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Tangled up in prices going up and down: price determination in selected agri-food markets

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Summary

The dissertation includes four papers altogether. Three papers examine mechanisms for price transmission between markets. Among them, one deals with price transmission along the supply chain, and the other two deal with price transmission across spatially distinct markets. The fourth paper measures the effects of demand shifts on the market price equilibrium.

Paper 1 investigates price transmission relationships in the salmon value chain between the export market of Norway and retail markets of France and Spain. We assume processing costs have a significant impact on a vertical price linkage along the supply chain. Specifically, asymmetric price transmission is investigated along the supply chain of two salmon products, a relatively unprocessed product (fresh salmon) and a more processed product (smoked salmon). Results indicate a price transmission relationship in both markets along the fresh salmon supply chain but not along the smoked salmon chain in either market. Furthermore, for the fresh salmon supply chain, asymmetric adjustment is observed in both markets. Processing salmon into value-added consumer products involves additional inputs, including labor, capital, marketing, and packaging costs and time (Landazuri-Tveteraas et al., 2018). The greater the share of these non-raw materials costs in the final consumer prices, the less the price influence of farmed fish on the final product price. Hence, we can argue that price transmission between the export and retail salmon markets decreases as more processing is involved.

The second paper explores market integration and spatial price transmission in the regional grain markets in Ethiopia. Since 2006, persistent increases in food-crop prices have become a critical challenge in Ethiopia. This study investigates price transmission between major regional grain markets to see if our study can add more evidence to the compelling argument that grain traders' price adjustments have contributed to the price increase in Ethiopia. Results indicate that central-market and local-commodity-market prices are cointegrated. In addition, the results do not suggest sufficient empirical evidence in support of asymmetric pricing behavior in grain trade between the central and local markets. This implies that the study does not support the argument that Ethiopian traders' marketing behavior has contributed to grain-price increases.

The literature indicates an increasing level of concentration in recent years in the European Union (EU) pork markets both in producing countries as well as at the retail level.

The third paper thus investigates spatial price transmission among the EU pork markets. The study uses the recent price data of the major pork markets in the EU (Germany, France, Spain, Denmark, the Netherland, and Poland). The results indicate the presence of long-run relationships between the price pairs. Furthermore, the study finds empirical evidence in support of asymmetric pricing behavior in the EU pork markets. This might imply that the growing level of concentration in the EU pork markets has an impact on the price transmission process among spatially separate EU pork markets.

The fourth paper takes a different approach by studying market price formation. Growing global demand for seafood has shifted the demand curves in the world seafood markets upward. The paper first estimates seafood demand growth in 107 countries from 1984 to 2013. The estimated results are then used to calculate aggregate demand growth by income level, regionally, and over time. Results indicate that demand growth varies considerably across countries, regions, income groups, and over time. Furthermore, while the literature show seafood production has more than doubled since the mid-1980s, our results indicate that global demand for seafood has exceeded the global seafood supply, which explains the increasing global seafood prices.

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List of papers

Paper 1

Kidane, D.G., Myrland, Ø, & Xie, J. (2021). Asymmetric price transmission in a changing food supply chain, *Aquaculture Economics & Management*, 25 (1), 89–105

Paper 2

Kidane, D.G. (2021). Market integration and price transmission in the regional grain markets in Ethiopia. Submitted to *Journal of Applied Economics*

Paper 3

Kidane, D.G., Myrland, Ø, & Xie, J. (2021). Spatial price transmission in EU pork markets: using threshold autoregressive and non-parametric local polynomial techniques. Submitted to *Journal of Commodity Markets*

Paper 4

Kidane, D.G. & Brækkan, E.H. (2021). Global seafood demand growth differences across regions, income levels, and time, *Marine Resource Economics*, 36 (3), 289–305.

1. Introduction

Price theory is a fundamental part of economic theory since price drives resource allocation (Stigler, 1969) and determines output-mix decisions by economic agents (Gardner, 1975). Price signal transmission integrates markets vertically along the supply chain and horizontally across spaces (von Cramon-Taubadel & Meyer, 2004). Thus, economists who study market competition, performance, efficiency, and integration, are interested in the price transmission processes.

We have witnessed dramatic price changes for many commodities in the world in the last two decades. Agricultural commodities have displayed extreme price fluctuation, reaching exceptional levels in 2007 and 2008, falling sharply in 2009 and then increasing again in 2011 (Tadesse et al., 2014; Baquedano & Liefert, 2014; Ceballos et al., 2017; Brümmer et al., 2013; Minot, 2011). Some agricultural commodities such as wheat and maize have also shown an incredibly high price growth in the last two years (FAO, 2020). To understand the commodity price changes, one must understand the determinants of demand and supply shifts (Brækkan et al., 2018; Asche et al., 2011) in the formation of a new market equilibrium price. In any commodity market, there can be multiple simultaneous shifts in demand and supply in distinct geographical locations (Brækkan & Thyholdt, 2014). For a globally traded commodity, each of these shifts can affect the market equilibrium price in regional markets.

Primary agricultural commodities undergo, to varying degrees, a substantial series of intermediary alterations before they are sold as final food products to consumers (Kilmer, 1986; Sexton & Lavoie, 2001) by intermediaries such as the food processing industry and the distribution sectors (wholesalers and retailers). The level of efficiency at each marketing stage and the nature and extent of market shocks transmitted through the various stages of the supply chain or horizontally related markets can significantly influence commodity price behavior. Thus, it is important to study the impact of the performance of agri-food markets on commodity prices via the transmission of shocks between markets (Vavra & Goodwin 2005; von Cramon-Taubadel & Meyer, 2004; Abdulai, 2000; Fackler & Goodwin, 2001).

The focus of this dissertation is thus on price dynamics in the agri-food markets. Specifically, using price data from selected agri-food markets and countries, three papers investigate the transmission of price signals along the supply chain and spatially related markets, and the fourth paper focuses on approximating shifts in global seafood demand.

The remainder of this dissertation is organized as follows. Section 2 provides a discussion of market price determination. This is followed by a discussion on price signal transmission between markets in Section 3 and a summary of the research papers included in the dissertation in Section 4. In the final section, I offer a general discussion and offer implications of the price transmission studies.

2. Market price determination

In this chapter, I start with a brief discussion on equilibrium price determination. This is followed by a discussion on determinants of demand and supply shifters following price changes in both vertically and horizontally interrelated markets.

2.1. Market equilibrium

The concepts of demand and supply and their relationship extend back a fairly long time in the economics literature. Marshall (1920) establishes the general law of demand, stating that quantity demanded increases with a fall in price and diminishes with a price rise. Furthermore, Marshall (1920) illustrates the importance of demand and supply in price determination by likening these to the blades on a pair of scissors. The significance of this relationship was also emphasized by Henderson (1922), who states that “price tends to the level at which demand equals supply.” The fundamental principle of supply is that a fall in the price causes a reduction in the quantity supplied (Smith, 1863). At the same time, the price of a product determines how much a consumer is willing to buy. Demand and supply thus became an integral part of the price formulation theory in the late nineteenth century (Stigler, 1966). Graphically, this is shown through the downward sloping demand curve and the upward-sloping supply curve, as illustrated in Figure 1 below.

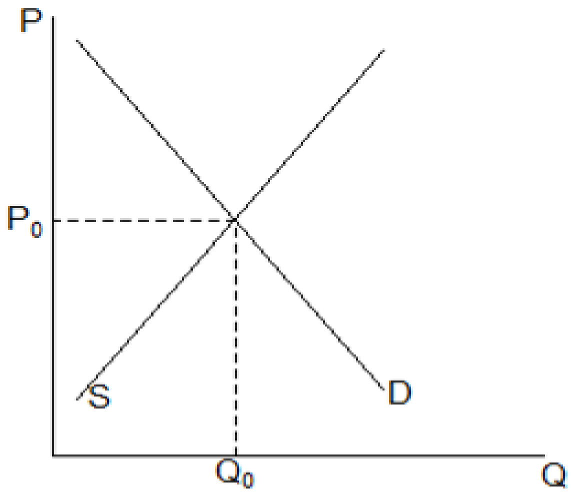

Figure 1. Illustration of market price equilibrium

Using simple mathematical equations, we can express Figure 1 as follows:

$$Q_d = D(P) \qquad \text{Demand} \qquad (1)$$

$$Q_s = S(P) \qquad \text{Supply} \qquad (2)$$

$$Q_d \equiv Q_s = Q \qquad \text{Equilibrium condition} \qquad (3)$$

where $D(.)$ and $S(.)$ are the demand and supply functions, respectively. Under the equilibrium conditions given by Equation (3), the market has the equilibrium price (P_0) and quantity (Q_0).

2.2. Determinants of price changes

2.2.1. Demand shifters

If the demand and supply condition changes, the market will not be in a state of equilibrium until these forces restore it and a new equilibrium occurs. Demand shifters are the factors that cause the demand curve to shift leftward or rightward.

Changes in income and prices of substitute or complementary products are the main economic factors responsible for demand shifts (Brækkan et al., 2018). Demand can also be affected by changes in non-economic factors such as social demographics, product attributes, and consumer preference. For example, older, more educated customers (Tomek, 1985), population growth (Nguyen & Kinnucan, 2018), and socioeconomic factors such as income distribution (Brown & Deaton, 1972) may also cause shifts in the demand function. Examples of consumer-preference-related changes include the appearance of new product information (Tomek, 1985) and changes in product attributes (such as product form and quality) (Ladd & Suvannunt, 1976).

Economic and non-economic factors can sometimes have a compounding effect. There is documented evidence that income and population growth in emerging markets, particularly China and India, are the key factors behind the post-2007 food-price increases (Krugman, 2008; Wolf, 2008; Bourne, 2009; Nguyen & Kinnucan, 2018).

Figure 2 illustrates how a demand shifter (e.g., income) affects equilibrium price. If the economy expands in such a way that many people's income is increased, more people are able to afford particular commodities, and this leads to an increase in their price and quantity demanded. Due to the increases in price and quantity demanded, the original demand curve D_0 shifts to the right to the new demand curve D_1, leading to a price increase from P_0 to P_1. Similarly, when the economy slows, the original demand curve could shift inwards to the new demand curve D_2, leading to a price decrease from P_0 to P_2. It is important to note here that a shift in demand captures a pattern for the market.

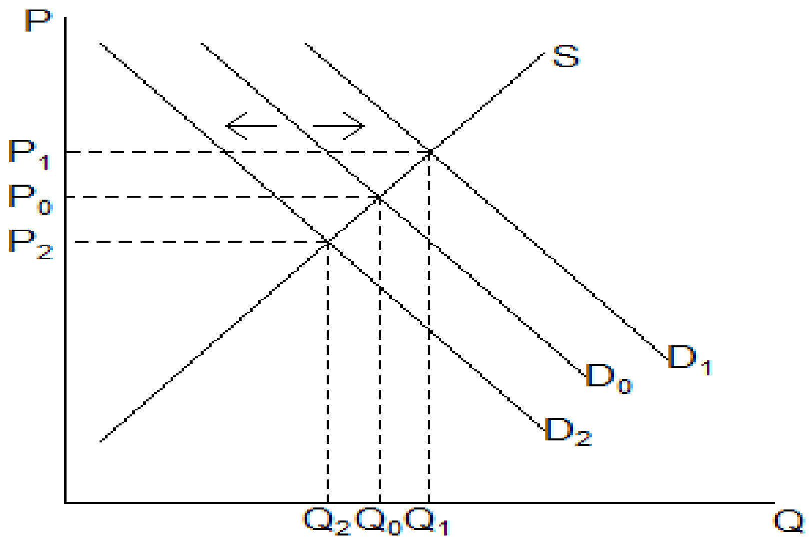

Figure 2. Illustration of the effects of demand shifts on market price equilibrium

2.2.2. Supply shifters

As discussed, the market equilibrium price is jointly decided by demand and supply. Changes in costs of production and improvements in technology are the two factors that most commonly cause the supply curve to move inward or outward (Muth, 1964; Tomek, 1985). Supply can also be affected by changes in the number of suppliers in the market and weather conditions (Fisher et al., 2012).

Figure 3 illustrates how a supply shifter (change in the price of factors of production) affects the market equilibrium price. When the price of factors of production falls, the cost of production of a commodity also falls, increasing the supply. An increase in supply will cause the original supply curve S_0 to shift outward to S_2, leading to a price decrease from P_0 to P_2. Similarly, a rise in the price of factors of production increases the cost of production and reduces the supply of a commodity. A fall in supply will cause the original supply curve S_0 to shift inward to S_1, leading to a price increase from P_0 to P_1. Improvements in technology can also cause an outward shift of the supply curve because more output can be produced with the same level of inputs (Tomek, 1985).

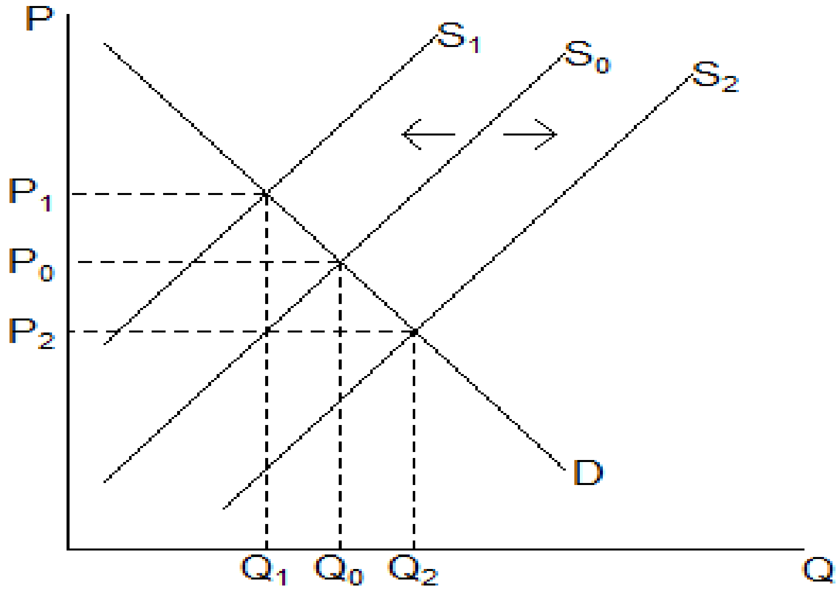

Figure 3. Illustration of the effects of supply shifts on market price equilibrium

2.3. Price determination in the case of interrelated markets

Agricultural product markets can be interrelated vertically along supply chains or horizontally across spaces (Barrett, 1996). In vertically interrelated markets (for example, between farm and retail markets), demand and supply shocks in the farm market may affect prices and quantities in the retail market or vice versa. Similarly, in spatially interrelated markets, shifts in demand and supply in one regional market may impact prices and quantities in another.

2.3.1. Vertically interrelated markets: Farm-Retail prices

Price determination in vertically related markets has been studied by agricultural economists for many decades. Gardner (1975) develops the modern theory of farm and retail price linkages in the food industry, examining the implications of simultaneous equilibrium in three related markets: farm output, marketing services, and retail food. They consider a competitive food marketing industry that uses two input factors of production – a farm product and marketing inputs – to produce a processed product sold in retail-level markets.

The basic structure of Gardner's (1975) model indicates an equilibrium relationship between the farm and retail markets. Any external shocks in the retail or farm market should call forth a response in both markets, which restores the new equilibrium among the farm and retail markets. Figure 4 illustrates the effects of a retail-market demand shock on the equilibrium prices in each market. A positive demand shock in the retail market shifts the demand curve in the market upward, leading to a price increase from P_r^0 to P_r^1. This will further increase the derived demand for production inputs in the farm market. As illustrated in Panel B of Figure 4, the demand curve in the farm market moves upward. Consequently, market equilibrium price in the farm market increases from P_f^0 to P_f^1. Any farm-level shock can also cause the formation of new equilibriums in both markets through a mechanism similar to that illustrated in Figure 4.

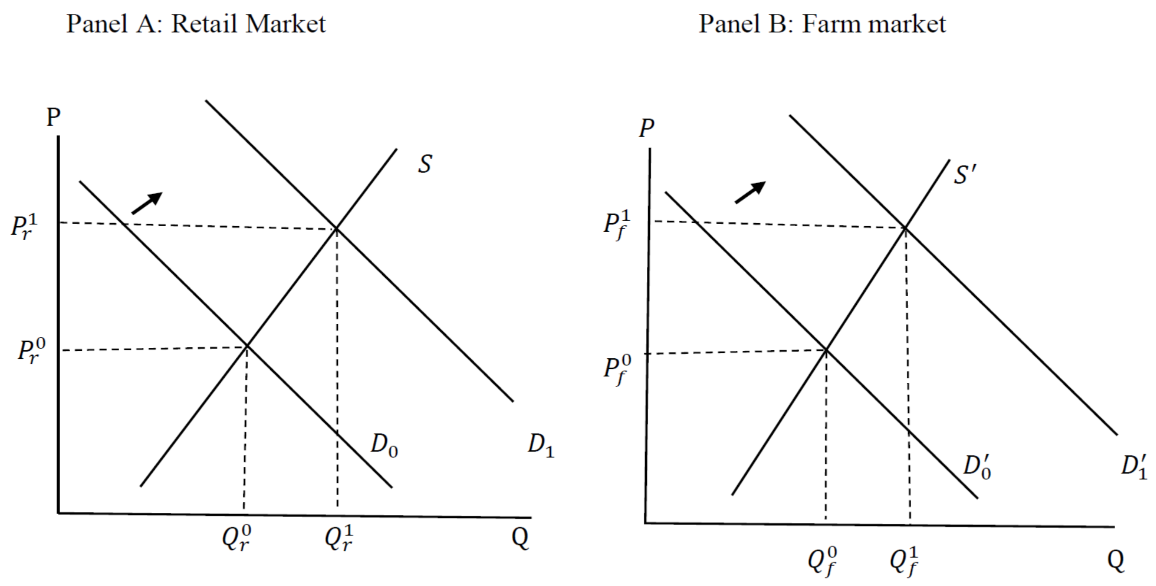

Figure 4. Illustration of the effects of a retail-market demand shock on equilibrium prices in both retail and farm markets

Gardner's (1975) original model is reformulated by Kinnucan and Zhang (2015) as follows:[1]

$$P_r^* = \frac{1}{\eta} Q_r^* + \alpha \qquad \text{(Retail demand)} \qquad (4)$$

$$P_r^* = S_f P_f^* + S_m P_m^* \qquad \text{(Retail supply)} \qquad (5)$$

$$Q_f^* = -S_m \, \sigma \, P_f^* + S_m \, \sigma \, P_m^* + Q_r^* \text{ (Demand for farm-based input)} \quad (6)$$

$$Q_m^* = S_f \, \sigma \, P_f^* - S_f \, \sigma \, P_m^* + Q_r^* \text{ (Demand for marketing input)} \quad (7)$$

$$Q_f^* = \varepsilon_f \left(P_f^* + \beta_f \right) \qquad \text{(Supply of farmed-based input)} \qquad (8)$$

$$Q_m^* = \varepsilon_m (P_m^* + \beta_m) \qquad \text{(Supply of marketing input)} \qquad (9)$$

The variables are expressed as proportionate changes (e.g., $P_r^* = \frac{dP_r^*}{P_r^*}$ represents the proportionate change in the retail price), and their coefficients represent elasticities or cost shares. Specifically, $\eta(< 0)$ represents the own-price elasticity of demand for the retail product Q_r; σ (≥ 0) is the elasticity of substitution between the farm-based input Q_f and the bundle of marketing inputs Q_m; $S_f = \frac{P_f \, Q_f}{P_r \, Q_r}$ and $S_m = \frac{P_m \, Q_m}{P_r \, Q_r}$ are the cost shares, which sum to one. In the cost shares, P_f is the price of the farm-based input, and P_m is the price of the bundle of marketing inputs; $\varepsilon_f(> 0)$ is the own-price elasticity of supply for the farm-based input; and $\varepsilon_m(> 0)$ is the own-price elasticity of supply for the marketing inputs. The remaining terms are the vertical shift parameters. Specifically, α indicates a proportionate shift in the price direction in the retail demand curve from the initial equilibrium point due to an exogenous retail-demand shifter, and β_f and β_m indicate proportionate shifts in the price direction in the input supply curves from their initial equilibrium points due to exogenous input supply shifters.

Equations (4) – (9) can be solved simultaneously to determine the expression of the required variables. For instance, Kinnucan and Zhang (2015) express the reduced-form equations for the retail price (P_r^*) and farm price (P_f^*) as follows:

[1] I use different notations than Kinnucan and Zhang (2015) to match the notations in the dissertation.

$$P_r^* = -\left(\frac{\eta(\sigma+S_f\varepsilon_m+S_m\varepsilon_f)}{D}\right)\alpha - \left(\frac{\varepsilon_f S_f(\sigma+\varepsilon_m)}{D}\right)\beta_f - \left(\frac{\varepsilon_m S_m(\sigma+\varepsilon_f)}{D}\right)\beta_m \tag{10}$$

$$P_f^* = -\left(\frac{\eta(\sigma+\varepsilon_m)}{D}\right)\alpha - \left(\frac{\varepsilon_f(S_f\sigma+\varepsilon_m-S_m\eta)}{D}\right)\beta_f - \left(\frac{S_m\varepsilon_m(\sigma+\eta)}{D}\right)\beta_m \tag{11}$$

where $D = (\varepsilon_f\varepsilon_m + \sigma(S_f\varepsilon_f + S_m\varepsilon_m - \eta) - \eta(S_f\varepsilon_m + S_m\varepsilon_f)) > 0.$

As Kinnucan and Zhang (2015) discuss, the reduced-form elasticities in Equations (10) and (11) imply that in vertically integrated retail and farm markets, price changes in one market can transmit to the other market. More specifically, under the stated parametric assumptions (i.e., retail demand is downward-sloping, input supplies are upward sloping, and $\sigma \geq 0$), an isolated increase in retail demand ($\alpha > 0$) causes both retail and farm prices to increase, as illustrated in Figure 4. At the same time, an isolated increase in the supply of the farm-based input ($\beta_f > 0$) causes the retail and farm prices to decrease.

2.3.2. Spatially interrelated markets

Price determination in spatially interrelated markets dates back more than a century (Fackler & Goodwin, 2001). Figure 5 represents an export market and an import market that produce a homogeneous commodity x. When there is no trade between these two markets, P^e and P^I are the equilibrium prices in each market. In this setup, the export market is assumed to be the low-cost provider of commodity x and has the cost advantage.

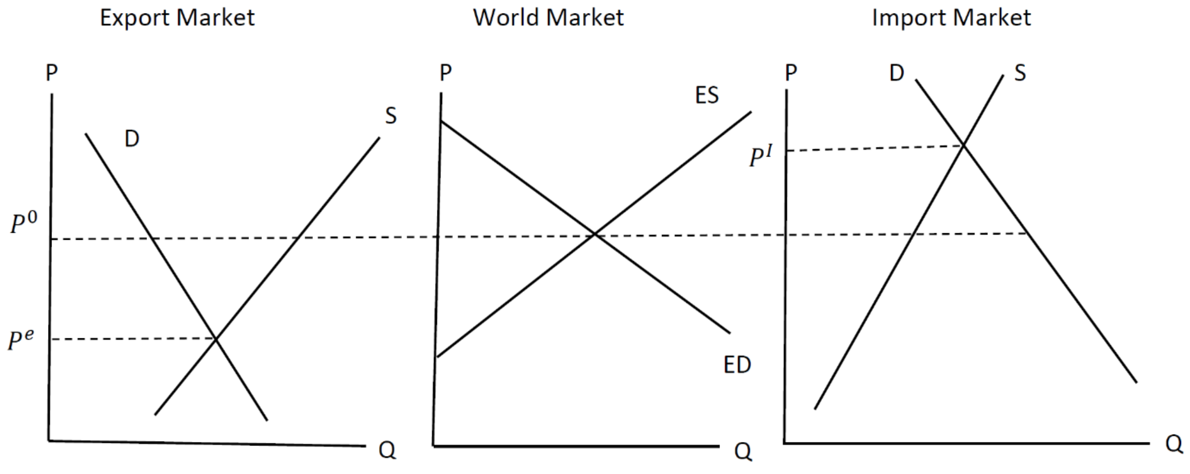

Figure 5. Illustration of price determination in spatially interrelated markets

Where an open market is assumed, and trade is free between the two markets, there are opportunities for profit-seeking firms (also called arbitrageurs): buy commodity x cheaply from the export market and sell it at a higher price in the import market. This means that the extra production of commodity x in the export market is shipped to the import market to supply the extra consumption that cannot be met by domestic production. This will reduce the supply in the exporter's domestic market and lead to a price rise in that market. By contrast, import of commodity x will increase the supply in the import market, lowering its price in that market. When the price is the same in both countries, there is no incentive to trade further; hence, in this case, we say the markets are at their equilibrium. In Figure 5, the equilibrium price is represented by P^0. In the literature, such an equilibrium is often referred to as the Law of One Price (LOP), implying that, in a competitive market structure, abstracting transaction costs, homogenous products sold in several markets should be sold at the same prices due to arbitrage (Fackler & Goodwin, 2001).

Now consider the two markets being hit by a supply shock in the export market (for instance, due to change in technology). As illustrated in Figure 6, this leads the supply curve in the export market to shift from S_1 to S_1', leading the initial equilibrium in the export market at a price P^0 and quantity Q_1^0 to move to the new equilibrium price P' and quantity Q_1'. The supply shock in the export market can affect the demand in the import market. However, this effect depends on the degree of integration between the two markets.[2] If the two markets are perfectly integrated, the demand curve in the import market shifts from D_2 to D_2' equilibrating the two markets (or ensuring the LOP holds). However, if they are only partially integrated, the demand curve will shift from D_2 to D_2'', which is insufficient to equate the price in the two markets. If the two markets are unrelated, then no price change in either of the markets will result in a response on the other.

In a spatially related market system, as Asche et al. (2012) discuss, there is sometimes a leading market that has a dominant influence on the other markets. The price changes in the leading market can affect the market prices and quantities in the other markets; the reverse is not the case.

[2] The definition of market integration and its related concepts will be discussed in the next section. However, at least at this point, market integration implies the extent to which the export and import markets are related.

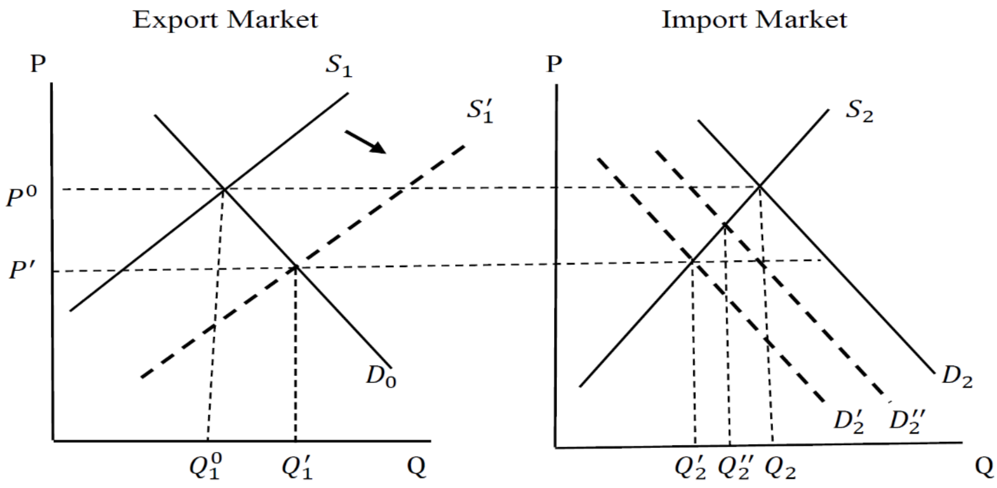

Figure 6. Illustration of price shifts between markets. *Source*: Asche et al. (2004).

The relationship between the two spatial markets described above can be analyzed mathematically (ignoring exogenous demand and supply shifters for simplification) as follows:

$$Q_1 = D_1(P_1) \text{ (Demand for home consumption)} \quad (12)$$

$$Q_1 = S_1(P_1) \text{ (Supply of home consumption)} \quad (13)$$

$$Q_2 = D_2(P_2) \text{ (Demand for export)} \quad (14)$$

$$Q_2 = S_2(P_1, P_2) \text{ (Supply for export)} \quad (15)$$

$$P_1 = P_2.R \text{ (Foreign currency price relations)} \quad (16)$$

where $D(.)$ and $S(.)$ are demand and supply functions, respectively. In the given equations, it is assumed that the export market produces for home consumption market Q_1 and sells at price P_1, and produces for export market Q_2 and sells at price P_2. It is also assumed that the two markets have currencies that differ from each other, and their relationship is given by Equation (16). In Equation (16), R denotes the bilateral exchange rate.

Equations (12) – (16) can be used to explore the most basic form of trade dependency between the two markets. In general, the equations given above show that, in spatially interrelated markets, price shocks in one market can transmit to the other market and affect commodity prices.

3. Transmission of price signals between markets

In the previous section, we have briefly seen the modes of market price determination. In this chapter, I discuss the basic theoretical concepts related to price transmission between markets and empirical models used by agricultural economists to study the mechanisms of price transmission.

3.1. Concepts of market integration and price transmission

There is no one definition of market integration in the literature; different studies use the term in different contexts based on their area of focus and method of analysis. The earlier literature defines integrated markets as locations connected by trade and displaying high price correlations (Harris, 1979; Ravallion, 1986). The contemporary definition of market integration takes in a broader context and concerns the flow of tradable commodities, information, and prices across space, form, and time between markets linked directly or indirectly by trade (Barrett, 1996; Fackler & Goodwin, 2001). It is also closely related to the concept of price efficiency and the LOP.

Fackler and Goodwin (2001) define market integration as a measure of the degree to which demand and supply shocks of a commodity in a given market are transmitted to another market. Specifically, they define market integration by the “price ratio” (denoted R_{AB}) associated with a market shock. Assuming a hypothetical shock in Market A (denoted ε_A), R_{AB} is mathematically expressed as follows:

$$R_{AB} = \frac{\partial P_B / \partial \varepsilon_A}{\partial P_A / \partial \varepsilon_A} \tag{17}$$

where P_A and P_B denote prices in Markets A and B, respectively, and ∂ denotes the first order derivative of the respective price to the market shock.

Fackler and Goodwin’s (2001) definition of market integration indicates that market integration is a degree varying between zero and one.[3] One implies perfectly integrated markets (or LOP), while zero implies completely segmented markets. A value between zero and one suggests partially integrated markets. In this regard, market integration is

[3] The unit of measurement can be obtained by taking the expected value of the price ratio R_{AB}.

synonymous with price transmission. In perfectly integrated markets, price change in one market is completely and immediately transmitted to another market. Whereas in the presence of market failure, price transmission will reflect the inefficiency and welfare losses in the economic system by incompletely and slowly transmitting between markets.

3.2. Types of price transmission

As discussed in Section 2, the direction of the price transmission between markets can be categorized as vertical or horizontal (spatial). Vertical price transmission refers to price linkages between different marketing stages (for instance, between farm and retail markets) along the value chain of a commodity, whereas spatial price transmission refers to price linkages across space for the same commodity. The interdependence between the Norwegian salmon export price with different salmon-product prices in the retail markets of France and Spain, as in Paper 1, can be taken as an example of vertical price transmission. An example of spatial price transmission is the interdependence between prices in the major regional grain markets in Ethiopia, as in Paper 2, or the interdependence between the European pork market prices, as in Paper 3.

3.3. Symmetric- asymmetric price transmission

Price transmission can be symmetric or asymmetric in terms of differences in transmission speed and/or magnitude between markets. If a price increase and decrease are completely and rapidly transformed from one market to another, the transmission is symmetric; otherwise, it is asymmetric (Vavra & Goodwin, 2005; von Cramon-Taubadel & Meyer, 2004; Abdulia, 2000). For instance, in the case of vertically related farm and retail markets, a price shock in the farm market should trigger appropriate and quick changes in the retail market and vice versa. Similarly, between spatially related markets, price shocks in the central market would transmit instantly and change correctly in local markets (Abdulia, 2000; Serra et al., 2006a). In the literature, symmetric price adjustment is often taken as an indication of efficient price signal transmission between markets (e.g., von Cramon-Taubadel & Meyer, 2004).

However, in practice, the adjustment to a market shock may not be symmetric but asymmetric (von Cramon-Taubadel & Meyer, 2004; Abdulia, 2000, Kinnucan & Forker, 1987). Asymmetric price transmission occurs when the price response in a market varies based on price increases and decreases in another market. For instance, retailers may respond more completely and rapidly to price increases than to price decreases in the farm market (Kinnucan & Forker, 1987). An example of asymmetric price transmission between markets horizontally separated in space would be when the increase in central-market prices more completely and rapidly transmit to local markets than equivalent price decreases (Abdulia, 2000). Hence, asymmetric price transmission can be classified as either vertical or horizontal.

Asymmetric price transmission can be further classified into positive and negative asymmetries depending on how prices in one market react to the negative/positive price changes in another market (Vavra & Goodwin, 2005). Positive asymmetric price transmission is when price increases in a market more fully and rapidly transmit to other markets than is the case for corresponding price reductions. If the converse holds, it is called negative asymmetric price transmission. The presence of asymmetric price transmission is often considered to be evidence of market failure or the abuse of market power (von Cramon-Taubadel & Meyer, 2004).

The asymmetric price transmission described above has been found to exist in many input and output markets, including in the agricultural, finance, and energy sectors. Due to the importance of this phenomenon, researchers have investigated what causes asymmetric price transmission. These reasons are discussed next.

3.4. Causes of asymmetric price transmission

Several factors have been identified in the literature as causes of asymmetric price transmission. Among them, non-competitive markets (market power) and adjustment (menu) costs are identified as the most important whether the transmission is along the supply chain or between spatially separated markets (von Cramon-Taubadel & Meyer, 2004; Abdulia, 2000; Bailey & Brorsen, 1989; Kinnucan & Forker, 1987). Some other causes include government intervention, search costs, and inventory management. To ease the discussion and for the sake of brevity, the following exploration of each factor is given for the vertical context only. Nevertheless, the same explanation is also applicable in the context of spatial transmission (von Cramon-Taubadel & Meyer, 2004).

Economic theory suggests that, under competitive market conditions, output price responses to input price increases and decreases are expected to be the same (Vavra & Goodwin, 2005). Both upstream price increases and decreases should transmit to the downstream market completely and rapidly along the supply chain. However, in the case of agricultural products, the upstream market is generally less concentrated than the downstream market (Vavra & Goodwin, 2005). This implies that the downstream market is more likely to exercise market power in the upstream market. According to von Cramon-Taubadel and

Meyer (2004), market power can lead to positive asymmetries when input price increases pass more completely and rapidly than input price decreases. Market power may also lead to negative asymmetry if a monopoly responds less rapidly to input price changes that squeeze their margin for the risk of having spoiled goods (Ward, 1982) or fear of losing customers (Heien, 1980). Positive and negative asymmetries may also result if firms face a kinked demand curve depending on their price expectation as input and output price changes (Bailey & Brorsen, 1989). If an individual oligopoly firm anticipated that other firms would match an output price increase caused by an input price increase, but not the reverse, a kinked demand curve could result in a positive asymmetry. In contrast, if a firm assumes that other firms are unlikely to match an output price increase than decrease, negative asymmetry will lead to a concave kinked demand curve.

Adjustment (menu) costs are the second notable cause of asymmetric price transmission (Bailey & Brorsen, 1989; Kinnucan & Forker, 1987). Adjustment costs are the costs incurred due to changing market conditions; that is, they arise when firms change the quantities and/or prices of inputs and/or outputs. If the costs of adjustment with respect to cost increases and decreases are different, asymmetric price transmission can occur. Adjustment costs are sometimes also called menu costs (von Cramon-Taubadel & Meyer, 2004). A straightforward example of menu costs is when a restaurant assumes the price decrease of raw materials is temporary and is reluctant to adjust the prices of the dishes on their physical menu because repricing has a cost. However, a restaurant may quickly adjust its menu when the raw material price increases since at least part of its repricing costs will be covered by the higher menu prices.

Search costs are also important in asymmetric price transmission (Abdulia, 2000). Search costs are those incurred by consumers when searching for market information about a product they are interested in buying. Market information is freely available in a perfectly competitive market. However, in a non-competitive market dominated by only a few firms, it is costly to obtain full market information. Hence, firms in a local market may exercise market power over their customers who may not have enough information on what firms in other local markets charge for the same product. Thus, a firm with local market power may completely and rapidly transmit farm-price increases while remaining reluctant to pass through decreases in input prices, leading to positive price asymmetry.

Kinnucan and Forker (1987) argue that government intervention can also lead to asymmetric farm-retail price transmission. They note that intermediaries (such as wholesalers and retailers) may face uncertainty in establishing prices for their products due to changes in input costs. If intermediaries think that the change in costs is temporary, they may not change their product prices because later menu repricing incurs additional costs. However, if the government intervenes in the market by establishing floor prices, part of intermediaries' pricing uncertainty is reduced. In this case, if farm prices increase, intermediaries assume that the increase is permanent and respond completely and rapidly.

Finally, the perishability of a product and certain inventory management decisions can also cause asymmetric price transmission. Ward (1982) argues that commodity perishability may lead to asymmetric price transmission as retailers of perishable products might hesitate to raise prices for fear of a reduction in sales that leads to spoilage. Frey and Manera (2005) argue that during a period of low demand, firms may increase their inventories rather than decrease their product prices (Reagan & Weitzman, 1982). However, when there is high demand for their products, they increase their prices. Furthermore, firms may adopt accounting criteria, such as first-in-first-out or last-in-first-out (Frey & Manera, 2005), and may, as a result, not adjust prices quickly in response to input price changes.

3.5. Modeling price transmission

A wide variety of empirical techniques are used in the literature to study market integration and price transmission. This section reviews the most common of these methods. Some of the reviewed techniques are applied in the papers included in this dissertation.

3.5.1. The basic framework

Early market integration studies use correlation and simple regression models to investigate the relationship between prices in different markets. Fackler and Goodwin (2001) discuss these methods in considering forms of price transmission analysis and identify them as the oldest approach. Commodity prices in different markets may have a high degree of correlation due to exogenous factors such as weather patterns, pricing by monopolists, and the actions of public agencies (von Cramon-Taubadel, 2017). Thus, a high degree of correlation does not necessarily reflect market integration in the sense of price transmission, which occurs due to flows of goods and/or information between markets by market participants such as traders and arbitrageurs. If $p_{1,t}$ and $p_{2,t}$ are prices in two distinct markets at time t (expressed in log form), the basic representation of a price transmission model can be expressed as follows:[4]

$$p_{1,t} = \beta_0 + \beta_1 p_{2,t} + u_t \tag{18}$$

where u_t is the error term and is assumed to be independent and identically distributed with mean zero. The parameter β_1 is interpreted as elasticity of price transmission and defines the relationship between the prices (or whether the markets are integrated). If $\beta_1 = 1$, the markets are said to be perfectly integrated (LOP holds), whereas $\beta_1 = 0$ indicates completely segmented markets. If $0 < \beta_1 < 1$, the markets are said to be integrated but not perfectly integrated.

The use of Equation (18) in price transmission analysis raises two major conceptual and practical concerns. The first is that $p_{1,t}$ and $p_{2,t}$ are interdependent; $p_{2,t}$ cannot be assumed exogenous with respect to $p_{1,t}$ (Fackler & Goodwin, 2001). The second concern is that the model expressed through Equation (18) is static. Since price adjustment toward the long-run equilibrium may take time, temporary deviations from this equilibrium are inevitable.

In response to the shortcomings of the earlier approaches, researchers in the 1980s turned to the use of dynamic regression models (Fackler & Goodwin, 2001). Dynamic regression models take price endogeneity into account and allow modeling of both contemporaneous and lagged price effects. As a result, these models are considered an improvement over earlier techniques of studying price transmission. The basic framework of the dynamic regression model is a vector autoregressive (VAR) model given by Equation (19) as follows:

$$p_t = \theta + \sum_{i=1}^{k} \Pi_i p_{t-i} + DX_t + u_t \tag{19}$$

where p_t is the $n \times 1$ vector of prices (where n is the number of price series considered in the analysis) at time t, θ is a constant term, Π_i are the $(n \times n)$ matrices of the k-th lagged prices. X_t is an $(m \times 1)$ vector of m possible exogenous factors with the associated $(n \times m)$ parameter matrix D, and u_t is an $(n \times 1)$ vector of disturbance terms, which is assumed to be serially independent.

[4] The prices can be considered as prices in two different locations (if transmission is across spatial markets) or prices in two distinct stages (if transmission is along a supply chain).

Fackler and Goodwin (2001) provide a common template embedding all dynamic regression models. Granger causality tests (also known as Ravallion tests; Ravallion, 1986), various extensions of the VAR model, impulse response functions, and cointegration analysis are empirical tools considered appropriate to analyze price transmission within the basic framework of the dynamic regression model.

3.5.2. Cointegration analysis

Despite dynamic regression models representing an important improvement over earlier techniques, they did not solve all the problems confronting price transmission analysis. In the 1980s, researchers identified that agricultural product price series appear to be non-stationary (or to contain a unit root). Application of regression models (such as that given in Equations (18) and (19)) to non-stationary data may generate spurious results, implying estimated models may indicate a relationship between prices when in fact, no theoretical relationship exists between them.

Cointegration models, which were first introduced in the seminal work of Ardeni (1989), provide a means to distinguish true from spurious price relationships. Cointegration suggests that only a stationary linear combination of non-stationary price series represents a true relationship.

In cointegration analysis, the first step is to ensure that all the price series under investigation are non-stationary. This can be checked using unit root tests such as the Dickey-Fuller and Augmented Dickey-Fuller tests (Dickey & Fuller, 1981). Once the price series are identified as non-stationary in levels while stationary at their first differences, the stationarity of a linear combination of the prices can be checked. If there exists a stationary relationship among the prices, the prices are said to have a cointegration relationship or long-run relationship; otherwise, they do not have a long-run relationship. The cointegrating relationship between prices can be investigated using the Engle and Granger (1987) approach or the Johansen cointegration test (Johansen & Juselius, 1990).

The Engle and Granger (E-G) cointegration test follows a two-step approach. Given the non-stationarity of the prices, the first stage involves estimating a long-run relationship between them. This relationship can be obtained by estimating Equation (18) using the ordinary least square (OLS) approach. The second stage involves extracting the estimated regression residuals (u_t) from the first stage (i.e., from Equation (18)) and then estimating Equation (20) using OLS.

$$\Delta u_t = \rho u_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta u_{t-i} + v_t \tag{20}$$

where v_t is a white-noise error term. The E-G cointegration test depends on the estimated parameter ρ. If the null hypothesis of no cointegration (i.e., $\rho = 0$) is rejected, then we can conclude that the residual series (u_t) is stationary. This further implies the two-price series are cointegrated; otherwise, they are not cointegrated. The presence of cointegration implies the existence of market integration or a price relationship (Barret & Li, 2002; Asche et al., 1999).

Once the prices are found to be cointegrated, in the E-G approach, an error correction model (ECM) can further be estimated to obtain both the long- and short-run price dynamics. According to the Granger representation theorem (Engle & Granger, 1987), an ECM can be given as follows:

$$\Delta p_{i,t} = \alpha_0 + \alpha_1 ECT_{t-1} + \sum_{m=1}^{L_1} \delta_m \Delta p_{i,t-m} + \sum_{n=1}^{L_2} \delta_n \Delta p_{j,t-n} + v_t \tag{21}$$

where $ECT_{t-1} = p_{1,t-1} - \beta_0 - \beta_1 p_{2,t-1}$ is the first lag of the cointegrating vector (also called the error correction term) obtained from the first stage of the E-G approach. The parameter α_1 captures the speed of correction of the short-run deviations towards long-run equilibrium, and the parameter δ captures the short-run price dynamics.

Although the E-G model is appropriate to capture both long- and short-run dynamics of non-stationary variables, it has certain limitations. First, it is limited to bivariate variables; it does not take a systematic approach to the separate estimation of multiple cointegration vectors. Second, the E-G approach relies on a two-step process; that is, a residual series is generated in the first step, and the residuals are then used to estimate an ECM in the second step. Such an approach may result in estimation errors from the first step being carried to the second step. Third, the E-G approach does not take price endogeneity into account.

To cope with the limitations of the E-G approach, Johansen and Juselius (1990) propose the Johansen vector error correction model (VECM), as presented by Equation (22). The approach is a multivariate generalization of the E-G approach and therefore overcomes its limitations.

$$\Delta p_t = \mu_0 + \Pi p_{t-1} + \textstyle\sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + u_t \tag{22}$$

In the equation p_t is a vector of the prices in the system, and p_{t-1} is the first lag of the vector of the prices. μ_0 is a parameter (or constant), Π and Γ_i are the matrices that capture estimated coefficients of the long-and short-run price dynamics, respectively, and u_t is the error term, where $u_t \sim niid(0, \Omega)$.

The number of cointegrating relationships among the prices considered can be obtained using the Johansen trace and maximum eigenvalue tests (Johansen & Juselius, 1990). These tests essentially check the rank of the long-run matrix (i.e., Π) to determine the number of stationary combinations of prices. If $r = rank(\Pi) = 0$, we say that there exists no cointegration relationship between the prices. If $r = n$, where n is the number of prices in the system (this case implies the matrix has full rank), each of the price series in the system is stationary. The use of cointegration analysis is not relevant when the long-run matrix is full rank. Cointegrating relationships between prices exist only if $0 < r < 1$. However, inference of an efficient integration of markets using the cointegration method can be made when $r = n - 1$; it is only in this case that the prices will have a common integrating factor. In the case where $r = 0$ or $r = n$, the VAR model for the first differences of prices and the price levels (respectively) can be estimated to capture the short-run price dynamics.

Whenever there is a cointegrating relationship between the prices (i.e., when $0 < r < 1$), the long-run matrix Π in Equation (22) can be written as a product of two matrices, $\Pi = \alpha\beta'$, where α denotes the adjustment matrix, which captures the speed of adjustments of the price series towards the long-run equilibrium, and β denotes the matrix of the cointegrating vector. In this case, the Johansen cointegration approach in Equation (22) can be rewritten as follows:

$$\Delta p_t = \mu_0 + \alpha ECT_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta p_{t-i} + u_t \tag{23}$$

where ECT_{t-1} $\left(= \beta' p_{t-1}\right)$ is the error correction term and can be defined in the same way as for the E-G model given in Equation (21).

By imposing restrictions on the parameters of the matrices α and β in Equation (23), we can perform various hypothesis tests to better understand the nature of the price transmission process. By restricting the parameters of the cointegration matrix $\beta = 1$, we can test whether the LOP holds. Similarly, by restricting the parameters of the matrix α, we can perform weak exogeneity tests to identify the leading market in the system. The tests can be performed using the Johansen and Juselius (1990) likelihood test.

Cointegration, with its core ECM, has become the dominant model in the agricultural economics empirical literature on price transmission over the last three decades (Goodwin & Schroeder, 1991; von Cramon-Taubadel & Meyer, 2004; Goodwin & Piggott, 2001; Fackler & Goodwin, 2001). Despite its dominancy and success in price transmission and market integration analysis, the cointegration approach still suffers from one major limitation. The cointegration approach implicitly assumes that price adjustment toward long-run equilibrium takes place in a linear form. The literature (such as Balke & Fomby, 1997) indicates that adjustment costs prevent economic agents from adjusting prices linearly and continuously. However, economic agents often act when deviations from equilibrium exceed a critical threshold, as it is only in this case that the benefits of adjustment outweigh its costs. Thus, estimating a linear cointegration model when a price relationship is nonlinear may lead to misleading results (Barrett & Li, 2002).

Several variants of the linear cointegration approaches that allow for modeling price transmission in nonlinear (and asymmetric) forms have been proposed in the literature. Some of these variant models include the threshold cointegration model (Balke & Fomby, 1997), the asymmetric error correction model (von Cramon-Taubadel, 1998), and the threshold vector error correction model (TVECM) (Goodwin & Piggott, 2001).

Before we discuss these nonlinear models, it is helpful to look at the models for asymmetric price transmission since such models have prevailed in the literature due to the growing asymmetric price transmission in both vertical and horizontal markets.

3.5.3. Asymmetric price transmission models

The asymmetric price transmission models in the literature can be categorized as “pre-cointegration” or “threshold cointegration” models.

Pre-cointegration methods

Empirical estimation of asymmetric adjustment goes back to the early 1950s. As reviewed in von Cramon-Taubadel and Meyer (2004), early studies on asymmetry focus on analyzing the irreversible behavior of demand and supply functions. Farell (1952) is the first to investigate the irreversible behavior of demand function for goods such as beer, spirits, and tobacco in the United Kingdom. In agriculture markets, Tweeten and Quance (1969) estimate an irreversible supply function of farm products in the US. By splitting the independent (input) variable (i.e., $p_{2,t}$) in Equation (18) into positive and negative components, Equation (24) given by Tweeten and Quance (1969) can be applied to estimate asymmetric price transmission.

$$p_{1,t} = \alpha + \beta_1^{+} D^{+}\, p_{2,t} + \beta_1^{-} D^{-}\, p_{2,t} + u_t \tag{24}$$

where D^{+} and D^{-} are dummy variables with $D^{+} = 1$ if $p_{2,t} - p_{2,t-1}$ is positive, otherwise zero, and $D^{-} = 1$ if $p_{2,t} - p_{2,t-1}$ is negative, otherwise zero. The dummy variables are used to split the independent variable $p_{2,t}$ into a variable that includes only increasing input prices with adjustment coefficient β_1^{+} and a variable that includes only decreasing input prices with adjustment coefficient β_1^{-}. Asymmetric adjustment is identified if ${\beta_1}^{+}$ and ${\beta_1}^{-}$are significantly different from one another, which can be tested using a standard F-test.

Later, Wolffram (1971) criticizes the approach in Tweeten and Quance (1969), stating the latter is only applicable if the influence of the independent variable over the whole study period is constant (linear); otherwise, it could lead to incorrect parameter estimates. By redefining the increasing and decrease components of the independent variable as the summation of first differences over the whole study period (T), Wolffram (1971) extended the Tweeten and Quance (1969) approach as follows:

$$p_{1,t} = \alpha + \beta_1^{+} \left(p_{2,0} + \sum_{t=1}^{T} D^{+} \Delta p_{2,t} \right) + \beta_1^{-} \left(p_{2,0} + \sum_{t=1}^{T} D^{-} \Delta p_{2,t} \right) + u_t \tag{25}$$

where $\Delta p_{2,t} = p_{2,t} - p_{2,t-1}$, $p_{2,0}$ is the initial value of $p_{2,t}$ at time $t = 0$, D^{+} and D^{-}, as defined earlier, are dummy variables. In Equation (25), the recursive sum of all the positive and negative changes is included as explanatory variables. This implies that Wolfram's model considers the effect of cumulative variation in the explanatory variable. This is distinct from the approach of Tweeten and Quance (1969), who only account for period-to-period variations in the explanatory variable.

Houck (1977) argues that Wolffram (1971) did not take into account the initial observation of the dependent variable since when considering the first differences of the independent variable, the first observation of the dependent variable will have no explanatory power. By defining a dependent variable as $p_{1,t}^{*} = p_{1,t} - p_{1,0}$, Houck (1977) presents a specification similar to Wolffram's model, as follows:

$$p^*_{1,t} = \alpha t + \beta_1^+ \sum_{t=1}^{T} D^+ \Delta p_{2,t} + \beta_1^- \sum_{t=1}^{T} D^- \Delta p_{2,t} + u_t \quad (26)$$

In addition, Houck proposes a specification that includes using the first differences of the dependent and explanatory variables without summing them, which is given by Equation (27):

$$\Delta\, p_{1,t} = \alpha + \beta_1^+ D^+ \Delta\, p_{2,t} + \beta_1^- D^- \Delta\, p_{2,t} + u_t \quad (27)$$

Next, Ward (1982) extended Houck's approach by incorporating lag terms of explanatory variables as follows:

$$p^*_{1,t} = \alpha t + \sum_{j=1}^{K} (\beta_j^+ \sum_{t=1}^{T} D^+ \Delta\, p_{2,t-j+1}) + \sum_{j=1}^{K} (\beta_j^- \sum_{t=1}^{T} D^- p_{2,t-j+1}) + u_t \quad (28)$$

$$\Delta p_{1,t} = \alpha + \sum_{t=1}^{K} (\beta_j^+ D^+ \Delta p_{2,t-j+1}) + \sum_{t=1}^{L} (\beta_j^- D^- \Delta p_{2,t-j+1}) + u_t \quad (29)$$

where K and L are the number of lags for the increasing and decreasing phases of the explanatory variables.

There are a lot of early studies of asymmetric price transmission that employed the pre-cointegration approach in different agricultural markets. Among these studies, Kinnucan and Forker (1987) employ Ward's (1982) model to study farm-retail price transmission of major dairy products in the USA.

Despite the popularity of the model in early studies and even in some of the most recent studies, the pre-cointegration methods for asymmetric price transmission have shortcomings. First, they disregard the time-series properties of price series. Second, they only focus on short-run contemporaneous effects and ignore long-run cointegration.

Threshold cointegration models

As noted earlier, the cointegration method (i.e., the Engle-Granger and Johansen cointegration approaches) is the best approach to identify both short-run and long-run price adjustments. Nevertheless, Granger and Lee (1989) indicate that the cointegration test could lead to a potential inconsistency if there is asymmetric price adjustment. This is because they rely on an assumption of symmetric (linear) price adjustment, and the independent variables in the model are not decomposed as discussed above into positive and negative components. To account for asymmetric price adjustments, Granger and Lee (1989) extend the E-G error correction model given by Equation (21) to an asymmetric error correction model (AECM) by splitting the error correction term ECT_{t-1} into positive deviations (ECT_{t-1}^{+}) and negative deviations (ECT_{t-1}^{-}). The AECM takes the following form:

$$\Delta p_{i,t} = \beta_0 + \beta_1^{+} ECT_{t-1}^{+} + \beta_1^{-} ECT_{t-1}^{-} + \sum_{m=1}^{L_1} \delta_m \Delta p_{i,t-m} + \sum_{n=1}^{L_2} \delta_n \Delta p_{j,t-n} + u_t \tag{30}$$

where ECT_{t-1}=$ECT_{t-1}^{+} + ECT_{t-1}^{-}$. In Equation (30), the asymmetric price transmission is confirmed if the null hypothesis ($\beta_1^{+} = \beta_1^{-}$) is rejected. When the null hypothesis is not rejected, the Granger and Lee (1989) model given in Equation (30) nests the Engle-Granger model given in Equation (21).

Von Cramon-Taubadel (1998) further extends the Granger and Lee (1989) model by splitting $\Delta p_{j,t-n}$ into positive and negative components (Equation 31) to make it possible for the model's use to test the short-run asymmetry adjustment.

$$\Delta p_{i,t} = \beta_0 + \beta_1^{+} ECT_{t-1}^{+} + \beta_1^{-} ECT_{t-1}^{-} + \sum_{m=1}^{L_1} \delta_m \Delta p_{i,t-m} + \sum_{n=1}^{L_2} \delta_n^{+} \Delta p_{j,t-m} + \sum_{n=1}^{L_3} \delta_n^{-} \Delta p_{j,t-m} + u_t \tag{31}$$

In Equation (31), the null hypothesis of short-run asymmetry takes either the strong form $(\text{i.e.}, \delta_n^{+} = \delta_n^{-})$ or the weak form $(\text{i.e.}, \sum_{n=1}^{L_2} \delta_n^{+} = \sum_{n=1}^{L_3} \delta_n^{-})$. Rejection of the null hypothesis of either form implies the presence of short-run asymmetric price transmission.

As noted earlier, Balke and Fomby (1997) argue that economic agents only start to adjust prices when the price deviations from long-run equilibrium exceed a critical threshold. Taking both the asymmetry and threshold effect into consideration, Enders and Siklos (2001) extend the E-G procedure given by Equation (20) to a threshold autoregressive (TAR) model as given in Equation (32) below.

$$\Delta u_t = I_t \rho_1 u_{t-1} + (1 - I_t)\rho_2 u_{t-1} + \textstyle\sum_{i=1}^{p} \gamma_i \Delta u_{t-i} + v_t \tag{32}$$

where I_t is an indicator variable defined by Equation (33).

$$I_t = \begin{cases} 1 \ if \ u_{t-1} \geq \tau \\ 0 \ if \ u_{t-1} < \tau \end{cases} \tag{33}$$

where τ is a threshold value, I_t is equal to one when u_{t-1} is larger than the threshold value and zero otherwise. The model in Equation (32) is known as the momentum-threshold autoregressive (M-TAR) model when I_t in Equation (33) is replaced by its differences Δu_{t-1}.

$$I_t = \begin{cases} 1 \; if \; \Delta u_{t-1} \geq \tau \\ 0 \; if \; \Delta u_{t-1} < \tau \end{cases} \tag{34}$$

As indicated by Equations (20) and (32), the Engle-Granger test is a special case of the Enders and Siklos (2001) test when $\rho_1 = \rho_2$. From Equation (32), cointegration exists if the hypothesis $\rho_1 = \rho_2$=0 is rejected. This leads to the test of symmetry $\rho_1 = \rho_2$. If both hypotheses are rejected, the process is called threshold cointegration. According to Enders and Siklos (2001), the cointegration test follows a non-standard F-test; hence, the results are compared to the critical values given in Enders and Siklos (2001). The symmetry test, however, is a standard F-test. Chan (1993) proposes a method to estimate the best threshold value. Chan (1993) suggests that the estimated residuals (in the case of the TAR model) or the first differences of the estimated residuals (in the case of the M-TAR model) are first sorted in ascending order. The largest and smallest 15% of values are excluded before the best threshold value, that yielding the lowest residual sum of squares, is selected.

If asymmetric price transmission is detected through Equation (32), according to Balke and Fomby (1997) and Enders and Granger (1998), the conventional ECM can be extended to a threshold error correction model. This is achieved by decomposing the error correction terms into positive and negative components to analyze the asymmetric effects on the dynamic behavior of prices, as presented by Equation (35). The cointegration model is preferable since it accounts for both short- and long-run price adjustments, as discussed above.

$$\Delta p_{i,t} = \beta_0 + \beta_1^+ ECT_{t-1}^+ + \beta_1^- ECT_{t-1}^- + \sum_{m=1}^{L_1} \delta_m \Delta p_{i,t-m} + \sum_{n=1}^{L_2} \delta_n \Delta p_{j-n,t} + u_t \tag{35}$$

where $i = 1, 2$, L_1 and L_2 are the lag lengths to be determined based on lag selection criteria. The error correction terms ECT_{t-1}^+ and ECT_{t-1}^+ in Equation (35) can be expressed as:

$ECT_{t-1}^+ = I_t(I_t \geq \tau)\big(p_{1,t} - \beta_0 - \beta_1 p_{2,t}\big)$ and

$ECT_{t-1}^- = I_t(I_t \leq \tau)(p_{1,t} - \beta_0 - \beta_1 p_{2,t})$

In Equation (35), the parameters β_1^+ and β_1^- capture the adjustment when the equilibrium is above and below the threshold value, respectively. Furthermore, in Equation (35), it is possible to investigate the presence of short-run asymmetric price transmission by decomposing the short-run coefficients into positive and negative components.

The threshold cointegration model discussed above follows a single equation specification to make the discussion easy to follow. However, if the causation is bi-directional, estimation results obtained using a single equation specification may be misleading. Goodwin and Holt (1999) and Goodwin and Piggott (2001) propose a threshold cointegration model that follows a system of equations known as the TVECM. The TVECM extends the linear VECM in Equation (23) to create non-linearity in the model. The two-threshold, three-regime TVECM takes the following form:

$$\Delta p_t = \begin{cases} \mu_1 + \boldsymbol{\alpha_1} ECT_{t-1} + \sum_{i=1}^{k} \boldsymbol{\Gamma_i}^{\boldsymbol{1}} \Delta p_{t-i} + u_{1t}, \text{if } -\infty < ECT_{t-1} < \tau_1 \\ \mu_2 + \boldsymbol{\alpha_2} ECT_{t-1} + \sum_{i=1}^{k} \boldsymbol{\Gamma_i}^{\boldsymbol{2}} \Delta p_{t-i} + u_{2t}, \text{if } \tau_1 \leq ECT_{t-1} \leq \tau_2 \\ \mu_3 + \boldsymbol{\alpha_3} ECT_{t-1} + \sum_{i=1}^{k} \boldsymbol{\Gamma_i}^{\boldsymbol{3}} \Delta p_{t-i} + u_{3t}, \text{if } \tau_2 < ECT_{t-1} < \infty \end{cases} \tag{36}$$

where τ_1 and τ_2 are the threshold values, and all the other notations have the same meanings as in Equation (23). In Equation (36), the long-run adjustment coefficients (i.e., α) and short-run adjustment matrix (i.e., Γ) may take different values depending on the relative magnitude of the deviation from the equilibrium of the threshold values, which characterize different adjustments toward the long-run equilibrium. Between the threshold values in Equation (36), a “neutral band” may form, within which, as a result of transaction costs, prices might not be linked to one another.

Estimation using a TVECM requires first determining whether the data at hand can be best estimated using a linear VECM or a TVECM. Empirical studies rely on different statistical tests to identify this. The most used statistical procedure in the literature is introduced by Hansen and Seo (2002). The results given by the procedure suggest those obtained estimating a linear VECM or TVECM. The number of threshold values in an empirical estimation is guided by the nature of the data at hand and price adjustment in a specific market (or trade flows), although most empirical studies estimate either one-threshold (two-regime TVECM) or two-thresholds (three-regime TVECM).

Other threshold models in a similar spirit to the previous models can be deduced from the standard autoregressive model of price differentials as follows:

$$d_t = \beta d_{t-1} + \varepsilon_t \tag{37}$$

where d_t represents the price differentials (i.e., $d_t = p_{1,t} - p_{2,t}$). When the estimated value of β is closer to one, it implies that a shock has a permanent effect on price differentials. By contrast, if $\beta = 0$, a shock tends to quickly die out over time. Rearranging Equation (37), we have:

$$\Delta d_t = \rho d_{t-1} + \varepsilon_t \tag{38}$$

where $\rho = \beta - 1$.

The above models do not account for price non-linearity. According to Balke and Fomby (1997), a nonlinear or TAR model occurs when the size of the lagged price differentials leads to different behaviors in the adjustment process in a regime fashion. This means that the value of the parameter ρ varies according to whether the price differentials are bigger or smaller than certain threshold values. A three-regime TAR model, as in the cointegration version of Equation (36), can be represented as follows:

$$\Delta d_t = \begin{cases} \rho_1\, d_{t-1} + \varepsilon_{1,t}, & if -\infty < d_{t-1} < \tau_1 \\ \rho_2 d_{t-1} + \varepsilon_{2,t}, & if\ \tau_1 \leq d_{t-1} \leq \tau_2 \\ \rho_3 d_{t-1} + \varepsilon_{3,t}, & if\ \tau_2 < d_{t-1} < \infty \end{cases} \tag{39}$$

where τ_1 and τ_2 are the threshold parameters, and a neutral band is formed between the threshold parameters (Serra et al., 2006a).

In general, the linear and threshold cointegration approaches discussed above have dominated the market-integration and price transmission literature. This has been particularly true in recent decades. However, several other approaches not analyzed in this dissertation have been documented in the literature. These include the parity bounds model (Baulch, 1997; Barrett & Li, 2002) and non-parametric techniques such as time-varying copula (Goodwin et al., 2011; Qiu & Goodwin, 2012; Emmanouilides, & Fousekis, 2015) and local polynomial techniques (Serra et al., 2006a; Serra et al., 2006b).

4. Summary of the papers in the dissertation

This dissertation comprises four papers on price determination in selected agri-food markets. Table 1 summarizes the markets considered, the countries investigated, and the data sources utilized. The first paper uses the threshold cointegration approach (Enders & Siklos, 2001) and the AECM (Blake & Fomby, 1997; Enders & Granger, 1998) to investigate asymmetric price transmission along the salmon supply chain. The second paper uses the same approach to explore spatial price transmission in the Ethiopian grain markets. The third paper employs two distinct approaches – threshold cointegration and local polynomial approaches – to investigate spatial price transmission in EU pork markets. The fourth paper adopts the demand index approach of Marsh (2003) for measuring global seafood-price formation.

Table 1. Selected agri-food products, markets, topics, and data sources considered

Papers	Agricultural product	Market	Topic	Data source
1	Salmon	Norway, France, Spain	Vertical Price Transmission	Norwegian Seafood council
2	Grains	Ethiopia	Market integration and Spatial Price Transmission	Ethiopian Grain Enterprise
3	Pork	EU	Market integration and Spatial Price Transmission	European commission
4	Aggregated seafood	Global	Demand index	FAO STATA database

4.1. Paper 1: Asymmetric price transmission in a changing food supply chain

The global agri-food market is transforming. There has been an increasing level of market concentration in retail-level markets in recent years. Policymakers and researchers are concerned that an increased level of concentration might yield firms with market power. Intermediaries with market power may follow pricing strategies that transmit margin-squeezing increases in input prices faster and/or more completely than the corresponding margin-stretching price changes (Vavra & Goodwin, 2005). Moreover, due to the increasing product differentiation in the seafood market (for instance, fresh and smoked in the case of salmon), different processes of price determination could apply in various product sub-markets. In this study, we investigate asymmetric price transmission for two salmon products: a relatively unprocessed salmon product (fresh salmon) and a more processed salmon (smoked salmon). We look at Norwegian exports to the retail markets of France and Spain.

The TAR model and the AECM are estimated to examine the patterns of price adjustment. The results of the TAR model indicate that the pairs of prices are cointegrated (in a long-run relationship) along the fresh salmon supply chain in both markets but not along the supply chain of smoked salmon. Moreover, along the fresh salmon supply chain, the results indicate asymmetric price transmission in both markets, showing that the rate of adjustment to the long-run equilibrium following negative price shocks tends to occur more rapidly than for positive shocks. Furthermore, the estimates of the AECM suggest that the French market plays the leading role in export prices in the supply chain of the fresh whole-salmon market; it was the retail prices that adjusted to changes in the export prices, not vice versa.

However, in the Spanish market, bi-directional price adjustment is observed. Processing salmon into value-added consumer products involves additional inputs that include labor, capital, marketing, and packaging costs and time (Landazuri-Tveteraas et al., 2018). The greater the share of these non-raw-materials costs in the final consumer prices, the less the price influence of the farmed fish on the final product price. Coupled with other factors that influence the price transmission process (e.g., storage, menu costs, market power, and so on), the influence on the price transmission would be minimal. Hence, our results may indicate that price transmission to retail prices decreases as more processing is involved.

4.2. Paper 2: Market integration and price transmission in the regional grain markets in Ethiopia

Since mid-2006, prices of basic food crops (such as wheat and maize) have persistently increased and have become a critical challenge in Ethiopia. The literature indicates that food prices in any market may change as a result of several factors, including price shock diffusion from international markets (Baquedano & Liefert, 2014; Ceballos et al., 2017; Minot, 2011), changes in the exchange rate (Hazel et al., 1990; Dawe, 2008), changes in domestic demand and supply conditions (Brækkan et al., 2018), and in the market structure (Abdulai, 2000). In this paper, the objective was to investigate whether the Ethiopian grain market structure has contributed to price increases in the Ethiopian grain market. The study uses monthly wholesale prices of wheat, maize, and teff from major regional grain markets in Ethiopia. The Engle-Granger and threshold cointegration models were the main methods employed to investigate the extent, speed, and nature of price signal transmission between the central and major regional grain markets. The estimated results given by both models suggest the presence of long-run relationships between the central and major regional grain markets for each commodity. Moreover, the results obtained from the threshold cointegration model indicate the general absence of asymmetric price adjustment in each of the markets. This might imply the presence of efficient price transmission in the Ethiopian grain market. Hence, I argue that this study provides insufficient evidence to support the hypothesis that the price adjustment behavior of grain traders (or market structure) leads to sustained grain-price increases in Ethiopia.

4.3. Paper 3: Spatial price transmission in EU pork markets: using threshold autoregressive and non-parametric local polynomial techniques

The literature indicates that the EU pork industry has shown a high level of concentration in countries of production and at the retail level (e.g., Fousekis, 2015; Holst and von Cramon-Taubadel, 2013; Skrzymowska, 2012). The high level of concentration may affect the EU pork-price transmission between markets both along the pork supply chain and spatially across EU pork markets. This study investigates spatial price transmission in EU pork markets. Specifically, using updated weekly pigmeat prices from six major EU markets including, Germany, France, Spain, Denmark, the Netherlands, and Poland, we replicate a price transmission study performed by Serra et al. (2006a) seventeen years ago. However, we rely on a statistical procedure to identify the presence of nonlinear price adjustment when using the non-parametric local polynomial approach; the latter was one of the approaches of Serra et al. (2006a), contrary to Serra et al. (2006a), that relies on visual inspection. At the same time, the replication of an old study empirically contributes to understanding market development. The other primary methods used are the Johansen cointegration test and the TAR model. The results from the Johansen cointegration test, in general, suggest the presence of long-run relationships between the pairs of prices investigated. The results of the TAR and local polynomial techniques also indicate price adjustment between the EU pork markets following price shocks. Nevertheless, the estimated speed of adjustments obtained from the TAR model is generally low, suggesting price adjustments take time. Furthermore, the TAR and local polynomial approaches provide some evidence of asymmetric adjustment within the EU pork markets. In general, the results in this study are in line with the results of Serra et al. (2006a).

4.4. Paper 4: Global seafood demand growth differences across regions, income levels, and time

The seafood sector is one of the rapidly growing food sectors globally (FAO, 2016). Since 2000, global seafood prices have displayed an upward trend (Tveterås et al., 2012), although the literature indicates the supply of seafood has been increasing globally over the past four decades. The upward price trend suggests demand growth has been outpacing supply growth. While there exist several studies on the supply side of the seafood market, the literature on the growth in seafood demand is limited. In this paper, we use the demand index approach of Marsh (2003) to measure the growth in seafood demand of 107 countries between 1984 and 2013. The approach allows us to measure demand shifts caused by all factors (both known and unknown), in contrast to measuring demand shifts using a smooth operator, such as a time trend, which is a widely used approach in the literature (Deaton & Muellbauer, 1980; Kinnucan et al., 1997; Stone, 1945). We further aggregate the demand shifters across countries and regions (and globally), and across income groups, and time. The results show substantial variation in demand growth across countries, regions, income groups, and over time. The average annual seafood demand growth across countries varies between −6 and 7.5%. The global growth in the demand for seafood has steadily declined since the 1980s, mainly caused by a slowdown of demand in Asia. South America and Africa saw the highest demand growth from 2004 to 2013, while both North America and Oceania saw demand decline in this period. High-income countries show consistently low demand growth from 1984 to 2013, while demand growth in low-income countries is substantial. Our results show that although seafood production has more than doubled since the mid-1980s, global demand for seafood has been higher than the global seafood supply, which explains the growing global seafood prices.

5. General discussion and implications

The causes of changes in agri-food prices will always be an issue of concern for producers, consumers, policymakers, and of course, economists. Often producers worry about price reductions for their products in the farm-level markets, consumers worry about commodity price increases in the retail-level markets, and governments are concerned about welfare distribution. Economists, however, try to understand the causes behind commodity price changes in a market.

This dissertation highlights the fact that while commodity price development can be attributed to demand and supply dynamics, price signal transmission between markets vertically along the supply chain and horizontally in space can play a major role in commodity price dynamics. Inefficient price signal transmission between markets can distort producer decisions and lead to inefficient product movement, and ultimately increase consumer prices (Goodwin & Schroeder, 1991). Thus, when the price signal transmission between markets is inefficient, it requires policy support for the improvement of the process.

Price transmission analysis focuses on the speed, extent, and nature of price signal transmission between markets. The speed, extent, and nature of price signal transmission are determined by, among other factors, the degree of market power and adjustment costs in the market (e.g., Abdulia, 2000). Thus, by devising appropriate policies that affect these two factors, it is possible to attain efficient price transmission between markets. For instance, if there is inefficiency because of market power, devising policies that encourage competition among market actors (both along the supply chain and between spatially related markets) may facilitate efficient price signal transmission. By contrast, if the inefficiency is due to adjustment costs in the market, price transmission can be facilitated by improving the quality of market infrastructure and enhancing coordination between markets.

The issue of asymmetric price transmission has been an issue of considerable importance in the literature on price transmissions. There are various possible reasons for this. The first is the assumption that the presence of asymmetric price transmission in a market represents market failure or market power (Vavra & Goodwin, 2005). The second reason could result from the assumption that asymmetric price transmission could lead to a welfare transfer, for instance, from producers to retailers (von Carmon-Taubadel & Meyer, 2004). Additional insight about the pricing strategies of market actors can be obtained by investigating asymmetric price transmission.

The literature often indicates that, along the supply chain, margin-squeezing increases in input prices will be transmitted faster and/or more completely than the corresponding margin-stretching price changes (Vavra & Goodwin, 2005). Similarly, between spatially related distinct markets, positive and negative price changes in some markets may not correctly and appropriately transmit to other markets (e.g., Serra et al., 2006a; Abdulia, 2000). This dissertation provides further evidence in support of this argument. For instance, in Paper 1, it was found that the retailers transmitted increases in export prices (a proxy for farm prices) more completely and quickly than the corresponding price reductions along the fresh salmon supply chain. Such results may imply that chain actors have market power or there is a high adjustment cost along the salmon chain.

There is a huge body of literature on price transmission between markets both vertically along supply chains and horizontally in spaces (von Carmon-Taubadel & Meyer, 2004; Peltzman, 2000; Serra et al., 2006a; Serra et al., 2006b; Abdulia, 2000). However, despite the availability of a huge literature on price transmission, it is hard to generalize the results due to the diversity of commodities, countries, regions, and time frequencies or periods analyzed and the methodologies employed. It is thus difficult to draw general conclusions on which policy decisions can be based (Vavra & Goodwin, 2005). In this regard, the three empirical papers presented for this dissertation (Papers 1–3) contribute to the understanding of price signal transmission mainly in the respective markets studied.

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Asymmetric price transmission in a changing food supply chain

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ABSTRACT

The farmed salmon supply chain in Europe is changing. There is a growing concentration at intermediary levels in the supply chain and more product differentiation in the market. This means that different price determination processes could apply in various product sub-markets. In this study, price transmission relationships in the salmon value chain were investigated at two different product levels. Specifically, a relatively unprocessed salmon product (fresh salmon) and a more processed product (smoked salmon) were investigated since processing costs might have a significant impact on a vertical price linkage. A threshold cointegration model was applied to estimate the price transmission between the Norwegian export market and the retail markets of France and Spain. The results indicated a price transmission relationship along the fresh salmon chain in both markets; but not along the smoked salmon chains in either market. Furthermore, for the fresh salmon value chain, asymmetric adjustment was observed in both markets.

KEYWORDS

Asymmetry; farmed salmon food chain; price transmission; threshold

Introduction

Due to the recent increased levels of firm concentration at intermediary levels in the supply chain (i.e., wholesalers and retailers); both researchers and the industry have raised concerns about the potential market power of intermediaries in the supply chain of seafood (Fernández-Polanco & Llorente, 2019; Guillen & Franquesa, 2015; Simioni et al., 2013). Intermediaries with market power are likely to employ pricing strategies that result in a relatively complete and rapid pass-through when farm price increases and an incomplete and slow pass-through when farm price decreases. At the same time, due to the increasing development of new product forms and transaction methods, the seafood value chain has become more sophisticated in recent years (Asche et al., 2014; Asche &

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Smith, 2018), and often these innovations require economies of scale or coordination (Anderson et al., 2018; Bergesen & Tveterås, 2019; Kvaløy & Tveteras, 2008).

In the study of price transmission, asymmetric price transmission is of particular interest for the understanding of seafood markets. It is expected that firms' marketing strategies for each product form are different. For instance, in the case of salmon, for more processed products such as smoked salmon, retailers would probably like to be engaged in a long-term contract (usually for six months or longer) to fix the price of raw fish. However, for a less processed product such as fresh salmon, retailers would prefer shorter contracts and more flexible pricing to avoid price risks (Asche et al., 2014; Larsen & Asche, 2011). The diversity of firms' marketing behaviors applied to different product forms might influence the price transmission of the same species along its value chain.[1]

Earlier asymmetric price transmission studies on seafood focused on different species (e.g., salmon, cod) but in an aggregated product form. To the best of our knowledge, there are only two studies in the literature of salmon asymmetric price transmission, Simioni et al. (2013) and Ankamah-Yeboah and Bronnmann (2017), both focusing on aggregate salmon products. However, understanding the different price transmission mechanisms among different salmon products in a disaggregated product level is important. Asche et al. (2014) and Landazuri-Tveteraas et al. (2018) investigated price transmission for a broader set of salmon products in the UK and French retail markets; they showed that the degree of price transmission varied with product forms, and was higher for relatively unprocessed product forms. This indicates that it is important to also consider product form when investigating asymmetric price transmission.

The main objective of this study was to examine the asymmetric price transmission in the value chain of whole fresh salmon and smoked salmon from the export market of Norway to the retail markets of France and Spain. Fresh whole salmon and smoked salmon were selected because fresh whole salmon is an unprocessed salmon product while smoked salmon is one of the most processed salmon products being exported. The reasons for selecting the French and Spanish markets are twofold: France and Spain are the main export markets for Norwegian salmon, and earlier studies on asymmetric price transmission have not covered these two markets. The main method used in this study is the threshold cointegration model with both zero and non-zero threshold. We use the threshold cointegration model because it allows us to investigate the presence of any price asymmetry in the value chain.

The rest of this paper is organized as follows. The next section contains background and a literature review. The theoretical framework of price transmission is presented in the third section. The fourth and fifth sections present the methodology and data used in the study, followed by the empirical results in the sixth section and the concluding remarks in the last section.

Background and literature review

The relationship between commodity prices is an important research area of agricultural product markets. In general, there are two common forms of price transmission, vertical and horizontal price transmission (von Cramon-Taubadel & Meyer, 2004). Studies in vertical price transmission look at the vertical price linkage in a commodity's value chain, while studies in horizontal price transmission look at the price linkage across marketplaces and different commodities and are also known as market integration studies. The theory of derived demand predicts that in the case of two vertically integrated market levels, a price change that occurs at one stage will create an impact on the price of the other stage for at least one input factor. Horizontal price transmission includes spatial and cross-commodity price transmission. The theoretical foundation of spatial price transmission is the spatial arbitrage and consequence of the Law of One Price (LOP), while the basis for cross-price transmission is the substitutability between and complementary relations among commodities (Singh et al., 2015).

The study of price transmission in a commodity's value chain gives insights about market efficiency, and the size and distribution of producer and consumer welfare (von Cramon-Taubadel & Meyer, 2004). As a result, economists who study market integration and market efficiency investigate the price transmission process. Asymmetric price transmission, which implies that increases and decreases in prices at one level of a value chain of a commodity are transmitted at different rates to other levels, has received considerable attention in agricultural commodity value chain research in recent years (Frey & Manera, 2007; Peltzman, 2000; Simioni et al., 2013; von Cramon-Taubadel & Meyer, 2004).

For a long time, asymmetric price transmission was considered a manifestation of market failure in microeconomic theory. Any exogenous shock to a price system, whether a price shock was negative or positive, should result in symmetric adjustments to the long-run market equilibrium (von Cramon-Taubadel & Meyer, 2004; Frey & Manera, 2007). However, in recent literature, it became clear that asymmetric price transmission can arise in a perfectly competitive market and is therefore incorrect to generalize asymmetric price transmission as a manifestation of market failure. Due to the importance of this phenomenon, researchers have also investigated what causes asymmetric price transmission. In the review papers of von Cramon-Taubadel and Meyer (2004) and Frey and Manera (2007), the documented causes are market power, search costs, adjustment costs of food menus, the nature of the agricultural products, and inventory storage.

Although the literature on asymmetric price transmission on seafood is limited, most of the existing studies confirm asymmetric price transmission (e.g., Ankamah-Yeboah & Bronnmann, 2017; Bittmann et al., 2019; Guillen & Franquesa, 2015; Jaffry, 2004; Simioni et al., 2013). This illustrates the fact that in the study of price transmission across market chains of seafood, price asymmetry is important. Without considering asymmetric price transmission, the estimated results of price transmission might be biased.

The salmon supply chain is more developed compared to any other farmed species (Asche et al., 2018). Some of the reasons for this level of sophistication are the development of diverse salmon product forms, the presence of long-run contracts, and salmon futures markets (Asche et al., 2014; Asche & Smith, 2018). The development of the salmon supply chain has resulted in increased interest in price transmission studies. A few studies have investigated the relationship between different levels in the supply chain. Larsen and Kinnucan (2009) found that price transmission for fresh salmon is complete. Asche et al. (2014) obtained a similar result for fresh salmon but found incomplete price transmission for smoked salmon. Out of 17 retail salmon products examined, Landazuri-Tveteraas et al. (2018) found full-price transmission in only one product value chain. Furthermore, their results showed that price transmission to retail prices decreased, as more processing was involved. As we have discussed above, there are only two studies that investigated price asymmetry along the salmon supply chain (i.e., Ankamah-Yeboah & Bronnmann, 2017; Simioni et al., 2013), but on aggregated product prices. Their conclusion of asymmetric price transmission might not hold when disaggregated data is used.

The European Union (EU) is the world's largest market for farmed salmon, with rapid demand growth (Braekkan et al., 2018). Norway is the primary supplier in the region (Asche et al., 2014; Guillotreau et al., 2005). For instance, Norway exported one million tonnes of salmon in 2017, of which 80% was exported to the EU (EUMOFA, 2017). The largest single markets for Norwegian salmon in 2017 were Poland (18%), France (13%), and Denmark (12%). Poland and Denmark are the hub markets where salmon is reexported to other countries within the EU. For the other European markets, Spain, the UK, the Netherlands, and Italy had market shares of 9%, 8%, 8%, and 7%, respectively.

Theoretical framework of price transmission

Following Asche, Menezes, et al. (2007) and Larsen and Kinnucan (2009), for a specific product in an international marketing channel, the fundamental relationship between retail level and farm level prices and the exchange rate variable can be given as follows:

$$P_R^f = f\left(P_F^f,\ P_C^f\right) \text{ (International Price Linkage Relation)} \tag{1}$$

$$P_F^f = P_F^d.Z \text{ (Exchange Rate Identity)} \tag{2}$$

P_R^f denotes the retail level price of a given product in a foreign market measured in the foreign currency, P_F^f is the farm price in the domestic market but measured in the foreign-currency, and P_C^f is the cost necessary to convert the farm product into a retail product (e.g., marketing services, transportation costs, and menu costs) measured in the foreign currency. P_F^d is the farm price in the domestic market measured in the domestic currency, and Z is the bilateral exchange rate expressed as units of foreign currency per unit of domestic currency.

Taking the logarithmic total differential of Equations (1) and (2) yields:

$$d\ lnP_R^f = B_F\ d\ lnP_F^f + B_C\ d\ lnP_C^f \tag{3}$$

$$d\ lnP_F^f = d\ lnP_F^d + d\ lnZ \tag{4}$$

where B_F is the farm-retail international price transmission elasticity, with both prices expressed in the same currency, and B_C is the cost price transmission elasticity when costs are priced in the foreign currency. Substituting Equation (4) into Equation (3), we get the following:

$$d\ lnP_R^f = B_F\ d\ lnP_F^d + B_Z\ d\ lnZ + B_C\ d\ lnP_C^f \tag{5}$$

In Equation (5), if the market is efficient such that changes in domestic exchange rates are perfectly reflected in foreign prices, then $B_F = B_Z$. Assuming a perfect pass-through of the exchange rate in the value chain of salmon in this study, Equation (5) reduces to:

$$d\ lnP_R^f = B_F\ d\ ln\ \tilde{P}_F^d + B_C\ d\ lnP_C^f \tag{6}$$

where $\tilde{P}_F^d$ $(= P_F^d * Z)$ is the farm price in the domestic market measured in the foreign currency. Moreover, in most empirical analyses of a supply chain, P_C^f is assumed to be constant so that it can be included in the constant term.[2] Taking this assumption in our case and appending time subscripts to the variable and incorporating a random error term (μ_t) after simplifying Equation (6), the long-run equilibrium relationship between the upstream and the downstream market stages can be given by:

$$lnP_{R,t}^f = \varphi + B_F\ \ln P_{F,t}^d + \mu_t \tag{7}$$

where $P_{R,t}^f$ as discussed denotes the retail level price at time t expressed in foreign currency, which is Euro/kg in our study. Different from above, $P_{F,t}^d$

now denotes the farm price at time t expressed in foreign currency units. Estimating Equation (7) creates a simultaneity problem because economic theory does not indicate the direction of the relationships. However, in several studies, the direction of causality is identified based on certain characteristics of the market (Kinnucan & Forker, 1987). It is usually assumed that the price is established at the farm level and it flows forward to the retail level market (Kinnucan & Forker, 1987). A common explanation for the choice has been that supply shock is more common than demand shocks and that retailers follow a fixed markup pricing. Another common approach followed by several other researchers and which is used in this study is to identify the causal market by employing exogeneity tests.

Methodology

Our empirical procedure comprises a series of tests and model estimations. First, we performed a stationarity test using the Augmented Dickey–Fuller (ADF) test (Dickey & Fuller, 1981) on individual price series. Then, we estimated the threshold cointegration models to investigate the presence of possible asymmetric price adjustment between the upstream and downstream prices. For price pairs demonstrating asymmetric adjustment, we estimated the threshold asymmetric error correction model.

Due to the presence of adjustment costs (e.g., menu costs), economic agents often react to a price change only when deviations from the equilibrium exceed a certain critical limit, which is called a threshold value (Blake & Fomby, 1997; Enders & Granger, 1998). The behavior of agents leads to asymmetric price adjustment in a commodity value chain. Therefore, to study price transmission in a commodity value chain, using models that allow asymmetric adjustment is often necessary. In this study, we used the threshold cointegration proposed by Enders and Siklos (2001) because it allowed us to investigate the presence of possible asymmetric price adjustment between the upstream and downstream markets.

The threshold cointegration approach proposed by Enders and Siklos (2001) is an extension of the Engle and Granger (1987) procedure. We start first with a discussion of the Engle-Granger procedure. The procedure relies on the ordinary least squares (OLS) estimate of ρ in the following regression Equation (8):

$$\Delta\mu_t = \rho\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-i} + \upsilon_t \tag{8}$$

where μ_t is the estimated regression residuals extracted from the price linkage Equation (7), and υ_t is a white-noise disturbance. Equation (8) implies an assumption of symmetric price adjustment since ρ is estimated as an average effect of the lagged error term μ_{t-1} regardless of whether μ_{t-1} is positive or negative. Taking the asymmetry into consideration, Enders and Siklos (2001) extended the Engle-Granger procedure to a Threshold Autoregressive (TAR) model given by Equation (9):

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p}\gamma_i\Delta\mu_{t-i} + \upsilon_t \tag{9}$$

I_t is an indicator variable defined by Equation (10):

$$I_t = \begin{cases} 1 \ \ if \ \ \mu_{t-1} \geq \tau \\ 0 \ \ if \ \ \mu_{t-1} < \tau \end{cases} \tag{10}$$

where τ is a threshold value and, I_t is equal to 1 when μ_{t-1} is larger than the threshold value, otherwise it is 0. The adjustment is modeled by $\rho_1\mu_{t-1}$, if μ_{t-1} is above the threshold and by the term $\rho_2\mu_{t-1}$, if μ_{t-1} is below the threshold. The model in Equation (9) is known as a Momentum-Threshold Autoregressive (M-TAR) model when I_t in the above Equation (10) is replaced by its differences $\Delta\mu_{t-1}$.

$$I_t = \begin{cases} 1 \ \ if \ \ \Delta\mu_{t-1} \geq \tau \\ 0 \ \ if \ \ \Delta\mu_{t-1} < \tau \end{cases} \tag{11}$$

Furthermore, TAR and M-TAR models are different in that the TAR model can capture a deep cycle process if, for instance, the variation above the threshold level is more prolonged than below the threshold level. Meanwhile, the M-TAR is capable of capturing sharp sequential movement and is especially valuable when the series exhibits more momentum in one direction than the other (Enders & Siklos, 2001).

As indicated by Equations (8) and (9), the Engle-Granger procedure is a special case of Enders and Siklos's test when $\rho_1 = \rho_2$. From Equation (9), cointegration exists if the hypothesis $\rho_1 = \rho_2 = 0$ is rejected. This leads to the test of symmetry $\rho_1 = \rho_2$. If both hypotheses are rejected, the process is called threshold cointegration. According to Enders and Siklos (2001), the cointegration test follows a nonstandard *F*-test; hence, results are compared to the critical values given in Enders and Siklos (2001). The symmetry test, however, is a standard *F*-test. Following the procedure proposed by Chan (1993), the best threshold value is used. The estimated residuals (in the case of the TAR model) or the first differences of the estimated residuals (in the case of the M-TAR model) is first sorted in ascending order, then 15% of the largest and smallest values are excluded before the best threshold value that yields the lowest residual sum of squares is selected.

The Granger representation theorem (Engle & Granger, 1987) states that a vector error correction model (VECM) can be estimated when variables are cointegrated. However, a conventional VECM cannot consider the asymmetric transmission issue since the error term has not been decomposed into positive and negative components. In this study, by following Blake and Fomby (1997) and Enders and Granger (1998), we extended the conventional VECM by decomposing the error correction terms into positive and negative components, which allowed us to analyze the asymmetric effects on the dynamic behavior of the prices as presented by Equations (12) and (13).

In Equations (12) and (13), the parameters β_2^+ and δ_2^+ capture the adjustment of the retail and export level prices, respectively, when the equilibrium deviation is above the threshold value. On the other hand, the parameters β_2^- and δ_2^- capture the adjustment of the retail and export prices, respectively, when the equilibrium deviation is below the threshold value. In the empirical results section, we call the estimated model of Equation (12) the retail equation because the dependent variable is the retail level price (i.e., $\Delta lnP^f_{R,t}$), while we call the estimated model of Equation (13) the export equation because the dependent variable is the export level price (i.e., $\Delta lnP^d_{F,t}$).

$$\Delta lnP^f_{R,t} = \beta_1 + \beta_2^+ ECT^+_{t-1} + \beta_2^- ECT^-_{t-1} + \sum_{m=1}^{L_1} \delta_m \Delta lnP^f_{R,t-m} + \sum_{n=1}^{L_2} \delta_n \Delta lnP^d_{F,t-n} + \varepsilon_{R,t} \tag{12}$$

$$\Delta lnP^d_{F,t} = \delta_1 + \delta_2^+ ECT^+_{t-1} + \delta_2^- ECT^-_{t-1} + \sum_{m=1}^{K_1} \alpha_m \Delta lnP^f_{R,t-m} + \sum_{n=1}^{K_2} \alpha_n \Delta lnP^d_{F,\ t-n} + \varepsilon_{F,t} \tag{13}$$

where L_1, L_2, K_1 and K_2 are the lag-lengths to be selected by the Schwarz information criteria (SC). The error correction terms ECT^+_{t-1} and ECT^+_{t-1} in Equations (12) and (13) in the case of the M-TAR model can be expressed as:

$$ECT^+_{t-1} = I_t(\Delta\mu_{t-1} \geq \tau)\left(ln\ P^f_{R,t-1} - \varphi - B_F ln\ P^d_{F,t-1}\right)$$

and

$$ECT^-_{t-1} = I_t(\Delta\mu_{t-1} < \tau)(ln\ P^f_{R,t-1} - \varphi - B_F ln\ P^d_{F,t-1})$$

Data

Retail prices were used as the prices in downstream markets and export price was used to represent the upstream market (the farm) price since the

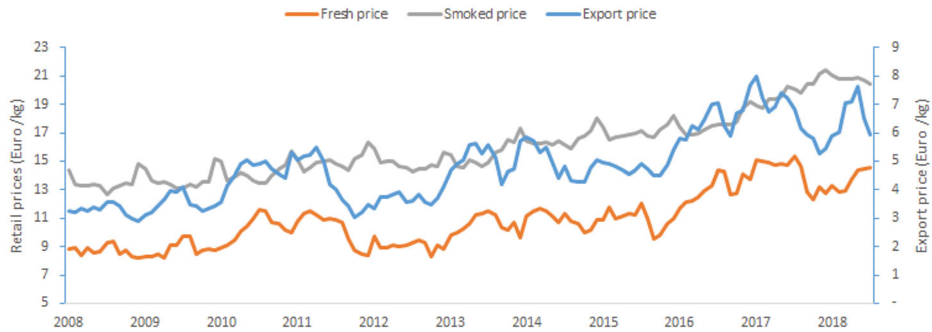

Figure 1. French retail fresh and smoked salmon prices & Norwegian export price. Source: Europanel and Statistics Norway.

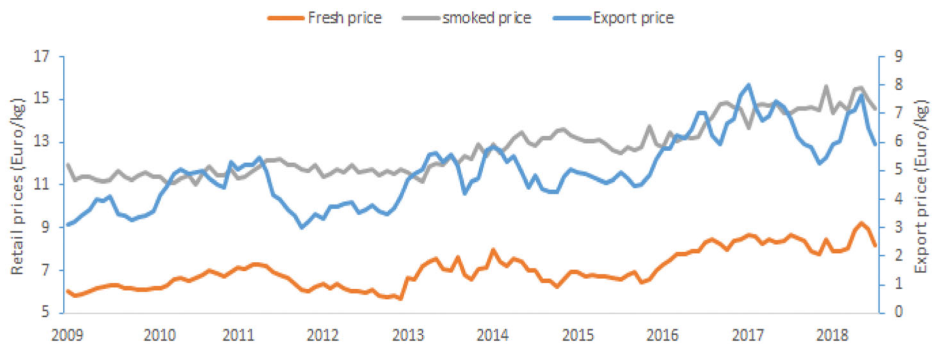

Figure 2. Spanish retail fresh and smoked salmon prices & Norwegian export price. Source: Europanel and Statistics Norway.

export price of the Norwegian whole fresh salmon is quite close to farm gate price with a price transmission elasticity of 1 (Asche et al., 2014). The sample period were January 2008 to December 2018 in the case of the French market, and January 2009 to December 2018 for the Spanish market; periods were selected based on the availability of the data. The retail salmon data was obtained from Europanel (2018) via the Norwegian Seafood Council (NSC) and the export prices were obtained from Statistics Norway (SSB) via NSC. Exchange rates were directly obtained from SSB. Table A1 in the appendix reports the summary of the prices.

Figure 1 illustrates the retail prices for fresh and smoked salmon in the French market together with the Norwegian export price. Figure 2 shows the Spanish market. The figures suggest that for fresh whole salmon, the retail price and export price follow each other closely in both the French and Spanish markets. In contrast, the retail price of smoked salmon and the export price of Norwegian fresh whole salmon behave differently in both markets. As a result, we expect the price transmission along the

Table 1. ADF test.

	France		Spain	
Prices	Constant and trend	First differences	Constant and trend	First differences
Export	−3.35 (1)	−9.286 (0)**	−3.09 (1)	−8.79 (0)**
Fresh salmon	−3.13 (1)	−11.78 (0)**	−2.69 (1)	−10.87 (0)**
Smoked salmon	−2.41 (12)	−1.9 (11)*	−3.28 (1)	−15.82 (0)**

Note. Asterisks ** and * denote significance at 1% and 5%, respectively. Price series are expressed in logarithm. Numbers of lags in ADF tests in parenthesis.

supply chain of fresh salmon and smoked salmon to show different relationships in each market.

Estimated results

The time series properties of the prices were investigated individually using the Augmented Dickey-Fuller (ADF) test, and in line with the general literature all processes were found to be non-stationary in levels but stationary in their first differences. Table 1 reports the results of the unit root tests.

The estimated residuals from Equation (7) in each product chain and market were estimated as a threshold model using both zero threshold ($\tau = 0$) and nonzero threshold ($\tau \neq 0$) values. To save space, we only report the results of the best-selected model based on the value of the Akaike information criterion (AIC). Table 2 reports the results of the selected models.[3]

The estimated statistics for cointegration (i.e., $\rho_1 = \rho_2 = 0$) suggest that the upstream and the downstream prices are cointegrated in the value chain of fresh salmon in both markets, but not cointegrated in the value chain of smoked salmon in either market. For fresh salmon, the estimated F-statistics for the symmetry test ($\rho_1 = \rho_2$) suggest the presence of asymmetric price transmission between the upstream and downstream prices in both markets. This implies that biased results may be obtained by reporting the equilibrium adjustment relationships between the upstream and downstream markets following the results from, for instance, the widely used Johansen cointegration approach in the literature, which assumes symmetric adjustment.

In general, the estimates of ρ_1 and ρ_2 were significantly different from zero and satisfied the conditions of convergence (i.e., $\rho_1 < 0$, $\rho_2 < 0$, and $(1 + \rho_1)(1 + \rho_2) < 1$). The estimate ρ_1 is the retail price adjustment when the retail price is "too high" with respect to the export price (i.e., when the margin is above its long-run equilibrium value), while the estimate ρ_2 is the adjustment when the retail price is "too low" with respect to the export price (i.e., when the margin is below its long-run equilibrium

Table 2. Consistent threshold cointegration and asymmetry test in France and Spain.

	ρ_1	ρ_2	γ_1	γ_2	τ	AIC	$\rho_1 = \rho_2 = 0^1$.	$\rho_1 = \rho_2{}^2$	Model
France									
Fresh	−0.35 (−3.86)	−0.84 (−5.65)	NA	NA	−0.037	−371.92	23.43	8.11 (0.005)	MTAR
Smoked	−0.16 (−2.56)	−0.24 (−3.52)	0.25 (2.81)	−0.16 (1.78)	−0.097	−268.61	5.52	0.948 (0.332)	TAR
Spain									
Fresh	−0.20 (−2.10)	−0.69 (−4.73)	−0.16 (−1.77)	NA	−0.04	−425.09	13.96	8.59 (0.004)	MTAR
Smoked	−0.25 (−2.49)	−0.10 (−1.25)	−0.02 (−0.21)	0.20 (2.14)	−0.07	−423.99	4.21	1.95 (0.165)	TAR

Notes. Numbers in brackets in columns from (1) to (4) are the t-statistics. [1] Entries in this column are the sample values of the M-TAR & TAR-statistics. [2] Entries in this column are the sample *F*-statistics for $\rho_1 = \rho_2$ & significance levels are in parentheses below. Enders and Siklos (2001) critical values for M-TAR for two variables and no lagged are approx. 5.45, 6.51 and 8.78 for 10%, 5% and 1%, respectively. Critical values of M-TAR for two variables and one lagged are approx. 5.47, 6.51 and 8.85 for 10%, 5% and 1%, respectively. Critical values of TAR for two variables and two lagged are approx. 5.80, 6.82, and 9.04 for 10%, 5% and 1%, respectively. τ represents the threshold value, and γ_1 and γ_2 are parameters included in the models to account for autocorrelation.

Table 3. Estimates of asymmetric error correction models in the fresh salmon value chain.

	France		Spain	
Independent variables	$\Delta lnP^f_{R,t}$	$\Delta lnP^d_{F,t}$	$\Delta lnP^f_{R,t}$	$\Delta lnP^d_{F,t}$
ECT^+_{t-1}	−0.43***	−0.13	−0.32***	−0.34
	(−4.91)	(−0.85)	(−3.11)	(−1.45)
ECT^-_{t-1}	−0.77***	0.09	−0.34**	0.71**
	(−6.16)	(0.39)	(−2.52)	(2.28)
$\Delta lnP^f_{R,t-1}$	−0.04	0.06	−0.24**	−0.25
	(−0.59)	(0.47)	(−2.24)	(−0.99)
$\Delta lnP^f_{R,t-2}$	NA	NA	−0.15*	−0.38*
			(−1.75)	(−1.89)
$\Delta lnP^d_{F,t-1}$	0.08	0.12	0.19***	0.16
	(1.15)	(0.95)	(3.52)	(1.24)
$\Delta lnP^d_{F,t-2}$	NA	NA	0.05	0.14
			(0.927)	(1.09)
R^2	0.47	0.04	0.43	0.14
Breusch–Godfrey serial correlation test				
Lag 1	0.016	0.008	0.24	0.002
	(0.898)	(0.928)	(0.622)	(0.967)
Lag 2	0.25	0.149	1.832	1.96
	(0.882)	(0.928)	(0.400)	(0.376)
Lag 3	0.391	1.35	1.99	2.18
	(0.942)	(0.717)	(0.575)	(0.537)
Lag 4	0.673	2.35	2.09	3.16
	(0.955)	(0.672)	(0.719)	(0.532)
Lag 5	5.26	4.18	2.78	3.16
	(0.385)	(0.524)	(0.734)	(0.676)
Lag 10	17.67	13.20	10.49	10.01
	(0.162)	(0.213)	(0.398)	(0.439)
Breusch–Pagan test	3.79	2.88	2.54	3.79
	(0.285)	(0.411)	(0.77)	(0.581)
Jarque–Bera test	5.84	0.11	10.65	0.51
	(0.054)	(0.94)	(0.01)	(0.78)

Notes. Asterisks ***, ** and * denote significance level at 1%, 5%, and 10%, respectively. Numbers in parentheses under coefficients are *t*-values, while significance levels under tests.

value). The result that the estimated magnitudes of ρ_1 were smaller overall than those of ρ_2 suggests that retail prices react more rapidly when the margin is squeezed than when it is stretched.

The estimated threshold values were about −0.04 for the fresh salmon value chain and on average −0.084 for the smoked salmon value chain in both markets. The negative threshold value means that a new adjustment takes place after a substantial reduction of the margin (Simioni et al., 2013). The approximately equal threshold values in the two countries might indicate the pricing strategies of retailers in the two countries are similar for the same product form. The fact that the magnitudes of threshold values were greater for smoked salmon than for fresh salmon indicates market response takes a longer time for those products that need further processing before they are sold in the retail market.

Next, we estimated the threshold asymmetric error correction model (ECM) along the fresh salmon value chains in both countries. Estimates of the ECM (Equations [12] and [13]) are reported in Table 3. In the table, consistent with our previous notations, $P^f_{R,t}$ denotes the retail level price while $P^d_{F,t}$ denotes the export level price. The estimated models were checked using various diagnostic tests (such as the Breusch-Godfrey serial Correlation test, Breusch-Pagan test, and Jarque-Bera test) and the results of the tests (reported in the bottom section of Table 3) confirmed the absence of any major misspecification problems except some issues of normality.

In the French market, the estimated parameters of the coefficient ECT_{t-1}^{+} and ECT_{t-1}^{-} were statistically significant for the retail equation but not for the export equation. This shows that the export prices do not adjust to changes in the retail prices, while on the contrary, the retail prices adjust to changes in the export prices following a deviation from the equilibrium. This suggests the leading role of the export price in the value chain of the fresh whole salmon market, and is a common result in the literature (Asche, Jaffry, et al., 2007).

Furthermore, the adjustment of the retail price to changes in the export price is faster when the deviation from the equilibrium is below the threshold. Specifically, the error correction rate of the retail prices in one month period following a shock in the export price is 43% and 77% for above and below the threshold value of -0.037, respectively. That is, the retail prices react more quickly when the margin is squeezed than when it is stretched. More specifically, the results suggest that when the margin is low for retailers; retailers will make a fast adjustment of their price according to the export price, while if the margin is high, the adjustment speed is slower. Retailers usually prefer a relatively stable price in their market to avoid transaction costs and possibly to exploit market power.[4]

In the Spanish market, as in the French market, the adjustment parameters ECT_{t-1}^{+} and ECT_{t-1}^{-} are statistically significant for the retail model. This implies the retail prices adjust to changes in the export prices following a deviation from the equilibrium. However, in contrast to the French market, in the export model of the Spanish market; the adjustment parameter ECT_{t-1}^{-} is statistically significant. This shows that the export prices also adjust to changes in the retail prices following a deviation from the equilibrium. However, the export prices adjust only if the deviation from the equilibrium is below the critical threshold.

Looking at the estimates of the parameters of the short-run coefficients (i.e., $\Delta lnP_{R,t-m}^{f}$ and $\Delta lnP_{F,t-1}^{d}$), in the Spanish market, a significant cross-price effect that goes from the retail price to the export price is observed at two lagged, while a significant cross-price effect that goes from the export to the retail price is detected at one period lagged. This further supports the bi-directional adjustment between the prices in Spain. No short-run cross-price effects were observed in the case of France.

Conclusion and discussion

The nature of price adjustment along the seafood value chain has gained substantial attention in recent years. In this study, threshold cointegration and asymmetric error correction models (ECMs) were estimated. The findings of asymmetric price transmission given by threshold cointegration suggest possible biased results given by the conventional Johansen approach in the previous literature. However, this bias appears to be small as our results where applicable corroborates the results from earlier price transmission studies for salmon.

Our results suggest that price transmission is more complete and quicker in the value chain of products that are less processed than the value chain of more processed products. Specifically, we found price transmission between the marketing chains along the fresh salmon chain in both the French and Spanish markets, but no price transmission was detected in the value chain of smoked salmon in either market.

Further investigation is needed to identify the reasons for the lack of price transmission in the value chain of smoked salmon. There are four main factors that may contribute to this phenomenon. First, if we look at Figures 1 and 2, there is evidence that the retail price of the smoked salmon is just a markup of the raw fish price after including processing costs. Second, there is a substantial time lag between the import of raw fish and retail sales of the final smoked salmon in the markets. For different processing companies, the period of the time lag varies significantly in relation to their production capacities and marketing strategies. Third, given the higher degree of processing for smoked salmon, the cost share of the raw fish is smaller, a feature that generally will tend to make the price transmission weaker. Finally, there appears to be a higher use of contracts where prices are fixed for longer periods for more processed products, making price transmission very slow as prices are infrequently updated. Consequently, compared to the fresh salmon market, it is more difficult to identify the price linkage between imported raw fish and smoked salmon in a retail market.

Asymmetric price transmission in the value chain of fresh salmon was detected, where agents reacted more quickly to shocks when the deviation from equilibrium was below the equilibrium level. Approximately equal threshold values in both the French and Spanish markets indicated that the marketing strategies of retailers in the two markets are somehow similar. The French market results suggest the leading role of the export price in the value chain of the fresh whole salmon market since it was the retail prices that adjusted to changes in the export prices, not vice versa. However, in the Spanish market, bi-directional adjustment was observed.

Noncompetitive behavior and adjustment costs (or transaction methods) are the two most widely cited causes of asymmetry in the price transmission literature (Frey & Manera, 2007; von Cramon-Taubadel & Meyer, 2004). However, since the global salmon market is assumed to be competitive (e.g., Larsen & Kinnucan, 2009), we argue that adjustment costs (transaction methods) are the most likely explanations for the price asymmetries along the salmon value chain.

Notes

1. Futures markets are an alternative mechanism for hedging short-and intermediate-term price risk (Asche et al., 2016; Oglend & Straume, 2020).
2. However, Larsen and Kinnucan (2009) include transportation costs measured in an index of retail auto diesel prices.
3. Results of all estimated models are available upon request from the corresponding author.
4. This is an issue that has received limited attention. However, Sogn-Gruntvåg et al. (2019) show significant differences in product longevity by product form and labelling for whitefish.

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APPENDIX

Table A1. Summary statistics of monthly retail prices from France and Spain and farm (or export) price from Norway (price in levels).

	France				Spain			
	Mean	Median	St. Dev.	No. of obs.	Mean	Median	St. Dev	No. of obs.
Export	4.83	4.79	1.24	127	4.99	4.89	1.196	115
Fresh	10.86	10.69	1.93	127	7.06	6.92	0.864	115
Smoked	16.05	15.61	2.28	127	12.67	12.36	1.26	115

Notes. Data for France are from January 2008 to June 2018, while data for Spain are from January 2009 to June 2018.

Market integration and price transmission in the regional grain markets in Ethiopia

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Abstract

Persistent increases in basic food prices have become a critical challenge in Ethiopia since 2006. This paper assesses whether the structure of the grain market has contributed to the price increases. Traders having market power could create commodity price stickiness, implying that what goes up does not come down, leading to price increases. The study examines price linkages between principal grain markets in Ethiopia, using monthly prices from the wheat, maize, and teff markets. The Engle-Granger cointegration test is used to check for cointegration, while the Threshold Aurogressive Model is employed to investigate potential asymmetric price transmission. The findings indicate that major grain markets in Ethiopia are well integrated. Moreover, the threshold cointegration model reveals that they are characterized by symmetric adjustment, implying positive and negative price shocks in the central market are equally transmitted to local markets. Hence, I argue that there is insufficient evidence to support the claim that market structure contributes to the price increase in Ethiopian grain markets.

Keywords: Asymmetry, spatial grain prices, Ethiopia, price transmission, threshold cointegration

1. Introduction

In Ethiopia, persistent increases in basic food grain prices have become a critical challenge since 2006. For instance, as illustrated in Figure 1 below, the 2020 national average wholesale prices of both wheat and maize have increased by about four times compared to their 2010 levels. Moreover, the rate of growth of food price inflation in Ethiopia has been among the highest in sub-Saharan Africa (SSA). For instance, according to estimates from the World Bank, the average annual consumer price inflation in Ethiopia grew from 8% to 16% between 2010 and 2020, while generally, it declined from 4.0% to 3.3% in the SSA region over the same period. Understanding what explains the persistent increase in food prices in Ethiopia might contribute to designing targeted government interventions that stabilize the market.

According to the literature, food prices in any market may change as a result of various factors including, price-shock diffusion from international markets, changes in the exchange rate and the domestic demand and supply conditions, and market structure. If a country is a net importer of food commodities, high world-price-shock transmission to the domestic market may affect commodity prices (Baquedano & Liefert, 2014; Ceballos et al., 2017; Minot, 2011; Conforti, 2004). However, the extent and speed of price shock diffusion from international to domestic market depends on several factors, including the exchange rate, border policies (such as import tariffs), and transfer costs (Hazel et al., 1990; Dawe, 2008). For instance, when a local currency appreciates (depreciates) against the US dollar, an increase in the commodity price in the local currency would be less (more) than an increase in the international price in dollars (e.g., Ozturk, 2020).

Domestic demand and supply dynamics are also important. In a country with a self-sufficient position, domestic demand, and supply factors (such as income growth, population growth, weather shocks, and production input costs) determine commodity price formation and stability (Brækkan et al., 2018). The other important factor in commodity price formation, and what motivated this study, is market structure (Abdulai, 2000). In a non-competitive market structure, middlemen with high market power may dominate the pricing of the commodities along the supply chain or across spatially distinct markets. In such circumstances, price increases in some markets may completely and quickly transmit to other markets, whereas price reduction might remain sticky, implying, what goes up does not come down, hence leading to price increases.

In Ethiopia, a profound agricultural market liberalization process started in the early 1990s, when all the restrictions on official prices, quotas, and private trade were removed (Gabre-Madhin & Goggin, 2005). The reform has massively changed the structure of the grain market in the country (Shahidur &Asfaw, 2011). After the reform, private sector participation in grain marketing activities has increased, and this contributed to an increase in market integration and grain production (Kindie, 2008; Negassa & Meyer, 2007). On the other side, however, the reform created a grain market structure where market power is concentrated around a few dominant firms because of a shortage of initial capital, which is the major obstacle to entering the market (Gebremeskel et al., 1998; Sassi & Mamo, 2019). This may create a substantial impact on the price transmission process, implying price shocks in some markets affect price behavior in others (e.g., von Cramon-Taubadel & Meyer, 2004). In fact, there is a growing perception in the Ethiopian grain market that once prices have started to increase, they never come down, even during the harvest seasons when producer prices are falling (Tadesse & Guttormsen, 2011; Sassi & Mamo, 2019).

Often economists who assess overall market performance investigate price transmission mechanisms in the market. In perfectly integrated markets, price changes in one market completely and quickly transmit to another market (Fackler & Goodwin, 2001). Moreover, in well-integrated markets, price shocks in one market would elicit the same response in other markets, regardless of whether the shock reflected a decrease or an increase in prices (e.g., von Cramon-Taubadel & Meyer, 2004). Such a price adjustment is called symmetric price transmission. However, as documented in the literature on price transmission, certain characteristics associated with imperfect competition (e.g., market concentration, government intervention), transaction cost, and inventory behavior of traders can contribute to asymmetric price responses (Abdulia, 2000). Moreover, the presence of transaction costs may prevent economic agents from continually adjusting. Only when the deviation from the equilibrium surpasses a critical threshold do the benefits of adjustment exceed the costs, and economic agents act to move the system back to equilibrium (Abdulia, 2000). The above implies that while investigating price transmission, it is relevant to consider price asymmetry and/or nonlinearity.

Previous price transmission studies in the Ethiopian grain market include explorations of price transmission along the vertical chain (Sassi & Mamo, 2019; Usman & Haile, 2017) and across spatially distinct markets (Yami, 2020; Kifle, 2015; Negassa & Meyer, 2007; Getnet et al., 2007; Getnet et al. 2005). While these available studies might be sufficient to illustrate the presence or lack of effective price transmission in the Ethiopian grain market, they generally have two shortcomings.

First, the data used by most of these studies are old. Thus, a new study using the most up-to-date data may help us understand the causes of the price increases observed in the Ethiopian grain market in recent years. Second, most of the studies mentioned do not consider price asymmetry and/or nonlinearity in their analysis. Results obtained without considering price asymmetry and/or nonlinearity might be biased when an asymmetric response is misspecified as symmetric (e.g., Barrett & Li, 2002). Moreover, while three studies (i.e., Usman & Haile, 2017; Sassi & Mamo, 2019; Kifle, 2015) investigate price asymmetries, results from these are inconclusive due to the diversity of the good analyzed, methodology, and time periods considered.

The purpose of this article is thus to investigate the possible existence of asymmetric price transmission between major regional grain markets in Ethiopia. Asymmetric price transmission is investigated in the three important grain markets: wheat, maize, and teff. The main method used is the threshold cointegration model with both zero and non-zero thresholds. The threshold cointegration model is used because it is suitable for investigating the presence of any asymmetric price transmission between distinct markets across space.

The remainder of this paper is organized as follows. Section 2 is the literature review. Section 3 covers my methodology and the data used in the study is presented in Section 4. The empirical results are contained in section 5 and the concluding remarks in Section 6.

2. Literature review

This section presents theoretical concepts of market integration and spatial price transmission, followed by a general overview of grain markets in Ethiopia, and finally, a review of previous empirical evidence concerning these markets.

2.1. Market integration, spatial price transmission, and asymmetric price transmission

Economic theory postulates that the proper functioning of markets and marketing channels is vital for the optimal allocation of resources (Abdulia, 2000). Price transmission has become a common tool used to assess the proper functioning (or integration) of spatially separated markets. According to the well-known Law of One Price (LOP), under free trade, price differences of a homogeneous good in two distinct markets separated in space will be, at most, equal to the transaction costs involved in transferring the goods from one market to the another (Fackler & Goodwin, 2001; Serra et al., 2006a). If the price spreads between the markets exceed the transaction costs, the activity of profit-seeking arbitrageurs will reduce the spread, allowing prices to move toward the LOP condition.

However, certain characteristics of agricultural production, marketing, and consumption, such as inadequate infrastructure, market-entry barriers, and unreliable market and price information, may render arbitrage a risky activity for traders (Abdulia, 2000). In such a circumstance, spatial markets may be partially integrated or completely segmented. When spatially separated markets are not well integrated, profitability opportunities will not be fully exploited by spatial arbitrageurs, thus resulting in efficiency losses (e.g., Fackler & Goodwin, 2001). Inefficient price signal transmission between markets can distort producer decisions and lead to inefficient product movement, and ultimately increase consumer prices (Goodwin & Schroeder, 1991). Hence, assessing whether or not spatial price transmission is efficient is highly relevant and has policy implications.

Spatial price transmission is characterized by the speed, extent, and nature of price signal transmission between markets. In the case of perfectly integrated markets, price changes in one market are completely and immediately transmitted to other markets. However, when markets are not well integrated, price transmission will be incomplete, and prices will take time to adjust. This might reflect inefficiency and welfare losses in the economic system. Another concern that has driven the interest of price analysts when dealing with markets' responses to one another is whether they adjust symmetrically or asymmetrically. If a shock to a market (for instance, to the central market) would elicit the same magnitude of response in local markets, regardless of whether the shock reflected an increase or decrease in prices, the transmission is called symmetric; otherwise, it is called asymmetric (von Cramon-Taubadel & Meyer, 2004; Serra et al., 2006). Von Cramon-Taubadel & Meyer (2004) indicate that symmetric adjustments are often assumed to be representative of competitive markets, whereas asymmetric price responses are linked with the existence of certain market imperfections (e.g., market power) that cause rational market participants to deviate from their preferred risk level. Moreover, the presence of asymmetry in price transmission implies potential welfare loss for some groups of market participants; welfare distribution could be different under asymmetry (von Cramon-Taubadel & Meyer, 2004).

As documented in the literature on price transmission, several factors can contribute to asymmetric price responses. Some of the most common causes include imperfect competition; adjustment, search, and menu costs; government intervention; and inefficient inventory management (von Cramon-Taubadel & Meyer, 2004). Non-competitive behavior (or market power) is the main cause of asymmetric price transmission identified in the literature (von Cramon-Taubadel & Meyer, 2004; Sexton et al., 1991). Oligopolistic intermediaries in spatial trade may act more quickly to shocks that squeeze their profit margin than to those that stretch their margin. Such marketing behavior can lead to asymmetric price transmission in the short run (Abdulai, 2000). In this case, increases in central market prices may more quickly and completely transmit to the local markets than will a corresponding decrease in prices. Asymmetric price transmission in the spatial market may also occur if local-market traders assume that competitors in the local market will follow a price rise but not a price decrease in the central market (Abdulai, 2000).

Adjustment costs are the second important cause of asymmetric price transmission (Abdulia, 2000). Adjustment (or transaction) costs are those incurred due to changing market conditions; these arise when firms change the quantities and/or outputs. If the cost of adjustment is different with respect to cost increases and decreases, asymmetric price transmission can occur (Abdulia, 2000). Search costs are another cause of asymmetric price transmission (von Cramon-Taubadel & Meyer, 2004). These are costs incurred by consumers when searching for market information about a product they are interested in buying; customers may have incomplete information about a particular market since obtaining information is costly. Thus, a firm in a local market may possess local market power over their customers, who may have insufficient information on what other firms in other local markets charge for the same product.

Inventory management can also lead to asymmetric price transmission. For storable commodities such as grains, farmers (particularly in developing countries) may lack the storage and capital needed to get their goods to distant markets. They are thus left selling locally to intermediaries who now have more suppliers from which to choose (Abdulai, 2000). Moreover, if firms think the prices in the central market will increase, they will hold their products, and if they think the prices in the central markets will decrease, they sell their products (Abdulai, 2000). Finally, government policy intervention can also lead to asymmetric price transmission (von Cramon-Taubadel & Meyer, 2004; Kinnucan & Forker, 1987). If the government intervenes in the market, for instance, by establishing price floors, firms may assume that the increase is permanent and respond asymmetrically.

2.2. An overview of grain market in Ethiopia and previous empirical evidence

Grain is a staple food crop in Ethiopia, accounting for more than 40% of a typical household's food expenditure, more than 60% of total caloric intake, and about 73% of employment (Rashid, 2010). Grain production in Ethiopia is highly specialized due to the heterogeneous nature of the country's agro-ecologies. Wheat, maize, and teff are the most-produced grains within the country.

In terms of production regions, the main wheat production zones (provinces) are Arsi and Bale (Negassa & Meyer, 2007), located southeast of the capital city of Ethiopia (Addis Ababa). The major maize-producing zones are west Gojjam (in the north), Jimma (in the west), East Shewa (in the center), and East Wellega (in the west). Teff production is mainly concentrated in the center and the northwest of the country. In terms of consumption, wheat is preferred in three regions, Afar (in the northeast), Tigray, and Amhara (in the north), whereas teff is the preferred commodity in Addis Ababa, Amhara, and Tigray. Maize is preferred in the center, eastern, and southern parts of the country.

Differences in grain production and consumption regions in Ethiopia make spatial domestic-grain trade highly relevant. In general, grain trade in Ethiopia has a radial structure, where grain trades typically flow from the surplus areas to Addis Ababa either for consumption or transshipment to deficit areas (Negassa & Myers, 2007).[1] While minor retailers and small wholesalers dominate the local and regional markets, larger firms engage in spatial arbitrage. However, these firms are relatively few due to a shortage of capital, creating a barrier to entry (Osborne, 2005; Gebremeskel et al., 1998). This, coupled with the poor market infrastructure in the country, makes it highly likely that the country's spatial grain markets are inefficient.

In the empirical literature on price transmission in the Ethiopian grain market mentioned earlier, there are only three studies that consider asymmetric price transmission. Kifle (2015) investigates asymmetry in the prices of white teff, red teff, and maize between the central market (Addis Ababa) and two local deficit markets (Mekelle and Dire Dawa). Usman and Haile (2017) find that symmetric adjustment generally characterizes the price transmission along the supply chains of teff, wheat, and maize in Amhara and Oromia, the two major cereal markets in Ethiopia. Sassi and Mamo (2019) identify asymmetric price transmission in the white teff market between producers and the wholesale market in Oromia and Tigray regions but identify symmetric price adjustment in the Amhara region and the Southern Nations, Nationalities, and People's region. The remaining two studies, in addition to Kifle (2015), focus on price transmission asymmetries along grain supply chains. In this study, I focus on spatial asymmetric price transmission in three important grain markets (wheat, maize, and teff) using more recent data and the threshold cointegration model proposed by Enders and Siklos (2001) and discussed below.

3. Methodology

The empirical procedure followed in this study comprises a series of tests and model estimations. First, stationarity of prices is confirmed using the Augmented Dickey-Fuller (ADF) (Dickey & Fuller, 1981) and Perron (1997) tests on individual price series. The Engle-Granger cointegration test is then employed to check the presence of linear long-run

[1] However, there are also direct trade flows from the local surplus areas to the deficit areas without passing through Addis Ababa.

relationships between the central and local grain-market prices. Then, the threshold cointegration model is estimated to investigate the presence of possible asymmetric price adjustment between central- and local-market prices. For price pairs demonstrating asymmetric adjustment, the threshold asymmetric error correction model is estimated; otherwise, the traditional error correction model is estimated.

In a system of spatially related markets, as discussed in Asche et al. (2012), there is sometimes a leading market that has a dominant influence on the other markets. The price changes in the leading market can affect the prices and quantities in the other markets, but the opposite is not the case. Based on the characteristics of the Ethiopian grain markets discussed above, the market in the capital Addis Ababa can be taken as the central market. Given the central wholesale price (${p_t}^c$) and local market price (${p_t}^l$) (expressed in log form), the basic price transmission model can be expressed as follows:

$$ {p_t}^l = \alpha + \beta\, {p_t}^c + \mu_t \qquad (1) $$

where μ_t is the error term, which is assumed to be independent and identically distributed with mean zero. Parameters α and β define the relationship between the prices or whether the markets are integrated. If $\beta = 1$, the LOP holds (i.e., there is complete price transmission), while $\beta = 0$ implies there is no relationship between the prices.[2] If $0 < \beta < 1$, there is a relationship between the prices, but their transmission is not complete.

Using Equation (1) for price transmission analysis raises two major conceptual and practical concerns. First, Equation (1) is a static model. However, price adjustment is a dynamic process, and hence temporary deviations from the long-run equilibrium are inevitable. Second, agricultural product prices appear to be non-stationary. The application of Equation (1) to a non-stationary price series may generate spurious results; estimated models will indicate a relationship between the prices when in fact, no theoretical relationship exists. A cointegration test provides a means to distinguish a true relationship from one that is spurious.[3]

[2] As it is common in most price transmission literature, transportation costs and quality differences are treated as constant in this study. Nevertheless, if transportation costs are not constant, this assumption may cause rejections of the LOP.

[3] Cointegration implies if the prices are non-stationary of same order (i.e., order one, denoted, $I(1)$), only a stationary linear combination of them represents the true price relationship. Specifically, the prices in Equation (1) are said to be cointegrated if the error term, μ_t, is stationarity (i.e., $I(0)$).

There are two commonly used approaches to investigating a cointegrating relationship between prices: the Engle-Granger approach (Engle & Granger, 1987) and the Johansen cointegration approach (Johansen & Juselius, 1990). In this study, we use the Engle-Granger cointegration approach, of which the threshold cointegration model is an extension. The Engle-Granger approach involves extraction of the estimated residuals μ_t from the OLS regression in Equation (1) to estimate the parameter ρ in Equation (2):

$$\Delta\mu_t = \rho\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-i} + \upsilon_t \tag{2}$$

where υ_t is a white-noise disturbance, and γ_i are parameters included in the model to account for serial correlation. According to the Engle-Granger approach, rejection of the null hypothesis of no cointegration ($\rho = 0$) implies stationarity of the estimated residuals (i.e., μ_t), and hence cointegration of the prices.

The Engle-Granger model given in Equation (2) relies on an assumption of symmetric (or linear) price adjustment since ρ is estimated as an average effect of the lagged error term μ_{t-1} regardless of whether μ_{t-1} is positive or negative. Taking asymmetry into consideration, Enders and Siklos (2001) extend the Engle-Granger procedure to a Threshold Autoregressive (TAR) model given by Equation (3):

$$\Delta\mu_t = I_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta\mu_{t-i} + \upsilon_t \tag{3}$$

where I_t is an indicator function that can be defined by either Equation (4) or Equation (5) as:

$$I_t = \begin{cases} 1 \ if \ \mu_{t-1} \geq \tau \\ 0 \ if \ \mu_{t-1} < \tau \end{cases} \tag{4}$$

$$I_t = \begin{cases} 1 \ if \ \Delta\mu_{t-1} \geq \tau \\ 0 \ if \ \Delta\mu_{t-1} < \tau \end{cases} \tag{5}$$

where τ is a threshold value, and ρ_1 and ρ_2 are the parameters to be estimated. If the indicator function is defined by Equation (4), the model in Equation (3) is said to be a TAR model, whereas if the indicator function is defined by Equation (5), the model in Equation (3) is known as a Momentum-Threshold Autoregressive (MTAR) model. In the literature, if $\tau \neq 0$, the TAR model is said to be consistent TAR, whereas the MTAR model is called consistent MTAR.

In Equation (3), the parameter ρ_1 captures the local market price adjustment when the local market price is “too high” with respect to the central market price (i.e., when the price differential is above its long-run equilibrium value), while the parameter ρ_2 is the adjustment when the local market price is “too low” with respect to the central market price (i.e., when the price differential is below its long-run equilibrium value).

As indicated by Equations (2) and (3), the Engle-Granger procedure is a special case of the Enders and Siklos’s (2001) model when $\rho_1 = \rho_2$. From Equation (3), cointegration exists if the null hypothesis $\rho_1 = \rho_2 = 0$ is rejected. This leads to the test of symmetry $\rho_1 = \rho_2$. If both hypotheses are rejected, the process is called threshold cointegration. According to Enders and Siklos (2001), the cointegration test follows a non-standard F-test; hence, results are compared to the critical values given in Enders and Siklos (2001). The symmetry test, however, is a standard F-test. Following the procedure proposed by Chan (1993), the best threshold value is used. The estimated residuals (in the case of the TAR model) or the first differences of the estimated residuals (in the case of the MTAR model) are first sorted in ascending order, then 15% of the largest and smallest values are excluded before selecting the best threshold value, that is, the value that yields the lowest residual sum of squares.

The Granger representation theorem (Engle& Granger, 1987) states that an error correction model can be estimated when variables are cointegrated. The standard vector error correction model (VECM) can be written as:

$$\Delta p_t{}^l = \beta_0 + \beta_1 ECT_{t-1} + \sum_{i=1}^{p} \delta_i \Delta p^l{}_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta p^c{}_{t-j} + e_{1,t} \tag{6}$$

$$\Delta p_t{}^c = \eta_0 + \eta_2 ECT_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta p^l{}_{t-j} + \sum_{i=1}^{q} \delta_i \Delta\, p^c{}_{t-i} + e_{2,t} \tag{7}$$

where $ECT_{t-1} = p^l{}_{t-1} - \alpha - \beta\, p^c{}_{t-1}$ is the error correction term and the coefficients β_1 and η_2 capture the speed of error correction. Parameters β_0 and η_0 are intercepts, γ_j capture short-run price dynamics, and δ_i are added to capture serial correlation. The parameters p and q are optimal lag lengths to be chosen using the Schwarz information criteria (SC).

However, a conventional error correction model as given by Equations (6) and (7) cannot be used to consider the issue of asymmetric transmission since the error term has not been decomposed into positive and negative components. In this study, by following Balke and Fomby (1997) and Enders and Granger (1998), the conventional VECM is extended to a threshold VECM, which is then capable of analyzing the asymmetric price transmission by decomposing the error correction terms into positive and negative components, as presented by Equations (8) and (9).

$$\Delta p^{l}{}_{t} = \beta_0 + \beta_1^{+} ECT_{t-1}^{+} + \beta_1^{-} ECT_{t-1}^{-} + \sum_{i=1}^{p} \delta_i \Delta p^{l}{}_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta p^{c}{}_{t-j} + e_{1,t} \tag{8}$$

$$\Delta p_t{}^{c} = \eta_0 + \eta_1^{+} ECT_{t-1}^{+} + \eta_1^{-} ECT_{t-1}^{-} + \sum_{j=1}^{p} \gamma_j \Delta p^{l}{}_{t-j} + \sum_{i=1}^{q} \delta_i \Delta p^{c}{}_{t-i} + e_{2,t} \tag{9}$$

where the error correction terms ECT_{t-1}^{+} and ECT_{t-1}^{-} are defined as follows:

$$ECT_{t-1}^{+} = I_t(p^{l}{}_{t-1} - \alpha - \beta\, p^{c}{}_{t-1})$$

$$ECT_{t-1}^{-} = (1 - I_t)(p^{l}{}_{t-1} - \alpha - \beta\, p^{c}{}_{t-1})$$

where I_t is either the TAR or MTAR indicator function with the consistent threshold.

4. Data

The price series used in this study were obtained from the Food and Agriculture Organization Global Information and Early Warning System database.[4] Based on data availability, monthly data covering the period from January 2010 to October 2020 were used for the wheat and maize markets, whereas monthly data that ranges from January 2010 to July 2018 were used for the teff market.[5] Prices are measured in local currency per 100 kilograms. All the prices are deflated using the Ethiopian consumer price index (obtained from the International Monetary Fund database) to account for inflation.

Regional markets with the most populous towns are selected for this study, and the extent and nature of price transmission between the central market (i.e., Addis Ababa) and local markets are then considered. Based on data availability, five market pairs are considered for the wheat market. Three of these pairs represent trade flows to Addis Ababa from the

[4] The FAO GIEWS collects monthly grain prices from the national grain board of several countries to support its technical activities. The FAO source in the case of Ethiopia is the Ethiopian Grain Trade Enterprise (EGTE).

[5] There are three types of teff in Ethiopia, white, mixed, and red. White teff and mixed teff are the types widely consumed in Ethiopia. Although both have the same purpose, white teff has superior quality, and is preferred by consumers with relatively high purchasing power, whereas mixed teff is consumed mainly by consumers who have low purchasing power. In this study, mixed teff markets are considered since no previous studies have done so.

wheat surplus regions surrounding Robe and Shashemene (located south of Addis Ababa) and Debre Markos (to the northwest). Two pairs represent trade flows from Addis Ababa to deficit areas surrounding Dire Dawa (in the east) and Jimma (in the west). Four market pairs are investigated for the teff market. The first pair represents trade flows from the teff surplus region surrounding Bahirdar (in the northwest) to Addis Ababa. The other three pairs represent trade flows from Addis Ababa to the teff deficit areas surrounding Mekelle (in the north) and Shashemene and Jimma. In the maize market, three price pairs are investigated.

Table 1 below presents summary statistics of the prices considered in this study. As Table 1 indicates, there is, in general, no noticeable difference in volatility between the markets in each of the commodity markets, measured with the standard deviation and coefficient of variation. Figure 1 illustrates the national trends for the nominal monthly average prices for wheat, teff, and maize at the wholesale level. Figure 1 indicates that each commodity prices show an increasing trend.

Table 1. Descriptive statistics of monthly real grain prices, Birr/kg (prices in log form)

Commodity	Mean	Min	Max	St. Dev.	CV[1]	Obs.
Wheat						
Addis	1.54	1.23	1.84	0.125	0.081	130
Robe	1.39	1.03	1.76	0.147	0.106	130
Dire Dawa	1.62	1.40	1.86	0.093	0.056	130
Debre Markos	1.44	0.86	1.73	0.167	0.115	130
Shashemene	1.48	1.06	1.86	0.149	0.101	130
Jimma	1.60	1.13	1.89	0.116	0.073	130
Teff						
Addis	1.93	1.59	2.12	0.09	0.051	103
Bahirdar	1.98	1.74	2.21	0.116	0.059	103
Mekelle	1.87	1.61	2.11	0.114	0.061	103
Shashemene	1.99	1.69	2.26	0.130	0.065	103
Jimma	1.89	1.65	2.15	0.106	0.056	103
Maize						
Addis	1.01	0.71	1.50	0.171	0.171	130
Dire Dawa	1.11	0.64	1.59	0.182	0.164	130
Mekelle	1.07	0.73	1.45	0.171	0.160	130
Bahirdar	1.01	0.53	1.44	0.193	0.191	130

Notes: Data for wheat and maize are from January 2010 to October 2020, while data for teff are from January 2010 to July 2018. CV[1] denotes the coefficient of variation, which measures relative variability or dispersion around the mean.

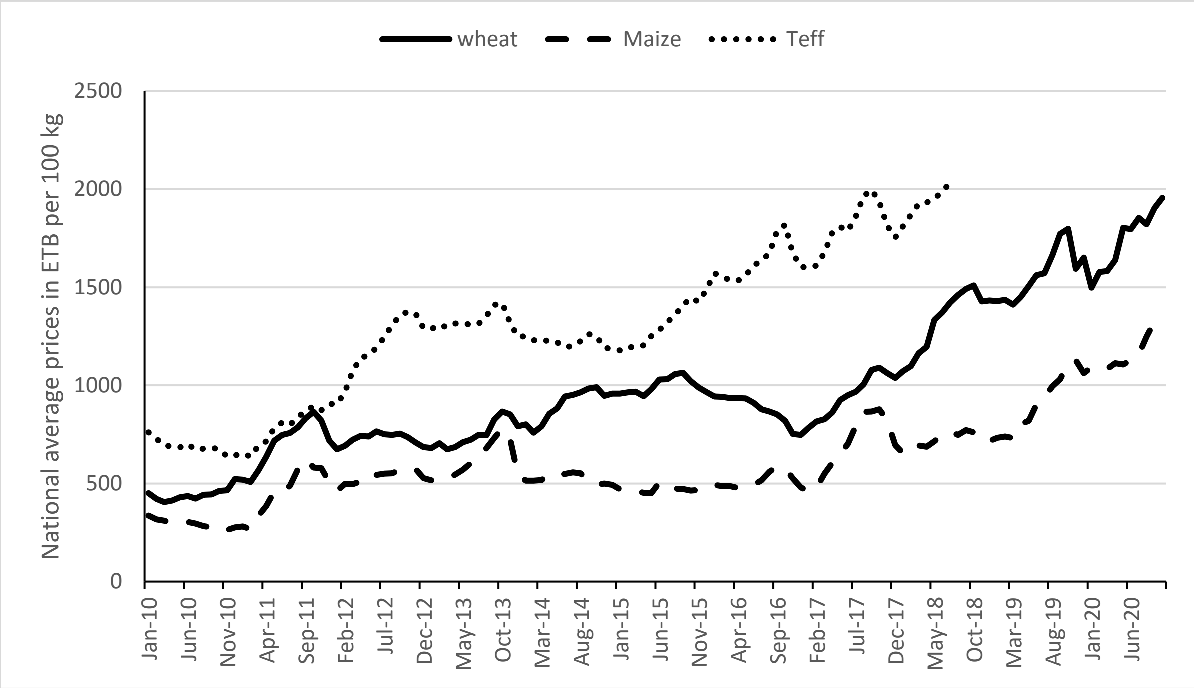

Figure 1. National nominal average monthly commodity price trends in Ethiopia (Jan 2010 – Oct 2020).

5. Empirical results

The time-series properties of the prices were investigated individually using the ADF and Perron tests. The test statistics indicated that the prices are non-stationary in levels but stationary in their first differences.[6] Hence, a cointegration test can be applied to determine the presence of long-run relationships between the central and local market prices in the Ethiopian grain market. Table 2 reports the results of the Engle-Granger cointegration tests, including the results of the LOP tests.

The Engle-Granger cointegration test results (reported in column 3 of Table 2) rejected the null hypothesis of no cointegration between each pair of prices investigated in each commodity market, suggesting the presence of a linear long-run relationship between the central and local market prices. However, the Johansen likelihood test results (reported in column 4 of Table 2) rejected the LOP, indicating the presence of incomplete price transmission between the central and local grain markets in Ethiopia. The incomplete price transmission might be resulted from inefficient arbitrage due to market power in the grain

[6] For brevity, the unit root test results are not reported here but can be obtained from the author upon request.

markets or the presence of high transportation costs due to poor market infrastructure in Ethiopia.

Table 2. Engle-Granger (E-G) cointegration test results ($p_t{}^l = \alpha + \beta\, p_t{}^c + \mu_t$)

	α	β	E-G test on the residuals	LOP [1]
Wheat				
Addis & Robe	−0.07 (−0.788)	0.95 (15.519)	−4.737**	15.52(0.000)
Addis & Dire Dawa	0.81 (11.23)	0.52 (11.12)	−4.939**	25.19(0.00)
Addis & Debre Markos	−0.15 (−1.268)	1.03 (13.874)	−4.067**	7.85(0.02)
Addis & Jimma	0.99 (8.672)	0.389 (5.240)	−4.684**	11.19(0.000)
Addis & Shashemene	−0.12 (−1.48)	1.04 (19.948)	−4.479**	9.52(0.01)
Teff				
Addis & Bahir Dar	0.23 (1.574)	0.90 (12.00)	−3.824**	10.99(0.00)
Addis & Shashemene	0.04 (0.238)	1.01 (11.97)	−4.878**	12.71(0.000)
Addis & Mekelle	0.24 (1.551)	0.85 (10.744)	−3.345*	5.3(0.07)
Addis & Jimma	0.40 (2.747)	0.77 (10.249)	−4.535**	5.08(0.08)
Maize				
Addis & Dire Dawa	0.30 (4.802)	0.80 (13.029)	−5.455**	23.06(0.000)
Addis & Mekelle	0.19 (4.462)	0.87 (20.18)	−4.468**	11.73(0.000)
Addis & Bahirdar	0.05 (0.928)	0.96 (18.09)	−4.705**	34.19(0.00)

Note ***, **, and * denotes significance level at the 1%, 5% and 10% levels, respectively. Critical values for the cointegration test with a constant are −3.96, −3.37, and −3.07 for the 1%, 5%, and 10% levels, respectively. Numbers in brackets in the estimated coefficients are t-values. 1. LOP denotes the Law of One Price test result using the likelihood ratio statistics.

Next, the estimated residuals from Equation (1) are estimated as a threshold model using both zero ($\tau = 0$) and non-zero ($\tau \neq 0$) threshold values in each commodity market. For brevity, only the model with the best fit, based on the value of the Akaike information criterion (AIC), are reported. Table 3 reports the results of the selected model.[7]

As reported in the table, the estimates of ρ_1 and ρ_2 are significantly different from zero in general and satisfy the conditions of convergence (i.e., $\rho_1 < 0, \rho_2 < 0,$ and $(1 + \rho_1)(1 + \rho_2) < 1$). Hence, we can proceed with the cointegration test. As reported in column 6 in Table 3, the estimated statistics for cointegration ($\rho_1 = \rho_2 = 0$) suggest that each pair of

[7] Results of all estimated models are available upon request from the author.

prices investigated for each commodity market is cointegrated, which confirms the Engle-Granger cointegration test results.

Next, as cointegration was not rejected in each commodity market, symmetric versus asymmetric price adjustment is tested. Column 7 in Table 3 reports the estimated F-statistic and the p-value of corresponding significance for the symmetry test ($\rho_1 = \rho_2$). The results in the wheat market show that the null hypothesis of symmetric price transmission is not rejected in any of the market pairs investigated. In the teff market, symmetric price adjustment is rejected only in one out of the four pairs of prices investigated. In the maize market, however, asymmetric price transmission is confirmed in two out of the three pairs of prices investigated. This result indicates the possibility of asymmetric adjustment in the maize market, although it is necessary to consider more regional maize markets for a more robust conclusion about the nature of the price adjustment.

Table 3. Consistent threshold cointegration & asymmetry test in grain market in Ethiopia.

Columns	(1)	(2)	(3)	(4)	(5)	(6)	(7)	
	ρ_1	ρ_2	γ_1	τ	AIC	$\rho_1 = \rho_2 = 0^1$	$\rho_1 = {\rho_2}^2$	Model
Wheat								
Addis & Robe	−0.31 (−3.17)	−0.49 (−3.80)	−0.20 (−2.32)	−0.059	−323.13	14.94	1.19 (0.276)	TAR
Addis & Dire Dawa	−0.26 (−3.54)	−0.39 (−3.49)	0.20 (2.27)	−0.056	−440.82	8.81	0.96 (0.328)	TAR
Addis & Debre Markos	−0.06 (−0.53)	−0.26 (−4.53)	-	0.062	−348.04	10.41	2.08 (0.152)	MTAR
Addis & Jimma	−0.32 (−3.47)	−0.49 (−3.54)	−0.08 (−0.94)	−0.105	−265.76	11.14	0.50 (0.48)	TAR
Addis & Shashemene	−0.23 (−2.57)	−0.48 (−4.36)	−0.09 (−1.11)	−0.013	−378.56	11.54	2.4039 (0.124)	TAR
Teff								
Addis & Bahir Dar	−0.23 (−2.56)	−0.37 (−3.12)	−0.08 (−0.81)	−0.074	−319.53	6.82	1.23 (0.271)	TAR
Addis & Shashemene	−0.35 (−2.48)	−0.46 (−4.36)	0.15 (1.44)	0.074	−258.54	8.27	0.27 (0.604)	TAR
Addis & Mekelle	−0.06 (−0.37)	−0.30 (−3.59)	−0.18 (−1.80)	0.046	−293.39	7.57	1.64 (0.203)	MTAR
Addis & Jimma	−0.28 (−2.99)	−0.69 (−3.93)	0.01 (0.09)	−0.072	−290.96	8.99	4.09 (0.046)	TAR
Maize								
Addis & Dire Dawa	−0.25 (−2.748)	−0.47 (−4.765)	-	−0.052	−252.71	15.13	2.73 (0.101)	MTAR
Addis & Mekelle	−0.61 (−4.487)	−0.25 (−3.523)	-	0.069	−357.89	16.27	5.49 (0.021)	MTAR
Addis & Bahirdar	−0.30 (−3.089)	−0.63 (−5.771)	-	−0.010	−268.74	21.42	5.267 (0.023)	MTAR

Notes. Numbers in brackets in columns (1)– (3) are the t-statistics. 1. Entries in this column are the sample values of the TAR & MTAR statistics. 2. Entries in this column are the sample F-statistics for $\rho_1 = \rho_2$ and significance levels are in parentheses below. Enders and Siklos's (2001) critical values for MTAR for two variables and none lagged are approx. 5.45, 6.51, & 8.78 for 10%, 5%, & 1% resp. Critical values of MTAR for two variables and one lagged are approx. 5.47, 6.51, & 8.85 for 10%, 5%, & 1% resp. Critical values of TAR for two variables and one lagged are approx. 4.99, 6.01, & 8.30 for 10%, 5% & 1% resp. Critical values for TAR for two variables and no lagged are approximately 5.01, 5.98, & 8.24 for 10%, 5%, & 1% resp. τ represents the threshold value and γ_1 is the parameter included in the models to account for autocorrelation.

In general, the null hypothesis of symmetric price adjustment is not rejected in nine out of the twelve pairs investigated, which implies that symmetric price adjustment generally characterizes the Ethiopian grain market. This means that, in the Ethiopian grain markets, increases and decreases in central-market prices transmit to local markets at equal speed. This could indicate the existence of efficient price transmission in the Ethiopian grain markets. However, given the discussion above, the presence of efficient price adjustment in these markets is unexpected. The active presence of the state-led trading enterprise, Ethiopian Grain Trade Enterprise (EGTE), and government restrictions on traders holding large stocks of grain might contribute to the efficient adjustment.

Based on these results, I estimate symmetric or threshold asymmetric ECMs for each commodity market. Tables 4, 5, and 6 present estimates of the ECMs from the wheat, teff, and maize markets, respectively. Results for the post-estimation tests in each estimated model (such as the Breusch-Godfrey serial correlation test, the Breusch-Pagan test, and the Jarque-Bera test) reveal that the estimated ECMs perform reasonably well, and there are no major misspecification problems.[8]

First, concerning the wheat market, symmetric ECMs are estimated for each pair of prices since symmetric transmission is not rejected. As shown in Table 4, the adjustment parameter, $\mathrm{ECT_t}$ for the central and local markets is statistically significant at a 5% level (at least) in each of the local-market models. By contrast, in the central market models $\mathrm{ECT_t}$ is statistically significant only in some of the models, and its significance level is not much stronger than in the local models. This implies that the local market adjusts to changes in central-market prices, whereas the central market might not adjust to changes in local-market prices. The central market not adjusting to changes in the local market price is an indication of the central market's leadership or market power; this is consistent with the central market being larger than local markets. Furthermore, the estimated speed-of-adjustment parameters for the local market models are low (ranging from 16% to 41%), indicating the presence of a slow correction speed following a shock in the central market. The joint significance of short-run, cross-price effects is indicated by $\sum_{i=1}^{s} \gamma_i = 0$ in Table 4 for each pair of prices estimated. The reported values are the F-statistics, and the significant parameter indicates that the central and local market prices for wheat are not generally integrated in the short run.

[8] To save space and for clarity, the diagnostic tests were not reported here. However, can be obtained upon the request from the authors.

Table 4. Symmetric Error correction Models (Wheat Markets)

Indep. Var.	Addis & Robe		Addis & Dire Dawa		Addis & Debre Markos		Addis & Jimma		Addis & Shashemene	
	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$
ECT_t	0.04	−0.29*	−0.20**	−0.36*	0.07***	−0.16**	−0.09**	−0.41*	0.03	−0.33*
$\sum_{i=1}^{s} \gamma_i = 0$	4.64	0.022	1.45	0.20	1.79	0.70	0.00	0.01	4.04**	0.14

Note: *, ** and *** denote significance at 1%, 5% and 10% levels, respectively. $\gamma_i = 0$ shows the joint significance of the short-run cross-price parameters.

In the teff market, the estimated ECMs (as results shown in Table 5) indicated that the error correction term is statistically significant in both the central and local market models in each estimated model, indicating the presence of bi-directional adjustment. The joint significance of the short-run cross coefficients indicated that only the Addis and Jimma pair are integrated in the short run.

Table 5. Symmetric & Asymmetric Error correction Models model (Teff Markets)

Indep. Var.	Addis & Bahirdar		Addis & Shashemene		Addis & Mekelle		Addis & Jimma	
	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$	$\Delta {p_t}^c$	$\Delta {p_t}^l$
ECT_t	0.16**	−0.14***	0.14**	−0.28*	0.11***	−0.21**	-	-
ECT_t^{+}	-	-	-	-	-	-	0.03	−0.29*
ECT_t^{-}	-	-	-	-	-	-	0.43**	−0.43**
$\sum_{i=1}^{s} \gamma_i = 0$	5.74**	0.09	0.88	0.00	10.37***	1.68	0.88	3.18***

Note: *, ** and *** denote significance at 1%, 5% and 105 levels, respectively. $\gamma_i = 0$ shows the joint significance of the short-run cross-price parameters.

In the maize market, the estimated ECMs (results reported in Table 6) indicate that in all pairs of prices estimated, the central market price does not adjust to local price changes, whereas the local market prices do adjust to changes in central market prices. This indicates the presence of uni-directional price adjustment in the maize market. The joint significance of the short-run cross-coefficients indicates that prices in the central and three local maize markets are not integrated in the short run.

Table 6. Symmetric & Asymmetric Error correction Models model (Maize Markets)

Indep. Var.	Addis & Bahirdar		Addis & Mekelle		Addis & Dire Daw	
	$\Delta p_t{}^c$	$\Delta p_t{}^l$	$\Delta p_t{}^c$	$\Delta p_t{}^l$	$\Delta p_t{}^c$	$\Delta p_t{}^l$
ECT_t	-	-	-	-	0.05	−0.33*
ECT_t^+	0.08	−0.10	0.21	−0.43**	-	-
ECT_t^-	0.13	−0.36*	0.11	−0.16***	-	-
$\sum_{i=1}^{s} \gamma_i = 0$	0.01	2.06	0.04	1.01	0.57	0.87

Note: *, ** and *** denote significance at 1%, 5% and 10% levels, respectively. $\gamma_i = 0$ shows the joint significance of the short-run cross-price parameters.

6. Conclusion and discussion

Basic food crop prices persistent increases have become a critical challenge in Ethiopia since 2006. This study examines the market dynamics of prominent grain categories and possible insights relevant to potential price-stabilization interventions by the government. Spatial markets for wheat, maize, and teff were evaluated, and symmetric and asymmetric convergences for disequilibria. Engle-Granger cointegration test, threshold cointegration, and asymmetric ECMs were estimated between the central (Addis Ababa) and major local grain markets.

The findings suggest that spatial market dependencies exist between the central and local markets and affect the formation of complete markets. Price transmission is similar for the wheat and teff markets as symmetric transmission characterizes the price transmission process. In the maize market, evidence of asymmetric price transmission is obtained, although considering more local markets is required to reach a more robust conclusion as only a few maize markets are investigated in this study. Overall, the findings of this study show that symmetric price adjustment characterizes the Ethiopian grain market, suggesting the existence of efficient price transmission between spatial grain markets in Ethiopia.

The presence of symmetric price adjustment in the Ethiopian grain market is surprising, given that the literature shows the existence of inefficiency in the market (Sassi & Mamo, 2019; Negassa et al., 2007). However, three main factors may contribute to this phenomenon. First, the state trading enterprise (the EGTE) plays an active role in the grain market. The EGTE occasionally buys and sells grains when there is a rise in price and bumpy harvest seasons due to drought (Kifle, 2015; Reshid, 2011). Second, the government imports grain (wheat particularly) from the international markets and distributes it at a subsidized price to wholesalers and millers to stabilize the markets (Reshid, 2011). Third, the government restricts traders from holding large stocks of grains on the assumption that such hoarding could exacerbate price increases.

Thus, this study argues that the sustained grain price increases that have been observed in the Ethiopian grain market have nothing to do with the grain market structure. Further investigation is needed to identify the contributing factors of the sustained increases in grain prices in Ethiopia. Possible contributing factors could be supply fluctuation due to drought, rainfall fluctuation, demand and supply factors, exchange rate fluctuation, world price shocks, and income and population growth.

Finally, the estimated results of the (asymmetric) ECMs suggest the leading role of the central wheat and maize markets since it was the local market prices that adjusted to changes in the central market, not vice versa. However, there exist bi-directional adjustments in the teff market.

Conflict of interest

The author declares that there are no conflicts of interest.

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Spatial price transmission in EU pork markets: using threshold autoregressive and non-parametric local polynomial techniques

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Abstract

The pork market in the EU has undergone important changes in recent years, such as the growing concentration of power in both producing and consuming countries, which might impact the price transmission process between markets. This study therefore investigated price transmission between major EU spatial pork markets. Two different techniques, namely threshold autoregressive and local polynomial techniques, were applied to obtain reliable results. The results suggest that price transmission mechanisms vary significantly between different EU markets. However, the transmission speeds between the markets are overall small, indicating that price adjustments take a long time in each market. Furthermore, the estimated results provide some evidence of asymmetric adjustment within the EU pork markets.

Key words: Spatial price transmission, EU pork markets, threshold autoregression, local polynomial fitting

1. Introduction

Spatially separated markets are said to be integrated when they are connected by the process of arbitrage (Fackler & Goodwin, 2001). If price differentials between two markets are greater than the transaction costs, arbitrageurs engage in spatial trade to exploit opportunities so that in the long run the prices in the two markets reach uniformity. This results in a common price known as the Law of One Price (LOP) (Barrett & Li, 2002). The LOP states that, given free trade, prices of a homogeneous commodity in two separate markets will differ at most by the cost of moving the commodity from a lower- to a higher-priced market. Two related concepts to LOP are market integration and market efficiency. While market integration is typically used as a measure of the extent of price transmission between spatially separate markets, market efficiency most often implies a condition in which no arbitrage opportunities remain to be exploited (Fackler & Goodwin, 2001).

Spatial price transmission is an important research topic since it is related to the functioning of geographically separated markets. The number of empirical studies on this subject has grown dramatically in the last two decades in response to concerns that rapid changes in market and business practices could influence price transmission processes, which further generate potentially important welfare and policy implications. In the empirical investigation of spatial price transmission, particular attention has been focused on how quickly prices adjust between markets when there are deviations from long-run equilibrium. If price differentials are arbitrated away quickly, it indicates a high degree of market integration (Fousekis, 2015). There is also a growing research interest in asymmetric price transmission, focusing on empirical validations of the existence of asymmetric price transmission in various markets (Serra et al., 2006a; von Cramon-Taubadel & Meyer, 2004; Abdulia, 2000).

Looking at the literature on price transmission between spatially separated agri-food markets (which forms the focus of our study), it is difficult to generalize a common conclusion despite a large number of hits on the topic. The reason for this is that studies generally differ in terms of the goods, countries, time frequencies, and periods analyzed, and in terms of their model specifications (Frey & Manera, 2007; von Cramon-Taubadel & Meyer, 2004; Abdulia, 2000, Serra et al., 2006a). Moreover, the global agri-food market structure is changing. There is an increasing level of concentration in the agri-food market due to vertical and horizontal integration (Borsellino et al., 2020; Saitone & Sexton, 2012).

A high level of market concentration in the agri-food market is concerning due to the possible implication of market power abuses on the overall market performance (Saitone & Sexton, 2012; European Competition Network, 2012; European Commission, 2009). For instance, Rezitis and Tsionas (2019) argue that the market power exercised by retailers is responsible for asymmetric price transmission in the EU food supply chain. In such a context, it is crucial to understand how the rapidly changing market structure in the EU agri-food sector influences the price transmission process between markets across space.

The objective of this study is thus to investigate the spatial price transmission in the EU agri-food markets, focusing on EU spatial pork markets. The analysis of spatial price transmission within the EU pork market is considered economically relevant for the following reasons. First, pork is the most produced and consumed commodity in many EU countries, accounting for approximately 8.5% of the total EU-27 agricultural output and 35% of the total EU meat output (European Commission, 2020). Secondly, an understanding of spatial price relationships is important because the structure of the pork industry in the EU has shown a high level of concentration in recent years (e.g., Fousekis, 2015), and thus big companies may wield unchecked market power. It is therefore interesting to investigate whether concentration influences price transmission in the markets (e.g., Sexton et al., 1991). Thirdly, the EU pork production and trade flow have changed significantly over the last two decades following the two eastern enlargements of the EU.[1] For example, the new member state, Poland, has substantially increased its imports from Germany (Skrzymowska, 2012). According to our analysis, the import share of Poland has grown from only 1% in 2003 to 13% in 2019. Thus, it is interesting to investigate whether the EU pork market changes have impacted the price transmission process in the markets.

Furthermore, to the best of our knowledge, in the empirical analysis of spatial price linkages in the EU pork markets, only two studies have considered price asymmetry (i.e., Serra et al., 2006a; Meyer, 2004), and both studies employ outdated data. Meanwhile, the results provided by studies that have not considered price asymmetry may be biased when an asymmetric response is misspecified as symmetric (Barrett & Li, 2002). Such outdated and unreliable data is a problem because, as discussed above, the EU pork market has been undergoing significant changes in recent years due to industry concentration and eastern enlargement. Therefore, a new study using the most up-to-date data is called for. The research methods used to achieve this goal are threshold autoregressive and local polynomial models, both of which follow Serra et al. (2006a). We used two different approaches to obtain reliable results.

The remainder of this paper is organized as follows. Section 2 outlines the structure of pork production and trade in the EU. Section 3 then covers the literature review, before section 4 looks at the data. After that, section 5 describes the methodology, section 6 presents the empirical results, and, finally, section 7 offers concluding remarks.

[1] The accession of ten countries to the EU in May 2004 (Cyprus, the Czech Republic, Estonia, Poland, Malta, Latvia, Hungary, Lithuania, Slovakia, and Slovenia), and two other countries in January 2007 (Romania and Bulgaria), eliminated trade barriers between old and new, and among new member states, which in turn has triggered many changes in EU pork production and trade flows (Holst & von Cramon-Taubadel, 2013).

2. The background of pork production and trade in the EU

The EU is the world's second-biggest producer of pork after China and the biggest exporter of pork and pork products in the world. The top six leading EU pork producers in 2019 were Germany, Spain, France, Denmark, the Netherlands, and Poland, which together represent more than 65% of the EU pigmeat production in 2019 (see Table 1 below). We find trade mainly in the form of fresh and frozen pigmeat, processed pigmeat, and live animals, among these six countries also represents a large share in the intra-EU trade flows.

EU pork production and trade flow have undergone many changes over the last two decades following the two eastern enlargements of the EU (in May 2003 and January 2007, respectively). Pork production in most of the new member states has fallen markedly, while production in the old member states has generally increased. As shown in column (1) of Table 1, the share of pigmeat production in all the selected new member states decreased between 2003 and 2019. By contrast, many of the new states increased production except for France and Denmark. Among the old members, meanwhile, Spain and Germany experienced the most significant growth.

Production change has also triggered change in trade flow in the EU pork market. We use the data for these six leading countries to illustrate the trade flow change. As presented in column (2), in terms of the biggest product category "meat of swine fresh, child or frozen", the total import share of the new member states increased, with the most significant growth in Poland by more than 10% between 2003 and 2019. The import growth among the new member states was offset by a decline in imports among the old members. Specifically, although Germany was still the largest importer in the EU in 2019, its import decreased by 10%. For exports, on the other hand, we find that shares of all the six member states increased except for Denmark, Netherlands, and France (column (3) in Table 1).

Table 2 illustrates the level of trade among these six countries. Trade activities were high, with Germany and Spain increasing their exports to the other EU countries. Notable examples of this were the 20% increase in exports from Germany to Poland and from Spain to France. Interestingly, despite the considerable increase in exports from Germany and Spain to the other markets, both Germany and Spain were also major importers as demonstrated by their import market shares in Table 2. This suggests the complex dynamics of the whole EU market, which might affect the price transmission between each market.

Table 1. Production and trade flow in the EU between 2003 and 2019

	Production (%)		Imports (%)		Exports (%)	
	2003	Change 2003-2019	2003	Change 2003-2019	2003	Change 2003-2019
Germany	19.3	+2.8	26.1	-9.9	15.4	+7.7
Spain	14.5	+5.1	2.1	-0.2	11.9	+5.84
France	10.6	-1.3	8.7	-3.7	10.0	-3.9
Netherlands	5.7	+1.2	5.2	-0.6	16.6	-4.45
Denmark	8.0	-1.7	1.1	0	19.9	-9.9
Poland	9.6	-1.3	1.1	+11	1.7	+4.34
Slovenia	0.2	-0.19	0.6	+0.2	0.01	+0.04
Czech Republic	2.1	-1.2	0.8	+4.2	0.25	+0.27
Hungary	2.3	-0.5	0.7	+2.2	1.6	+0.9
Slovakia	0.8	-0.5	0.3	+2	0.001	+0.44
Old member total.	58.1	+6.1	43.1	-14.4	73.8	-4.07
New member total.	15	-3.69	3.7	+19.4	3.56	+5.99
Total share	73.1	+2.4	46.7	+5.2	77.4	+1.94

Source: Own calculation based on data from Eurostat (2021).

Table 2. Intra-trade among the leading EU member states

Export from	Year	Imports by (in %)					
		Germany	Denmark	Spain	France	Netherlands	Poland
Germany	2003	-	36	9.9	5.9	49	0.4
	2019	-	38.3	16.6	11.2	42.8	22.6
Denmark	2003	29.6	-	3.9	13.5	6	58.8
	2019	29.7	-	5.7	2.7	2.1	15.7
Spain	2003	7.6	12	-	48.2	3	2.9
	2019	4.2	19.9	-	73.7	4.0	11.8
France	2003	3.5	10	16.7	-	10	5.1
	2019	1.9	3.7	24.8	-	8.5	2.9
Netherlands	2003	19.8	17	19.4	15.2	-	20.3
	2019	18.5	10.0	19.8	3.2	-	11.5
Poland	2003	0.4	1	3.7	0.3	1	-
	2019	8.6	5.0	6.8	0.3	7.9	-

Source: Own calculation based on data from Eurostat (2021).

3. Literature review

Study of spatial price transmission measures the degree to which spatially separated markets are integrated. In integrated markets, price shocks in one market trigger responses in other markets due to profit-seeking arbitrage. Arbitrage activities ensure that prices of a homogeneous commodity in two spatially separated markets will differ by at most the cost of moving the commodity from the low- to the high-price market (e.g., Emmanoulides & Fousekis, 2012; Van Campenhout, 2007; Serra et al., 2006a; Goodwin & Piggott, 2001). On the other hand, when markets are not well integrated, profitability opportunities will not be fully exploited, which results in efficiency losses (e.g., Fackler & Goodwin, 2001). Inefficient spatial price relationships between markets can distort producer decisions and lead to inefficient product movement, and ultimately increase consumer prices (Goodwin & Schroeder, 1991).

As noted earlier, spatial price transmission on agricultural product markets has been studied for a long time; hence, there is an abundance of literature on it (e.g., Chen et al., 2011; Van Campenhout, 2007; Serra et al., 2006a; Goodwin & Piggott, 2001; Baulch, 1997). The literature is diverse in terms of commodities, methodologies, periods, and regions, and thus it is difficult to draw generalized conclusions from it. However, a common approach taken by the literature is to investigate how quickly prices adjust between separated markets when there are deviations from a long-run equilibrium and whether this adjustment is asymmetric. As stated, if price differentials are quickly adjusted, it indicates a high degree of market integration (Serra et al., 2006a).

Price asymmetry is frequently observed in agricultural commodity markets. Asymmetric price transmission occurs when price responses in a market vary based on positive and negative price changes in other markets, otherwise it is called symmetric (e.g., Abdulia, 2000). Taking price transmission between import and export markets as an example, no trading partner enjoys an advantage over others in the case of symmetric price transmission. That means that consumers in an importing country are equally likely to experience the same increase or decrease in price as consumers in the country where the goods were exported (Fousekis, 2015; Serra et al., 2006a). In the case of asymmetric price transmission, though, welfare can be transferred from exporters to importers, or vice versa, depending on who has the leading role in the markets (von Cramon-Taubadel & Meyer, 2004). The main causes of asymmetric price transmission identified in the literature include market power, product nature, transaction cost, and inventory storage (von Cramon-Taubadel & Meyer, 2004).

Regarding the research methods, cointegration with error correction models have been the traditional method used to study price linkages between spatial markets. However, this approach is widely criticized in the literature as it implicitly assumes that price linkage is linear or symmetric. As noted earlier, price relationships are likely to be non-linear and/or asymmetric (Barrett & Li, 2002).[2] Several econometric techniques that are capable of capturing non-linearity have been proposed in the literature. Among them, the threshold autoregressive (TAR)

[2] We use the three terms (i.e., asymmetric, nonlinear, and threshold) synonymously in this study, in contrast to symmetric (or linear) adjustment, although their meanings could differ slightly in the literature.

model is widely used (e.g., Serra et al., 2006a; Van Campenhout, 2007). TAR models are useful models to assess price transmission across spatially separate markets in the presence of transaction costs. They are based on the idea that the presence of transaction costs creates a neutral band where spatial price relationships are weak or even non-existent because price differentials do not exceed transaction costs, thus making spatial trade unprofitable (Serra et al., 2006a). If price differentials between markets exceed transaction costs, profit-seeking arbitrage activities drive price relationship towards the equilibrium (i.e., towards the neutral bands). This non-linear pattern of price adjustment is represented through a combination of different regimes.

The discussion above suggests non-linear price transmission models are preferred to capture non-linear price linkage. However, most non-linear models, including the TAR models, as used in Serra et al. (2006a), are parametric models, which require the specification of an exact functional form before estimation. Moreover, these models are piecewise linear; that is, they allow for abrupt and discontinuous transmissions from one price transmission regime to the other. Thus, unless the number of regimes is correctly specified, they might lead to biased and inconsistent results (e.g., Serra et al., 2006a; Fousekis, 2015). One possible solution to this is to estimate non-parametric models. Non-parametric techniques, which are data-driven and hence do not require any functional form a priori, offer a highly flexible way to analyze spatial price transmission.

To combine the benefits of both parametric and non-parametric techniques for reliable results, Serra et al. (2006a) employed parametric TAR and non-parametric local polynomial techniques to investigate spatial price transmission between major EU pork markets. Serra et al. (2006b) followed the same approach to study spatial price transmission in the U.S. egg markets. In this study, we follow these two studies to study price transmission between major EU pork markets. However, in a departure from the other studies, where the presence of nonlinear price adjustment relies on visual inspection of the regression fits,[3] our study uses a statistical procedure to identify the presence of nonlinear price adjustment when using the non-parametric local-polynomial approach. More discussion on both parametric TAR and non-parametric local polynomial techniques is provided in the methodology section.

In the EU pork markets, relatively few studies have investigated spatial price transmission considering its importance to the EU economy and most of these studies are old.

[3] Results from non-parametric models are often best described by using a graph of the regression fits.

The available research starts with Sanjuàn and Gil (2001), who confirmed a higher degree of pork market integration in the EU by analyzing price transmission between seven member states. After that, different market dimensions were covered by the literature (Meyer, 2004; Serra et al., 2006a; Fousekis, 2007; Emmanouilides & Fousekis, 2012; Cramon-Taubadel, 2013). The most recent research is by Fousekis (2015), who studied EU pork market integration using a nonparametric technique in five markets, Germany, Spain, France, Denmark, and Poland. These studies overall show that trade volume is positively related to the speed of price transmission. As noted earlier, there are only two empirical studies of price asymmetric transmission in the EU pork markets in the literature. Of these, Meyer (2004) shows that pork markets between Germany and the Netherlands are asymmetrically integrated, while Serra et al. (2006a) indicate that price transmission among the four major EU pork markets (i.e., Germany, Spain, Denmark, and France) are generally asymmetric. However, since these two studies used old data, it is necessary to revisit them using more updated data.

4. Data

Our analysis of price transmission in the EU pork markets focuses on the six leading EU pork markets, namely Germany, Spain, Denmark, France, Netherlands, and Poland. As noted above, these countries together represent more than 65% of the total pork production and cover a significant portion of the intra-EU pork trade in 2019.

We used weekly national average slaughter pigmeat prices from each market considered in this study, covering the period from week 1, 2006 to week 52, 2020. The datasets were obtained from the European Commission (2020). Prices are measured by euro/100 kg. Using more disaggregated data is recommended in the study of price transmission to achieve more robust empirical results that can illustrate the true nature of the markets studied. However, our analysis using disaggregated data in this study was constrained by data availability. Moreover, previous price transmission studies that focused on EU pork markets (e.g., Serra et al., 2006a; Fousekis, 2015) used national average slaughter prices to investigate price transmission between EU spatial pork markets. Hence, using average slaughter pigmeat prices from each market to study spatial price transmission in this study is considered reasonable.

Table A1 in the Appendix presents a summary of the price series. Graphs of the price series presented in Figure A1 in the Appendix generally show that the pigmeat price series in the six major EU pork markets follow very similar patterns, suggesting the possibility of a price transmission process across EU pork markets.

5. Methodology

Our empirical procedure comprises a series of tests and model estimations. First, we performed a stationary test using the Augmented Dickey-Fuller (ADF) test (Dickey-Fuller, 1981) and Perron's (1997) test on individual price series. The use of two different stationarity tests helps to avoid the risk of wrongly accepting or rejecting the null hypothesis of non-stationarity. We then used the Johansen cointegration test to investigate long-run relationships between price pairs. Finally, we used a parametric threshold autoregressive (TAR) model and a non-parametric local polynomial model to investigate asymmetric and/or non-linear price transmission process among the EU pork markets. As stated above, the use of two different methods in this study may allow us to obtain more reliable results.

5.1. Threshold Autoregressive Models

The TAR model is constructed as follows. Suppose p_{it} and p_{jt} (expressed in log form) denote the prices at time t of a homogeneous commodity (pigmeat in our case) in spatially separated markets i and j, respectively. According to spatial price transmission literature (e.g., Serra et al., 2006a; Balke & Fomby, 1997), the relationship between the prices can be explored by analyzing the relationship between the price differential in time $t-1$, denoted by $X_{t-1} = p_{i,t-1} - p_{j,t-1}$, and the adjustment of that differential in time t, denoted by $\Delta X_t = X_t - X_{t-1}$. If we assume this relationship is given by a linear form, X_{t-1} and ΔX_t can be related through the standard autoregressive AR (1) model of the form:

$$Y_t = \beta X_{t-1} + e_t \qquad (1)$$

where $Y_t = \Delta X_t$, and $e_t \sim N(0, \sigma^2)$ is the estimated residual. The parameter β is the speed of price adjustment, which indicates the response of the price difference at time t to the price difference at time $t-1$. For stationary price differentials, β is expected to be a negative number implying that “high” (i.e., above the long-run equilibrium value) price differentials in time $t-1$ tend to adjust downwards in time t, while “low” (i.e., below the long-run equilibrium value) price differentials in $t-1$ tend to adjust upwards to ensure mean reversion (e.g., Fousekis, 2015).

As discussed above, the price transmission could be both linear and non-linear and/or asymmetric due to transaction costs or market imperfections. The model given in Equation (1) excludes the possible non-linear relationship. To account for the possibility of non-linearity, Balke and Fomby (1997) extended the AR model (1) to a TAR model. The TAR model allows the relationship between the price differentials to vary in a regime fashion, depending on the degree of the price difference relative to the transaction costs. A three-regime TAR model is given as follows:

$$Y_t = \begin{cases} \beta_1 X_{t-1} + \varepsilon_{1,t} & if\ -\infty < X_{t-1} < c_1 \quad \text{(Regime 1)} \\ \beta_2 X_{t-1} + \varepsilon_{2,t} & if\ \ c_1 \leq X_{t-1} \leq c_2 \quad \text{(Regime 2)} \\ \beta_3 X_{t-1} + \varepsilon_{3,t} & if\ c_2 < X_{t-1} < +\infty \quad \text{(Regime 3)} \end{cases} \qquad (2)$$

where the model in Regimes 1, 2, and 3 captures the price adjustment when the price differentials are less than, fall between, and exceed the estimated transaction costs of moving the commodity from market i to j, i to j and j to i, and j to i, respectively. The parameters c_1 and c_2 denote the threshold values that define the regimes (Serra et al., 2006a).[4] The band $[c_1, c_2]$ represents the so-called “inactive (neutral) band” where no arbitrage activity takes place. The model assumes that arbitrage opportunities derive prices toward the inactive band edges, where LOP is satisfied. Thus, the speed of adjustment coefficients in the outer band (in the two outer-regimes), that is, β_1 and β_3 are expected to be statistically significant and negative in sign, whereas β_2 is expected to be statistically insignificant. From Equation (2), asymmetric price transmission exists if the hypothesis $\beta_1 = \beta_3$ is rejected.

A grid search approach is used to determine the appropriate threshold values. To ensure an adequate number of observations in each regime, following Serra et al. (2006a), after sorting X_{t-1} in ascending order, the first threshold value (c_1) was searched over the space defined by the minimum and median of X_{t-1}, and the second threshold (c_2) was searched between the median and the maximum of X_{t-1}. For a given threshold values c_1 and c_2, ordinary least squares (OLS) regression can be used to estimate the speed of price adjustment (β_j) in each regime (Serra et al., 2006a). From this estimation, the residual sum of squares (denoted $\hat{\sigma}$) can be obtained using $\hat{\sigma}(c_1,c_2) = \sum_{t=1}^{n} \hat{e}_t(c_1, c_2)^2$. Then a standard *F*-statistic is calculated using Equation (3) to test a linear AR model against the alternative of a TAR model.

[4] It is assumed in the literature that a variable TAR model provides a more accurate set of estimates than a constant threshold model (Van Campenhout, 2007). However, to investigate price asymmetry, and to compare our results with the local polynomial model (discussed below), we stick to the constant threshold model.

$$F = \frac{\tilde{\sigma}^2 - \hat{\sigma}(c_1, c_2)}{\hat{\sigma}(c_1, c_2)} n \tag{3}$$

where n represents the number of observations, $\hat{\sigma}(c_1,c_2)$ denotes the error variance of the TAR model, and $\tilde{\sigma}^2$ denotes the error variance of the AR model (e.g., Serra et al., 2006a).

Since the *F*-test above does not have a standard distribution, its *p*-value can be determined based on the method provided by Hansen (1997). We chose between the two-or three-regime TAR models based on Akaike Information Criterion (AIC).

5.2. Local polynomial fitting

As noted earlier, nonparametric techniques such as local polynomial (LP) modeling do not require any restrictive assumptions about the functional form of the relationship between X_{t-1} and Y_t; hence, they offer a flexible way to investigate the nature of price transmission. The general form of a nonparametric regression model describing the relationship between price differentials is given as follows:

$$Y_t = m(X_{t-1}) + e_t \tag{4}$$

where $m(.)$ is a smooth function relating Y_t and X_{t-1}, and it is assumed to be differentiable at X_{t-1}. The objective of nonparametric regression is to estimate the regression function $m(.)$ directly, rather than to estimate parameters like in the parametric AR or TAR models above.

As discussed in Serra et al. (2006a), the fundamental idea behind local polynomial fitting is to use observations (X_{t-1}, Y_t) for $t = 2,, \dots, n$, relatively close to a local point x_k to estimate the function $m(X_{t-1})$. In most cases, the literature recommends using local linear regression (LLR) to approximate m (e.g., Fan & Gijbels, 1999; Serra et al., 2006a). Thus, using Taylor series, at a given point x_k, a local linear regression function $m(x)$ can be written as: $m(x) = \beta_0 + \beta_1(x - x_k)$, where β_0 and β_1 are regression parameters to be estimated. Then, the estimator for $m(x_k)$ will become $\widehat{m}(x_k) = \widehat{\beta_0}$. The regression coefficients β_0 and β_1 at a point x_k can be obtained by solving the weighted least squares problem, which is given by:

$$\min_{\beta_0, \beta_1} \sum_{t=1}^{n} \left(Y_t - \beta_0 - \beta_1(X_{t-1} - x_k)\right)^2 K_t \left(\frac{X_{t-1} - x_k}{h_k}\right) \tag{5}$$

where K_t denotes weights (kernels) and h_k is a smoothing parameter or bandwidth.

Solving Equation (5) requires the selection of a kernel function and a bandwidth. The choice of a kernel function is not especially relevant as far as it is a continuous and smooth function (Serra et al., 2006a). However, using an appropriate bandwidth is required since it controls the smoothness of a regression fit (Loader, 1997). The simplest choice is to take a constant bandwidth, but often it is desirable to vary with the fitting point. The optimal bandwidth that minimizes the Generalized Cross-Validation (GCV) criteria is used in this study.

A solution of the local linear regression (5) "nests" the globally linear model (1). Thus, we can test the null hypothesis of the global linear model (1) (call it the restricted model) against the alternative potentially non-linear model (5) (call it the unrestricted model) to correctly identify which model captures the price linkages. According to Fox (2005), the relevant test statistic is given as follows:

$$F = \frac{(RSS_R - RSS_U)/(df_{MOD} - 2)}{RSS_U/df_{RES}} \tag{6}$$

where RSS_R and RSS_U are the residual sum of squares from the restricted and the unrestricted models, respectively. df_{MOD} and df_{RES} are the degrees of freedom of the local fit and residual degrees of freedom, respectively.

In this study, we follow a pair-wise analysis to investigate the price transmission process in the EU pork markets by following the mainstream of the literature (e.g., Serra et al., 2006a; Van Campenhout, 2007). The popularity of pair-wise analysis is primarily a result of it being relatively easy to understand and interpret. Furthermore, the LLR model favors relatively simple model specifications since it is best interpreted by graphical representation. Nevertheless, in a pair-wise investigation of price relationships, identifying the causality direction is a highly relevant concept in price transmission analysis. The leading market is often identified based on certain characteristics of the market. Goodwin and Piggott (2001) recommended considering the largest market in terms of volume as the central (or leading) market.

In their empirical study of the EU pork markets price transmission, Serra et al. (2006a) used Germany as the leading market. They argue that Germany as the most populated country in the EU is the largest importer of pigmeat, a relevant exporter, and has one of the highest per-capita consumption levels of pork. Besides, despite substantial net imports, Germany also exports pork to other EU members, which might affect its price transmission with other pork countries. Our weak exogeneity tests (results reported in Table 4 below) also generally confirm this hypothesis. Hence, we choose Germany as the central (reference) market and investigate the price transmission between the price pairs consisting of the central market price (p_{it}) and alternative market price (p_{jt}). In our empirical analysis below, the prices are expressed in logarithmic form.

6. Estimation results

Tests assessing the time series properties of the price series are reported in Table 3. As indicated in the table, results of the ADF test cannot reject the null hypothesis of a unit root in levels for all price series at least at 1% significance level, except for France and Spain, where the null hypothesis is rejected at 1% level. The ADF test rejected the unit root hypotheses of Spain and France may be due to confusion with a structural break of the series, which is a well-known weakness of the ADF unit root tests. To confirm our results, we used Perron (1997) test. The Perron (1997) test can allow us to identify whether price series are non-stationary or whether the apparent non-stationarity is due to structural breaks. As indicated in column (2) of Table 3, the Perron (1997) test suggests that the null hypothesis of a unit root cannot be rejected for all price series at least at a 5% level. Results of unit root tests on the first differences of the prices also suggest that all price series are stationary at least at a 5% level[5]. We can thus conclude that all price series are non-stationary in levels but stationary in their first differences. Hence, we can proceed with cointegration tests to investigate long-run relationships between pairs of prices.

Johansen cointegration tests (results reported in Table 4) provide evidence in favor of stationary long-run relationships between all pairs of prices investigated, except between Germany and Spain. Given the existence of a considerable amount of pork and pork products trade among these countries, the presence of long-run relationships between the pairs of prices investigated is not surprising. Nevertheless, what is surprising here is the absence of a long-run relationship between Germany and Spain, given that both countries are the leading pork producers and traders within the EU. However, since the Johansen cointegration test relies on the assumption of a linear price relationship, the absence of a long-run relationship between them might be due to a non-linear and asymmetric price relationship between them.

[5] Results of unit root tests on the first differences of the prices are not reported here but are available upon request.

Table 3. Unit root test

Prices (log form)	ADF test	Perron test	
		t-value	Break date
Germany	-3.03	-2.75	2020:06:25
Spain	-4.61***	-2.82	2014:06:27
Denmark	-2.37	-2.55	2013:10:44
France	-4.18***	-2.68	2013:08:36
Poland	-3.18	-2.71	2020:06:24
Netherlands	-3.38	-3.15	2014:06:26

Notes: ***, **, and ** denote significance level at 1%, 5%, and 10% level, respectively. Critical values for ADF test statistics are -3.43, -2.86, and -2.57 at 1%, 5%, and 10%, respectively. Critical values for Perron test statistics are -5.34, -4.80, and -4.58 at 1%, 5%, and 10% level, respectively.

Table 4. Johansen cointegration test

Markets	λ max, $r = 0$ (Critical value at 10%)	λ max, $r = 1$ (Critical value at 1%)	Weak exogeneity test (p-value)	
Germany – Poland	36.81 (13.75)	9.44 (12.97)	1.64 (0.200)	16.85 (0.000)
Germany – Denmark	17.66 (13.75)	7.30 (12.97)	5.37 (0.020)	3.27 (0.070)
Germany – France	54.90 (13.75)	12.55 (12.97)	3.22 (0.070)	22.49 (0.000)
Germany – Spain	28.72 (13.75)	14.33 (12.97)	0.09 (0.760)	13.29 (0.000)
Germany – Netherlands	27.66 (13.75)	10.87 (12.97)	0.79 (0.370)	6.81 (0.010)

Results obtained from the application of the TAR model are presented in Table 5. The *F*-test suggests that threshold effects are statistically significant for all pairs of prices investigated, except between Germany and Spain. Moreover, a three-regime TAR model is selected for all pairs of prices based on AIC. As shown in the table, the transaction cost bands between each pair of prices investigated are small, indicating the absence of large price differentials between the reference and other major EU pork markets.

However, generally, the middle-regime parameter estimates for all pairs of prices investigated are statistically insignificant. In the outer regimes, most of the estimated speed of adjustments are negative and statistically significant, suggesting that price differentials exceeding the threshold values arbitrage away. Nevertheless, estimated speeds of price transmission are generally low, showing the presence of slow price differentials adjustment in the EU pork markets. Since we considered pigmeat prices in this study and pigmeat is a relatively homogeneous commodity, we expected a higher speed of adjustment in the EU pork markets. One possible reason for getting a low speed of adjustment might be the imbalance of power in the EU pork markets. For instance, as noted earlier, only 5% of abattoirs conduct 65% of the total pig slaughtering in the EU-27 (e.g., Fousekis, 2015).

Specifically looking at the results for each pair of prices investigated, from Table 5 we can see that the speed of adjustment for the pair Germany and Poland is moderately larger than any of the other pairs of prices investigated. More specifically, TAR parameter estimates provide evidence of asymmetric adjustment that confers some advantage to Germany over Poland. While negative price deviations (i.e., in Regime 1, $-\infty < X_{t-1} < c_1$) quickly adjust (with an adjustment speed of 41%), positive price deviations (i.e., in Regime 3, $c_2 < X_{t-1} < +\infty$) do not adjust. As noted earlier, Poland's pork import is rapidly growing, and this import is supplied primarily from Germany. The price asymmetry results obtained here might be due to the market power that German exporters wield over Poland's importers. The results further confirm the hypothesis that trade flow that occurs primarily in one direction causes price asymmetry.

A three-regime TAR model estimates between the reference market (Germany) and other local markets (i.e., Denmark, the Netherlands, and France) provide evidence that price asymmetries grant a certain advantage to local markets over Germany, since positive price deviations ($c_2 < X_{t-1} < +\infty$) adjust, while negative price deviations ($-\infty < X_{t-1} < c_1$) do not. The advantage that Denmark and the Netherlands have over Germany might be due to the market power that exporters from these two countries wield over German importers. As we discussed above, both Denmark and the Netherlands command a considerable share of Germany's imports. However, the advantage that France has over Germany is unexpected given that France's share of Germany's imports is not that large.

Table 5. TAR model parameter estimates

Markets	Thresholds and F-test			TAR parameters		
	Lower threshold c_1	Upper threshold c_2	F-test (p-value)	1st regime β_1	2nd regime β_2	3rd regime β_3
Germany – Poland	-0.011	0.128	27.70 (0.000)	-0.41***	-0.04*	-0.03
Germany – Denmark	0.028	0.209	5.00 (0.022)	0.009	-0.005	-0.03***
Germany – France	0.027	0.152	4.92 (0.013)	-0.017	-0.009	-0.05***
Germany – Spain	-0.075		0.34 (0.358)	-0.04**	-0.02	-0.03**
Germany – Netherlands	0.092	0.160	51.16 (0.000)	-0.10***	-0.002	-0.06***

Notes: ***, **, and * denotes significance at 1%, 5% and 10% respectively. Half-lives are expressed in weeks.

Next, the results obtained from the LLR models are discussed. Figures 1–5 present the results of the LLR model, together with the approximated pointwise 95% confidence band. The simple visual inspection of the regression fits shows some evidence of non-linear price transmission in the EU pork markets. The F-test (results reported in column (1) of Table 6) also failed to accept the null hypothesis of global linearity (or symmetric adjustment) between the reference market (Germany) and the alternative markets at a 10% significance level or lower, except between Germany and Spain. This implies that, consistent with the results of the TAR model, the results of the LLR model suggest that price deviations from long-run equilibrium relationships between the reference market and alternative markets are corrected generally in a non-linear fashion.

We also estimated linear AR models to compare with our results discussed above. Column (2) in Table 6 reports estimates of AR models together with their estimated corresponding half-lives.[6] The results show that the estimated speed of adjustments in each pair of prices investigated is negative and statistically significant at a significance level of at least 10%, suggesting the existence of price adjustment between the reference and alternative markets. Just as in the results from the TAR models, the estimated speed of price adjustments is low, and the estimated half-lives are very large, suggesting that adjustments of deviations

[6] Given the estimated speed of price adjustment β, a half-life can be obtained using the formula $-\ln(2)/\ln(1+\beta)$. Half-lives imply a time period required for half of the price deviations to be corrected.

take several weeks. Indeed, it generally takes from 17 to 69 weeks for half (or 50%) of price deviations to be corrected.

We expected a larger speed of adjustment with associated small half-lives in the EU pork markets due to the following reasons. First, following the literature (Fousekis, 2007; Emmanouilides and Fousekis, 2012; Holst and von Cramon-Taubadel, 2013), trade volume should be positively related to price transmission speed. Secondly, the EU policymakers have intervened less in the pork market compared to any other EU agricultural market (such as dairy) (Bartova et al., 2009). However, as has been discussed, there is an extremely high concentration in EU pork production. Five percent of abattoirs accounted for 65% of the total pig slaughtering in the EU-27 (Fousekis, 2015) and four nations (Germany, Spain, Denmark, and the Netherlands) were responsible for more than 64% of the total EU pork production in 2019. Therefore, we suspect market power from the farm level in the exporting countries might explain the lower price transmission between the EU markets.

Table 6. Test for global linearity, and AR model parameter estimates

Price pairs	*F*-statistics (p-value)	AR model parameter estimates & half-lives	
		β	Half-lives
Germany – Poland	5.56 (0.005)	-0.04***	17
Germany – Denmark	3.86 (0.060)	-0.01**	69
Germany – France	5.19 (0.009)	-0.02***	34.3
Germany – Spain	0.87 (0.669)	-0.03***	17
Germany – Netherlands	11.94 (0.000)	-0.01*	69

Note: ***, **, and * indicate significance at 1%, 5%, and 10%, respectively.

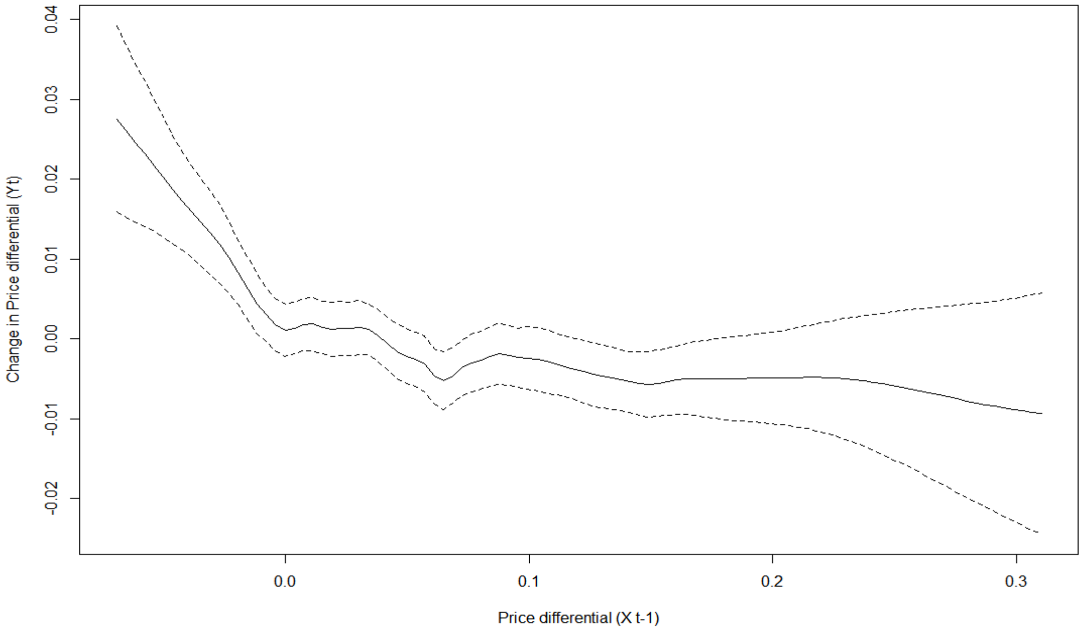

Figure 1. LLR model fit between Germany-Poland with 95% confidence bands

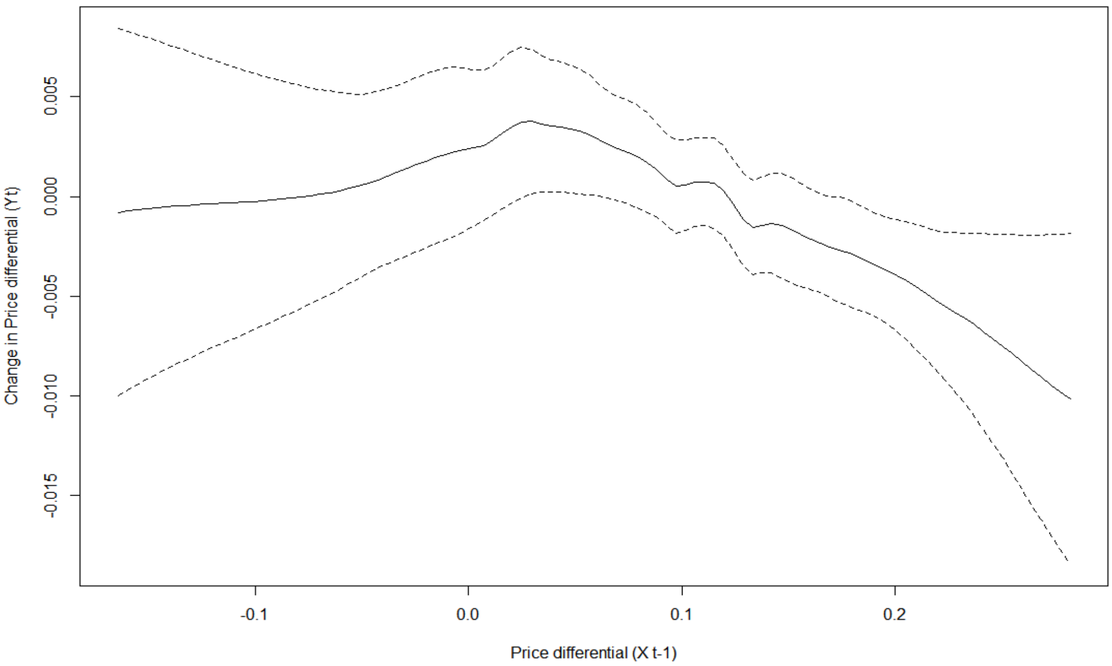

Figure 2. LLR model fit between Germany-Denmark with 95% confidence bands

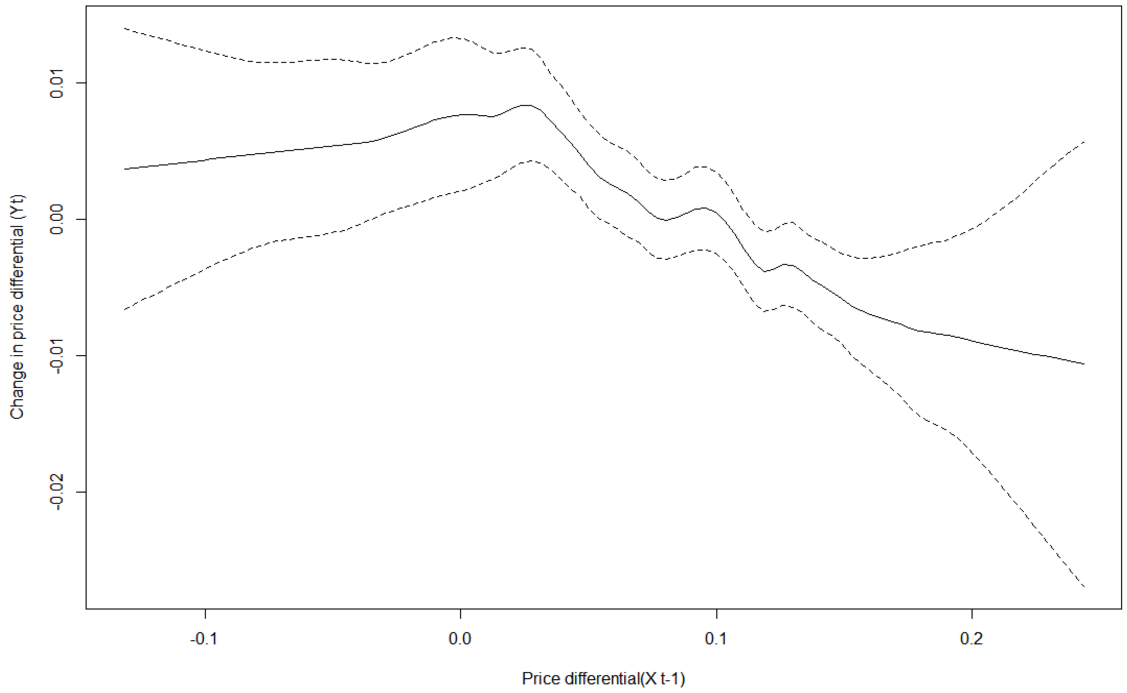

Figure 3. LLR model fit between Germany-France with 95% confidence bands

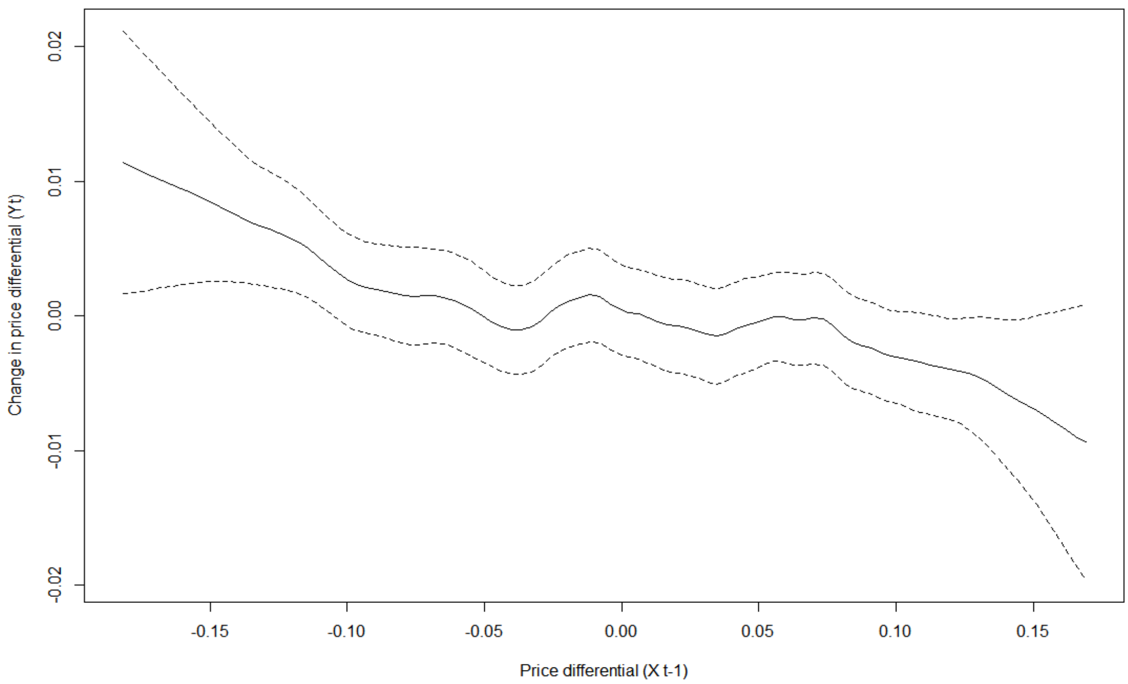

Figure 4. LLR model fit between Germany-Spain with 95% confidence bands

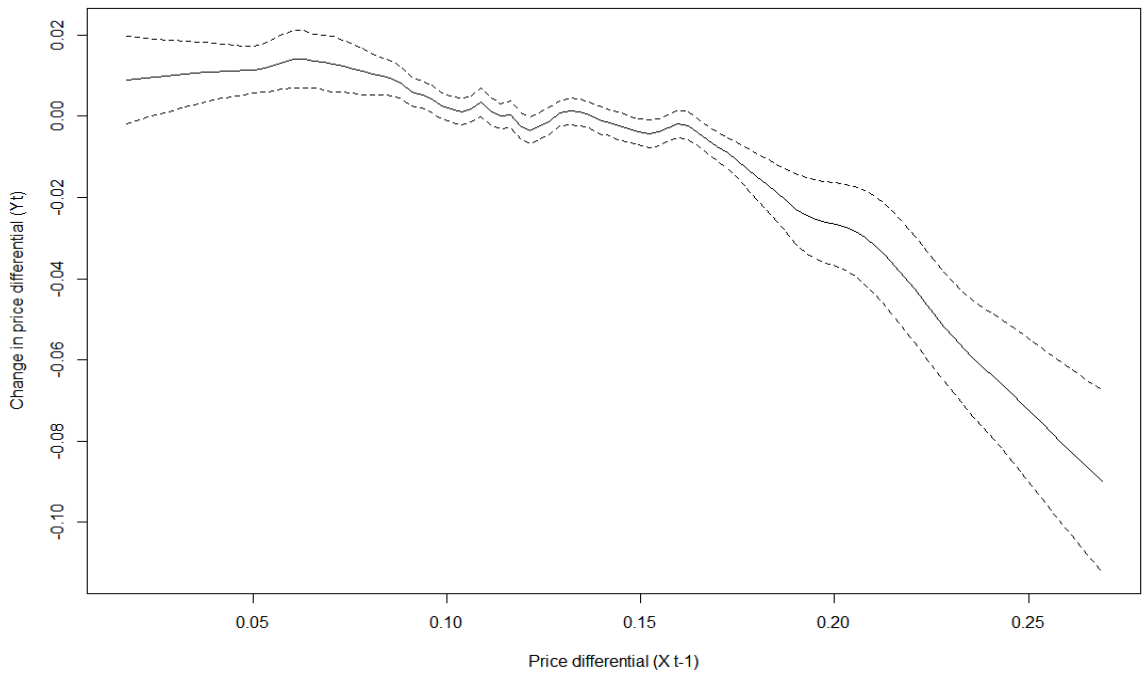

Figure 5. LLR model fit between Germany-Netherlands with 95% confidence bands

7. Conclusions and discussion

The EU pork industry has shown a high level of concentration in countries of production and at the retail level. The high concentration level is likely to give exporting countries and slaughters market power and to respond more quickly to price shocks that involve a reduction in their marketing margins and slowly when the margins are increased. Asymmetric price transmission between markets can in turn distort producer decisions and lead to inefficient product movement. The primary objective of the present paper was to explore the asymmetric price transmission among spatially separated EU pork markets. The empirical analysis used weekly slaughter pigmeat prices from the six major EU pork markets (namely Germany, Poland, France, Denmark, Spain, and the Netherlands), covering the period from week 1, 2006 to week 52, 2020. We estimated spatial price transmission between the reference market (Germany) and each of the other markets. We used two different approaches, parametric threshold autoregressive and non-parametric local polynomial models, to estimate the price transmission process. Two different methods were employed to obtain robust results.

The estimated results provided by both methods are generally in agreement with each other. The estimated speed of price adjustments between the reference and other major EU pork markets is overall low. The low adjustment speeds and consequent large half-lives suggest that price adjustments between the main EU pork markets can take a long period to correct price deviations. The high production concentration both in terms of abattoirs and production countries may explain the low speeds of price adjustments. Concentration and the induced market power from the supply side may have enhanced market inefficiency.

Moreover, asymmetric and non-linear price adjustments are identified within the EU pork markets. Among the old member states, although the estimated results suggest the existence of price asymmetry between Germany and the alternative markets (i.e., Denmark, Netherlands, and France), the estimated speeds of adjustments obtained here are relatively small. The results suggest that positive deviations adjust, whereas negative deviations do not adjust in these markets. However, the relatively small speed of adjustment shows that the advantages obtained by the alternative markets over German importers from the price asymmetries are not large. Both models favor symmetry price adjustment between Germany and Spain. Since both Germany and Spain are leading producing and trading countries, we propose that the relative equal power of these two markets leads to this market efficiency.

Compared to the price transmissions between the EU old members, the price asymmetry between Germany and Poland is sufficiently pronounced to confer a certain advantage to Germany over Poland. This result indicates that the big old EU pork exporter has expanded their exports further to new members, taking advantage of their market power over the entrance of the new members. Given that, out of the new member states, our study only looks at Poland, it is clear that a price transmission analysis that considers more new member states is required to reach a more robust conclusion.

Although the results of this work appear to be generally in line with a previous study by Serra et al. (2006a) on non-linear price transmission among the leading EU pork markets, there is a difference in terms of which markets price asymmetry favors. Serra et al. (2006a) obtained evidence of non-linear price adjustment that appeared to grant a certain advantage to Germany over Denmark, Netherlands, and France, whereas, as noted above, we found the reverse in this study. The reasons for this could be as follows. First, this study covers a different period of study than Serra et al. (2006a). Secondly, Denmark and Netherlands are net exporters of pigmeat, while Germany is the largest EU importer, particularly from these two countries. However, given that France's share of Germany’s imports was not especially large, the result obtained between France and Germany was not expected. Between Spain and Germany, though, our result is in line with that of Serra et al. (2006a).

Our results generally suggest the possible market power granted by the production concentration and following market power of the big exporting country. The advantages enjoyed by the main player are more significant in relation to the new EU members compared to the old members. To continue to examine this phenomenon, we call for further studies on the market power of the upstream supply side in the value chain of the EU pork market.

Conflict of interest

The authors declare that they have no conflict of interest.

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Appendix

Table A1. Descriptive statistics of national average slaughter pigmeat prices (Euro/100 kg), from week 1, 2006 to week 52, 2020

Prices	Mean	Min	Max	St. Dev.	# Obs
Germany	156.95	123.15	208.15	17.84	780
Spain	155.75	116.22	217.10	21.03	780
France	143.71	116.00	190.00	16.18	780
Poland	150.16	112.18	199.41	20.23	780
Denmark	140.56	106.94	206.51	19.57	780
The Netherlands	139.13	107.43	190.01	17.10	780

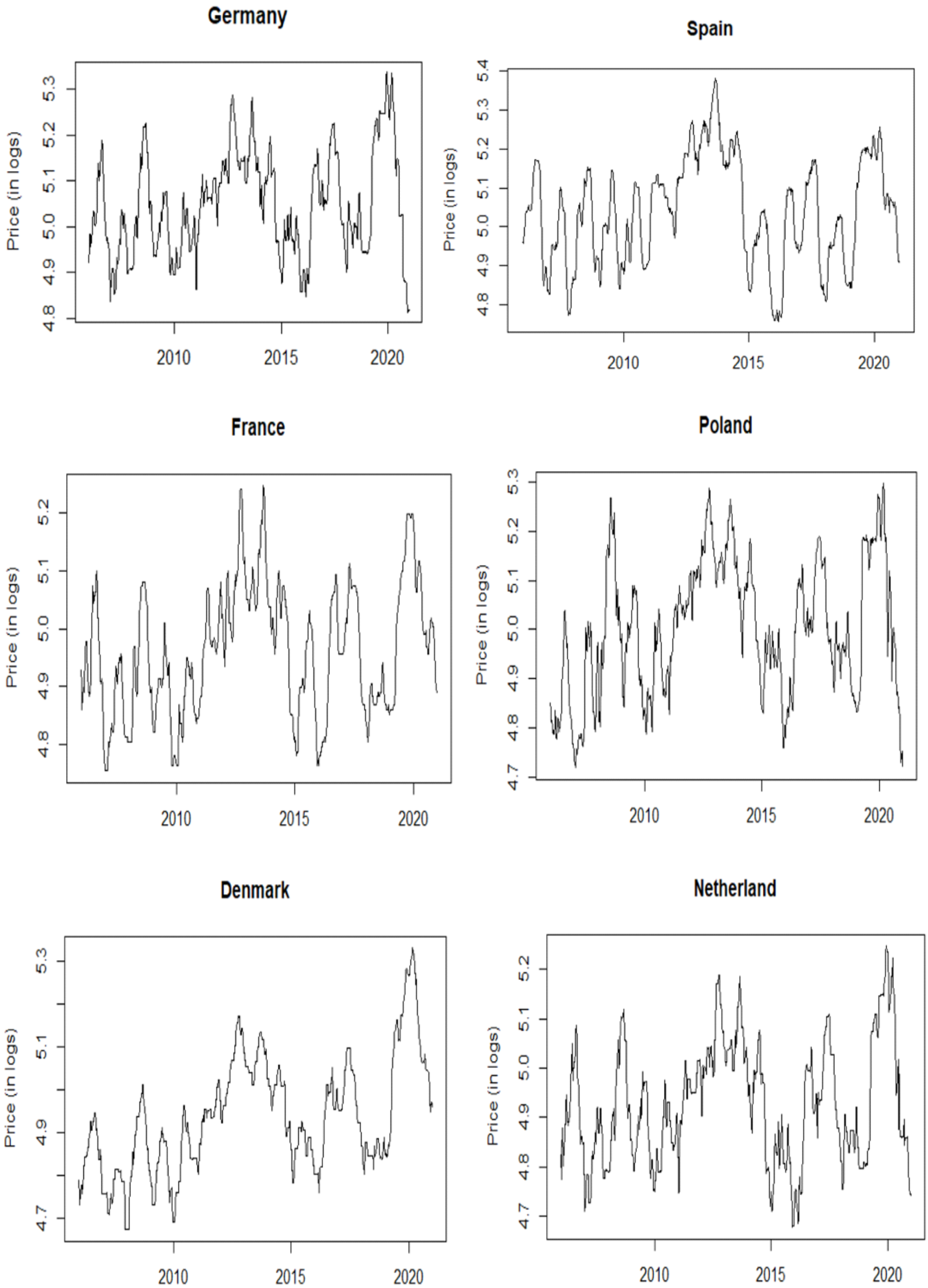

Figure A1. Price series Germany, Spain, France, Poland, Denmark, and Netherlands

PERSPECTIVES

Global Seafood Demand Growth Differences across Regions, Income Levels, and Time

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ABSTRACT

We used an index approach to calculate demand growth for seafood in 107 countries from 1984 to 2013. We used the results to calculate aggregate demand growth by income level, regionally, and globally. While seafood production has more than doubled since the mid-1980s, we showed that global demand for seafood has been higher than the global seafood supply. Demand growth for seafood varies across time, countries, regions, and income groups. The average annual seafood demand growth across countries varies between –6% and 7.5%. Global demand growth for seafood has steadily declined since the 1980s; a slowdown of demand in Asia is the main cause. South America and Africa had the highest demand growth from 2004 to 2013, while both North America and Oceania had negative demand growth in this period. High-income countries have had consistently low seafood demand growth from 1984 to 2013, while demand growth in all other income levels has been substantially larger.

Key words: Demand growth, seafood consumption, world.
JEL codes: C20, D11, Q11, Q18.

INTRODUCTION

The consumption of food of animal origin has shown significant growth over recent decades (FAO 2016). Between 1961 and 2013, global fish consumption per capita more than doubled, with its rate of growth faster than that of any other animal-based food products.[1] Fish consumption per capita increased from 9 kg to 20 kg, milk and cream consumption increased from 76 kg to 90 kg, and meat consumption increased from 23 kg to 43 kg (FAO 2016; World Cancer Research Fund 2018). Despite its growth, fish accounts for only roughly 17% of all animal protein and 6.7% of all protein consumed by humans (FAO 2018).

Changes in consumption in any market can in essence be attributed to two factors: a change in supply or a change in demand. Global fish production has more than doubled since the mid-1980s (Valderrama and Anderson 2010; FAO 2016). According to the FAO (2016), in 2013, the global supply of fish reached 141.5 million tonnes. The rapidly growing aquaculture sector has

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1. We sometimes use the term *fish* instead of *seafood*. When we say *fish*, we are specifically referring to fish rather than to seafood in general.

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been the main contributor to this supply growth (Anderson 2002; Kobayashi et al. 2015).[2] Productivity growth and increased control over the production process in the sector have played a significant role in reducing production costs and hence reducing fish prices (Asche, Roll, & Tveterås 2007). Besides the supply growth, the demand side of the sector has also played a significant role in global fish consumption and production expansion. Demand is a key determinant of aquaculture productions. In particular, the demand for specific species of seafood types is critical. In aquaculture production, species selection is determined by profit-maximizing firms subject to production costs and consumers' willingness to pay for various species. Without demand growth, any increase in consumption in the future needs to be caused by further productivity and supply growth. Demand growth leads to higher prices that increase the quantity supplied (and consumed), even if there is no productivity growth.

Empirical studies on seafood demand growth are limited. However, existing studies highlight the importance of demand growth for the expansion of both production and consumption (e.g., Asche, Roll, and Tveterås 2007; Roheim, Gardiner, and Asche 2007; Dey et al. 2008; Asche, Roll, and Trollvik 2009; Brækkan et al. 2018). For instance, Asche, Roll, and Tveterås (2007) using salmon and shrimp and Asche, Roll, and Trollvik (2009) using salmon and cod as examples argued that if there is no demand growth for a species, the production growth will be limited, even if productivity growth for that species is substantial. Asche et al. (2011) and Brækkan and Thyholdt (2014) also emphasized the role that the demand side of the market has played in the success of salmon aquaculture.

The study of demand growth has not received the same attention in the literature as supply (or productivity) growth. This could be due to several factors, including the major issue of methodological complexity. According to Asche et al. (2011) and Brækkan et al. (2018), the methodological framework to investigate demand growth is not as established in the literature as that of supply (or productivity) growth. Demand growth or contraction may occur for various reasons, including changes in consumer income, prices of substitute and complementary products, population growth, demographics, and the appearance of new information about a product (Dey et al. 2008; Brækkan and Thyholdt 2014; Brækkan et al. 2018). The existence of a multitude of factors affecting demand makes the methodological framework for studying demand growth complex and challenging.

Thus, most demand studies on seafood have focused on estimating demand elasticities for particular seafood species. The estimated demand elasticities can be used to evaluate issues such as the effects of changing prices, incomes, and the degree of substitutability between potentially competing seafood products (e.g., Dey et al. 2008; Gallet 2009; Bronnmann, Loy, and Schroeder 2016). This can help reveal how consumers respond to an increase in income, prices, and the price of substitute products. It is worthwhile to note here that, on average, seafood demand is more inelastic in high-income countries than in middle- and low-income countries (Muhammad et al. 2011). While elasticities are useful, they must be used to enable an understanding of what has happened in the past and what may happen in the future.

The main objective of this study was to estimate seafood demand growth across countries from all over the world. We used the demand index approach developed by Marsh (2003) to

2. While the supply from wild-capture production has remained stagnant, with an annual production of no more than 95 million tonnes since the 1990s, the contribution of the rapidly growing aquaculture sector to the global seafood supply has been growing. According to the FAO (2016), in 2014 the sector had a share of 44% of the total seafood supply, with a production level of 74 million tonnes.

estimate the demand growth. Based on data availability, demand growth across 107 countries is estimated using data from 1984 to 2013. To our knowledge, this paper is the first analysis of fish demand growth on a global scale.

The rest of this paper is organized as follows. The next section presents the trend of global seafood consumption and consumption differences across countries and regions. The method and data used are presented in the third and fourth sections, respectively. The fifth section presents the empirical results and discussion, with concluding remarks in the last section.

TREND OF GLOBAL SEAFOOD CONSUMPTION AND CONSUMPTION DISPARITY ACROSS COUNTRIES AND REGIONS

Significant seafood supply growth over the past few decades has enhanced the world's capacity to consume more seafood (FAO 2016). As illustrated in figure 1A, annual global seafood consumption increased from approximately 58.6 to 133 million tonnes between 1984 and 2013. This implies annual average consumption growth of 4.2% over this period. Similarly, the average global seafood consumption per capita showed an increasing trend, rising from 12.4 kg in 1984 to 20 kg in 2013 (figure 1B).

Despite the increase in global seafood consumption, the level of seafood consumption varies considerably among countries and regions. As shown in figure 2, in 2013, seafood consumption per capita across countries varied from 1 kg to more than 161 kg. Table A1 in the appendix also reports the per capita consumption in 2013.

Figure 3 shows the development of average seafood consumption per capita by region. From the figure, we can see that seafood consumption per capita grew most noticeably in East Asia and the Pacific (from approximately 28 kg to 37 kg) and South Asia (from 5.7 to 11.2 kg) from 1984 to 2013.

Over the same period, seafood consumption per capita remained static in sub-Saharan Africa (at around 11.5 kg) and in North America (at 30 kg). Between 1984 and 2013, seafood consumption per capita increased in Europe and Central Asia from 18.6 to 24.5 kg, in the Middle East and North Africa from 8.5 to 13.5 kg, and in Latin America and the Caribbean from 11.9 to 14.5 kg.

Over recent decades, China has shown remarkable growth in seafood consumption per capita, increasing from 14.4 kg in 1993 to 38 kg in 2013. Total fish consumption in China is also far above

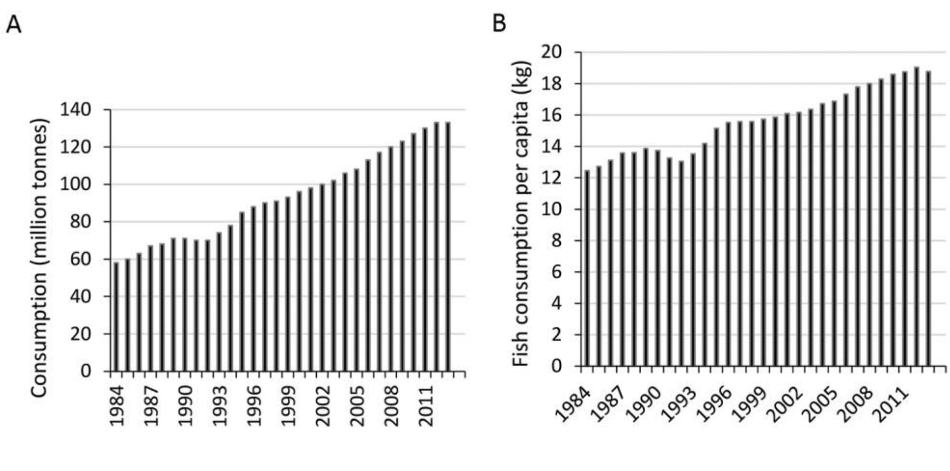

Figure 1. Trend of Global Seafood Consumption and Consumption Per Capita, from 1984 to 2013
Source: Authors' plots using data from the FAOSTAT database.

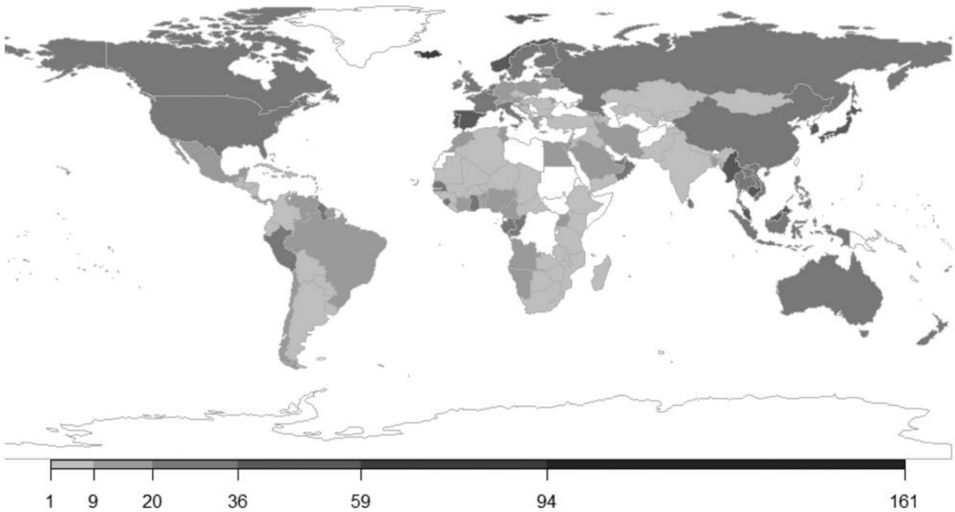

Figure 2. Seafood Consumption Per Capita for 157 Countries in the World, in 2013. Note that white in the figure represents missing data. A color version of this figure is available online.
Source: Authors' plots using data from the FAOSTAT database.

that of all other countries. For instance, in 2013, the volume of seafood consumed in China was approximately 49 million tonnes. This number is higher than the volume of seafood consumed by the top nine fish-consuming countries in the world (excluding China) combined in the same year (FAO 2016). It is worth mentioning here that China has also been responsible for most of the growth in the world per capita seafood supply in recent decades. This is owing to the drastic expansion in its seafood production, from aquaculture in particular, with a significant share of this production being exported (FAO 2016).

Various factors are cited in the literature as possible causes of the seafood consumption differences across countries and regions. Some of the main factors include food eating habits, culture, accessibility, income, population growth, and lifestyle changes (FAO 2016).

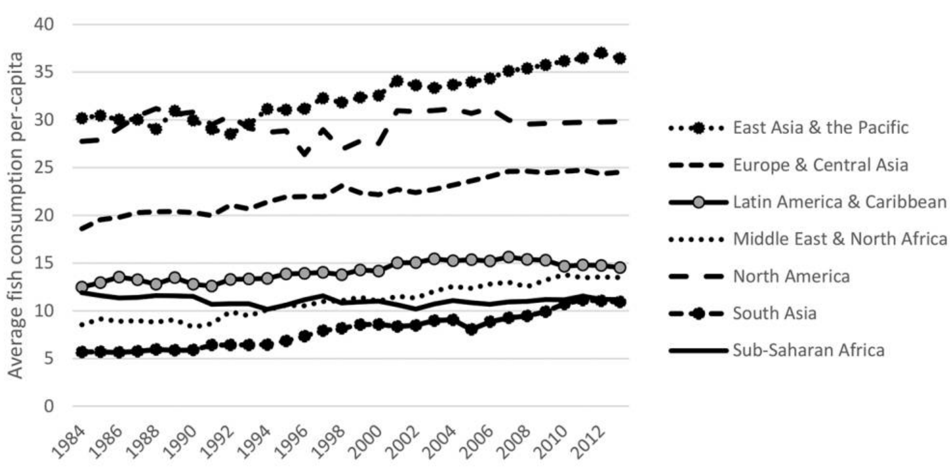

Figure 3. Seafood Consumption Per Capita, from 1984 to 2013
Source: Authors' plots using data from the FAOSTAT database.

METHOD

In this study, we use an index approach developed by Marsh (2003) for measuring demand growth. The approach measures demand shifts in the price direction, where the demand shift can be interpreted as a shift in consumers' willingness to pay for a given quantity of a product. However, depending on whether a price variable or a quantity variable is exogenous, the demand shift can be measured as either a quantity shift or a price shift (Asche et al. 2011).

Brækkan and Thyholdt (2014) and Brækkan et al. (2018) argued that the choice to measure demand shifts in the quantity or price direction is simply a matter of preference. They stated that any shift in demand is a movement of the demand schedule between two periods, and a demand shift measured using this approach is a local measure of the size of this movement. As a result, measuring this movement vertically or horizontally does not depend on price or quantity being exogenously given. For any shift in the quantity direction (horizontally), the corresponding shift in the price direction (vertically) can easily be computed (Sun and Kinnucan 2001; Brækkan and Thyholdt 2014).

Following Brækkan et al. (2018), the derivation of the approach is presented below. Figure 4 illustrates a demand shift in the quantity direction (i.e., horizontally). In the figure, suppose that the demand schedule in period t is denoted by D_t and the demand schedule in period $t + 1$ is denoted by D_{t+1}. Moreover, let Q_t and P_t be the quantity and price in period t, and Q_{t+1} and P_{t+1}. be the quantity and price in period $t + 1$.

If there is no shift in demand from period t to $t + 1$, the expected quantity demanded given the observed price P_{t+1} would be at point c. Denote this expected (or predicted) quantity demand at point c by $Q|_{E|D=D_t}$. The horizontal distance between $Q_{E|D=D_t}$. and the actual quantity demanded Q_{t+1} is the absolute shift in demand. That is, the absolute demand shift is the horizontal distance between the demand schedules in periods t and $t + 1$. As it is customary to express demand shifts in relative (percentage) terms, we express the absolute shifts in demand here in relative terms.

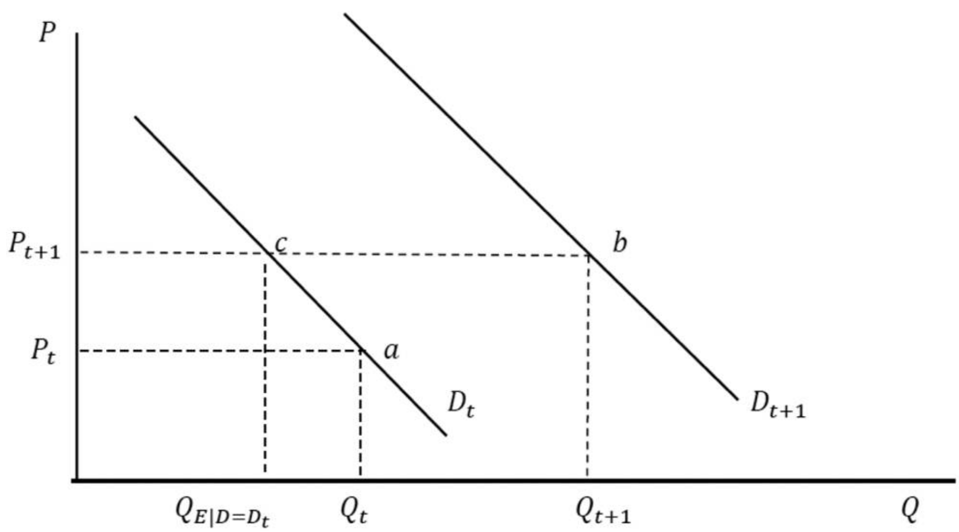

Figure 4. Horizontal Shift in Demand between Two Periods

Previous studies of demand shift using this approach express the shifts relative to the expected quantity (Asche et al. 2011; Marsh 2003). However, as argued in Brækkan and Thyholdt (2014), it is also possible to specify the shift in demand relative to the quantityn period t (i.e., Q_t). The latter implies that, for instance, a 10% increase in demand can be interpreted as a 10% increase in the quantity demanded relative to the quantity in period t, given the price in period t. As Brækkan and Thyholdt (2014) claimed, this calculation is consistent with the specification of horizontal shifts in demand in equilibrium displacement models (Muth 1964; Alston, Norton, and Pardey 1995). Following Brækkan and Thyholdt (2014), we specify the shift in demand relative to the quantity in period t. Therefore, theorizontal relative shift in demand from period t to period $t+1$, denoted here by η, can be given by the following:

$$\eta = \frac{Q_{t+1} - Q_{E|D=D_t}}{Q_t}. \tag{1}$$

Some adjustment of equation 1 yields

$$\eta = \frac{Q_{t+1} - Q_t}{Q_t} - \frac{Q_{E|D=D_t} - Q_t}{Q_t} \equiv Q_t^* - Q_E^*, \tag{2}$$

where $(Q_{t+1} - Q_t) / Q_t = Q_t^*$ is the relative change in quantity, and $(Q_{E|D=D_t} - Q_t) / |Q_t = Q_E^*$ is the relative difference between the expected quantity in period $t+1$ and the observed quantity in period t. Given a demand elasticity, denoted by ε, the expected quantity change Q_E^* can be obtained as the following:

$$Q_E^* = \frac{\varepsilon(P_{t+1} - P_t)}{P_t} \equiv \varepsilon P_t^* . \tag{3}$$

Now, by substituting equation 3 into equation 2, the relative horizontal shift in demand given as follows:

$$\eta = Q_t^* - \varepsilon P_t^* . \tag{4}$$

The demand shift in the price direction can be obtained by dividing the horizontal shift in demand by the negative of the corresponding elasticity of demand (Sun and Kinnucan 2001; Brækkan and Thyholdt 2014). This vertical demand shift can be expressed as the following:

$$\eta^v = \frac{\eta}{-\varepsilon} = -\frac{Q_t^*}{\varepsilon} + P_t^* . \tag{5}$$

For example, a vertical demand shift of 5% would imply a 5% increase in consumers' willingness to pay for a given quantity of a product. The price and quantity direction measures of demand shift will be identical if the elasticity of demand is equal to –1. In this study, as explained above, following Brækkan and Thyholdt (2014), we measure the shifts in demand in the quantity direction.

The above model has some advantages. First, the model is suitable for measuring the shift in demand between two different periods (Brækkan et al. 2018). Moreover, the model measures aggregated demand shifts caused by various factors (e.g., logistics, increased variety of products, income growth, changes in tastes and preferences), which are impossible or at least difficult to measure using other econometric demand models, because of limited data accessibility or model specification issues. Furthermore, the approach has been used extensively in the literature on demand growth (e.g., Sun and Kinnucan 2001; Marsh 2003; Asche et al. 2011).

Despite the above strengths, like any model, it has shortcomings. The results are highly dependent on the value of the demand elasticity. For instance, the true demand elasticity might not be constant over time. We perform sensitivity analysis using different elasticity values to check the robustness of the estimated results in this study.

DATA AND DEMAND ELASTICITY

The data required for the analysis are price and per capita seafood consumption. This means that the estimated demand growth should be interpreted as changes in per capita seafood demand.

The annual aggregate seafood consumption for each country, over the period from 1984 to 2013, is obtained from FAOSTAT (FAO database). The consumption data reported by the FAO are apparent consumption of fish and fishery products. They are measured as the total quantity of FAO-reported fish and fishery products produced in a country added to the total quantity imported and adjusted to any change in stocks minus exports and nonfood uses.[3] Then per capita consumption is calculated by dividing the aggregate consumption by the total population in each country. The population data are obtained from the World Bank database (World Bank 2016). Considering that seafood products are highly diversified in quality and price, using aggregate data has some drawbacks. However, given the lack of more detailed data, this is probably the best way to proxy country-level seafood consumption. Moreover, other studies (e.g., Muhammad et al. 2011; Nguyen and Kinnucan 2018) have followed a similar approach.

The data available to compute the consumer-level seafood prices in each country are aggregate import quantities and values, which are obtained from the FAOSTAT database (FAO 2015). In this study, we used import prices for three reasons. First, most previous demand studies have been carried out using trade data. As a result, many of the estimated demand elasticities in the literature are based on trade data (Asche et al. 2011). Second, it is relatively easy to get several years' worth of trade data for most countries. Third, at least at present, there is no better alternative to import price to proxy the domestic fish price in most countries.

Using import price to proxy consumer-level price obviously has some drawbacks. First, domestic consumption in most developing countries is mostly supplied by local production, not by imports. Second, developing countries, particularly countries in East and Southeast Asia, are mainly exporters of seafood. However, it is also true that developing countries export high-value seafood to developed countries, while retaining and importing lower value seafood products for their domestic supply (Tran et al. 2019). Nonetheless, because of rising consumer incomes, consumers in developing countries are diversifying the types of seafood they consume through import. This has caused a surge in seafood imports to developing countries (FAO 2016). For developed countries, a sizable and growing share of the seafood consumed is supplied through imports; using import prices in these countries therefore seems reasonable. Thus, while interpreting the results, especially those from developing countries, we must keep in mind the uncertainty regarding the use of import prices as a proxy for consumer seafood prices.

3. The FAO apparent consumption data are compiled from various sectors (e.g., production and trade). As a result, because of problems associated with variable or uncertain conversion factors and inadequate knowledge on stock changes, some uncertainties are more likely that apparent consumption might not reflect changes in consumption habits in a country. For more detail, please refer to http://www.fao.org/cwp-on-fishery-statistics/handbook/socio-economic-data/food-balance-sheets/en/.

For each country, the unit price is expressed in the local currency, with exchange rate data obtained from the World Bank database (World Bank 2016). The prices are deflated using the respective country's consumer price indices, which are also extracted from the World Bank database.

To compute the shifts in demand, we must have appropriate demand elasticity estimates for each country considered in this study. Muhammad et al. (2011) estimated demand elasticities for seafood and fishery products for most countries in the world using a consistent methodology and data. Although there are several other studies estimating demand elasticities, they are often limited to a specific species or country. Muhammad et al.'s (2011) elasticities of demand estimates for most countries of the world are appropriate to use in this study. They estimated price and income elasticity of demand for broad food categories, including seafood, based on cross-country demand analyses conducted using International Comparison Program data from 2005. The study used a two-stage demand system to estimate the elasticities for 144 countries. Three types of own-price elasticities are reported in the paper: the Frisch deflated own-price elasticity, the Slutsky (compensated) own-price elasticity, and the Cournot (uncompensated) own-price elasticity.

The Frisch deflated own-price elasticity of a good is computed when price changes and income are compensated for to keep the marginal utility of income constant. The Slutsky (compensated) own-price elasticity measures the change in demand for a good when the price of that good changes, while real income remains unchanged. The Cournot (uncompensated) own-price elasticity refers to the situation where own-price changes, nominal income remains constant, and real income changes. The use of each measure of elasticity depends on the needs of the researcher. Since there is no compensation to keep the marginal utility of income constant in real life, Cournot (uncompensated) own-price elasticity is appropriate to use in this study. However, since only Frisch own-price elasticity is reported in the food subcategories in Muhammad et al's (2011) paper, we use Frisch own-price elasticity in this study. The Cournot and Frisch elasticities are relatively close for high-income countries, but can be different for low- and middle-income countries. Thus, using Frisch own-price elasticity in low- and middle-income countries might affect the results. Table A2 in the appendix reports these price elasticities.

As one can observe from table A2 in the appendix, there are variations in the price elasticities across countries, ranging from 0.19 to 0.55 in absolute value. The elasticity of demand computed based on income level is also reported in Muhammad et al.'s (2011) paper. The reported elasticity estimates show that demand is more inelastic in high-income countries than in low- and middle-income countries. As we included some countries in this study whose own-price elasticities are not reported in Muhammad et al.'s (2011) paper, we use the own-price elasticity for these countries' income level. This permits us to estimate demand growth for seafood in 107 countries in the period from 1984 to 2013.

RESULTS AND DISCUSSION

Figure 5 reports the annual average seafood demand growth across countries over the study period.[4] As shown in figure 5, it ranges from –6% to 7.5%.

The estimated average seafood demand growth in China, which is by far the largest seafood-consuming country in the world, is 6.3%. In general, our results show that there are substantial differences in demand growth across countries.

4. The reported demand growth should be interpreted as per capita seafood demand growth, as it is calculated using per capita seafood consumption.

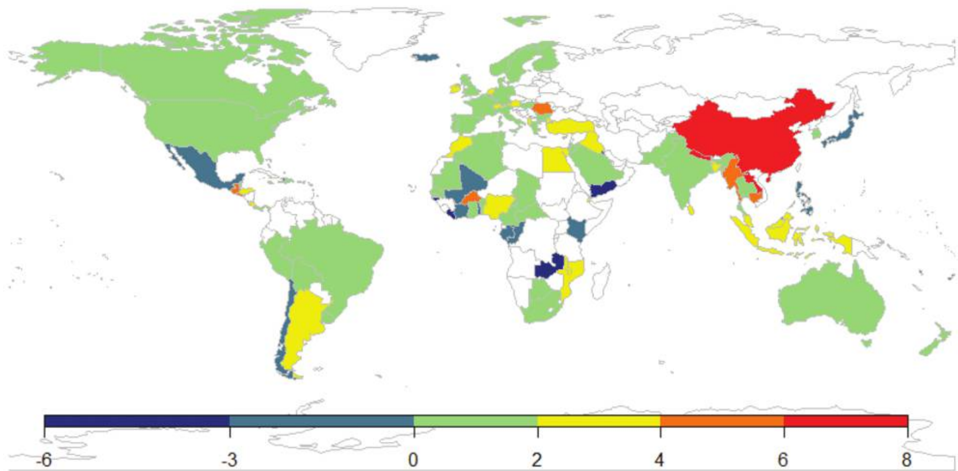

Figure 5. Average Annual Seafood Demand Growth (in %) across Countries, from 1984 to 2013. Note that white in the figure represents missing data.
Source: Authors' plot.

Next, we calculated aggregate demand growth based on the income categories of countries (high income, upper middle income, lower middle income, and low income), to see whether there are differences in seafood demand growth based on income category. To calculate the aggregate demand per income category, we used the total population in every country as weights.[5] Figure 6 reports the results.

As shown in the figure, the aggregate demand growth for the upper middle income category was higher than that of the other income categories. This result seems reasonable for at least two reasons. First, as we discussed in the theoretical section, income is among the main driving factors of demand growth. High-income growth over recent decades has mainly been observed in developing countries, particularly in Asia. Furthermore, the presence of China in the upper middle income category is a major contributor to the substantial growth.

Despite the effect of income on demand growth, it is slightly surprising to observe that the aggregate seafood demand index in countries in the low-income category was higher than in the lower middle income and higher income categories. Since there are factors other than income that can affect seafood demand growth, this result might still hold. However, it is quite surprising to see the slow growth of the aggregate seafood demand index for the high-income category. This might be an indication that consumers in high-income countries have diverse sources of protein other than seafood, or that seafood demand in higher income countries is reaching a point of saturation. Nevertheless, because of the relative sizes of different markets, 1% growth in demand from the high-income countries translates to a much larger increase in quantity demanded than 1% growth in demand from low- and middle-income countries.

Next, we calculated the aggregate seafood demand growth based on the countries' regional classification and continents. Figures 7 and 8 show the trends of the aggregate seafood demand

5. Since our demand growth calculation was based on per capita seafood consumption, it is reasonable to use the total population in every country as the weights. Here and in the subsequent section, demand indices are calculated relative to the base year 1984 (i.e., 1984 = 100).

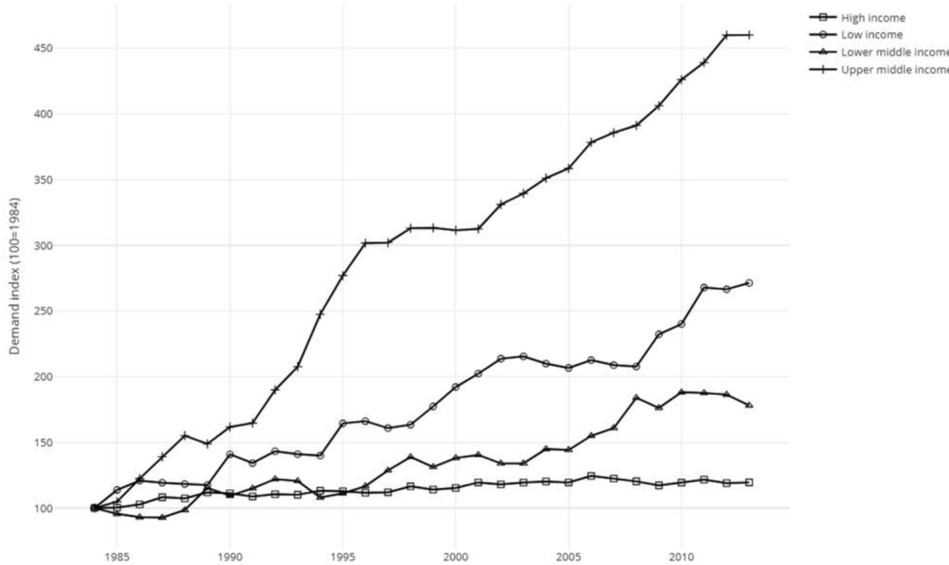

Figure 6. Aggregate Global Seafood Demand Index Calculated Based on Countries' Income Category, from 1984 to 2013
Source: Authors' plot.

indices based on regions and continents, respectively. As shown in figure 7, the aggregate seafood demand index in East Asia increased continuously over the study period, and the growth was far higher than in the other regions.

As shown in figure 8, the aggregate seafood demand index grew much faster in Asia than on any other continents. The trends of the aggregate seafood demand indices on other continents

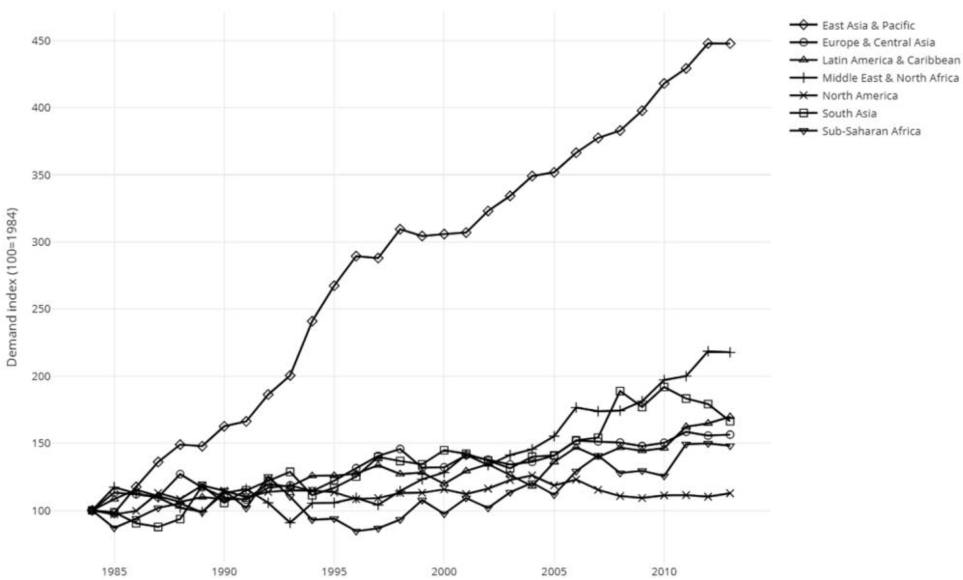

Figure 7. Aggregate Global Seafood Demand Index Calculated Based on Region, from 1984 to 2013
Source: Authors' plot.

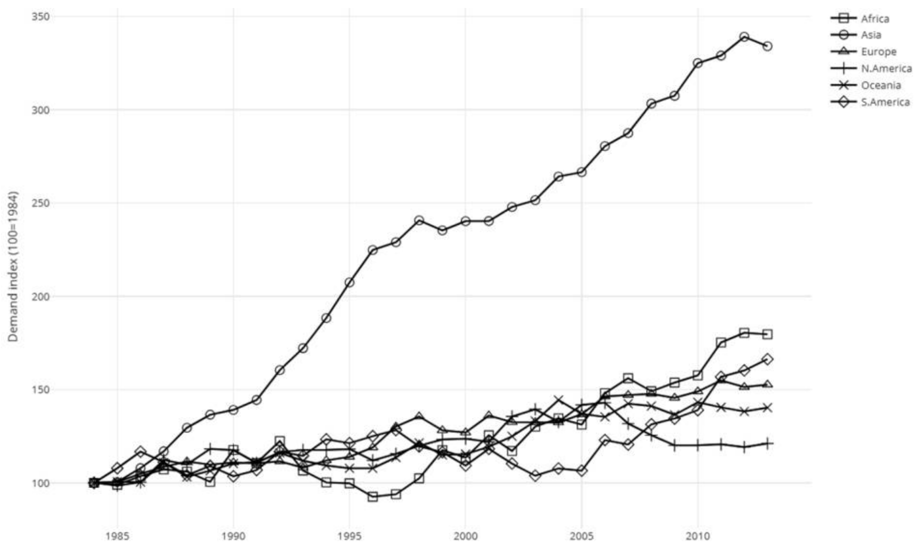

Figure 8. Aggregate Global Seafood Demand Index Calculated Based on Continents, from 1984 to 2013
Source: Authors' plot.

(except for North America) were similar. In North America, a downward trend has been observed, particularly after 2005.

Finally, we calculated the aggregate global demand growth. For the sake of clarity and for illustration purposes, we illustrate the ratio of the global demand index and quantity index (hereafter, global demand-quantity index). For the purpose of comparison, we also calculated a global seafood price index, where each country is weighted by population. In figure 9, the top graph is the global demand-quantity index, while the bottom graph is the price index.

As clearly shown in the figure, the global demand-quantity index increased over the study period. This indicates that seafood demand grew more than quantity supplied, since the global demand-quantity index could only increase if there had been higher growth in demand than in quantity supplied. Larger growth in demand than in quantity supplied necessitates an increase in

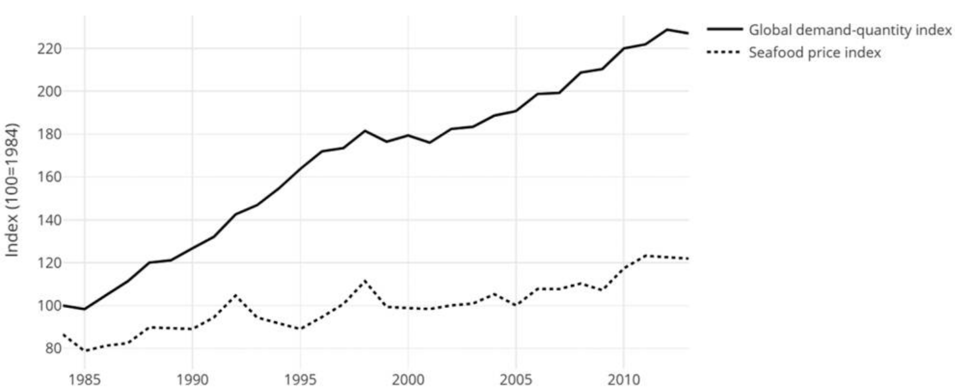

Figure 9. Global Demand-Quantity Index (the Ratio of Global Seafood Demand and Global Seafood Production Indices) and Global Seafood Price Index, from 1984 to 2013
Source: Authors' plot.

Table 1. Annual Average Demand Growth in Different Periods (%)

	1984–1993	1994–2003	2004–2013	1984–2013
Asia	5.64	3.05	2.39	4.22
Africa	0.59	2.70	2.96	2.00
Europe	0.90	1.67	1.34	1.44
North America	1.63	1.71	–0.91	0.64
South America	1.35	–1.70	4.44	1.71
Oceania	1.15	2.00	–0.30	1.16
World	4.28	2.69	2.27	3.50

price, as illustrated by the increase on our global seafood price index. In other words, the substantial global growth in seafood supply over recent decades has coincided with even larger growth in global seafood demand.

It might be interesting to observe the demand growth across continents and globally over different periods. Table 1 reports the average annual demand growth for each continent and globally over different periods.

As can be seen from the table, the global average seafood demand growth varied between periods, with the rate of growth decreasing over time. Specifically, the global annual average demand growth was 4.28%, 2.69%, and 2.27% over the periods 1984–93, 1994–2003, and 2004–13, respectively. Over the whole study period, the global annual average demand growth was approximately 3.5%. Likewise, the average demand growth across continents also varied over different periods. The average demand growth in Asia, which is historically the largest seafood-producing and -consuming continent, decreased over time.

Another interesting point to observe from the table is the rate of average growth in Africa. The average growth rate in Africa was very low over the period from 1984 to 1993 compared with the other continents, but the rate of growth increased over time. The annual average demand growth in Africa in the period 2004–13 was 2.96%. The average growth rate in South America was comparably higher in recent years than on the other continents. This result is consistent with Garlock et al.'s (2020) finding that aquaculture production in some non-Asian countries has been growing more rapidly than in the major Asian producers in recent years. This may have facilitated growth in demand in these countries. However, over the entire period from 1984 to 2013, only the average growth rate in Asia (4.22%) was higher than the global annual growth rate (3.50%).

In general, the above results show that there are substantial differences in demand growth across countries, income groups, regions, continents, and over time.

SENSITIVITY ANALYSIS

It may not be likely that the elasticity of demand for a commodity is constant over time. Therefore, checking the robustness of the above results using different elasticity values is necessary. We recomputed the shift in demand in every country by varying the elasticities by –0.2 and 0.2, where –0.2 is the difference between the average demand elasticity for seafood in low-income and high-income countries as reported in Muhammad et al. (2011). The argument for using this number is based on the assumption that, over time, seafood demand in developing countries may have become more inelastic as income increased.

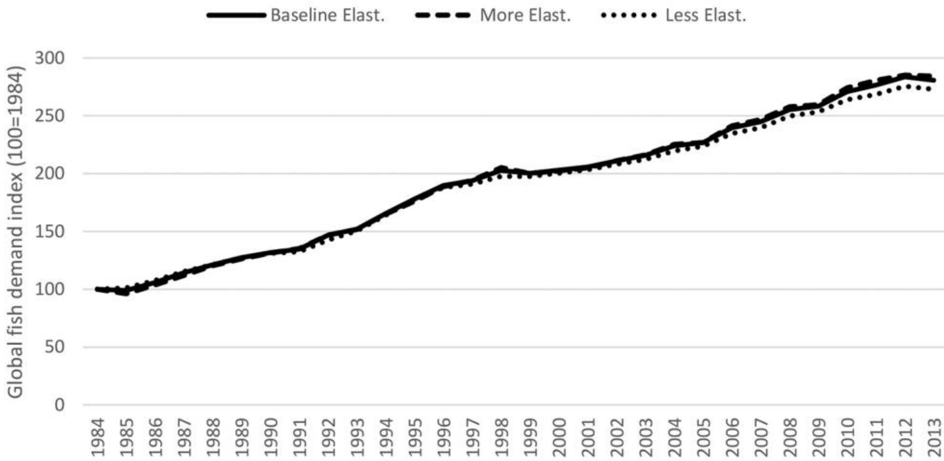

Figure 10. Global Seafood Demand Indices Calculated with Different Seafood Demand Elasticities. Baseline elasticity refers to the elasticity from Muhammad et al.'s (2011) paper, while the other two are the baseline elasticities plus or minus 0.2, from 1984 to 2013.

Source: Authors' plot.

Figure 10 shows the global demand index for each elasticity value. In the figure, the elasticity obtained from Muhammad et al. (2011) is referred to as the baseline elasticity, while the baseline elasticity minus and plus 0.2 are referred to as more elastic and less elastic demand, respectively. As shown in the figure, the estimated global demand indices are similar regardless of the elasticities used in the estimation. This indicates that the results are not sensitive to the choice of elasticity values.[6]

CONCLUSION AND IMPLICATIONS

Any growth in seafood consumption necessitates a growth in quantity produced (and/or caught). However, growth in quantity produced can be caused by both an increase in supply and an increase in demand. While the growth in seafood production in recent decades is well documented, the demand side of seafood consumption has received less attention. In this paper we show that global demand growth between 1984 and 2013 was higher than the global growth in supply. This implies that demand growth has been a vital contributor to increased global seafood consumption in recent decades.

While global demand growth for seafood has been substantial, it has also lost pace from average annual global demand growth of 4.28% between 1984 and 1993, to annual growth of 2.27% between 2004 and 2013. The slowdown of demand growth in Asia has been the main contributor to slower growth in global demand for seafood. South America and Africa, while still miniscule compared with Asia in terms of seafood consumption, had the largest demand growth between 2004 and 2013.

6. We also did a robustness check using different elasticities for some of the major seafood-producing and -consuming countries, and our analysis showed that the results are not particularly sensitive to the choice of elasticity values.

High-income countries show remarkably low demand growth, and may have reached a point of saturation. All other income categories had substantially higher demand growth than did high-income countries.

The results raise numerous questions, some of which are the following: Why has global seafood demand been slowing down since the 1980s? Why has demand in Africa and South America been so high in recent years? Why has seafood demand growth in high-income countries been consistently low for 30 years? Why is fish demand contracting in Oceania and North America?

In this study we show demand growth for seafood over time, globally, across regions, and across income groups. There is considerable variation on all measures, but it is not within the scope of this study to explain or understand the causes behind these. While one could therefore argue that the results of this study raise more questions than they answer, knowing which questions to ask can also be valuable knowledge. Hopefully this study has contributed in terms of both providing a thorough picture of demand growth for seafood, and inspiring future research to answer some of the questions that the results have raised.

APPENDIX

Table A1. Seafood Consumption Per Capita Based on Countries, in 2013

Country	Per Capita	Country	Per Capita	Country	Per Capita	Country	Per Capita
Albania	5.3	Czech Rep.	8.8	Kazakhstan	5.1	Portugal	54.5
Algeria	4.0	Denmark	23.2	Kenya	4.2	Romania	6.8
American Samoa	44.4	Djibouti	3.7	Kuwait	12.9	Russia	22.8
Angola	12.7	Dominica	20.4	Kyrgyzstan	2.2	Rwanda	4.1
Argentina	6.9	Dominican Rep.	8.2	Laos	20.7	Saudi Arabia	13.0
Armenia	4.5	Ecuador	8.2	Latvia	24.1	Senegal	24.0
Australia	26.3	Egypt	20.2	Lebanon	9.7	Sierra Leone	28.4
Austria	13.9	El Salvador	6.9	Lesotho	0.8	Solomon Islands	33.5
Azerbaijan	2.1	Estonia	14.4	Liberia	4.2	South Africa	6.1
Bahamas	27.6	Ethiopia	0.2	Madagascar	4.6	South Korea	51.6
Bangladesh	19.1	Fiji	36.1	Malawi	7.2	Spain	42.7
Barbados	40.5	Finland	36.3	Malaysia	59.0	Sri Lanka	26.5
Belarus	16.3	France	32.6	Maldives	160.5	Suriname	16.4
Belgium	24.9	Gabon	32.8	Mali	7.2	Swaziland	1.3
Belize	13.2	Gambia	23.9	Malta	33.0	Sweden	31.9
Benin	13.8	Georgia	12.2	Mauritania	9.0	Switzerland	17.8
Bolivia	2.3	Germany	12.9	Mauritius	23.2	Tanzania	5.4
Bosnia and Herzegovina	4.5	Ghana	25.8	Mexico	10.4	Thailand	24.4
Botswana	3.8	Greece	19.6	Moldova	12.7	Togo	11.3
Brazil	10.8	Grenada	28.4	Mongolia	0.7	Trinidad and Tobago	23.8
Brunei Darussalam	48.1	Guatemala	1.3	Morocco	17.6	Tunisia	13.6
Bulgaria	6.9	Guinea	9.4	Mozambique	7.9	Turkey	6.0
Burkina Faso	6.7	Guinea-Bissau	1.4	Myanmar	56.3	Uganda	12.5
Cabo Verde	11.5	Guyana	31.2	Namibia	11.4	United Arab Em.	23.6
Cambodia	41.7	Haiti	4.8	Nepal	2.2	United Kingdom	20.5
Cameroon	15.8	Honduras	3.7	Netherlands	22.1	United States	21.8
Canada	22.5	Hungary	5.1	New Zealand	25.3	Uruguay	7.5

Table A1. (*Continued*)

Country	Per Capita	Country	Per Capita	Country	Per Capita	Country	Per Capita
Central African Rep.	8.0	Iceland	93.7	Nicaragua	4.9	Uzbekistan	0.7
Chad	4.6	India	4.9	Niger	2.6	Vanuatu	32.0
Chile	12.6	Indonesia	27.9	Nigeria	16.4	Venezuela	9.6
China	36.2	Iran	10.0	Norway	51.7	Vietnam	32.7
Colombia	6.4	Iraq	3.3	Oman	22.0	Yemen	2.3
Congo	24.8	Ireland	22.1	Pakistan	1.9	Yemen	2.3
Costa Rica	13.4	Israel	22.3	Panama	13.1	Zambia	6.0
Cote d'Ivoire	14.3	Italy	25.4	Paraguay	3.9	Zimbabwe	2.7
Croatia	19.1	Jamaica	23.4	Peru	22.0		
Cuba	5.5	Japan	48.5	Philippines	31.6		
Cyprus	21.6	Jordan	4.6	Poland	10.7		

Source: Authors' computation based on data extracted from the FAOSTAT database.

Table A2. Frisch Own-Price Elasticities of Demand for Fish

Country	Elasticity	Country	Elasticity	Country	Elasticity
Albania	–0.431	Georgia	–0.435	Norway	–0.267
Angola	–0.512	Germany	–0.269	Oman	–0.386
Argentina	–0.389	Ghana	–0.500	Pakistan	–0.463
Armenia	–0.419	Guinea	–0.511	Paraguay	–0.440
Australia	–0.279	Guinea-Bissau	–0.523	Peru	–0.425
Azerbaijan	–0.444	Hungary	–0.352	Philippines	–0.455
Bahrain	–0.340	Iceland	–0.267	Poland	–0.363
Bangladesh	–0.490	India	–0.484	Portugal	–0.316
Belarus	–0.398	Indonesia	–0.456	Qatar	–0.320
Belgium	–0.278	Iran	–0.395	Romania	–0.399
Benin	–0.496	Iraq	–0.479	Russia	–0.390
Bhutan	–0.476	Ireland	–0.287	Rwanda	–0.512
Bolivia	–0.459	Israel	–0.328	Sao Tome and Principe	–0.479
Bosnia and Herzegovina	–0.399	Italy	–0.287	Saudi Arabia	–0.401
Botswana	–0.458	Japan	–0.279	Senegal	–0.486
Brazil	–0.419	Jordan	–0.430	Serbia	–0.402
Brunei Darussalam	–0.352	Kazakhstan	–0.403	Sierra Leone	–0.511
Bulgaria	–0.390	Kenya	–0.493	Singapore	–0.325
Burkina Faso	–0.504	Korea, Rep.	–0.351	Slovakia	–0.353
Burundi	–0.538	Kuwait	–0.308	Slovenia	–0.327
Cape Verde	–0.459	Kyrgyzstan	–0.462	South Africa	–0.415
Cambodia	–0.483	Laos	–0.490	Spain	–0.281
Cameron	–0.481	Latvia	–0.374	Sri Lanka	–0.454
Canada	–0.271	Lebanon	–0.364	Sudan	–0.470
Central African Rep.	–0.507	Lesotho	–0.476	Swaziland	–0.441
Chad	–0.512	Liberia	–0.540	Sweden	–0.286
Chile	–0.402	Lithuania	–0.356	Switzerland	–0.254
China	–0.480	Luxembourg	–0.208	Syrian Arab Rep.	–0.445
China, Hong Kong SAR	–0.285	Macedonia	–0.405	Taiwan	–0.297
Colombia	–0.432	Madagascar	–0.502	Tanzania, Rep.	–0.504
Congo	–0.501	Malawi	–0.527	Thailand	–0.433
Congo, Dem. Rep.	–0.551	Malaysia	–0.422	Togo	–0.500
Croatia	–0.363	Maldives	–0.469	Tunisia	–0.425
Cyprus	–0.272	Mali	–0.509	Turkey	–0.409
Czech Rep.	–0.333	Malta	–0.309	Uganda	–0.504

Table A2. (*Continued*)

Country	Elasticity	Country	Elasticity	Country	Elasticity
Côte d'Ivoire	–0.491	Mauritania	–0.491	Ukraine	–0.418
Denmark	–0.288	Mexico	–0.371	United Kingdom	–0.258
Djibouti	–0.485	Moldova, Rep.	–0.445	United States	–0.191
Ecuador	–0.434	Mongolia	–0.479	Uruguay	–0.398
Egypt	–0.434	Montenegro	–0.415	Venezuela	–0.417
Equatorial Guinea	–0.446	Morocco	–0.463	Vietnam	–0.484
Estonia	–0.356	Mozambique	–0.516	Yemen	–0.480
Ethiopia	–0.523	Namibia	–0.458	Zambia	–0.517
Fiji	–0.441	Nepal	–0.495	Zimbabwe	–0.528
Finland	–0.297	Netherlands	–0.266		
France	–0.273	New Zealand	–0.299		
Gabon	–0.451	Niger	–0.524		
Gambia	–0.518	Nigeria	–0.489		

	Elasticity Value
Low-income countries, average	–0.478
Middle-income countries, average	–0.378
High-income countries, average	–0.277

Source: Muhammad et al. (2011).

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