCH effects and no serial correlation. ARCH effects are tested using a multivariate ARCH-LM test on the VAR model residuals. Serial correlation is tested using a multivariate Breusch-Godfrey test on the residuals of the VAR model.

Table 5 reports test results for short-run Granger non-causality, long-run weak exogeneity, and forecast error variance decompositions for the three different error correction models. The short-run Granger non-causality test shows somewhat mixed results depending on specification. Overall, we can reject the null that domestic prices are Granger non-causal on the short-run movements of import prices. The import price also shows some evidence of short-run causality on the domestic price.

For the long-run weak exogeneity tests, the import price is long-run weakly exogenous. The short-run Granger causality effect of the domestic price has no persistent effect on the import price. The trend in the domestic shrimp price is driven by the trend in the import price.

The FEVD results largely confirm these findings. The numbers in the table refer to the share of forecast errors in each price series due to own shocks. We look at short-run movements (1 year) and long-run movements (10 years). The range of the numbers for each cell refers to the range implied by the two possible specifications of instantaneous price causality. In the short run, less than a year, most of the domestic and import price forecast errors are due to own price shocks. In

Table 4. Johansen Trace Test Statistics for Cointegration Rank

Dependent variables: $y_t = [\log(Fisher\_Index\_imp_t), \log(Fisher\_Index\_dom_t)]$			
Deterministic term:	Constant + Trend No Seasonality	Constant + Trend No Seasonality	Constant + Trend Seasonal Dummies
Lag order:	VAR(2)	VAR(13)	VAR(2)
Cointegration rank			
H_0: Rank = 0	44.72**	30.1**	32.39**
H_0: Rank ≤ 1	0.07	0.56	0.03

Note: The cointegrated VAR model assumes an unrestricted constant in the VECM representation. ** $p < 0.01$.

Table 5. VECM Model Test Results for Dynamic Causalities in Real Fisher Price Indices

Dependent variables: $\Delta \boldsymbol{y}_t = [\Delta\log(\mathit{Fisher\_Index\_imp}_t), \Delta\log(\mathit{Fisher\_Index\_dom}_t)]$			
Deterministic term:	Constant, No Seasonality	Constant, No Seasonality	Constant, Seasonal Dummies
Lag order:	VECM(1)	VECM(12)	VECM(1)
Short-run Granger causality (*F*-statistics)			
H_0: Imp. do not Granger cause dom.	2.65	3.71***	6.33*
H_0: Dom. prices do not Granger cause imp.	16.17***	2.75**	11.34***
Long-run weak exogeneity (est. (SE))			
Dom. price-adjustment coeff., α_{DOM}	0.192***	0.223***	0.128***
	(0.032)	(0.042)	(0.026)
Imp. price-adjustment coeff., α_{IMP}	–0.029*	0.004	–0.025*
	(0.012)	(0.017)	(0.012)
FEVD (share due to own price shock) (%)			
Dom. price (short run, 12 months ahead)	68.9–77.5	57.8–67.8	69.5–80.0
Dom. price (long run, 10 years ahead)	21.4–31.5	11.5–13.9	23.1–35.4
Imp. price (short run, 12 months ahead)	82.4–91.3	86.0–93.9	82.1–92.0
Imp. price (long run, 10 years ahead)	78.1–88.3	97.6–99.1	74.6–86.8

Note: VECM model assumes an unconstrained intercept. For the FEVD, the range of percentages refers to the range implied by two specifications of the instantaneous price relationships: (1) when the import price structural shock has no instantaneous impact on the domestic price, and (2) when the domestic price structural shock has no instantaneous impact on the import price. *** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

the long run, the majority of the variation in the system is determined by shocks to the import price, consistent with the short-run Granger non-causality and long-run weak exogeneity results.

Table 6 replicates the table 5 analysis using equal-weighted average price measures instead of the Fisher indices. The results are essentially the same; the index constructions do not drive the results.

Table 7 reports *p*-values for the law of one price restriction on the long-run equilibrium relationship. The estimated β parameter is close to but slightly less than unity. As is clear from figure 1, the domestic price is more volatile than the import price in the short run, fluctuating with shorter periods below and above the import price. While such fluctuations are consistent with domestic landings impacting the domestic price in the very short run (as suggested, for instance, in Smith et al. [2017], in which size-based prices respond to ecological shocks), it is more plausible that the higher domestic price volatility is due to the change in size composition of the US landings over the year. Changing size composition is part of the natural seasonality of the capture fishery (Smith et al. 2014). This will bias the β parameter to be less than unity. However, we fail to reject a β parameter that is different from unity and statistically significant, implying that the law of one price holds.

DISCUSSION AND CONCLUSIONS

In the movie *Forrest Gump*, the main character gets rich because his vessel was the only one that survived a big storm, and he became the sole supplier of shrimp with associated high prices. That story may have been possible in the 1970s, but our results indicate that Forrest Gump would not have gotten rich today. Rather, the reduction in the US supply would be offset by an increase in imports with limited or no pressure on prices.

Our results are clear in that not only is the US shrimp market a fully integrated part of the global market, the price of shrimp in the US is determined in this global market. With the increasing

Table 6. VECM Model Test Results for Dynamic Causalities in Real Equal-Weighted Unit Values of Import and Domestic Unit Values Across Size Categories

Dependent variables: $\Delta \boldsymbol{y}_t = [\Delta\log(Unit\_Value\_imp_t), \Delta\log(Unit\_Value\_dom_t)]$			
Deterministic term:	Constant, No Seasonality	Constant, No Seasonality	Constant, Seasonal Dummies
Lag order:	VECM(1)	VECM(12)	VECM(1)
Short-run Granger causality (*F*-statistics)			
H_0: Imp. do not Granger cause dom.	4.05*	2.61**	5.31*
H_0: Dom. prices do not Granger cause imp.	12.46***	1.33	12.36***
Long-run weak exogeneity (est. (SE))			
Dom. price-adjustment coeff., α_{DOM}	0.145***	0.181***	0.120***
	(0.0264)	(0.0360)	(0.0262)
Imp. price-adjustment coeff., α_{IMP}	–0.021	0.0089	–0.022
	(0.0132)	(0.0187)	(0.0134)
FEVD (share due to own price shock) (%)			
Dom. price (short run, 1 year ahead)	65.0–78.1	59.3–72.5	66.9–82.0
Dom. price (long run, 10 years ahead)	16.2–28.9	10.8–12.3	18.9–34.7
Imp. price (short run, 1 year ahead)	82.7–94.2	83.8–94.3	80.1–93.5
Imp. price (long run, 10 years ahead)	80.6–93.1	97.6–98.5	75.3–90.7

Note: VECM model assumes an unconstrained intercept. For the FEVD, the range of percentages refers to the range implied by two specifications of the instantaneous price relationships: (1) when the import price structural shock has no instantaneous impact on the domestic price, and (2) when the domestic price structural shock has no instantaneous impact on the import price. *** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

dominance of farmed shrimp in the global market, the production cost of farmed shrimp most plausibly drives the domestic price of wild-caught shrimp in the US. Domestic supply shocks are then having no impact on price beyond short-run transitory effects, as the imported quantity is adjusted in response to the supply shock, and whatever changes the global shrimp price also changes the US shrimp price. The implication is that US shrimper revenues increase when the global price increases and vice versa, but revenues decrease when there are negative shocks to domestic landings because there is no strong price compensation to shield fishers from the quantity losses. Furthermore, trade policies in other markets, such as higher tariffs for imported shrimp in major markets such as Europe, can reduce US import prices and the price of domestically produced shrimp (Debaere 2010).

Although any analysis of a single market can raise questions of external validity, our results are similar to what was reported for salmon by Asche et al. (1999). Namely, the domestic US salmon price is determined in the global market, and this price is determined by production costs in salmon farming, although here also diseases and other environmental shocks may cause short-run shocks (Asche et al. 2017; Iversen et al. 2020; Oglend and Soini 2020; Oglend et al. 2024a, 2024b). The fact that most seafood consumed in the US is imported (NMFS 2024) and that

Table 7. Tests for the Law of One Price

Cointegration relationship: $\log(Fisher\_Index\_imp_t) = \beta \log(Fisher\_Index\_dom_t)$			
Deterministic term:	Constant, No Seasonality	Constant, No Seasonality	Constant, Seasonal Dummies
Lag order:	VECM(1)	VECM(12)	VECM(1)
β (p-value H_0: $\beta = 1$)	0.945 (0.32)	1.001 (0.97)	0.942 (0.35)

the market share of imports is substantial for most major species (Love et al. 2023) suggest that a similar structure may be present for many, if not most, other major species produced in the US independently of whether they are wild-caught or farmed. That antidumping tariffs against several named countries for species like catfish, salmon, and shrimp had little impact on the domestic US prices, but rather just led to a reallocation of trade patterns (Engle et al. 2022; Anderson et al. 2019; Keithly and Poudel 2008), provides corroborating evidence. That prices in other countries can be used as a control when investigating the impact of changes in domestic fisheries management (Birkenbach et al. 2023) is another indication. However, there exist few formal tests for other species and seafood markets such that more empirical testing in future research would improve our understanding of the price determination process facing US seafood producers.

There are some candidates for species where the impact of the international market may be more limited. New England lobster, which is now the most valuable US fishery, is one candidate, as the US is by far the largest-producing nation globally (Gordon 2021). Lobster exemplifies a species that has not experienced the same declining trend in the price as, for example, salmon and shrimp. Stone crab is another example as a unique Florida fishery with virtually no import competition (Kehoe et al. 2023). However, these species are likely to be exceptions. For a number of other small regional species, import competition is strong at least when measured by market share (Love et al. 2023).[4]

Fear of increased competition from a domestic aquaculture industry and associated downward price pressure is claimed to be a main reason domestic fishers in the US are generally opposed to the development of a US aquaculture industry (Knapp and Rubino 2016; Garlock, Nguyen, et al. 2020; Rubino 2023). If the domestic seafood prices are determined at a global market, this opposition appears misguided. As with shrimp, with this market structure the only effect of increased domestic production would be lower imports. There is virtually no price effect because price effects will only be due to the impact of an increased US supply on global production. Moreover, a local aquaculture production may enhance local supply chains (Garlock, Nguyen, et al. 2020), an effect that is lost with continued opposition to the industry.

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4. For export-oriented species like spiny lobster, similar relationships exist but with export competition (Garlock, Asche, et al. 2025).

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