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PRICE TRANSMISSION FOR ORGANIC AND CONVENTIONAL MILK PRODUCTS IN SWEDEN

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Abstract

Although much empirical work addresses the efficiency of food supply chains by studying price transmission, studies on quality-differentiated food are scarce, and particularly for organic food vis-á-vis conventional food. This study adds to this scarce literature by analysing wholesale to retail price transmission for organic and conventional milk in the Swedish milk sector, using time-series analysis applied to monthly price data for the period Jan 2007–Nov 2017. Estimations are performed using the non-linear ARDL model which allows for asymmetric cointegration of prices and a simultaneous analysis of short- and long-run asymmetry, the latter of which has been largely overlooked in previous studies. In the case of conventional milk, results indicate positive asymmetries both in the short-run and the long-run. For organic milk, the long-run positive asymmetry is smaller and not statistically significant in all specifications. Organic consumers are therefore likely to experience smaller differences between surplus losses and gains, following positive and negative wholesale price changes, respectively.

JEL: C22, L11, Q13

Keywords: Price transmission, organic food, non-linear ARDL.

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1. Introduction

The aim of this paper is to analyse wholesale to retail price transmission in the Swedish milk sector and whether this transmission differs for organic and conventional milk products. Specifically, the paper answers questions regarding the extent to which price changes are transferred along the supply chain, the speed of these transfers, and whether there are asymmetries with regards to positive and negative price changes. While a majority of empirical studies on agricultural markets show evidence of positive asymmetries, meaning that a price increase is transmitted faster or to a larger extent to the next stage of the supply chain compared to a price decrease (Meyer & Von Cramon-Taubadel, 2004), most previous studies focus on conventional, i.e., non-organic, food. Since organic and conventional food products are qualitatively differentiated, price transmissions may differ (Würriehausen et al., 2015), meaning that price shocks (due to changes in, e.g., subsidies, climate) may affect organic and conventional markets differently.

Organic food production is commonly viewed as a system capable of providing a public good in terms of increased environmental protection and animal welfare, as well as a private good by responding to an increase in consumer demand for a more sustainable diet (European Commission, 2018). Increased demand for organic food is a global trend, with shares of organic retail sales reaching 5% in Germany and the US in 2017, and 9–13% in Austria, Sweden and Denmark (Willer & Lernoud, 2019). In the EU, labelling standards and measures to increase organic production have been in place since the 1990's (European Commission, 2007). Costs pertaining to organic conversion and production are generally financed by subsidies within the Common Agricultural Policy (CAP), and by private consumers willing to pay an organic price premium (European Commission, 2004). Further, there has also been discussions on EU level to increase demand for organic food by reducing organic price premiums through differentiated VAT rates for organic and conventional products or input factors (Oosterhuis et al., 2008).

The effects of changes in subsidies or taxes on demand depend on how changes in price are passed through the supply chain. To this end, a well-functioning and competitive supply-chain is necessary for demand to increase. However, agricultural markets tend to be characterised by high levels of concentration further down the supply chain, with the Swedish milk sector being no exception (Persson, 2011). Retailers may therefore enjoy considerable market power, resulting in market inefficiencies regarding price transmission (McCorriston et al., 2001). Importantly, the price transmission may differ for conventional and organic products. From a policy perspective, it is thus important to consider price transmission for both organic and conventional milk when designing policies aimed at these markets. Analysing price transmission can also yield insights into the efficiency of a supply chain, affecting not only the welfare of consumers, but also other actors within the supply chain, requiring a tenable distribution of the value-added along the chain (European Commission, 2009).

A few studies address price transmission for food products that are differentiated in terms of, e.g., quality-grades (Surathkal & Chung, 2017), or pricing schemes (Loy et al., 2015). However, studies on price transmission for organic vs non-organic food are generally lacking (Antonioli et al., 2019; Darbandi & Saghaian, 2018), and the few existing studies do not address asymmetries (see, e.g., Darbandi and Saghaian (2018)), or study markets where organic market shares are small and organic food is mainly sold in specialized organic stores (see, e.g., Antonioli et al. (2019)). Other market characteristics such as consumer demand and concentration levels may also differ between countries, thus hampering generalisation. The Swedish market for organic food is considered highly mature in terms of sales and distribution compared to many other countries (Furemar, 2004; Willer & Lernoud, 2019). This paper therefore adds to the scarce literature on price transmission for quality-differentiated food products, and for organic vs non-organic products particularly. Further, the method used provides a detailed analysis of price transmission by allowing for asymmetric cointegration. This permits a simultaneous analysis of both short-run asymmetry (in speed) and long-run asymmetry (in magnitude), the latter of which has been largely overlooked in previous studies that assume a linear cointegration framework.

The focus on milk when studying price transmission for organic and conventional food is justified for many reasons. First, the milk sector is a large agroeconomic sector in many countries. For example, about 18% of the total value of agricultural goods output in Sweden comes from milk production, compared to an average of 14% in the EU28 (EUROSTAT, 2021b). About 20% of Sweden's dairy cows are organic, making it the 6th largest organic milk producer within the EU28 (EUROSTAT, 2021a; Swedish Board of Agriculture, 2020). Second, dairy is often among the top selling categories within the organic food segment (Willer & Lernoud, 2019). In Sweden, the organic share of total milk sales amount to about 24% (Statistics Sweden, 2020a, 2020b). Household demand may therefore have a tangible impact on domestic organic production and the potential of reaching related environmental targets. Third, milk is a perishable good, purchased frequently, and with few substitutes (Dhar & Foltz, 2005). Retailers' pricing strategies can therefore be important matters for households.

Concentration levels are high among Swedish dairies, and food retailers. Three cooperatively owned dairies (Arla, Skånemejerier, Norrmejerier) account for about 90% of the total fluid milk production, and one of them for almost 60% (Swedish Competition Authority, 2016). These dairies concurrently act as wholesalers of both organic and conventional milk under their own brands. The high concentration levels are mainly due to a long period of regional dairy monopolies. Although monopolies were abolished in 2000, with retail private labels entering the market in 2011, catchment areas still prevail to a large extent, and consumers tend to prefer the "local" brand, suggesting strong regional preferences (Swedish Competition Authority, 2016). Within the Swedish food retail sector, one chain accounted for about 50% of the market in 2013, and the joint market share of the three largest chains amounted to 86%, compared to, e.g., 60% in Germany, and 20% in Italy (Swedish Competition Authority, 2018). Due to a relatively large supply of raw milk, and the dairies' obligation to purchase all milk supplied by its members, Swedish retailers have a fairly good bargaining position towards dairies (Swedish Competition Authority, 2016). Importantly, concentration levels in wholesale and retail are similar for both organic and conventional milk, since all larger dairies produce both types of milk, and since general retailers represent practically all organic (and conventional) sales.

The dataset employed in this paper covers the period 2007–2017 and consists of monthly wholesale and retail prices for organic and conventional milk. Wholesale prices are provided by a large Swedish dairy, and retail prices are averages from a region in Sweden, encompassing 45 municipalities (out of 290), and results should be interpreted with this in mind. In particular, I make no claim to study the whole of Sweden's milk market. Estimations of price transmission are performed using a non-linear autoregressive distributed lag model (NARDL) which allows for asymmetric cointegration. Results show that wholesale to retail price transmission for conventional milk is characterised by positive asymmetries both in the long-run and the short-run, and that the long-run asymmetry is quite substantial. For the consumer, this means that a loss in surplus due to a positive shock to wholesale price is larger than the gain in surplus due to a negative shock. In the case of organic milk, the long-run positive asymmetry is less pronounced and organic consumers are therefore likely to experience smaller differences between surplus losses and gains, following positive and negative wholesale price changes, respectively.

The remaining part of the paper is structured as follows: Section 2 describes the concept of price transmission with a particular focus on agricultural food products. Section 3 presents previous literature, while Section 4 provides a description of the data. The empirical approach is described in Section 5, with results presented in Section 6. Section 7 concludes the paper.

2. Price transmission

This paper studies vertical price transmission, which is the study of how price changes at one level of the supply chain affect prices at another level.[1] Price transmission is characterised by speed of adjustment, magnitude, and symmetry/asymmetry (Frey & Manera, 2007). Symmetric price transmission occurs when a price increase is transmitted to other parts of the supply chain with the same speed and magnitude as would a price decrease. Positive asymmetry is found when a price increase is transmitted faster or to a larger extent, described by Peltzman (2000) as a 'rockets and feathers' effect, whereas negative asymmetry implies that a price decrease is transmitted faster or to a larger extent (Meyer & Von Cramon-Taubadel, 2004). In the case of downstream causality, meaning that prices in one part of the supply chain affect prices further down the chain, positive asymmetry is not beneficial for consumers. Negative asymmetry would benefit consumers but is less beneficial for actors further up in the supply chain.

The main factors identified to affect the magnitude of price transmission in agricultural markets relate to market power and product characteristics. A perfectly competitive market implies perfect price transmission in the sense that price changes in one stage are transmitted directly, completely, and symmetrically to the next stage (Lloyd et al., 2009). McCorriston et al. (1998, 2001) show that increased market power reduces the degree of price transmission between farm and retail level. Supply chain actors with strong bargaining power can therefore reduce the degree of price transmission (European Commission, 2009). Regarding product characteristics, McCorriston et al. (2001) show that an increase (decrease) in the price elasticity of retail demand for a good increases (decreases) price transmission magnitude, and that price transmission is larger in industries with increasing returns to scale,[2] such as in many agricultural markets, and for products with low value added in the supply chain. As noted in the introduction, concentration levels among Swedish dairies and retailers are quite high, and empirical work shows that the Swedish food retail sector is characterised by increasing returns to scale (Maican & Orth, 2017). Further, the added-value is relatively low for fluid milk in the Swedish market, with few (if any) middle hands between dairies and retailers (Persson, 2011).

[1] Horizontal price transmission refers to how price changes are transmitted across markets or countries.

[2] In the case of increasing returns to scale, output must expand by relatively more for equilibrium to be restored, compared to the constant cost case. Therefore, a fall in upstream prices will induce a larger magnitude of price transmission compared to the case of constant marginal cost, even though market power due to increasing returns to scale will most likely dampen the degree of price transmission.

The causes of asymmetric price transmission are not as straightforward, but are commonly associated with market power and adjustment costs (Meyer & Von Cramon-Taubadel, 2004). Increased market power is usually argued to cause positive asymmetries (Romain et al., 2002; Zachariasse & Bunte, 2003), although the type of asymmetry depends on market characteristics such as competitors' behavior. If an oligopolistic retailer expects its competitor(s) to match a price increase, but not a price decrease, positive asymmetry occurs, whereas the opposite case implies negative asymmetry (Bailey & Brorsen, 1989; Ward, 1982). Adjustment costs can also explain asymmetries. If a business incurs costs when changing prices, it may wait to adjust prices following a change in input price. When implementing the adjustment, the new price incorporates adjustment costs making the price increase (decrease) higher (lower) compared to the change in input price (Azzam, 1999). If constant price fluctuations reduce customer loyalty, a business may similarly choose to avoid instant adjustments (Persson, 2011). Asymmetries due to adjustment costs can also be linked to the perishability of a good. For perishable goods, such as fluid milk, a wholesaler and/or retailer may be more likely to adjust prices following a fall in input price, compared to a rise in input price, in order to avoid spoiled stock (Ward, 1982). Importantly, adjustment costs are typically associated with asymmetries in the short-run, i.e., in adjustment speed (Meyer & Von Cramon-Taubadel, 2004).

3. Previous literature

Agricultural and food markets have been subject to various empirical studies on price transmission (for an overview see, e.g., Meyer and Von Cramon-Taubadel (2004) and Frey and Manera (2007)). A majority of these studies reveal asymmetric price transmission, and mainly of the positive type, e.g., Goodwin and Holt (1999) on beef, Miller and Hayenga (2001) on pork, Loy et al. (2015) on milk and butter. Although few studies perform any formal tests for the cause of asymmetry, evidence point to the relation between market power in the supply chain and incomplete or asymmetric price transmission for food products (Hong & Li, 2017; Romain et al., 2002). Despite relatively high concentration levels in Swedish agricultural markets, especially within retail, price transmission studies for the Swedish food market are scarce. Two exceptions are Persson (2011) and Karantininis et al. (2011) both studying price transmission in the Swedish conventional food market.

Studies addressing price transmission for conventional fluid milk report somewhat mixed results. Romain et al. (2002) find positive asymmetries in farm to retail price transmission for the New York fluid milk market, whereas Awokuse and Wang (2009) find negative asymmetric price transmission using a threshold approach and national average milk prices for the US market. Serra and Goodwin (2003) find symmetric farm to retail price transmission for highly perishable fluid milk in the Spanish milk market, despite high retail concentration levels. Antonioli et al. (2019) use Italian processor and retail prices for a specific milk product, much in line with the present paper, and find symmetric wholesale to retail price transmission, which is mainly attributed to the existence of long-term price contracts. Using price indices for the Swedish conventional fluid milk market, Persson (2011) finds symmetric price transmission both in the short-run and long-run and attribute this finding to the high perishability of milk, and its role as a basic good facing an almost constant demand.

Studies on price transmission for quality-differentiated products are more scarce. Surathkal and Chung (2017) study wholesale to retail price transmission for quality-differentiated beef products in the US, using a threshold approach, and find positive asymmetric price transmission that is more pronounced for higher quality grades. However, when studying the German market for differentiated milk products, Loy et al. (2015) find evidence of positive asymmetric price adjustment for all milk types, but a quicker and, unexpectedly, more asymmetric adjustment for the low price (low margin) private labels than for high priced national brands. Although these studies do not consider the organic dimension, results are of interest as consumers often regard organic products as representing higher prices and quality (Shafie & Rennie, 2012).

The organic quality dimension is addressed by Darbandi and Saghaian (2018), and Antonioli et al. (2019) who study wholesale to retail price transmission in the US carrot market, and the Italian milk market, respectively. Darbandi and Saghaian (2018) do not address asymmetries but conclude that organic carrot prices adjust slower in response to an upstream price change, compared to conventional carrot prices. Similar results are found in Antonioli et al. (2019) with retail prices for organic milk adjusting symmetrically, but slower compared to conventional milk prices. The slower adjustment is explained by higher search costs for organic consumers, which may invoke a less price elastic demand, allowing organic retailers to dampen the speed of adjustment. This is somewhat in line with the theoretical findings in McCorriston et al. (2001) who show that the magnitude of price transmission in an imperfectly competitive market decreases with a fall in the absolute value of the price elasticity of demand.

Compared to Italy, the Swedish organic market is characterised by large organic market shares, small organic price premiums, few organic labels, and low search costs for organic products, which are widely available at general retailers (Furemar, 2004; Willer & Lernoud, 2019).

Although demand for organic food tends to be more price elastic than for conventional food, elasticities vary depending on these market characteristics and the sample used (Bunte et al., 2007; Lindström, 2021; Schröck, 2012). Antonioli et al. (2019) argue that the specialized organic food stores in their sample likely face consumers with are relatively price insensitive with regards to organic milk. In Sweden, these stores represent a marginal portion of the organic retail sales and are not included in the sample used for this study. Demand for organic milk is therefore, probably, more elastic in this study's sample compared to the sample used in Antonioli et al. (2019). Using a similar sample as in the present study, Lindström (2021) finds that demand for organic milk in Sweden is more price elastic than for conventional milk, and that demand for regional milk brands is less price elastic than demand for private label milk, indicating strong preferences for regional brands.

The empirical models typically used to study asymmetric price transmission are categorized by Meyer and Von Cramon-Taubadel (2004) as pre- and post-cointegration techniques.[3] Early work typically employed the former, and analysed price transmission without testing for cointegration between prices. However, without testing for cointegration, one cannot be sure that prices are related, and results indicating (asymmetric) price transmission could be due to spurious regression (von Cramon-Taubadel & Loy, 1996). The first model to draw on cointegration techniques when testing for asymmetric price transmission was the asymmetric Error Correction Model (ECM) introduced by von Cramon-Taubadel and Fahlbusch (1994) and widely applied thereafter. The ECM distinguishes between positive and negative price changes, and among its many variants is the Threshold Error Correction Model (TECM) (Loy et al., 2015; Serra & Goodwin, 2003) which also takes into account the size of the price changes. These models generate a short-run adjustment parameter and a long-run multiplier, which can be interpreted as measuring the speed and magnitude of price transmission, respectively (Prakash, 1999). However, both the ECM and TECM assume a linear cointegration relationship between prices, which means that price series are restricted to not diverge and the models can thus only test for asymmetry in terms of speed (Meyer & Von Cramon-Taubadel, 2004). Asymmetric price transmission in magnitude means that responses to negative and positive price changes differ, implying that price series must diverge over time and cannot share a linear long-run relationship (Fousekis et al., 2016). The present study applies an asymmetric cointegration approach introduced by Granger and Yoon (2002), Schorderet

[3] An extensive overview of the different methods applied in price transmission studies can be found in Meyer and Von Cramon-Taubadel (2004).

(2003), and advanced by Shin et al. (2014), which allows for a simultaneous analysis of long-run and short-run asymmetries.

4. Data

The data used in this study consist of monthly wholesale prices provided by a large Swedish dairy, paired with barcode level retail prices, for the period Jan 2007 – Nov 2017. The data on retail prices is sourced from the market research company AC Nielsen, which collects weekly observations on sales value and sales volume from a representative sample of stores, as well as detailed product level data, such as manufacturer and organic label. The average retail price of a product is obtained by dividing sales value by volume. AC Nielsen provides aggregated barcode level data for six Swedish regions, encompassing all Swedish municipalities.[4] However, since not all products are sampled in all regions, observations used in this study stem from only one of these six regions, which encompasses 45 Swedish municipalities. These observations are aggregated to monthly level, to match the level of wholesale prices, and to avoid gaps in the time series, since not all products are sampled in all weeks.[5]

The products studied are organic and conventional branded milk of 1.5% fat, in 1 litre cartons. The 1.5% fat segment is chosen due to its large market share, more than half of all milk consumed in Sweden belongs to this fat segment (Swedish Board of Agriculture, 2019). For the specific region and period studied, the average share of sales volume for the branded conventional (organic) product amounts to 43% (49%) of all conventional (organic) 1 litre milk sales, and 72% (67%) of all 1.5% fat conventional (organic) 1 litre milk sales. While previous studies generally consider milk to be a homogenous product, this paper employs product-level data for a specific fat segment, which may capture important features of differentiation, present in most modern agricultural markets (Sexton, 2013). Both wholesale and retail prices are expressed in SEK/litre, net of VAT, and deflated using monthly CPI for food products with base month January 2007. Prices are logarithmic, meaning that price changes can be interpreted as percentage changes. Time series for conventional wholesale and retail prices ($lnpc\_w$ and $lnpc\_r$), and organic wholesale and retail prices ($lnpo\_w$ and $lnpo\_r$) are shown in Figure 1.

[4] These regions are created by AC Nielsen, and do not constitute any administrative categories, or regions pertaining to a specific dairy. A map of the regions is provided in the Appendix.
[5] Monthly prices are common in agricultural markets, and according to Frey and Manera (2007), higher frequency data does not affect price transmission results to a large extent.

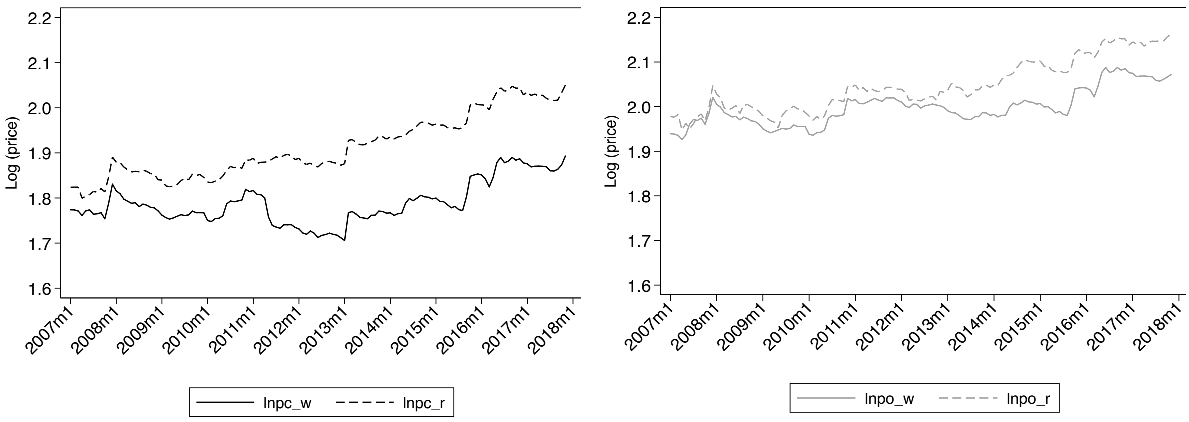


Figure 1. Conventional milk wholesale ($lnpc\_w$) and retail ($lnpc\_r$) prices in left panel. Organic milk wholesale ($lnpo\_w$) and retail ($lnpo\_r$) prices in right panel. Source: Author's elaboration.

As visible in Figure 1, wholesale and retail prices move together to quite similar extents for conventional and organic milk over the period studied. Both the organic and conventional price pairs show spikes at the start of 2008, and an upward trend can be noted from around 2013. Price margins, derived from the vertical distance between wholesale and retail price, are generally somewhat larger for conventional milk, although margins appear to increase for both products over time. The lower price margins for organic milk, could indicate a more competitive setting for the organic milk product, although low margins in the early periods could also be linked to aggressive pricing during the launch of organic milk. By graphical inspection, it appears to be a structural break, i.e., a break in the mean of the series, for conventional wholesale milk price between early 2011 and early 2013, which is not mirrored in any other price series. For organic milk, a break in the mean of retail price appears around early 2013 when retail and wholesale prices start to diverge. Conversations with the dairy providing wholesale prices confirm that no apparent exogenous shocks, e.g., extreme weather, sanctions, etcetera, occurred during the period that can aid in explaining these shifts.[6] The presence of structural breaks can be controlled for in the analysis by including dummy variables in the estimations.

The analysis focuses on retailers' response to a product-specific change in the wholesale price. In line with Antonioli et al. (2019), I only have wholesale prices from one dairy. Whether a price change is industry-wide or not is thus unknown. Moreover, since retail prices only cover one region, results can differ for other regions with other competition characteristics. When

[6] The sanctions against Russia, issued by EU in 2015, affected Swedish prices of butter and cheese but not milk.

using scanner data, one may face promotional sales, i.e., significant temporary price reductions, that are unrelated to cost changes but might affect the efficiency of estimations by adding unexplained price variation (Loy et al., 2015). However, the scanner data used in this study provide average prices by region, not by store, and do not contain information on price promotions. Since promotional sales are not necessarily synchronised across stores, temporary price reductions would only have a minor effect on estimations over a period of 11 years.

Price setting strategies of Swedish wholesalers (dairies) and retailers are not openly disclosed (Swedish Board of Agriculture, 2009), but some general information is available. For example, Sweden's largest dairy keeps a price list, from which individual retail chains and stores can negotiate prices based on, e.g., sales volume and store size. Prices are normally set for a couple of months, but can remain for longer periods if no parties negotiate (Jordbruksaktuellt, 2015). As noted in the introduction, retailers have a strong bargaining position towards dairies. The retail price is also influenced by the size, location, and sales volume of an individual store, as well as the price of competitors, private label prices, import competition and price development in other parts of the EU (Jordbruksaktuellt, 2015; Swedish Board of Agriculture, 2009).

5. Empirical framework

As described in Section 3, a majority of previous studies analysing price transmission either refrain from accounting for cointegration or assume that the long-run cointegration relationship between two price variables can be described as a linear combination (Meyer & Von Cramon-Taubadel, 2004). The current paper accounts for the possibility of asymmetric cointegration, using the non-linear autoregressive distributed lag (NARDL) model (Shin et al., 2014). The NARDL cointegration approach has been used in studies on, e.g., housing price dynamics (Katrakilidis & Trachanas, 2012), energy markets price transmission (Greenwood-Nimmo & Shin, 2013), and in some studies on food markets, e.g., price transmission in the US beef market (Fousekis et al., 2016) and in the Swedish market for milk, cereals, beef (Persson, 2011), and pork (Karantininis et al., 2011). The main advantage of NARDL is its simultaneous analysis of asymmetries in the long-run cointegrating relationship, as well as in the short-run dynamic error corrections. This enables estimation of asymmetries in both speed and magnitude and provides a deeper analysis, as short- and long-run asymmetries can have different causes. E.g., adjustment costs usually explain most of the short-run asymmetry, while long-run asymmetry is usually linked to market power (Karantininis et al., 2011).

Since the NARDL model is an extension of the standard linear ARDL model (Pesaran & Shin, 1999; Pesaran et al., 2001), the ARDL*(p,q)* model is briefly explained here. It can be described by the following equation with two time series, y_t and x_t, with $t = 1,2,\dots,T$:

$$\Delta y_t = a_0 + \rho y_{t-1} + \theta x_{t-1} + \gamma \boldsymbol{z_t} + \sum_{j=1}^{p-1} a_j \, \Delta y_{t-j} + \sum_{j=0}^{q-1} \delta_j \, \Delta x_{t-j} + e_t \quad (1)$$

with $\boldsymbol{z_t}$ being a vector of deterministic regressors (e.g., trends and structural breaks), and e_t being an *iid* stochastic error. p and q denote number of lags of the independent and dependent variable, respectively. In the context of this study, y_t corresponds to retail price, and x_t corresponds to wholesale price. Under the null hypothesis that y_t and x_t are not cointegrated, i.e., there is no long-run relationship in levels, it follows that $\rho = \theta = 0$. Compared to other linear cointegration procedures, such as the Engle and Granger (Engle & Granger, 1987) two-step, and Johansen (Johansen, 1988) test, ARDL tends to perform better in small samples, and is more flexible as it does not require time series to be strictly I(1) (Pesaran et al., 2001).

To account for potential asymmetries, Shin et al. (2014) developed the NARDL model which introduces short-run and long-run asymmetries by decomposing the independent variable x_t into positive and negative partial sums, according to:

$$x_t = x_0 + x_t^+ + x_t^- \quad (2)$$

where $x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} \max(\Delta x_j, 0)$, and $x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} \min(\Delta x_j, 0)$.

The asymmetric long-run equilibrium relationship can thus be expressed as:

$$y_t = \beta^+ x_t^+ + \beta^- x_t^- + u_t \quad (3)$$

with β^+ and β^- denoting asymmetric long-run parameters for positive and negative changes in x_t. By combining (3) and (1), the NARDL*(p,q)* model is obtained according to:

$$\Delta y_t = a_0 + \rho y_{t-1} + \theta^+ x_{t-1}^+ + \theta^- x_{t-1}^- + \gamma \boldsymbol{z_t} + \sum_{j=1}^{p-1} a_j \, \Delta y_{t-j} + \sum_{j=0}^{q-1} (\delta_j^+ \, \Delta x_{t-j}^+ + \delta_j^- \Delta x_{t-j}^-) + e_t \quad (4)$$

where $\theta^+ = -\rho\beta^+$ and $\theta^- = -\rho\beta^-$. After estimating (4) by standard OLS, the existence of cointegration in levels is tested using either one of two tests. The bounds test approach proposed by Pesaran et al. (2001) tests the joint null hypothesis of no cointegration ($\rho = \theta^+ = \theta^- = 0$) using the F_{PSS} statistic with critical values in Pesaran et al. (2001). Alternatively, one may use the t_{BDM} statistic with the null hypothesis of $\rho = 0$ against $\rho < 0$ with critical values in Banerjee et al. (1998). The null hypothesis of long-run symmetry ($\beta^+ = \beta^-$, i.e. $-\theta^+/\rho =$

$-\theta^-/\rho$), and short-run symmetry ($\sum_{j=0}^{q-1} \delta_j^+ = \sum_{j=0}^{q-1} \delta_j^-$) are tested by standard Wald tests.[7] If there is no evidence of asymmetry, one can impose constraints of short-run symmetry, long-run symmetry, or both, in the model. Imposing both constraints will result in expression (1). The cumulative dynamic multipliers, m_h^+ and m_h^-, associated with unit changes in x_t^+ and x_t^-, can provide an illustration and analysis of the path and duration of adjustment following negative and positive shocks to wholesale prices. These dynamic multipliers are calculated as:

$$m_h^+ = \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial x_t^+}, \text{ and } m_h^- = \sum_{j=0}^{h} \frac{\partial y_{t+j}}{\partial x_t^-} \text{ for } h = 0,1,2 \ldots \qquad (5)$$

Before testing for cointegration, i.e., if there is any price transmission at all or if prices are set independently of each other, unit root tests in levels and in first differences are performed in order to establish the order of integration of the individual time series. Even if (N)ARDL model(s) are relatively flexible and permit testing for cointegration regardless if series are I(0) or I(1), they do not permit series to be I(2). Further, standard linear tests for cointegration that are also performed in the analysis require that individual time series are strictly I(1).

The cointegration tests employed include the standard linear tests of Engle and Granger (1987) and Johansen (1988), followed by a Gregory-Hansen (1996) test allowing for one structural break, and a NARDL cointegration test allowing for structural breaks and asymmetric cointegration. When testing for cointegration, a relationship of downstream causality between prices is assumed, meaning that retail prices are regressed on wholesale prices. Downstream causality is shown to hold for the price relationship between wholesalers and retailers in the Swedish milk sector (Persson, 2011), as well as for many other agricultural products (see, e.g., Fousekis et al. (2016)).

6. Results

In what follows, results from unit root tests are presented, followed by results from cointegration tests. Time series are in logarithmic form, to mitigate fluctuation and increase the likelihood of stationarity after first-differencing (Hamilton, 1994). After showing that cointegration holds, results from NARDL estimations of price transmission are then presented.

6.1 Unit root tests

[7] The null hypothesis of short-run symmetry can either take the additive (weak) form described here, or the pairwise (strong) form requiring that $\delta_j^+ = \delta_j^-$ for all *j=0,...,q-1*

Based on graphical inspection after first-differencing the time series, a trend does not appear to be present, and is therefore not included in the unit root test. Although the first-differenced time series appear to be fluctuating around a mean of zero, a constant is included in the unit root tests, to improve test stability (Antonioli et al., 2019).

Table 1 presents results from unit root tests, including the augmented Dickey-Fuller (ADF) (Dickey & Fuller, 1979, 1981) test, the Phillips-Perron (PP) (Phillips & Perron, 1988) test, and the Kwiatkowski-Phillips-Schmidt-Schin (KPSS) (Kwiatkowski et al., 1992) test. The latter is used as a complementary test, since both the ADF and PP test may suffer from low-power and size distortions, which may lead to over-rejection of the null hypothesis of a unit root (DeJong et al., 1992). Table 2 presents results from the Zivot-Andrews (1992) unit root test with one structural break. Test statistics in Tables 1 and 2 show that variables are non-stationary in levels, and stationary in first differences. It can thus be concluded that individual time series are integrated of order 1, i.e., I(1).

Table 1. ADF, PP and KPSS unit root tests.

	ADF		PP		KPSS	
	Level	1st diff.	Level	1st diff.	Level	1st diff.
lnpc_w	-0.686	-8.939***	-0.865 (4)	-8.767***(4)	0.778*** (7)	0.152 (1)
lnpc_r	-0.113	-9.768***	-0.145 (4)	-9.673***(4)	1.610*** (7)	0.086 (4)
lnpo_w	-1.178	-9.148***	-1.285 (4)	-8.988***(4)	1.140*** (7)	0.051 (2)
lnpo_r	-0.541	-11.337***	-0.458 (4)	-11.357***(4)	1.580*** (7)	0.103 (9)

Note: ADF test: H0=Variable is non-stationary. Z(t) statistic reported. Critical values are -3.500, -2.888, and -2.578 for 1%, 5% and 10% significance level, respectively. PP test: H0=Variable is non-stationary. Z(t) statistic reported. Critical values are -3.500, -2.888, and -2.578 for 1%, 5% and 10% significance level, respectively. Optimal lag length for PP in parenthesis is selected within the test based on Newey and West (1994). KPSS test: H0=Variable is stationary. Critical values are 0.739, 0.463 and 0.347, for 1%, 5% and 10% significance level, respectively. Optimal lag length for KPSS in parenthesis is selected within the test based on Newey and West (1994).

Table 2. Zivot-Andrews unit root test with one structural break.

	Level	Break date	1st diff	Break date
lnpc_w	-3.761 (2)	2015m9	-8.539*** (1)	2013m2
lnpc_r	-4.152 (1)	2015m9	-9.851*** (0)	2013m1
lnpo_w	-4.217 (1)	2015m9	-8.557*** (1)	2015m9
lnpo_r	-3.950 (0)	2015m9	-11.380*** (0)	2009m6

Note: H0=Variable is non-stationary. Critical values are -5.34, -4.80, and -4.58 for 1%, 5% and 10% significance level, respectively. Optimal lag length in parenthesis is selected based on the Akaike Information Criteria (AIC).

6.2 Cointegration tests

Results from cointegration tests are presented in Table 3 below. The number of lags included in each test is shown in parenthesis and chosen based on the AIC.[8] Results from the Engle-Granger and Johansen tests are presented in the first two columns and do not show any evidence of linear cointegration. However, as noted above, these tests do not account for non-linear cointegration and do not include the presence of any structural breaks. The presence of a structural break, i.e., a change in the mean of the series, can affect coefficients to differ before and after the onset of the change, and cointegration tests for time series with structural breaks tend to have low power unless these breaks are controlled for (Gregory & Hansen, 1996).

Table 3. Results from cointegration tests.

	Engle Granger test	Johansen test	Gregory-Hansen test		NARDL*(2,3)*		NARDL*(max p=max q=6)*	
			Z(t)	Break date	F_{PSS}	t_{BDM}	F_{PSS}	t_{BDM}
lnpc_r= lnpc_w	-1.643 (2)	4.694 (2)	-5.21***(0)	2011m6	6.379**	-4.052***	8.775***	-5.072***
lnpo_r= lnpo_w	-1.975 (2)	7.526 (2)	-4.76**(0)	2013m2	7.074**	-4.583***	8.068***	-4.675***

Note: Critical values in Engle-Granger test are -3.077, -3.384 and -3.982 for the 10%, 5% and 1% significance level, respectively. The 5% critical value in Johansen test is 15.41. Critical values for the Gregory-Hansen test are -5.13, -4.61 and -4.34 for the 10%, 5% and 1% significance level, respectively. For $k=1$, critical values for the 10%, 5% and 1% significance level are 4.78, 5.73 and 7.84 for the F_{PSS} statistic (Pesaran et al., 2001), and -2.90, -3.23 and -3.82 for the t_{BDM} statistic (Banerjee et al., 1998)

Results from Gregory-Hansen cointegration tests are presented in the third and fourth columns of Table 3 and show that wholesale and retail prices are cointegrated in levels for both conventional and organic milk, when allowing for a structural break in level. The break dates are June 2011 for conventional milk series, and February 2013 for organic milk series. These findings are corroborated by a graphical inspection of Figure 1, which shows a sharp fall in conventional wholesale price in mid 2011, and that organic wholesale and retail prices diverge between early 2013 and mid 2015. To control for the structural breaks in the NARDL estimations, dummy variables are included taking the value 1 for all months prior to the break date, and 0 after. The equations to be estimated using NARDL are thus the following:

[8] The AIC criteria for optimal lag structure in Engle-Granger, Johansen, and NARDL cointegration tests, are retrieved using the *varsoc* command in Stata ver. 14.

$$\Delta \ln pc\_r_t = a_0 + \rho \ln pc\_r_{t-1} + \theta^+ \ln pc\_w^+_{t-1} + \theta^- \ln pc\_w^-_{t-1} + \sum_{j=1}^{p-1} a_j \Delta \ln pc\_r_{t-j} + \sum_{j=0}^{q-1} (\delta_j^+ \Delta \ln pc\_w^+_{t-j} + \delta_j^- \Delta \ln pc\_w^-_{t-j}) + \gamma_t jun2011_t + e_t \qquad (6)$$

$$\Delta \ln po\_r_t = a_0 + \rho \ln po\_r_{t-1} + \theta^+ \ln po\_w^+_{t-1} + \theta^- \ln po\_w^-_{t-1} + \sum_{j=1}^{p-1} a_j \Delta \ln po\_r_{t-j} + \sum_{j=0}^{q-1} (\delta_j^+ \Delta \ln po\_w^+_{t-j} + \delta_j^- \Delta \ln po\_w^-_{t-j}) + \gamma_t feb2013_t + e_t \qquad (7)$$

The last four columns in Table 3 show the results from the NARDL cointegrations tests,[9] using two different approaches regarding optimal lag structure. NARDL*(2,3)* refers to a model where the optimal lag structure for each individual variable is obtained using the AIC criteria, in line with previous cointegration tests, resulting in *p=2, q=3* for both the organic and conventional milk systems. NARDL*(max p=max q=6)* denotes a model when using the general-to-specific approach to obtain optimal lag structure. This approach is employed in several empirical applications of NARDL (see, e.g., Fousekis et al. (2016), Greenwood-Nimmo et al. (2013) and Shin et al. (2014)). Starting at a maximum lag of *p=6, q=6,* and dropping insignificant regressors based on a 10 % significance rule, results in a parsimonious model with little noise and highly significant regressors. Regardless of lag approach, the F_{PSS} and t_{BDM} test statistics in Table 3 indicate a long-run cointegrating relationship in levels between wholesale price and retail price for both conventional and organic milk, respectively. These results are in line with Antonioli et al. (2019), who find long-run cointegration between wholesale and retail price for both organic and conventional milk, respectively, in the Italian milk market.

6.3 Results from NARDL estimations

With a long-run cointegrating relationship in levels established, we turn to results regarding speed, magnitude and symmetry of price transmission. Tables 4 and 5 present parameter estimates, and asymmetry tests statistics based on NARDL estimations of equations (6) and (7), for the two different approaches regarding optimal lag structure as described above.

In the case of conventional milk, the null hypothesis of long-run symmetry is rejected with a test statistic significant at 1% level, and so is the null hypothesis of short-run symmetry. This holds for both lag choices employed. In terms of magnitude, i.e., long-run effects captured by $\beta^+$and β^-, a 1% increase (decrease) in conventional wholesale price is associated with a 0.702% (0.426%) increase (decrease) in conventional retail price with the model using AIC to

[9] Estimations are carried out in Stata ver. 14, using the *nardl* command written by Marco Sunder.

choose optimal lags. This suggests that in the long run, increases in wholesale price are transmitted to a substantially larger extent to retail price, than decreases in wholesale price. The same model reports short-run asymmetry indicating that an increase in wholesale price is more strongly transmitted compared to a decrease in the periods immediately following the price change, and that this asymmetry primarily appears on impact, i.e., in *t-0*. Results do not change much with the general-to-specific approach in Table 5. Long-run asymmetry prevails with $\beta^{+}$and β^{-} corresponding to 0.789 and 0.595, respectively, and short-run asymmetry mainly occurring in *t-0* is again confirmed.

Table 4. Wholesale to retail price transmission, conventional and organic milk, NARDL*(2,3)*.

Conventional milk			Organic milk		
Variable	Coefficient	Std. error	Variable	Coefficient	Std. error
$lnpc\_r_{t-1}$	-0.259***	0.064	$lnpo\_r_{t-1}$	-0.290***	0.063
$lnpc\_w^{+}_{t-1}$	0.182***	0.053	$lnpo\_w^{+}_{t-1}$	0.248***	0.060
$lnpc\_w^{-}_{t-1}$	0.110**	0.047	$lnpo\_w^{-}_{t-1}$	0.206***	0.063
$\Delta lnpc\_r_{t-1}$	-0.027	0.089	$\Delta lnpo\_r_{t-1}$	-0.009	0.087
$\Delta lnpc\_w^{+}_{t-0}$	0.809***	0.048	$\Delta lnpo\_w^{+}_{t-0}$	0.819***	0.101
$\Delta lnpc\_w^{+}_{t-1}$	0.126	0.087	$\Delta lnpo\_w^{+}_{t-1}$	0.096	0.124
$\Delta lnpc\_w^{+}_{t-2}$	0.019	0.050	$\Delta lnpo\_w^{+}_{t-2}$	-0.002	0.103
$\Delta lnpc\_w^{-}_{t-0}$	0.295***	0.086	$\Delta lnpo\_w^{-}_{t-0}$	0.894***	0.203
$\Delta lnpc\_w^{-}_{t-1}$	-0.250***	0.093	$\Delta lnpo\_w^{-}_{t-1}$	-0.093	0.214
$\Delta lnpc\_w^{-}_{t-2}$	0.010	0.088	$\Delta lnpo\_w^{-}_{t-2}$	0.427**	0.191
jun2011	0.008**	0.004	feb2013	0.011***	0.003
constant	0.468***	0.116	constant	0.566***	0.123
Long-run coefficients, p-values in parenthesis.					
$\beta^{+}_{lnpc\_w}$	0.702***	(0.000)	$\beta^{+}_{lnpo\_w}$	0.856***	(0.000)
$\beta^{-}_{lnpc\_w}$	-0.426***	(0.000)	$\beta^{-}_{lnpo\_w}$	-0.711***	(0.000)
Wald tests for long- and short-run symmetry, p-values in parenthesis					
$H_0: \beta^{+}_{lnpc\_w} = \beta^{-}_{lnpc\_w}$	27.29***	(0.000)	$H_0: \beta^{+}_{lnpo\_w} = \beta^{-}_{lnpo\_w}$	4.521**	(0.036)
$H_0: \sum_{j=0}^{q-1} \delta_j^{+} = \sum_{j=0}^{q-1} \delta_j^{-}$	22.60***	(0.000)	$H_0: \sum_{j=0}^{q-1} \delta_j^{+} = \sum_{j=0}^{q-1} \delta_j^{-}$	0.567	(0.453)
Model statistics					
N		128			128
R^2, adj. R^2	0.793	0.773		0.625	0.589
χ^2_{SC}	46.09	(0.235)		31.20	(0.839)
χ^2_{HET}	0.381	(0.537)		10.05***	(0.001)

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. χ^2_{SC} and χ^2_{HET} are Lagrange multiplier tests for serial correlation and heteroskedasticity, respectively.

For organic milk, the difference in long-run coefficients, i.e., difference between $\beta^{+}$and β^{-}, is generally smaller than in the case of conventional milk. Specifically, a 1% increase (decrease) in organic wholesale price is associated with a 0.856% (0.711%) increase (decrease) in organic retail price with the NARDL*(2,3)* model, and corresponding values at 0.908 and 0.859 when using a general-to-specific approach to determine lags. Although parameter estimates for organic milk are quite similar when comparing results for the different lag choices, the model NARDL*(2,3)* leads to long-run symmetry being rejected, and a failure to reject short-run symmetry, whereas the opposite holds for the model using a general-to-specific approach.[10] The null hypothesis of short-run asymmetry tests the presence of additive short-run asymmetries, i.e., $\sum_{j=0}^{q-1} \delta_j^+ = \sum_{j=0}^{q-1} \delta_j^-$, but does not reveal adjustment asymmetries in terms of speed. The dynamic multipliers illustrated and discussed below therefore adds important information regarding patterns and speed of adjustment.[11]

Table 5. Wholesale to retail price transmission, organic and conventional milk. NARDL*(max p=max q=6)*.

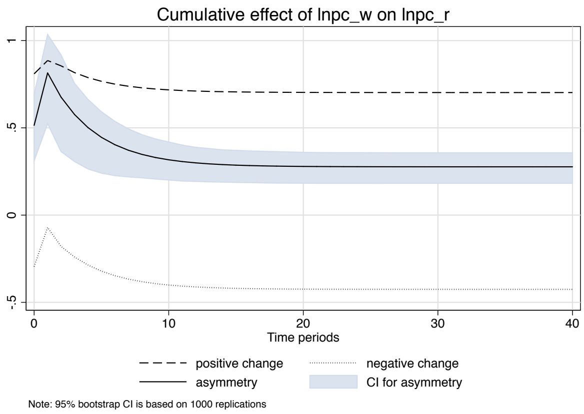
Wholesale to retail - Conventional milk			Wholesale to retail - Organic milk		
Variable	Coefficient	Std. error	Variable	Coefficient	Std. error
$lnpc\_r_{t-1}$	-0.313***	0.062	$lnpo\_r_{t-1}$	-0.259***	0.055
$lnpc\_w^+_{t-1}$	0.247***	0.071	$lnpo\_w^+_{t-1}$	0.235***	0.052
$lnpc\_w^-_{t-1}$	0.186***	0.047	$lnpo\_w^-_{t-1}$	0.223***	0.053
$\Delta lnpc\_r_{t-2}$	0.055	0.043			
$\Delta lnpc\_w^+_{t-0}$	0.844***	0.042	$\Delta lnpo\_w^+_{t-0}$	0.890***	0.084
$\Delta lnpc\_w^-_{t-0}$	0.297***	0.079	$\Delta lnpo\_w^-_{t-0}$	0.904***	0.178
$\Delta lnpc\_w^-_{t-1}$	-0.181**	0.077	$\Delta lnpo\_w^-_{t-2}$	0.391**	0.158
			$\Delta lnpo\_w^-_{t-3}$	0.193	0.155
			$\Delta lnpo\_w^-_{t-4}$	0.300*	0.156
jun2011	0.015***	0.004	feb2013	0.012***	0.003
constant	0.570***	0.113	constant	0.512***	0.108
Asymmetry statistics, p-values in parenthesis.					
$\beta^+_{lnpc\_w}$	0.789***	(0.000)	$\beta^+_{lnpo\_w}$	0.908***	(0.000)
$\beta^-_{lnpc\_w}$	-0.595***	(0.000)	$\beta^-_{lnpo\_w}$	-0.859***	(0.000)
$H_0: \beta^+_{lnpc\_w} = \beta^-_{lnpc\_w}$	24.54***	(0.000)	$H_0: \beta^+_{lnpo\_w} = \beta^-_{lnpo\_w}$	0.547	(0.461)
$H_0: \sum_{j=0}^{q-1} \delta_j^+ = \sum_{j=0}^{q-1} \delta_j^-$	40.42***	(0.000)	$H_0: \sum_{j=0}^{q-1} \delta_j^+ = \sum_{j=0}^{q-1} \delta_j^-$	6.808**	(0.010)
Model statistics					
N		125			125
R^2, adj. R^2	-	-		-	-
χ^2_{SC}	35.82	(0.659)		26.78	(0.946)
χ^2_{HET}	-	-		-	-

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. χ^2_{SC} and χ^2_{HET} are Lagrange multiplier tests for serial correlation and heteroskedasticity, respectively.

[10] Re-estimating a model with short-run symmetry constraint does not alter results to any large extent. These results are not presented, but are available upon request.

[11] The pattern of dynamic adjustment depends on a combination of long-run parameters $\beta^+$and β^-, the error correction coefficient ρ, and the coefficients of Δx^-_t and Δx^-_t.

Adjustment asymmetries are captured by the dynamic multipliers, which are depicted in Figures 2a, 2b, 3a and 3b, showing the paths and durations of adjustment in retail prices following a 1% increase (dashed line) and a 1% decrease (dotted line) in wholesale prices.[12] The solid line in the middle of the graphs depicts the difference between the impact of a positive and negative shock in wholesale price. Figure 2a shows the dynamic multipliers for the conventional milk system when the model is NARDL*(2,3).* Comparing the dashed and dotted lines indicates that the immediate response in retail price after a 1% wholesale price increase is significantly larger in magnitude compared to the response following a 1% wholesale price decrease. This positive short-run asymmetry is also indicated by the solid asymmetry line being above zero throughout the 40 periods (months) in the graph. The rate of adjustment is slightly larger following a wholesale price decrease, although reaching a new equilibrium takes the same time, about 11 months, given a price decrease as well as when given a price increase.



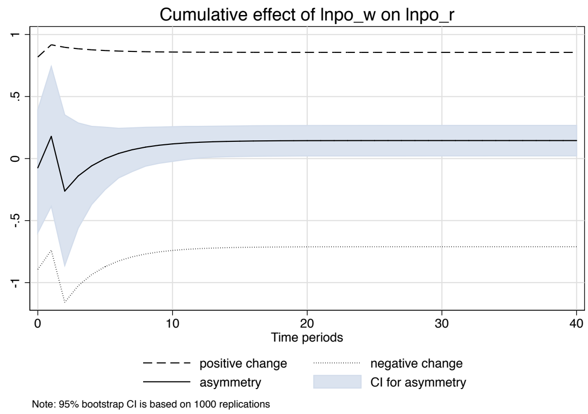


Figure 2a. Dynamic multipliers, conventional milk. NARDL*(2,3)*.

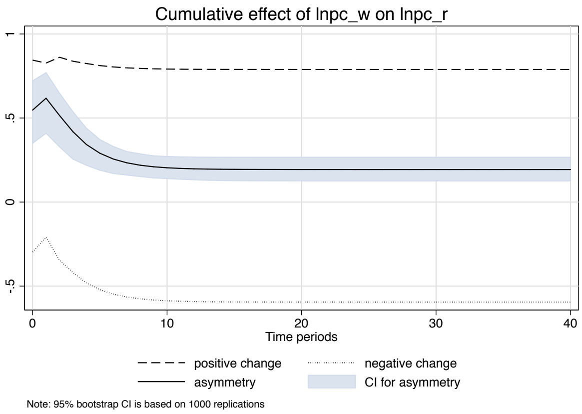
Figure 2b. Dynamic multipliers, organic milk. NARDL*(2,3)*

In the case of organic milk, results in Table 4 show no evidence of short-run asymmetries, i.e., no difference in the additive short-run impacts of positive and negative changes in $lnpo\_w$. The dashed and dotted lines in Figure 2b confirm that the immediate response in retail price is similar in magnitude for negative and positive shocks in wholesale price. This is also illustrated by the solid asymmetry line and its 95% confidence interval not being significantly different from zero in the short run. However, the adjustment is quicker given a positive change in the wholesale price compared to a negative change. The new equilibrium following a wholesale

[12] Plotting of dynamic multipliers is carried out in Stata ver. 14, using the *nardl* command written by Marco Sunder.

price increase is reached within a couple of months, whereas it takes around 10 months to reach the new equilibrium following a decrease. This points to positive asymmetry in adjustment speed. As noted in Table 4 the long-run asymmetry is not very pronounced, as illustrated by the asymmetry line in Figure 2b being just slightly above zero in the long-run.

When allowing for another type of lag structure, adjustment patterns for conventional milk do not change much. In Figure 3a, asymmetries in magnitude are positive in the long-run as well as in the short-run, just as in Figure 2a, and the duration of adjustment is relatively similar for increases and decreases in wholesale price. For organic milk, the graph in Figure 3b illustrates the differing results regarding asymmetries found in Tables 4 and 5. A strong reaction in retail price occurring some periods after a negative change in wholesale price results in a large overshooting of negative retail price changes. This leads to a negative asymmetry in magnitude from around the third period up until the seventh period, after which no asymmetries are detected. As in Figure 2b, retail prices adjust quicker following an increase in wholesale price compared to a decrease in wholesale price, indicating positive asymmetry in adjustment speed.



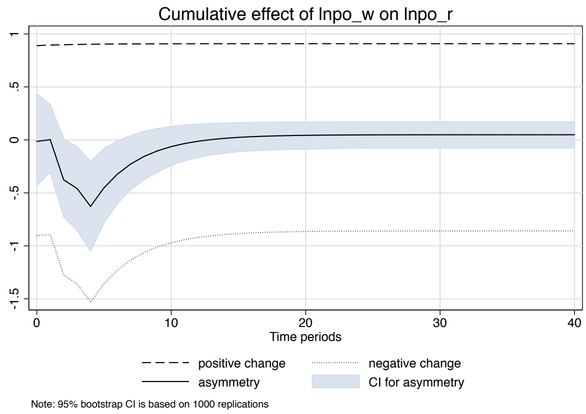


Figure 3a. Dynamic multipliers, conventional milk. NARDL*(max p=max q=6).*

Figure 3b. Dynamic multipliers, organic milk. NARDL*(max p=max q=6).*

To summarize, results show that wholesale to retail price transmission for conventional milk is characterised by both long-run and short-run asymmetries, whereas results are mixed regarding price transmission for organic milk depending on which lag structure is employed. Nonetheless, the following three conclusions may be drawn from the results above: (i) the magnitude of price transmission is larger within the organic milk system, (ii) the difference in price transmission magnitude following a wholesale price increase and decrease, respectively, is smaller for organic milk compared to conventional milk, i.e., positive asymmetry in the long-run is more pronounced for conventional milk, and (iii) based on the dynamic multipliers depicted in Figures 2a, 2b, 3a, and 3b, organic retail prices adjust slightly faster compared to conventional retail prices, at least following an increase in wholesale price, and a decrease in organic wholesale price leads to a strong response in organic retail price in the short-run.

The long-run asymmetry found for conventional milk is quite substantial, and consumers of conventional milk will thus lose considerably more surplus following an increase in wholesale price, than they will gain in surplus following a decrease in wholesale price. For organic consumers, this difference is less pronounced. When β^+ exceeds β^-, retailers will experience an increasing margin between wholesale and retail price. This development is confirmed by a comparison of margins over time for the sample used. For the first half of the sample, the average margin in the conventional milk case is around 0.50 SEK, and around 0.19 SEK in the case of organic milk. In the second half of the sample, average margins expand to 1.09 SEK for conventional milk, and 0.60 SEK for organic milk.

Since many empirical studies on price transmission of milk consider a linear long-run relationship and use industry wide price data, comparing results should be made with caution. Persson (2011) studies the Swedish market for foods using a NARDL approach, and national price indices for each level of the supply chain. He finds long-run symmetry for conventional milk which results in Tables 4 and 5 strongly reject for the current sample of one manufacturer (brand) and one region. However, using a sample similar to the present study in terms of coverage, Antonioli et al. (2019) do not find evidence of adjustment asymmetry for neither conventional milk or organic milk. This difference in results indicate that characteristics of the market, and possibly estimation methods, play a large role when estimating price transmission.

7. Conclusions and discussion

The objective of this paper is to analyse wholesale to retail price transmission for organic and conventional milk in the Swedish milk market. Studying price transmission can yield insights into the efficiency of a supply chain, by analysing the extent to which price changes are transferred along the chain, the speed of these transfers, and whether there are asymmetries with regards to positive and negative price changes. Knowing whether asymmetric price transmission is present is informative for policy makers, for example when evaluating effects on welfare or demand following changes in, e.g., subsidies or taxes. Agricultural markets are often characterised by high concentration levels further down the supply chain, which is also the case in the Swedish milk market, and most work on agricultural markets show evidence of positive asymmetric price transmission. This study contributes to previous literature by (i) adding new knowledge on price transmission for quality-differentiated products, (ii) studying price transmission for organic and conventional food in a highly mature organic market, and (iii) employing an asymmetric cointegration framework, allowing for a simultaneous analysis of short- and long-run asymmetry, the latter of which has been largely overlooked in previous studies. The paper employs monthly wholesale prices for the period Jan 2007–Nov 2017 for organic and conventional branded milk, retrieved from a large Swedish dairy, and corresponding retail prices covering one Swedish region retrieved from the market research company AC Nielsen. Estimations are carried out using the NARDL model, developed by Shin et al. (2014).

Results reveal incomplete price transmission, i.e., $\beta^{+} \neq 1, \beta^{-} \neq -1,$ for both organic and conventional milk. In the case of conventional milk, results indicate positive asymmetries in magnitude both in the short-run and in the long-run. The finding of long-run asymmetry is itself an important result, as it shows the importance of allowing for non-linear cointegration of prices when studying price transmission. For organic milk, the long-run positive asymmetry is less pronounced and not statistically significant in all specifications. Organic consumers will therefore experience smaller differences between surplus losses and gains following positive and negative wholesale price changes, respectively. From a policy perspective, this difference in asymmetry suggests that the long-run impact of a change in wholesale price due to, e.g., changes in subsidies or differentiated VAT rates, is likely to be larger and less asymmetric in the case of organic milk compared to conventional milk.

Incomplete price transmission from wholesale to retail may be indicative of imperfect competition or strong bargaining power within the retail level, as suggested in McCorriston et al. (1998). The long-run asymmetry detected for conventional milk, and in one case for organic milk, may strengthen this suspicion, since asymmetries due to adjustment costs are usually only present in the short-run (Meyer & Von Cramon-Taubadel, 2004). However, results showing incomplete or asymmetric price transmission should not be taken as a formal test for market power given that the data only contains wholesale prices from one dairy, and retail prices from one region. Although asymmetries in the short-run are commonly related to adjustment costs (Meyer & Von Cramon-Taubadel, 2004), the difference in short-run asymmetries between organic and conventional milk is not straightforward. However, a strong(er) downturn in organic retail price following a negative change in wholesale price could indicate a fear of spoiled stock among retailers, that is more pronounced for organic milk compared to conventional milk.

Based on Lindström (2020), price elasticities of demand are likely larger in absolute values for the organic milk product compared to the conventional alternative. The fact that results indicate a larger price transmission magnitude for organic milk is thus in line with the theoretical finding in McCorriston et al. (2001) which shows that the degree of price transmission increases with demand elasticity. In contrast to Antonioli et al. (2019) and Darbandi and Saghaian (2018), results from the present study indicate that retail prices adjust slightly faster for the organic product compared to the conventional one, at least when following a wholesale price increase. This contrasting result could be due to lower search costs for organic consumers in the Swedish market, combined with a, probably, more price elastic demand for the organic product. Results regarding adjustment speed differences are also in line with results in Loy et al. (2015) who study the German dairy market and find that costs are passed-through quicker for milk products with small margins compared to products with higher margins.

The results found in this paper suggest the need for future studies within this area. A natural extension would be to include data on farm-level prices for a notion of price transmission throughout the whole supply chain. Moreover, it would also be informative to widen the quality dimension and study price transmission for private labels and brands specifically since margins and pricing strategies may differ for these categories. Data availability is an issue here, as not all dairies provide information on producer and wholesale prices, and since national processor and consumer price indices are aggregated without discriminating between organic and conventional milk, or between private labels and brands.

Importantly, this paper's analysis provides measures of and tests for asymmetric and/or incomplete price transmission but does not test for the causes of it. The relatively large difference in long-run asymmetry between organic and conventional milk calls for studies that can address explanations to this finding more thoroughly. Some possible explanations to the study's findings are discussed, such as retailers' misuse of market power, differing margins and demand elasticities for organic and conventional milk, and lower organic search costs compared to markets with separate distribution channels for organic products. However, a more formal analysis of how these different factors affect price transmission is outside the scope of the current paper, and a suggestion for future research.

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Appendix

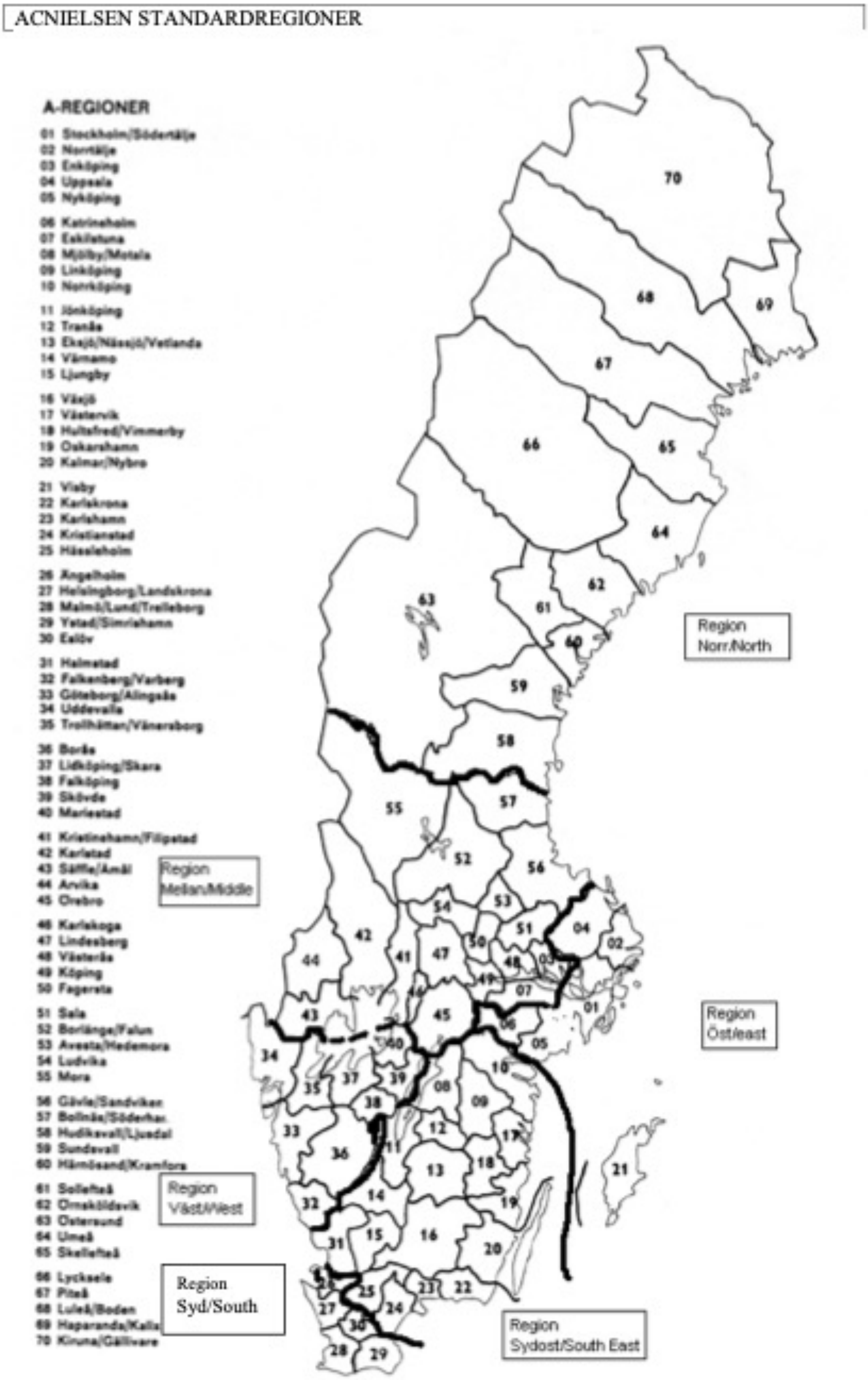

Figure 1: Map of Swedish Regions in AC Nielsen data.