Asymmetric price transmission along the food marketing chain: A focus on the recent price war.

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Abstract

This paper investigates the price transmission along the food marketing chain in France. We focus on this transmission at the upstream and downstream levels during the so-called "price war" waged by the retailers in the French market. To this aim, we rely on an asymmetric cointegration approach: the Nonlinear Autoregressive Distributed Lag (NARDL) model. We find that although the asymmetric price transmission is effective along the French food marketing chain, it is more pronounced at the downstream level, illustrating that agri-food companies are the main losers of the recent war price in the food sector.

Keywords: Food sector, Price transmission, Asymmetry, Nonlinear ARDL model

***Classification JEL*:** C13, C22, Q11, Q13

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1. Introduction

The food sector in France consists of a wide range of companies, from upstream agriculture up to processing industry and retail. The latter (food supply) chain, including around 83,500 companies and generating more than 3 million jobs throughout the country, is strategic for the French economy, as it represents France's second largest employer and its third largest trade surplus. Nevertheless, the food sector is subject to serious competitiveness challenges (fragmented economic structure, adaptation to international demand, cost-competitiveness, investments decrease...), which have become structural. Moreover, since the mid-2000s, regulatory changes (end of quotas under the Common Agricultural Policy), upheavals on world markets, which have become much more volatile, as well as several reforms of the French legislation ("Dutreil law" in 2005, "Chatel law" and "LME law" in 2008, and "EGAlim law" in 2018) increasingly destabilise the sector.

Concerning the market of agricultural goods, gradual abandon by Brussels authorities of any form of market regulation remains its most notable upheavals. Indeed, until 2006, prices of cereals, sugar, dairy products and, to a lesser extent, beef and veal benefited from administered systems and therefore showed a certain stability. Except for a few emergency mechanisms (intervention and storage), quotas abandon in the sugar sector in 2017 can be considered as the end of stabilising policies on the agricultural markets.This implies some new business opportunities, especially for exports, but also challenges such as competition between operators and increasing volatility in food commodity prices, forcing actors of the food chain to adapt to changes. For instance, food commodity prices have risen by almost 50% since 2009, while producer prices increased at a much slower rate (9%), and retail price rises have been even more limited (about 4%). The latter developments reflect an incomplete transmission of changes in prices along the value chain (Ferrucci, 2012). In particular, margins have shrunk at all levels, putting pressure on the most fragmented levels (agricultural and industrial producers).

In recent years, the instabilities of the price of agricultural goods have crystallized. This, coupled with the recent climatic hazards, has profoundly weakened business opportunities for the sector's stakeholders while highlighting the difficulties associated with managing market volatility. This agricultural crises have significantly increased the cost of production, in a context where retail prices remain sluggish and even falling when considering mass-retail consumer products. For several years now, the market for mass consumption food products is characterized by a "price war", fuelling a deflation of product prices on the shelves, in a context where some raw materials prices are reaching historically high levels (butter, eggs, oils, cocoa, cereals). This situation reinforces the disconnection between input prices and consumer prices, causing a value destructive spiral for the entire food chain.

The link between agricultural prices and consumer prices is not immediate. Indeed, commodities represent only a share of producing costs of food products. However, the relative disconnection between commodity prices rise and retail prices fall may reveal a market failure. This latter is, particularly, linked to the competitive structure at certain stages of the chain as well as inequalities in the balance of power between shareholders at different stages of the chain. In France, the most fragmented sectors concern production (450,000 farms) and processing (more than 17,000 agri-food industries). They find themselves confronted with four purchasing groups that represent more than 80% of food purchases on the French market. These structural differences deepen the divergences in terms of market power, in a sector where each of the major retail players strives to be the most competitive on prices to attract consumers, whose purchasing power has been eroded since the 2008 financial crisis.

The "price war" consequences on the various food chain actors is a point shared by stakeholders. In such a context of high market volatility and erosion of the competitiveness of economic actors, the "Food General Statements (FGS)" was launched on 20 July 2017 by the French Minister for Agriculture and Food. In return, the FGS have motivated the

"EGAlim" law of October 2018, aiming at rebalancing the initial relations between the various sector actors and ensuring a better value distribution.

This paper aims at investigating price transmission in the French food sector, by focusing more sophistically on the asymmetric aspect of price transmission at the downstream and upstream levels of the sector during the price war (over the period 01:2011-05:2019). Existing studies on the topic (very few) discuss the vertical price transmission, focusing on sub-sector activities rather than the whole French food sector. We intend to fill this gap considering the whole aggregated sector in one hand, and choosing the "war price" period as the one covered by our data sample. From a methodological viewpoint, in contrast to existing studies relying on usual linear approaches, we rely on a non-linear co-integration specification to investigate the price transmission according to increases or decreases in input prices. Therefore, our approach is based on a non-linear autoregressive distributed lag (NARDL) model, which allows us to test for asymmetry at both long- and short-run and to capture long-run asymmetry effects. Indeed, this specification, proposed by Shin et al. (2011), consists of a dynamic error correction representation integrating an asymmetric long-run co-integrating regression. More important, contrary to previous literature that estimates the price transmission in first differences and thus ignores its long-run behavior, our specification allows us to capture long-run asymmetry effects.

The remaining of this paper is organized as follows. Section 2 proposes a literature review focused on studies that justify the asymmetric nature of price transmission. Section 3 describes comprehensively the data we use in our empirical analyses. Sections 4 and 5 are respectively devoted to empirical methodology description and our main empirical results. in Section 6, we highlight some potential economic consequences of our results and conclude in Section 7.

2. Asymmetry in the price transmission

The vertical price transmission has motivated several studies in the agri-food literature, the existence of an asymmetric farm-retail price transmission being the focus of these studies. Thereby, Peltzman (2000) argues that the asymmetry in price transmission is the rule, not the exception. The term 'asymmetric', in what follows, refers to the dependence of a price reaction at one level of the chain on the sign of price change at another level.

To explain asymmetric price transmissions, several reasons, mostly related to the downward adjustment (retail level), are pointed out. One of these is the so-called *Menu Costs*.[1] Blinder (1982) argues that menu cost is the main reason firms do not react to temporary price changes, therefore inducing an asymmetric price transmission. For Ball and Mankiw (1994), firms facing menu costs and inflation may adjust their costs depending on whether prices are rising or falling. Indeed, a positive shock tends to enlarge the gap between the actual price and the desired price leading to a fast price adjustment, while regarding a negative shock, firms don't need to adjust the prices because the inflation offsets this price cut. In the same perspective, Kovenock and Widdows (1998) explain that menu cost drives the firm's choice of adjusting prices if this latter considers the change as transitory. For instance, to avoid re-engaging customers in search behaviour, firms choose not to inform their clients about changes in the market conditions. A further key element in the firm's adjustment process is inventory management. Indeed, the accounting methods and stock management can lead to an asymmetric price transmission (Balke et al., 1998; Blinder, 1982). Reagan and Weitzman (1982) explain that, when facing a low demand, firms tend to adjust the quantity and increase the inventory instead of decreasing prices. While in periods of high demand, firms will increase the price. This kind of behaviour creates an asymmetric price transmission.

[1] *"Menu cost"* is the cost to a firm resulting from changing its nominal prices, such as reprinting catalogs, labeling, etc. (Barro; 1972)

Public policies and political intervention can also lead to asymmetric price transmission (Gardner, 1975). In this vein, several studies have discussed the food chain price transmission in the Eurozone and among others the works by Ferrucci et al. (2007), Serra and Goodwin (2003) and Kinnucan and Forker (1987). Ferrucci et al. (2007) argue that the Common Agricultural Policy (CAP) plays a significant role in the price transmission, as the CAP subsidy system may lead firms to consider some shock transitory and then react differently to the price changes (increases/decreases). For instance, downstream level firms consider that a negative shock at the farm level will be offset by the CAP system, making this shock transitory. In this context, firms will respond faster to price increases than decreases leading to an asymmetric transmission. Also, governments implement some threshold prices, such as the resale threshold at a loss, inducing asymmetry in the firms' behaviour.[2] For their part, Kinnucan and Forker (1987) show that if the wholesalers and the retailers believe decreases are transitory (due to government intervention), but the increases appear to be more persistent, they will react differently to increases and decreases leading to a price asymmetry. Considering the Spanish Dairy sector, Serra and Goodwin (2003) argue that the quota system generates some scarcity of milk leading to an asymmetric price transmission. Thus, downstream firms compete to increase their market share, which causes incomplete transmission of increasing the farm prices.

A further explanation of the asymmetric price response particularly in the Agri-Food sector is market power. Along the food marketing chain, the stages are more and more concentrated, and firms enjoying market power influence the price adjustments to their advantage to capture profits and welfare. This market power effect is manifested through an oligopoly or oligopsony situation inducing distortions in the process of price adjustment (Meyer et al.,2004; Scherer, 1996; Beily and Brorsen, 1989). Studying retail gasoline

[2] For instance, the French government increased, in 2018, the resale threshold at loss, to assure a better price transmission and rebalance the trade negociation at the downstream level of the food marketing chain.

prices response to crude oil price changes, Borenstein et al. (1997) find results suggesting that oligopolistic co-ordination can cause a downward stickiness of retail prices, leading to an asymmetric price transmission. Another element associated with the market power effect is the collusive behaviour, discussed among others by Balke et al. (1998) and Brown and Yucel (2000). These authors explain how, in a non-competitive environment, firms aiming to maintain a high profits dynamic engage in collusive behaviour, leading to an asymmetric price transmission. In order to protect their agreement, firms tend to respond to price increases faster than price decreases.

For Blinder et al. (1998), the search costs in locally imperfect markets is also a driver of asymmetric adjustments. In the absence of any competition, firms can react more rapidly to increases than decreases in input prices because customers do not have complete information about price changes in other firms. In the same perspective, Benson and Faminow (1985) argue that the customers take into account the search costs, and their store choice is mainly based on location convenience leading to a locally non-competitive market and, therefore, to an asymmetric price transmission. In the dairy and meat markets, Gohin and Guyomard (2000) found that about one-fifth of the wholesale-retail price transmission is attributed to the oligopoly structure in the French food retail market. Their results reject the competitive behaviour of the French retail firms. Several studies suggest that the presence of market power leads to an asymmetric price transmission, but it is not clear how this price transmission will be skewed.

All these factors may cause the transmission of prices in the Agri-Food sector to be asymmetric in the short-run (asymmetry in speed of price transmission) as well in the long-run (asymmetry in magnitude of price transmission). In particular, the joint dynamics of prices along the marketing food chain can deviate from their historical relationships in the short-run due to menu costs or inventory management (among others...) and in the long-run owing to the structure of retail markets (power market).

3. Data and Stylized facts

As this paper aims to investigate the price transmission at different levels of the French food marketing chain, we focus on price indices at the agricultural, industrial, and consumer levels. Thereby, it is important to offer a clear insight into the data. This section is devoted to data description.

3.1. Data

In this analysis, we intend to use mainly three variables extracted from the Institut National de la Statistique et des Etudes Economiques (INSEE) and Eurostat databases, observed between January 2011 and April 2019 (year basis, 2015). These series are monthly farm producer price index, industrial product price index, and food consumption price index. Hereafter, we provide more details on these three series.

The Farm Producer Price Index (FPPI) relies on a monthly representative sample of transaction prices; the sample covers 158 agricultural products and 46 product groups. This index is meant to measure the change in prices received by farmers for their products. The Food Industrial Producer Price Index (FIPP) is calculated from monthly price surveys collected from a representative sample of companies. FIPP measures the change in prices of food goods resulting from industrial food activities. Finally, the Food Consumer Price Index (FCPI) of food and beverage products is used to measure the inflation of those products. Considering two given periods, it measures the average variation in the prices of food products consumed by households. It is based on the observation of a fixed basket of food goods, annually updated. Each product is weighted in proportion to its relative weight in household expenditure dedicated to food consumption.

3.2. Glance at the data

This subsection offers a first look at the evolution of price indices at different levels of the marketing food chain in France[3]. Observing Figure 1, we note that before the 2008 financial crisis, the price series showed an upward trend. During the crisis, the farm producer price index settled at a lower level, while the food consumption price index maintained a weak upward trend. For its part, processing prices stabilized over the same period.

Figure 1: Food prices at the three marketing chain levels (100 = 2015)

Source: INSEE and EUROSTAT

After the 2008 financial crisis, the different price series returned to their pre-crisis dynamics but with various evolutions. In 2013, the farm producer price index continued to move towards increasingly higher price levels, while the processing and food consumption price indices are part of a stable dynamic despite the steady evolution of farm producer prices. This episode of stability coincides with the beginning in France of the "price war" that retailers are delivering on the French market. In 2015, farm producer prices considerably fell before resuming an upward trend in 2016. Over the same period, processing and

[3] Table A.1 in the Appendix presents the descriptive statistics for price series as well as the number of negative and positive price changes. We can clearly see that the evolution of prices, along the agri-food marketing chain, shows as many increases as decreases over the price war period.

food consumption prices continue to move in a context of stability, slightly less pronounced for food consumption prices. The structure of the French food chain is characterized by a funnel shape, thus giving a specific market power growing along the chain. Downstream actors are taking advantage of this market power and waging a race for market shares by exerting pressure on the prices of their suppliers (farmers and processors) to keep their low pricing strategies.[4]

One of the most striking observations is the disconnection between the evolution of upstream prices and the rest of the actors' prices along the food marketing chain. The period 2013-2016 shows substantial variations in farm producer prices, while processing and food consumption prices remain stable or slightly upwards. This disconnection can have serious consequences on the economic health of the sector. Indeed, the absence of transmission or the low transmission of upstream prices leads to a deterioration in the margins of processors and farmers, since suppliers are charging prices disconnected from the reality of their production costs.

After a long period of stability, food consumption prices recovered slightly in 2017, marking the end of a long period of stability. This recovery is in line with the upward trend in farm producer prices, which also returned to an upward trend over the same period. Processing prices are slowing down and struggling to keep pace with the dynamics of the other players in the sector as they barely recover their upward trend. Moreover, the beginning of 2018 confirms this slowing recovery relative to that of the other players in the French agri-food chain.

[4] The concentration of downstream purchasing groups in France is very high, the four leading purchasing groups control more than 90% (compared to five purchasing groups for about 65% of the market in Germany).

4. Empirical methodology

4.1. Empirical literature

A relatively large literature, exploiting various model specification and methods, has been devoted to price transmission along the food marketing chain in the agricultural markets.

Meyer and von Cramon-Taubadel (2004) noticed the first attempt to assess asymmetric price transmission in Tweeten and Quance (1969). The latter used a dummy variable technique to distinguish the increasing and decreasing inputs prices, and in the case where the dummies' effects are statistically different, the symmetry hypothesis is to be rejected. From this work, a technique including first differences of the prices in the specification was developed (Wolfram; 1971, Houck; 1979) and extended by Ward (1982) to include lags of the exogenous variables. Thus, the delay effects can be different depending on whether the prices are decreasing or increasing. Thereby, a differentiation can be made between two types of asymmetry, the magnitude and the speed of transmission (Boyd and Borsen, 1988; Punyawadee et al., 1991; Mohanty et al., 1995; Aguiar and Santana, 2002).

Researches have started paying more attention to the time-series properties of the prices, as many of them are non-stationary. According to Granger and Newbold (1974), with non-stationary variables, a significant relationship can be misleading insofar as this relationship is, in fact, without any economic meaning. Therefore, testing for non-stationarity and eventual cointegration characterize the new generation of models used in this literature. Among others, studies by Von Cramon-Taubadel and Fahlbush (1994); Von Cramon-Taubadel and Loy (1996); Borenstein et al. (1998) and Frost and Bowden (1999) have relied on the cointegration approach in assessing price transmission.

The standard cointegration approach, however, assumes the series to have a symmetric long-run equilibrium relationship, and the impact of positive and negative components is to be the same (Schorderet, 2001; Shin et al., 2013), making the transmission asymmetry analysis possible only in the short run. Further, it is not possible to evaluate the price transmission asymmetry with respect to the magnitude. Indeed, asymmetry in the magnitude means that the positive and the negative price changes are permanently driven apart, meaning that the series cannot be co-integrated. An alternative approach is testing for asymmetric long-run adjustment instead of the usual cointegration tests. In this vein, Granger and Yoon (2002) introduced the "hidden cointegration" concept. Instead of considering the cointegration relationship between the main variables, one considers that the positive and negative components of the variables are co-integrated with the dependant variable, leading to an asymmetric cointegration. This approach makes it possible to test a long-run relationship between positive and negative input price changes and output price (Granger and Yoon, 2002; Schorderet, 2003).

4.2. The model

As previously mentioned, Granger and Young (2002) have shown that the symmetric cointegration is a particular case of the "hidden cointegration". In a first attempt, Schorderet (2003) proposes with a bivariate regression allowing for asymmetric cointegration. In 2011, Shin et al. (2011) proposed an asymmetric cointegration approach mainly based on the linear Autoregressive Distributed Lag (ARDL) model developed by Pesaran and Shin (1999) and Pesaran et al. (2001). The general form of linear ARDL is defined as follows:

$$\gamma(L)y_t = \alpha_0 + \beta'(L)x_t + u_t \tag{1}$$

where $\gamma(L) = 1 - \sum_{i=1}^{p} \gamma_i L^i$ and $\beta(L) = 1 - \sum_{j=1}^{q} \beta_j L^j$, with L being the lag operator and (p,q) respectively the lag order of the dependent variable (y_t) and the exogenous vector of variables (x_t) and u_t is the error term.

One of the main advantage of this approach is that it performs in detecting cointegration in small samples (Romilly et al., 2001), which is the case in the present paper.

Another advantage of the ARDL approach is that it can be applied without considering the order of integration (I(0) or I(1)), which is not the case of the other cointegration approaches (Pesaran and Shin; 1999). In our analysis, we will still control for the order of integration, since this technique is not valid in the case where the variables are I(2).

The approach proposed by Shin et al.(2011) is an extension of the linear ARDL allowing for the detection of non-linearities in the short and long-run asymmetries. The first step of Shin et al. (2011) method is to decompose the vector of exogenous variables into positive and negative components:

$$x_t = x_0 \ + \ x_t^+ \ + \ x_t^- \tag{2}$$

where x_t^+ and x_t^- represent partial sum of positive and negative changes of the exogenous variables, and are defined as follows:

$$x_t^+ = \sum_{j=1}^{t} \Delta x_j^+ = \sum_{j=1}^{t} max(\Delta x_j, 0) \tag{3}$$

$$x_t^- = \sum_{j=1}^{t} \Delta x_j^- = \sum_{j=1}^{t} min(\Delta x_j, 0) \tag{4}$$

For the state of presentation, we consider the following nonlinear long-term regression:

$$y_t = \alpha^+ x_t^+ + \alpha^- x_t^- + u_t \tag{5}$$

where α^+ and α^- are the long-run parameters, associated with the positive and negative changes in the independent variable. By combining the Eq. (5) and the ARDL specification (Eq. (1)), we get the asymmetric error correction model defined as follows:

$$\Delta y_t = \rho y_{t-1} \ + \ \eta^+ x_{t-1}^+ \ + \ \eta^- x_{t-1}^- \ + \ \sum_{j=1}^{p-1} \theta_j \Delta y_{t-j} \ + \ \sum_{j=0}^{q} (\omega_j^+ \Delta x_{t-j}^+ + \omega_j^- \Delta x_{t-j}^-) \ + \ e_t, \tag{6}$$

where $\eta^+ = -\rho\alpha^+$ and $\eta^- = -\rho\alpha^-$. The empirical procedure consists of four steps:

- The first step is the estimation of Eq. (6), using Ordinary Least Squares (OLS).
- The second step consists of controlling for the long-run relationship, by testing the joint null hypothesis of $\rho = \eta^+ = \eta^-$ using the Bounds-testing procedure proposed by Pesaran et al. (2001).
- Step three aims to test the long and short-run symmetry using a Wald test. The long-run symmetry test refers to $\eta^+ = \eta^- = \eta$ and the short-run symmetry to $\sum_{k=0}^{q} \omega_k^+ = \sum_{k=0}^{q} \omega_k^-$.
- In the final step, the asymmetric cumulative dynamic multiplier effects are derived of a unit change in the two components of the exogenous variable on the endogenous one. These effects are defined as follows:

$$m_h^+ = \sum_{j=0}^{h} \frac{\delta y_{t+j}}{\delta x_t^+} \quad and \quad m_h^- = \sum_{j=0}^{h} \frac{\delta y_{t+j}}{\delta x_t^-}, \qquad h = 0,\ 1,\ 2, ... \tag{7}$$

5. Empirical results

In this section, we estimate a non-linear autoregressive distributed lag (NARDL) model to assess price transmission along the food chain, considering the 2011-2019 period as it coincides with the "price war" in France. Indeed, this period is of a central interest insofar as the retailers have led a war of market shares, applying pressure on prices. Taking advantage of market power and in a race for low prices, retailers have pressured their suppliers, causing some non-linearity in the transmission of prices. Our aim thus is to evaluate this transmission upstream and downstream, particularly the asymmetric nature of the latter.

Although an ARDL specification can be used regardless of the integration order of the series (both I(0) and I(1)), we will perform unit root tests. It allows us to check that the series are not I(2), because in this case, the bounds testing test statistic is invalid.

We implement the Augmented Dickey-Fuller (ADF) and Phillip-Perron (PP) tests, to test for the null hypothesis of unit root. Table 1 summarises the results of unit root tests for the three series. They show that the farm producer price index and the food consumer price index are non-stationary in level but become stationary in the first difference. They are, thus, integrated of order 1. For the processing products price index, the ADF test indicates that it is stationary in level. The PP test confirms our results.[5] Since the unit-root tests suggest that the series are not I(2), we can now proceed to the bounds testing procedure.

Table 1: Augmented Dickey-Fuller unit root test

	Test statistic		Test statistic	5% critical value
		Trend		
lfpp	-1,43	Δlfpp	-7,09	
lipp	-4,135	Δlipp	-5,917	-3,452
lfcp	-2,733	Δlfcp	-8,929	
		No trend		
lfpp	-1,319	Δlfpp	-7,109	
lipp	-4,864	Δlipp	-5,852	-2,892
lfcp	-1,794	Δlfcp	-8,945	

Note: fpp, ipp and fcp stand respectively for Farm Producer Price, Industrial Product Price and Food Consumption Price. All the variables are in log. Lags orders are chosen using the Akaike Information Criterion

Next, we accordingly estimate an unrestricted non-linear ARDL specification, for the upstream and downstream levels, as described by the following equations:

[5]See Appendix A, Table A.2

$$\Delta lfcp_t = \mu + \rho lfcp_{t-1} + \eta^+ lipp^+_{t-1} + \eta^- lipp^-_{t-1} + \sum_{j=1}^{p-1} \theta_j \Delta lfcp_{t-j} + \sum_{j=0}^{q} (\omega_j^+ \Delta lipp^+_{t-j} + \omega_j^- \Delta lipp^-_{t-j}) + e_t, \quad (8)$$

$$\Delta lipp_t = \mu + \rho lipp_{t-1} + \eta^+ lfpp^+_{t-1} + \eta^- lfpp^-_{t-1} + \sum_{j=1}^{p-1} \theta_j \Delta lippi_{t-j} + \sum_{j=0}^{q} (\omega_j^+ \Delta lfpp^+_{t-j} + \omega_j^- \Delta lfpp^-_{t-j}) + e_t, \quad (9)$$

The equation (8) describes the downstream specification, with fcp_t and ipp_t standing for Food Consumption Price and Industrial Product Price, respectively. As for the equation (9), it describes the upstream level dynamics, with fpp_t standing for Farm Producer Price. All the variables are in logarithms.

The final specification is chosen, based on Akaike Information Criteria, by starting with max p=max and q =12 and dropping all non-significant regressors.[6] The cointegration test applied to the equations (8) et (9) consists of testing the joint null hypothesis that the coefficients of the lagged variables are equal to zero: $\rho = \eta^+ = \eta^- = 0$.

Table 2 reports the Bounds F-statistics for both models (upstream and downstream models). The results reveal that the two variables, in each specification, co-move in the long-run. Indeed, the F-statistics are respectively equal to 10.57 and 8.76 for the downstream and upstream equations, which are above the critical upper bound.[7] These results allow us to assess the relationship between the food consumption price and the positive and negative changes in industrial product price, on the one hand, and the industrial product price and the positive and negative changes in farm price, on the other hand.

[6] In practice, the inclusion of insignificant lags induce some inaccuracies and may introduce noise into the dynamic multipliers

[7] See Pesaran, M. H., Shin, Y., and Smith, R. J. (2001): Bounds testing approaches to the analysis of level relationships. Journal of applied econometrics, 16(3), 289-326.

Table 2: Bounds testing cointegration test

			95% lower bound	95% upper bound	Result
Downstream specification	F-statistic	10.5795	6.56	7.30	Cointegration
	t-statistic	-4.2365	-3.41	-3.69	
Upstream specification	F-statistic	8.7605	4.94	5.73	Cointegration
	t-statistic	-3.8113	-2.86	-3.22	

Note : The exact specification of the two models are reported in Tables 3 and 4.

Following these tests, we estimate the two specifications described by equations (8) and (9), representing the upstream and downstream dynamics of the agri-food sector.

Downstream specification

To test the non-linear specification adequacy of the relationship between the food consumer price index and the processing price index, we use the Wald test to assess the symmetry of the long and short-run coefficients. The lower part of Table 3 reports the results. In the long-run, The Wald test rejects the null hypothesis of symmetry between the coefficients of the positive and negative components of the processing price index. Indeed, the Wald test shows a p-value of 0.000 below the 5% threshold, suggesting rejection of the null hypothesis, and thus confirming the asymmetric nature of the relationship between the food consumer price index and the positive and negative components of the processing price index. This result supports the hypothesis that a linear specification is inadequate in modeling this relationship in the long-run.

For the short-run dynamic analysis, the Wald test suggests the rejection of the null hypothesis of short-run symmetry: $\sum_{i=0}^{q} \pi_i^+ = \sum_{i=0}^{q} \pi_i^-$ (Table 3). In other words, it rejects the null hypothesis of a weak form additive symmetric adjustment of the processing price index.[8]

[8]We retain the weak form additive symmetric adjustment in the short run asymmetry test, because it is

Once asymmetry has been established, we focus on the long-run dynamic of downstream price transmission. As indicated by the results reported in Table 3, the long-run coefficients of the changes in the processing price index are significant. The estimated long-run coefficients of the positive and negative changes in the processing price are -1.62 and -2.027, respectively. The latter are both negative, indicating that, regardless of changes in processing prices, food consumer prices experience a decreasing trend. Moreover, our estimations also show that the greater effect is sourcing from negative changes since the coefficient of negative changes is higher (-2.027) in magnitude than that of positive changes (-1.62). We can also note the presence of asymmetry in the speed transmission of food consumer prices to changes in supplier prices. Indeed, consumer prices react more quickly to the rise than to the fall in processing prices, as shown by the results in Table 3.

We can extend the analysis of the relationship between food consumption prices and processing prices with an analysis of dynamic multipliers. Figure 2 shows the dynamic response of food consumer prices to changes in the positive and negative components of processing prices. On the one hand, we observe that consumer prices react faster to the rise than to the fall in processing prices. On the other hand, we note that asymmetry is relatively unstable in the short-run. Indeed, during the first year, we observe that the asymmetry curve oscillates between positive and negative values.

Our results confirm the asymmetric dynamics of the transmission in the long-run. Indeed, we find that, over the 2011-2019 period, consumer prices always react downwardly to changes in processing prices, whether downward or upward. Our results also provide evidence of a dominant effect of downward trends, meaning that the different upstream price changes are not reflected symmetrically downstream. A possible explanation is the "price war" that the French agri-food environment has experienced. At the beginning of

more appropriate insofar as excluding the non-significant stationnary regressors introduces heterogeneity in the lags of the positive and negative components and the dependent variable.

Table 3: Downstream transmission specification

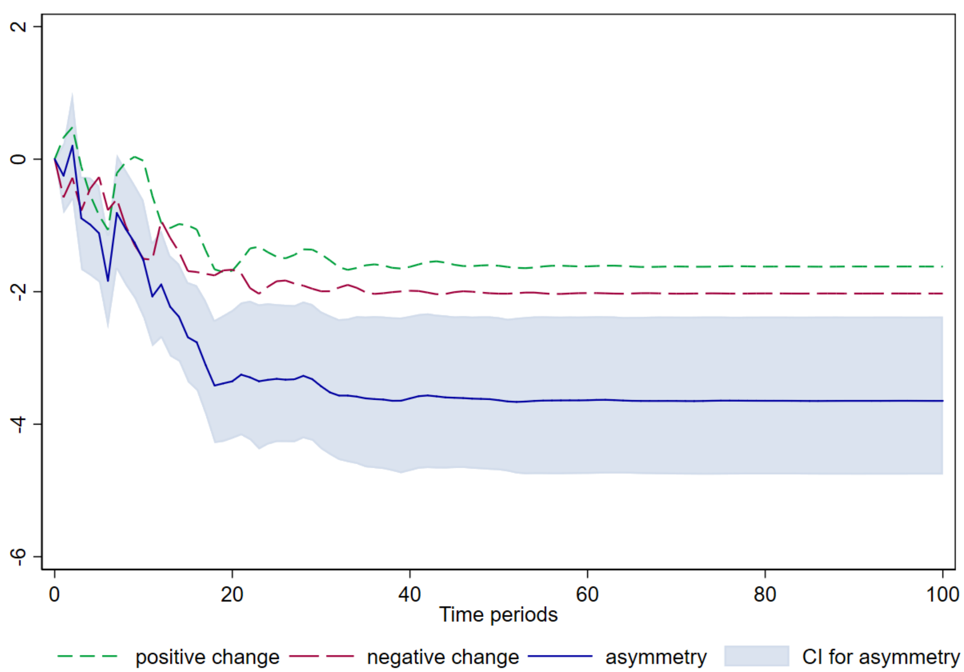
	Coefficients	t-value	p-value
Constant	0.564	4.24	0.000
Trend	0.0005	5.41	0.000
$lfcp_{t-1}$	-0.283	-4.24	0.000
$lipp^{+}_{t-1}$	-0.459	-5.24	0.000
$lipp^{-}_{t-1}$	0.573	4.60	0.000
$\Delta lfcp_{t-10}$	-0.235	-2.23	0.029
$\Delta lfcp_{t-11}$	-0.208	-2.10	0.040
$\Delta lipp^{+}_{t-1}$	0.787	3.01	0.004
$\Delta lipp^{+}_{t-2}$	0.706	2.60	0.011
$\Delta lipp^{+}_{t-7}$	1.013	3.93	0.000
$\Delta lipp^{+}_{t-8}$	0.564	2.29	0.025
$\Delta lipp^{+}_{t-9}$	0.529	2.12	0.038
$\Delta lipp^{+}_{t-10}$	0.409	2.01	0.048
$\Delta lipp^{-}_{t-2}$	-0.706	-2.88	0.005
$\Delta lipp^{-}_{t-4}$	-0.686	-2.93	0.005
$\Delta lipp^{-}_{t-5}$	-0.623	-2.66	0.010
$\Delta lipp^{-}_{t-7}$	-0.523	-2.03	0.047
$\Delta lipp^{-}_{t-12}$	-0.68	-3.14	0.002
Long-run coefficients			
	Coefficient	F-Stat	p-value
$lipp^{+}$	-1.620	20.18	0.000
$lipp^{-}$	-2.027	36.24	0.000

Asymmetry test			Diagnostics test		
Long-run	30.91	0.000	χ^2_{sc}	33.1	0.771
Short-run	43.26	0.000	JB	1.45	0.484

Note: This table reports the estimation of Eq.(9). The estimated long-run coefficients associated to the positive and negative components are defined by $\frac{\hat{\eta}}{\hat{\rho}}$. Long-run effect refers to a negative change in exogenous variable. The retained specification has been determined following the general-to-specific approach (Hendry, 1995).

2013, the retailers waged a market share war. In their races for low prices, they exerted downward pressure on consumer prices, explaining the disconnection between processing prices and food consumption prices. Benefiting from very significant market power, retailers reinforced their negotiation strategies with arguments based on their economic performance on the market, constrained by the aggressiveness of competition in the retail sector.[9] Indeed, under pressure from increased competition, retailers tend to lower their costs, i.e. the purchase prices of industrial products. This strategy puts considerable pressure on their suppliers and tariffs in an economic context already stressed by the high volatility of agricultural and energy input prices. This "price war" has several consequences for the various actors in the agri-food sector. On the one hand, consumers benefit from a relatively low and stable general level of food prices. On the other hand, it puts economic pressure on producers and processors. Indeed, agri-food companies have a declining margin rate that has reached its lowest level in 40 years.

Figure 2: Food processing prices cumulative effects on Food consumption prices.



Note: Figure 2 shows the predicted dynamic multipliers for the adjustment of food consumption prices to positive and negative changes in food processing prices. The asymmetry curve depicts the linear combination of the dynamic multipliers associated with positive and negative changes. Lower band and upper band for asymmetry indicate the 95% confidence interval. "CI" stands for Confidence Interval.

[9]Some recent studies have shown that retailers are enjoying increasing market power over food processors/farmers through the food marketing chain (European Commission 2009; Saitone and Sexton 2012).

Upstream specification

In the following, we investigate the upstream prices' transmission in the marketing chain. We assess the response of processing prices to positive and negative changes in farm prices. First, we assess the adequacy of the non-linear specification adopted. The Wald test reported in Table 4 indicates that in the long-run, the null hypothesis of symmetry is rejected at the 5% level suggesting asymmetry in the magnitude of price transmission. For the short-run behaviour, the Wald test shows a p-value of 0.063, above the 5% threshold. This last result suggests that at the 5% threshold, the null hypothesis of weak additive symmetric adjustment is not rejected. In other words, processing prices react symmetrically to changes in the positive and negative components of farm producer prices in the short-run, indicating a symmetric speed adjustment of prices. In the end, the dynamics upstream of the supply chain are symmetrical in the short-run, while in the long-run the transmission is asymmetric.

Discussing the long-run dynamics, the results in Table 4 show that the long-run coefficients are significant and also exhibit signs in line with the literature. Indeed, the estimated effects of the negative and positive components are respectively 0.33 and -0.376. As mentioned above, the Wald test indicates an asymmetry in the transmission of prices upstream of the sector, thus implying that decreases in farm producer prices are relatively more transmitted than increases to processing prices.

Figure 3 displays the dynamic effects of farm producer prices on the processing prices, and confirms the absence of asymmetry in the short-run. In the first 12 months, the absolute effect of a decrease in farm prices is greater than the effect of an increase. From the second year onwards, we notice that the effect of asymmetry tends to approach a situation of neutrality. Nevertheless, the persistence of asymmetry remains significant, and we observe that the absolute effect of a decrease dominates the effect of an increase in processing prices. Overall, we note that in the long-run, despite the persistence of asym-

Table 4: Upstream transmission specification

	Coefficients	t-value	p-value
Constant	0.236	3.83	0.000
$lipp_{t-1}$	-0.117	-3.81	0.000
$lfpp^{+}_{t-1}$	0.038	3.82	0.000
$lfpp^{-}_{t-1}$	0.044	4.37	0.000
$\Delta lipp_{t-2}$	0.161	1.74	0.087
$\Delta lfpp^{+}_{t}$	0.081	2.73	0.008
$\Delta lfpp^{+}_{t-2}$	-0.110	-3.48	0.001
$\Delta lfpp^{-}_{t}$	0.073	2.02	0.047
$\Delta lfpp^{-}_{t-2}$	0.135	3.31	0.001
$\Delta lfpp^{-}_{t-3}$	0.073	2.09	0.040
$\Delta lfpp^{-}_{t-4}$	0.084	2.12	0.037
$\Delta lfpp^{-}_{t-6}$	-0.118	-3.2	0.002
$\Delta lfpp^{-}_{t-11}$	-0.060	-1.98	0.051
Long-run coefficients			
	Coefficients	F-Stat	p-value
$lfpp^{+}$	0.330	31.11	0.000
$lfpp^{-}$	-0.376	41.95	0.000

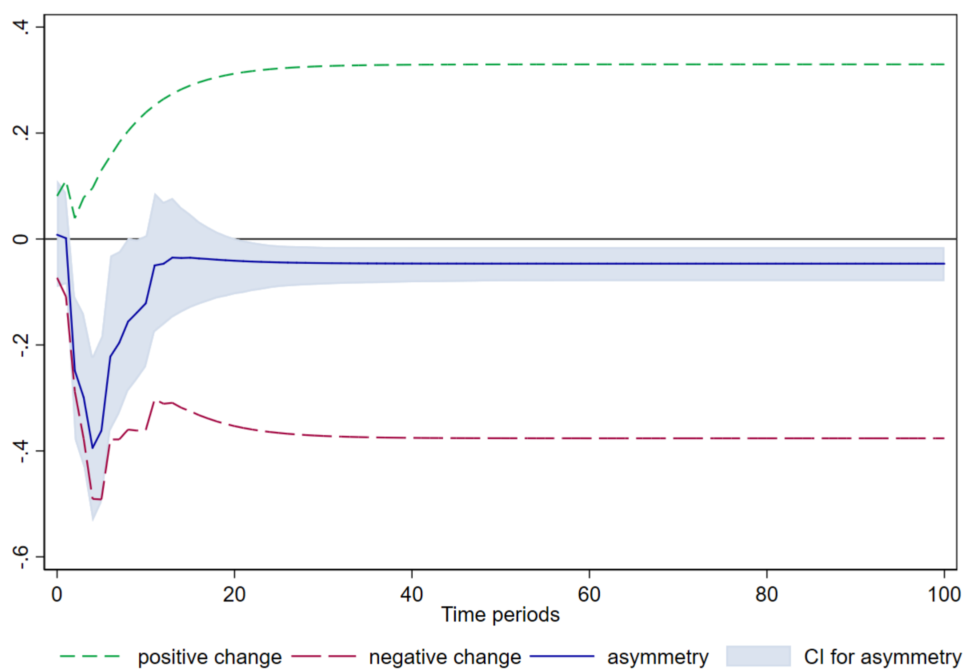
Asymmetry test			Diagnostics test		
Long-run	6.533	0.013	χ^2_{sc}	35.22	0.685
Short-run	3.573	0.063	JB	1.66	0.435

Note: This table repoorts the estimation of Eq.(9). The estimated long-run coefficients associated to the positive and negative components are defined by $\frac{\hat{\eta}}{\hat{\rho}}$. Long-run effect refers to a negative change in exogenous variable.The retained specification has been determined following the general-to-specific approach (Hendry, 1995)

metry in upstream price transmission, it remains relatively low. In other words, producers can track price changes for their products with slightly more difficulties when it comes to price increases.

During the "price war" period, despite the asymmetric nature of upstream price transmission, this upward rigidity remains relatively low. Thus, food processors, to a certain extent, take into account the economic conditions of their suppliers. At the same time, they are under enormous downstream pressure, which explains the low downward rigidity of the downstream price transmission. In a vice between a fragile upstream and a downstream price war, agri-food companies are, in a way, the adjustment variable in the price transmission process along the agri-food marketing chain.

Figure 3: Farm producer prices cumulative effects on processing prices



Note: Figure 3 shows the predicted dynamic multipliers for the adjustment of food processing prices to positive and negative changes in food farm producer prices. The asymmetry curve depicts the linear combination of the dynamic multipliers associated with positive and negative changes. Lower band and upper band for asymmetry indicate the 95% confidence interval. "CI" stands for Confidence Interval.

To check that the strong asymmetry in the downstream prices' transmission is specific to the price war period, we have estimated the same specifications over a broader period, namely 2005-2019.[10] First, we note that the upstream price transmission dynamic over the 2005-2019 period is similar to that observed during the price war, and in line with the literature (Meyer and von Cramon-Taubadel, (2004); Vavra and Goodwin, (2005)). Second, the results of the downstream level, on the broad sample, are utterly opposed to

[10]The estimation results are reported in Appendix B (tables B.3 and B.4).

our results for the price war period. These results confirm our central intuition. They confirm that the previous results are, indeed, specific to this period and that these effects are mainly due to the price war that characterized the French food sector over the 2011-2019 period.

6. Economic consequences

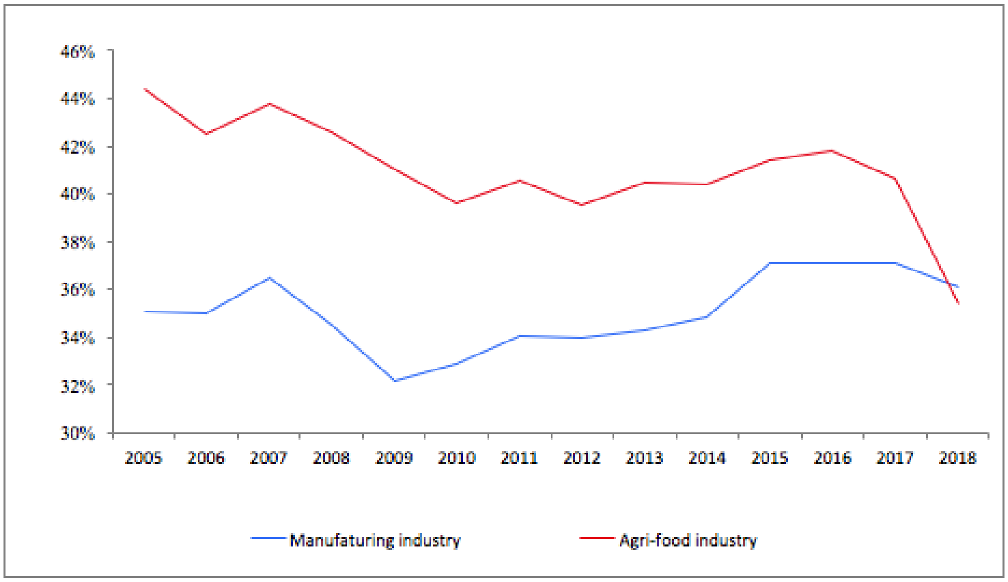
Asymmetry analysis in price transmission along the French food marketing chain, highlighted in this paper, reveals different economic consequences.

Our results show a disconnection between processing prices and food consumption prices downstream of the chain. Indeed, consumer prices are mostly adjusting downwards to the various variations in the prices of industrial suppliers. The price war that retailers have been waging since 2013 mainly explains this situation. Given the market power enjoyed by retailers,[11] on the French market, this asymmetry in price transmission may suggest an economic pressure on the industrial actors of the sector, characterised by a more fragmented structure.

With a buyer power benefiting to downstream actors, the conditions for commercial negotiations are unbalanced. Thus, in the scheme of an annual commercial negotiation, suppliers are unable to transmit their tariffs variations, leading to a disconnection between their negotiated tariffs and their production costs. In a context of high input price volatility, this disconnection has significant consequences on agri-food companies' margins. Indeed, in recent years, the margin rate of agri-food companies has shown a sharp downward trend, reaching historically low levels. Figure 4 compares the changes in margin rates in the manufacturing industry and the agri-food industry. Both series show a downward

[11] Competition authority, document No. 15-A-06 of 31 March 2015 on the purchasing groups and referencing in the mass retail sector.

Figure 4: Manuafturing and Food industries margin rate.



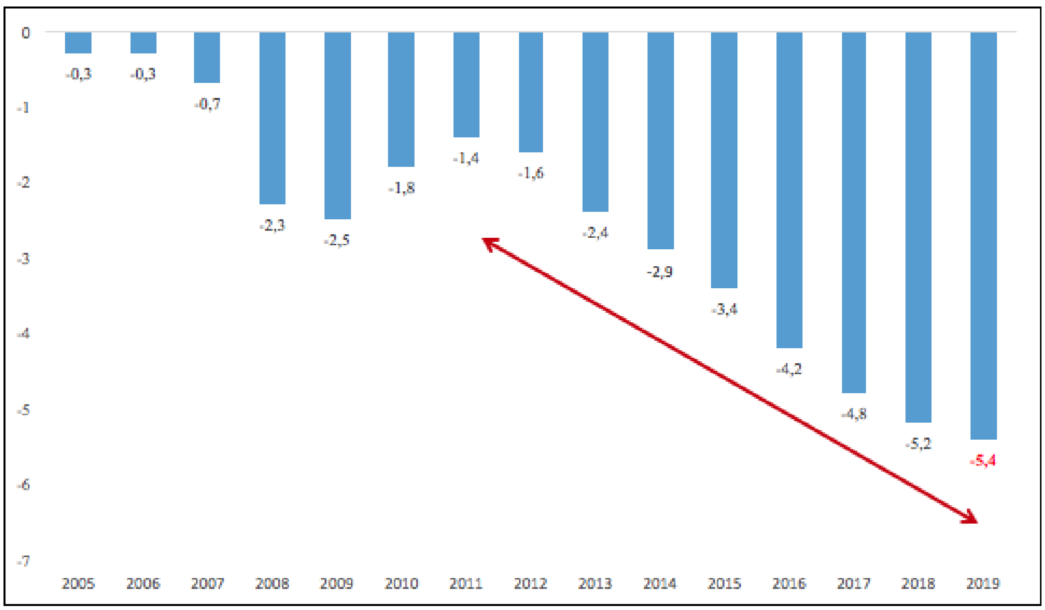
Note: The margin rate represents the ratio of gross operating margin to added value. The margin rate reflects what remains available to companies - the gross operating surplus - in particular as a return on capital, after deducting remuneration of employees.
Source: INSEE

trend during the crisis and post-crisis period, and they then recover an upward trend, slightly less pronounced for the agri-food industry. Towards the end of 2018, and after a long "price war" period, the margin rate of agri-food companies reached an unusually low level and for the first time, fell below the margin rate level of the manufacturing industry.

On the one hand, the contraction of agri-food companies' margins leads to a significant reduction in investments. In addition to this quantitative investment deficit, the investment orientation of agri-food companies may often appear defensive. In such conditions, companies would invest less in long-run projects such as innovation or modernization (robotics, digital...), investments that can ensure sustainable value creation. On the other hand, companies are more likely to make investments to comply with standards or to invest in the maintenance and upkeep of the production equipment, especially since the aging of the latter in France is a real brake for industrial growth. From a macroeconomic perspective, the renewal of production capacities is now the main reason for investment, in a context where the aging of the production tool increasingly limits the ability of the agri-food sector to rebound and, more generally, the competitiveness of companies.

The loss of export market share can measure this setback of competitiveness. While it is still largely positive (around 7 billion euros in 2019), the trade surplus of the agri-food industry has been declining steadily in recent years. Also, this surplus is sustained by a few sectors, in particular beverages or milk. Beverages excluded, the agri-food industry has a historically high trade deficit in exports (Fig. 5), particularly against their European competitors.

Figure 5: Trade deficit in the French food sector (Beverages excluded).



Source: INSEE

7. Conclusion

The present paper investigates the price transmission throughout the agri-food sector marketing chain in France, specifically the downstream transmission during the so-called "war price". This investigation contributes twofold to the literature, first by focusing on the recent price war in the French food sector and second by using a non-linear approach, more precisely the asymmetry, to evaluate the price transmission in the aggregate food sector in France. To this end, we exploit data on prices at the three levels of marketing chain: the food consumption price, the agricultural product price, and the industrial product price indices over the period from January 2011 to April 2019.

Firstly, our results suggest the presence of a non-linear downstream price transmission in both the short and long-run. The long-run reaction of food consumption prices to the industrial product price variations is asymmetric. Indeed, regardless of the sign of the industrial price variation, the consumption price reaction is downward, which is a consequence of the price war. Regarding the short-run dynamic, we also evidence an asymmetric behaviour of the price transmission, indicating an asymmetry in the speed of price transmission.

Secondly, we evaluate the upstream price transmission and find that the long-run effects are asymmetric and skewed to the left, indicating that there are increasing rigidities. In the short-run, there is no evidence of asymmetry, which means that the prices' pass-through is relatively the same during increases and decreases episodes. The difference in price transmission elasticities for positive and negative shocks is more significant for the downstream level than the upstream one, indicating that the long-run asymmetry is higher for shocks spread out from the processing level than the ones from the farm level. These results suggest that the main factor of this asymmetric dynamic is the price war along the food chain, induced by the abuse of market power of the downstream actors. It is also important to mention that other factors might have contributed to the asymmetric behavior of price transmission. Among these factors, we can name the increase of import prices of agricultural and food products, the cost of various structural changes in consumer demand (origin of products, additives, pesticides, etc.).

Finally, our results indicate that the linear hypothesis to describe the price transmission in the French food sector could be misleading. We shed light on the relevance of taking into account non-linearity in price behaviour for the aggregated sector. We are aware that a more disaggregated analysis could be interesting, as providing insights into specific sectors. Indeed, the heterogeneity of different food sectors could be of a central interest, and an investigation with sub-sectors-level data will probably be more informative.

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Appendix A

Table A.1: Descriptive statistics

Variable	Obs	Mean	Std.Deviation	Min	Max	Positive changes
			Price-war sample			
lfpp	172	1.977	0.053	1.866	2.047	47,0%
lipp	172	1.985	0.021	1.944	2.011	57,0%
lfcp	172	1.982	0.025	1.931	2.022	59,0%
			Whole sample			
lfpp	100	2.014	0.016	1.983	2.047	54,7%
lipp	100	1.999	0.007	1.974	2.011	58,1%
lfcp	100	2.001	0.010	1.973	2.022	62,2%

Note: The last column represents the percentage of price positive changes compared to negative one.

Table A.2: Phillips-Perron unit root test

		Test statistic		Test statistic	5% critical value
			Trend		
Z(rho) Z(t)	lfpp	-7,071 -1,83	Δlfpp	-66,279 -7,045	-20,466 -3,463
Z(rho) Z(t)	lipp	-8,928 -3,493	Δlipp	-46,596 -5,821	-20,466 -3,463
Z(rho) Z(t)	lfcp	-9,967 -2,747	Δlfcp	-81,085 -8,881	-20,466 -3,463
			No trend		
Z(rho) Z(t)	lfpp	-6,587 -1,733	Δlfpp	-66,047 -7,063	-13,684 -2,892
Z(rho) Z(t)	lipp	-9,784 -3,943	Δlipp	-43,202 -5,734	-13,684 -2,892
Z(rho) Z(t)	lfcp	-2,793 -1,797	Δlfcp	-80,585 -8,9	-13,684 -2,892

Appendix B

Preliminary tests

Table B.1: Augmented Dickey-Fuller unit root test (whole sample)

Test statistic		Test statistic		5% critical value
		Trend		
lfpp	-1.515	Δlfpp	-8.348	
lipp	-1.350	Δlipp	-8.706	-3,441
lfcp	-1.882	Δlfcp	-10.714	
		No trend		
lfpp	-1.618	Δlfpp	-8.351	
lipp	-1.714	Δlipp	-8.691	-2,885
lfcp	-1.167	Δlfcp	-10.731	

Note: fpp, ipp and fcp stand respectively for Farm Producer Price, Industrial Product Price and Food Consumption Price. All the variables are in log. Lags orders are chosen using the Akaike Information Criterion

Table B.2: Bounds testing cointegration test (whole sample)

			95% lower bound	95% upper bound	Result
Downstream specification	F-statistic	4.7434	6.56	7.30	Cointegration
	t-statistic	-3.5711	-3.41	-3.69	
Upstream specification	F-statistic	27.268	4.94	5.73	Cointegration
	t-statistic	-6.742	-2.86	-3.22	

Note : The exact specification of the two models are reported in Tables B.3 and B.4.

Whole sample estimation results

Table B.3: Downstream transmission specification (the whole sample)

	Coefficients	t-value	p-value
Constant	0.239	3.56	0.000
Trend	0.00002	2.08	0.039
$lfcp_{t-1}$	-0.124	-3.57	0.000
$lipp^{+}_{t-1}$	0.049	2.56	0.011
$lipp^{-}_{t-1}$	0.008	0.68	0.495
$\Delta lfcp_{t-1}$	0.122	1.78	0.076
$\Delta lfcp_{t-12}$	0.439	6.88	0.000
$\Delta lipp^{+}_{t-2}$	0.380	3.24	0.001
$\Delta lipp^{-}_{t-12}$	-0.123	-2.28	0.024
Long-run coefficients			
	Coefficient	F-Stat	p-value
$lipp^{+}$	0.402	22.56	0.000
$lipp^{-}$	-0.067	0.465	0.496

Asymmetry test			Diagnostics test		
Long-run	9.706	0.002	χ^2_{sc}	39.07	0.512
Short-run	14.59	0.000	JB	2.49	0.287

Note: This table reports the estimation of Eq.(9). The estimated long-run coefficients associated to the positive and negative components are defined by $\frac{\hat{\eta}}{\hat{\rho}}$. Long-run effect refers to a negative change in exogenous variable.The retained specification has been determined following the general-to-specific approach (Hendry, 1995)

Table B.4: Upstream transmission specification (the whole sample)

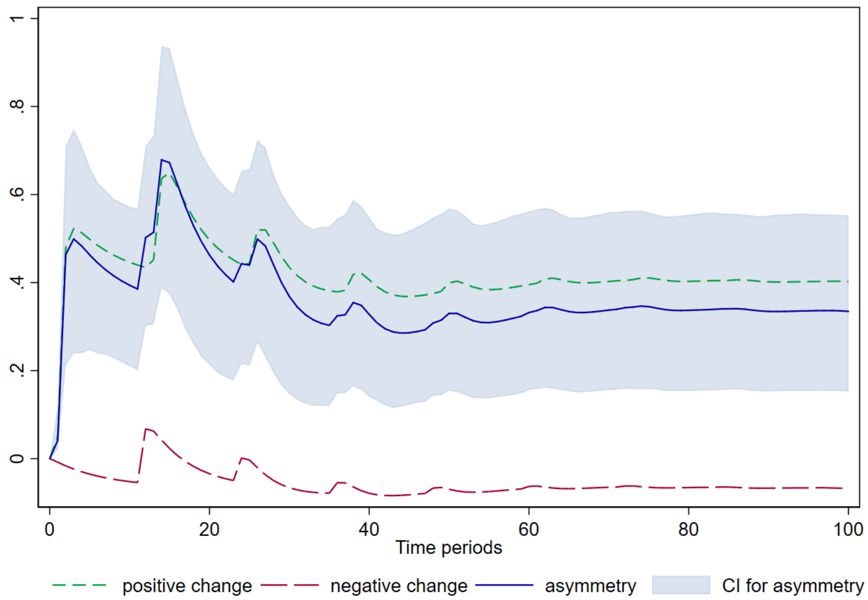
	Coefficients	t-value	p-value
Constant	0.263	6.78	0.000
$lipp_{t-1}$	-0.134	-6.74	0.000
$lfpp^{+}_{t-1}$	0.075	8.44	0.000
$lfpp^{-}_{t-1}$	0.087	8.71	0.000
$\Delta lipp_{t-2}$	0.224	3.95	0.000
$\Delta lipp_{t-6}$	-0.315	-4.93	0.000
$\Delta lipp_{t-11}$	-0.131	-2.31	0.022
$\Delta lipp_{t-12}$	-0.122	-2.19	0.030
$\Delta lfpp^{+}_{t-3}$	-0.120	-3.48	0.010
$\Delta lfpp^{+}_{t-9}$	-0.071	-2.28	0.024
$\Delta lfpp^{+}_{t-11}$	0.080	2.45	0.016
$\Delta lfpp^{+}_{t-12}$	0.067	2.03	0.045
$\Delta lfpp^{-}_{t-3}$	0.210	4.93	0.000
$\Delta lfpp^{-}_{t-7}$	-0.131	-3.36	0.001
$\Delta lfpp^{-}_{t-9}$	0.317	7.03	0.000
$\Delta lfpp^{-}_{t-10}$	-0.084	-2.30	0.023
Long-run coefficients			
	Coefficients	F-Stat	p-value
$lfpp^{+}$	0.562	131.4	0.000
$lfpp^{-}$	-0.652	88.95	0.000

Asymmetry test			Diagnostics test		
Long-run	16.92	0.000	χ^2_{sc}	35.9	0.655
Short-run	15.31	0.000	JB	547.5	0.000

Note: This table repoorts the estimation of Eq.(9). The estimated long-run coefficients associated to the positive and negative components are defined by $\frac{\hat{\eta}}{\hat{\rho}}$. Long-run effect refers to a negative change in exogenous variable.The retained specification has been determined following the general-to-specific approach (Hendry, 1995)

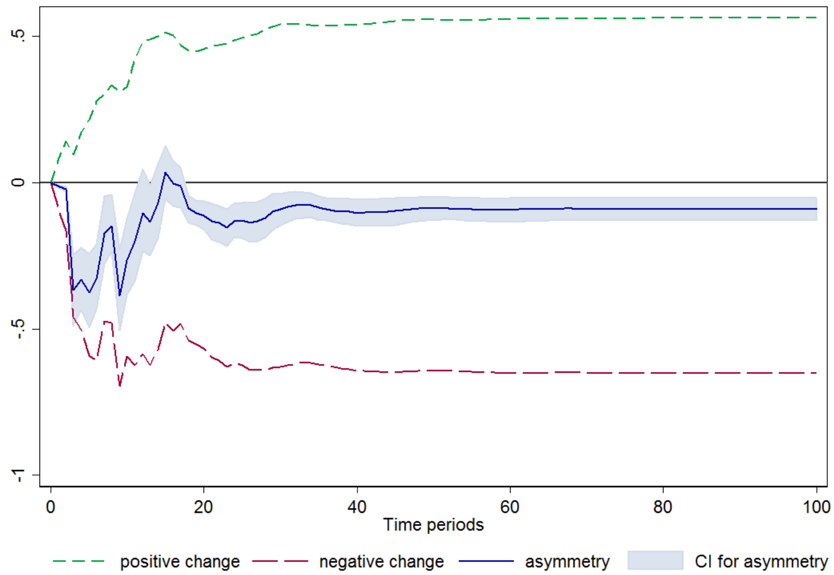
Dynamic cumulative effects

Figure B.1: Processing prices cumulative effects on food consumption prices



Note: Figure B.1 shows the predicted dynamic multipliers for the adjustment of food consumption prices to positive and negative changes in food processing prices. The asymmetry curve depicts the linear combination of the dynamic multipliers associated with positive and negative changes. Lower band and upper band for asymmetry indicate the 95% confidence interval. "CI" stands for Confidence Interval.

Figure B.2: Farm producer prices cumulative effects on processing prices



Note: Figure B.2 shows the predicted dynamic multipliers for the adjustment of food processing prices to positive and negative changes in food farm producer prices. The asymmetry curve depicts the linear combination of the dynamic multipliers associated with positive and negative changes. Lower band and upper band for asymmetry indicate the 95% confidence interval. "CI" stands for Confidence Interval.