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ARE PRICE ADJUSTMENTS ASYMMETRIC IN BASIC FOOD CATEGORIES? THE CASE OF THE GREEK FOOD MARKET*

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1 INTRODUCTION

Asymmetric price transmission is of particular relevance when it comes to commodities, food in particular, because of the incommensurate impact on low-income consumers, who tend to spend a higher share of their income on basic food items. To the extent that consumer prices respond faster to rises in producer prices than to decreases, profit margins for producers widen and a higher producer surplus is recorded to the detriment of consumers. The more competitive an industry/sector is, the more likely it is for price transmission from producers to final consumers to be direct and symmetric. In reality though, wholesalers and retailers, who represent the intermediate links of the food supply chain, strongly affect the effective functioning of the market and the price transmission mechanism.

By "positive asymmetry" it is understood that retail prices adjust to a greater degree or at a faster pace to a rise in (wholesale) producer prices than to a respective decline, while "negative asymmetry" refers to the case where retail prices adjust to a greater degree or at a faster pace to a decline in producer prices than to a respective rise. Thus, negative asymmetry favours final consumers, as it increases consumer surplus, whereas positive asymmetry harms them, reducing consumer surplus to the benefit of producers and/or wholesalers.

Most empirical studies suggest that price adjustment asymmetries are caused by the monopoly power of firms operating in the food supply chain, i.e. the strongly oligopolistic structure of the wholesale and retail markets (Peltzman 2000; Meyer and von Cramon-Taubadel 2004).

The main objective of this paper is to empirically investigate the existence of asymmetries in the speed of retail price adjustment to a long-run equilibrium, focusing on a group of basic food commodities. The food categories under examination are cereals, meat and meat products, dairy products, fruit, and vegetables. This group of foods was chosen for the following reasons: (a) there is evidence that these food industries are characterised by imperfect competition conditions and exhibit high concentration, leading to oligopoly pricing; and (b) it carries increased weight in the average consumer "basket". Thus, any changes in the prices of this food group greatly affect the disposable income and the living standards of the average consumer.

To this end, an error correction model (ECM) with threshold autoregressive (TAR) cointegration was applied. Data refer to the period from January 2002 to June 2016. This paper draws on the fullest and most recent sample of available monthly observations for the variables examined, compared with any other study on the Greek food market.

The remainder of the paper is structured as follows: Section 2 provides a brief literature review on price adjustment in the Greek food market. Section 3 outlines the main causes of asymmetric adjustment of food prices. Section 4 documents the importance of the basic food group for the average household and estimates the Herfindahl-Hirschman Index (HHI), measuring market concentration in the respective food industries. The data used are presented in Section 5. Section 6 describes the econometric methodology. Section 7 reports the empirical findings. Finally, Section 8 summarises and concludes.

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2 LITERATURE REVIEW

There is a relatively limited literature on the Greek food market. Such studies provide a brief description of the food industry or of individual food categories, explore the price transmission mechanism and note imperfect competitive conditions. More specifically:

Reziti (2004) studied empirically the market for lettuce, oranges, potatoes and meat in Greece over the period 2000-2003, using monthly data. Empirical results in the short run support the existence of asymmetries (with the exception of the orange market) in price transmission. In the event of falling producer prices, retailers delay lowering their prices, while conversely, when producer prices rise, retailers are quick to directly pass on increases to consumers.

Reziti and Panagopoulos (2006; 2008) studied empirically the food, vegetable and fruit markets, using monthly data for the period 1994-2005. Empirical results corroborate the existence of speed asymmetries in price transmission from producers to consumers in the cases of food and vegetables, but not in the case of fruit.

Tsakistara et al. (2008) studied the milk market, using annual data for the period 1990-2008, and concluded that oligopoly conditions prevail in this market, indirectly confirming an alleged milk cartel.

The Hellenic Competition Commission (2011) conducted a sectoral survey for the period 2005-2011 on fruit and vegetables, focusing on a sample of seven products (apples, oranges, peaches, potatoes, cucumbers, and lettuce) on the basis of their resepctive shares in the so-called "housewife's shopping basket". The results point to significant asymmetric response to both positive and negative price changes along the fruit and vegetable supply chain. Such asymmetries refer to the speed of changes' transmission as well as to the size and the persistence of adjustment to the new levels.

Rezitis and Reziti (2011) carried out an empirical study on the milk market, using monthly data for the period 1989-2009. Empirical findings suggest an asymmetric speed of transmission between producer and consumer prices. Consumer prices rise faster than producer prices to restore the long-run equilibrium between consumer and producer milk prices.

Reziti (2014; 2016) investigated the transmission mechanism between producer and consumer prices in the milk market, employing monthly data for the period from January 1998 to August 2014. Empirical evidence shows that the price transmission mechanism is asymmetric, implying that dairy companies and big retailers have the power to shape price dynamics at the expense of both milk producers and consumers.

3 CAUSES OF ASYMMETRIES IN THE FOOD MARKET

The literature has empirically shed light on possible factors behind the asymmetric adjustment of food retail prices.

The most plausible explanation is thought to be the monopoly power of firms, which urges them to engage in a tacit collusion (cartels) to systematically maintain higher profit margins (Brown and Yücel 2000; Miller and Hayenga 2001; Peltzman 2000; Meyer and von Cramon-Taubadel 2004).

Second, asymmetric price transmission is also due to the perishability of some food items (e.g. milk, fresh fruit and vegetables). Ward (1982) argues that producers-farmers and resellers (wholesalers and retailers) holding perishable products may choose not to raise their prices when input prices rise to avoid the risk of accumulating spoiled goods.

Third, adjustment or re-pricing costs as a result of movements in input prices play an important role for firms. Higher input prices are more likely to lead to adjustments in the retail price of the end product than lower input prices. Some inflation-induced increases are inevitable because of lower actual profit margins (Ball and Mankiw 1994; Buckle and Carlson 2000).

Fourth, government intervention setting price floors may lead to asymmetric price transmission, if it urges wholesalers or retailers to expect that the decrease in agricultural product prices will only be temporary. This occurs because they anticipate that the State will intervene to support producer prices (Kinnucan and Forker 1987).

Fifth, information asymmetries are also observed between firms. Larger firms are typically better informed of developments in input prices during the production process, such as developments in international oil or grain prices. Besides, they can acquire key raw materials faster and at a lower cost than their competitors, allowing them greater freedom of adjusting the retail prices of their products (Bailey and Brorsen 1989).

4 THE IMPORTANCE OF BASIC FOOD ITEMS AND THE DEGREE OF CONCENTRATION ACROSS SECTORS

This section attempts to highlight the particular relevance of the food group under examination (dairy products, meat, cereals, fruit, and vegetables) for the consumption pattern of the average household. Moreover, the Herfindahl-Hirschman Index (HHI), which measures the degree of concentration in the food industries under review, is estimated (Herfindahl 1950; Hirschman 1964; Rhoades 1993).

According to the results of the Household Budget Survey 2015 that were released in July 2016, households' average monthly expenditure on goods and services in 2015 amounted to EUR 1,419.57. Households' average monthly expenditure on food items (food and non-alcoholic beverages) was EUR 293.30, or 20.7% of total household average monthly expenditure, and constitutes the main expenditure of households over time (see Table 1). Expenditures on housing (EUR 189.21, or 13.3%) and transport (EUR 181.64, or 12.8%) follow suit.

It can be seen that, while the average monthly expenditure of households according to the Household Budget Survey 2010 stood at EUR 1,956.42, in 2015 it fell by 27.4% (or EUR 536.85) to EUR 1,419.57. On the other hand, household average monthly expenditure on food items over the same period (2010-2015) dropped by only 16.6% (or EUR 58.37), suggesting consumers' inelastic income demand. It should also be pointed out that the share of food items in average monthly expenditure grew over time, from 18% in 2010 to 20.7% in 2015.

In 2015, as well as throughout the period 2010-2015, this food group (dairy products, meat, cereals, fruit, and vegetables) has an increased weight in households' average monthly expenditure on food items (see Table 2).

Expenditure on the reviewed food group accounts for 75% (EUR 218) of households' average monthly expenditure on food items (EUR 293.30). Meat, with a share of 22.4% (EUR 65.56), accounts for the bulk of households' expenditure on food, followed by dairy products with 17.6% (EUR 51.52), flour, bread and cereals with 15.3% (EUR 44.84), vegetables with 12.1% (EUR 35.62), and fruit with 7.2% (EUR 21.26). Therefore, the degree of market efficiency for those basic food items strongly affects households' disposable income and living standards.

The Herfindahl-Hirschman Index (HHI),[1] which measures the degree to which a small number of firms represents a large segment of the market, confirms high market concentration in all food industries under review. The

1 The HHI is used as a potential indicator of market power or of competition among firms. When firms' market shares do not add up to 100, the ensuing deviation is taken into account in the calculation of the HHI as unknown market share.

Table 1 Household average monthly expenditure on goods and services

Goods and services	Household Budget Survey (HBS)													
	2010		2011		2012		2013		2014		2015		Changes 2010-2015	
	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)
Average monthly expenditure (total)	1,956.42	100.0	1,824.02	100.0	1,637.10	100.0	1,509.39	100.0	1,460.52	100.0	1,419.57	100.0	-536.85	-27.4
Food	351.67	18.0	355.05	19.5	328.57	20.1	307.33	20.4	299.79	20.5	293.30	20.7	-58.37	-16.6
Alcoholic beverages, tobacco	68.7	3.5	66.52	3.6	62.71	3.8	62.8	4.2	58.8	4.0	57.27	4.0	-11.43	-16.6
Clothing and footwear	140.84	7.2	112.51	6.2	95.34	5.8	87.38	5.8	85.7	5.9	83.06	5.9	-57.78	-41.0
Housing	228.82	11.7	230.16	12.6	227.07	13.9	206.99	13.7	195.29	13.4	189.21	13.3	-39.61	-17.3
Durables	130.47	6.7	110.05	6.0	94.97	5.8	83.94	5.6	72.76	5.0	66.49	4.7	-63.98	-49.0
Health	124.43	6.4	114.58	6.3	104.71	6.4	104.44	6.9	105.76	7.2	107.06	7.5	-17.37	-14.0
Transport	264.87	13.5	240.05	13.2	209.88	12.8	189.19	12.5	184.82	12.7	181.64	12.8	-83.23	-31.4
Communications	78.46	4.0	73.69	4.0	68.19	4.2	61.91	4.1	60.08	4.1	58.46	4.1	-20	-25.5
Recreation, culture	91.6	4.7	85.72	4.7	72.87	4.5	68.82	4.6	68.71	4.7	67.95	4.8	-23.65	-25.8
Education	64.21	3.3	63.71	3.5	57.33	3.5	50.83	3.4	50.84	3.5	46.7	3.3	-17.51	-27.3
Hotels, cafes, restaurants	209.75	10.7	189.11	10.4	160.47	9.8	145.55	9.6	143.49	9.8	141.05	9.9	-68.7	-32.8
Miscellaneous goods and services	202.61	10.4	182.89	10.0	154.98	9.5	140.19	9.3	134.49	9.2	127.37	9.0	-75.24	-37.1

Source: ELSTAT.

Table 2 Average monthly expenditure on food

Food and non-alcoholic beverages	HBS 2015		HBS 2014		HBS 2015/2014	HBS 2015/2014
	Value (in EUR)	Distribution (%)	Value (in EUR)	Distribution (%)	Changes in value (%)	Differences in percentage distribution (%)
Total	293.30	100	299.79	100	3.9	
Flour, bread, cereals	44.84	15.3	46.99	15.7	-4.6	-0.4
Meat	65.56	22.4	68.06	22.7	-3.7	-0.3
Fish	21.16	7.2	21.6	7.2	-2.2	0.0
Dairy products and eggs	51.52	17.6	54.24	18.8	-5.0	-0.5
Oils and fat	17.74	6.0	17.59	5.9	0.9	0.1
Fruit	21.26	7.2	21.39	7.1	-0.6	0.1
Vegetables	35.62	12.1	35.36	11.8	0.7	0.3
Sugar, jams, honey, chocolate, confectionery	14.57	5.0	13.76	4.6	5.9	0.4
Food products n.e.c.	6.10	2.1	4.92	1.6	24.0	0.5
Coffee, tea and cocoa	6.21	2.1	6.68	2.2	-7.0	-0.1
Mineral water, soft drinks, fruit and vegetable juices	8.72	3.0	9.21	3.1	-5.3	-0.1

Source: ELSTAT.

HHI is a measure of market concentration, calculated by squaring the market share of each firm in the sector/industry and then summing the resulting numbers. The formula for the HHI is the following:

$$\text{HHI} = \sum_{i=1}^{n} (s_i)^2 = s_1{}^{\wedge 2} + s_2{}^{\wedge 2} + s_3{}^{\wedge 2} + \dots + s_n{}^{\wedge 2}$$

where s_i is the market share of firm i, with a total number of firms n in the market. When the HHI is lower than 1,000, market concentration is characterised as low; when it ranges between 1,000 and 1,800, concentration is considered to be medium; and a HHI above 1,800 denotes high market concentration. The HHI takes the maximum value of $10{,}000=100^2$ if there is a monopoly, i.e. the firm has a 100% market share. Conversely, the HHI takes a very low value close to zero, if the market operates under perfect competition.

Table 3 presents the number of firms by food sector, the respective market shares and the Herfindahl-Hirschman Index (HHI) in 2014.

In the cereals sector, in the industry of bread and bakery products in particular, two firms held a combined market share of 24% in terms of turnover (sales) and of 8% in terms of employment, while eleven firms accounted for 42% of total sales and employed 21% of the sector's total workforce. The HHI was calculated at 2,926 (>1,800), indicating high market concentration. With regard to the flour mill industry, 17 firms accounted for 75% of total sales in 2013, with a high HHI of 8,200 (>1,800).

Turning to the sector of meat and meat products, market concentration was also strong, since four very large firms and 17 large firms were active, with the industry of cured meat products exhibiting a very high concentration index of 5,822 (>1,800).

In the sector of dairy products, only four firms accounted for more than 53% of total sales in 2013. The HHI was calculated at 4,094 (>1,800), indicating high market concentration.

Table 3 The concentration index Herfindahl-Hirschman (HHI)

Firms / Food sectors	Very large		Large		Small and medium-sized		Total of firms in the sample		Market concentration index Herfindahl-Hirschman (HHI)	Conclusion
	Sales over EUR 250 million		Sales between EUR 50 million and 250 million		Sales between EUR 2 million and 50 million					
	Number of firms	Market share	Number of firms	Market share	Number of firms	Market share	Number of firms	Market share		
Cereals and cereal products[1]										
Bakery and confectionery products	2	24%	11	42%	80	19%	93	85%	2926	High concentration
Flour mill industry			1		17		18	90%	8200	High concentration
Meat and meat products	4		17		1,514[2]					High concentration
Cured meat products			3	74%	5	11%	8	85%	5822	High concentration
Dairy products and eggs	4	53%	5		56		65	86%	4094	High concentration
Fruit and vegetables	2[3]	20%	59	25%		35%		90%	2350	High concentration

Source: Author's own calculations based on data from Grant Thornton and IELKA.
1 Four firms have a combined share of around 50% in the Greek bread and bakery market, while 17 firms accounted for 75% of the flour mill industry in 2013.
2 256 industrial firms and 927 commercial firms. of which 162 industrial and 179 commercial firms were medium-sized.
3 Central markets are organised wholesale points, mainly consisting of: (a) the Athens Central Market Organisation S.A. (OKAA), with 550 active wholesalers of fresh fruit and vegetables, and (b) the Thessaloniki Central Market S.A. (KATH), which comprises 280 shops. According to ICAP (2015), the two central markets hold a market share of 20%, wholesalers outside central markets a 35% share, super markets a 25% share and farmers' markets the remaining 20%.

With respect to the fruit and vegetable sector, the country's two central markets (OKAA and KATH)[2] held a market share of 20%, followed by wholesalers outside central markets with 35%, super markets with 25% and farmers' markets with the remaining 20%. The HHI was calculated at 4,050 (>1,800), indicating the sector's high concentration.

5 THE DATA

The data used in the empirical investigation are producer price indices and retail price indices for the main food categories: cereals and cereal products; meat and meat products; dairy products and eggs; fruit; and vegetables. Data refer to the period from January 2002 to June 2016. Data sources are Eurostat statistics, the Hellenic Statistical Authority (ELSTAT) and the Bank of Greece. In greater detail, food retail prices are reflected in the respective sub-indices of the Harmonised Index of Consumer Prices (HICP), while food producer prices are captured by the respective categories of the Producer Price Index (PPI) of the domestic market.

To ensure the robustness of empirical results, the HICP and the PPI must refer to the same basket of goods to the extent possible. As

2 OKAA: Athens Central Market Organisation S.A., with 550 active wholesalers of fresh fruit and vegetables; and KATH: Thessaloniki Central Market S.A., which comprises 280 shops.

Table 4 Consistency between HICP and PPI classification codes

HICP (2015=100)	PPI (2010=100)
Food categories	
Cereals and cereal products	Mill and cereal products, starches and starch products
Meat (in general)	Cured meat and meat products
Dairy products and eggs	Dairy products
Fruit (in general)	Prepared and preserved fruit and vegetables
Vegetables (in general)	

Source: ELSTAT.

shown in Table 4, there is indeed great consistency between classification codes referring to price indices for the specific food groups.

Nevertheless, the present study encountered a number of problems with regard to data, as described below:

First, in the fruit and vegetable category, the HICP comprises two separate series, whereas the PPI has only one consolidated series. Second, in the cereals category, in order to ensure greater consistency between the HICP and the PPI, two separate sub-indices of the PPI were consolidated into one weighted sub-index. The weighted sub-index comprises mill and cereal products, starches and starch products on the one hand and bread/bakery and flour products on the other hand. Third, the base year of the PPI is adjusted from 2010 to 2015 to ensure consistency with the HICP.[3] Fourth, while account is also taken of the value added tax (VAT) when measuring prices on the basis of the HICP, for the PPI basic prices are collected, i.e. prices minus VAT or similar deductible taxes that are directly associated with turnover and minus duties or other taxes on products, but plus subsidies on products. For this reason, VAT is excluded from the HICP to ensure harmonisation with the PPI.[4] Fifth, food indices (mainly fruit and vegetables) that are measured using the methodology for the compilation of either the HICP or the PPI are characterised by strong seasonality. Seasonality, which can be attributed to weather conditions and/or seasonal patterns of consumption behaviour, hinders the investigation into possible asymmetric price responses that are due to the oligopolistic structure of those sectors. Against this backdrop, all indices in the study are seasonally adjusted using the X-12 ARIMA filter.

6 THE ECONOMETRIC METHODOLOGY

To empirically investigate whether there is asymmetric adjustment of food prices to their long-run equilibrium, an error correction model (ECM) with threshold autoregressive (TAR) cointegration is employed, as proposed by Enders and Siklos (2001). In the event that retail prices and producer prices follow a long-run equilibrium relationship, the latter can be expressed as follows:

$$r_t^r = \alpha_0 + \beta_1 r_t^w + u_t \qquad (1)$$

where r_t^r denotes the retail price of a food item in period t, while r_t^w denotes the producer price of the item over the same period t. Relationship (1) is expressed in logs. The log-transformation of the series normalises the effect of any squared trends in the evolution of the series and, besides, the estimations of the equation parameters are directly expressed in elasticities. Parameter a_0 represents a constant cost, while parameter β_1 reflects the degree of retail price adjustment to producer price movements and indirectly captures the intensity of market competition. The more intense the competition, the larger the expected degree of adjustment and the narrower the expected profit margin. u_t symbolises the error term with constant mean and variation, which reflects short-term deviations from the long-run equilibrium relationship. When prices r_t^r and r_t^w reach their long-run equilibrium, residuals are

3 Different base years create problems in the estimation of models, as differences are transferred to the fixed term, distorting its estimation. This issue is addressed through base year adjustment. As a rule, the most recent reference year is opted for.
4 Comparing prices or indices excluding taxes with prices or indices at constant taxes distorts the outcomes (IOBE 2009).

expected to be zero. By contrast, if prices deviate from the long-run relationship, the error term u_t may be either positive, when the price r_t^r is higher than its long-run equilibrium, or negative otherwise.

If the adjustment to the long-run equilibrium is asymmetric, suggesting that the autoregressive behaviour of residuals in the long-run relationship $\hat{u}_t$ is conditional upon the sign of exogenous shocks, then the traditional Dickey-Fuller (1979) tests as well as their augmented versions for cointegration using the two-step procedure by Engle and Granger (1987) are misspecified, because they assume a symmetric adjustment of residuals to the long-run equilibrium (Enders and Granger 1998; Enders and Siklos 2001). In line with this technique, the typical Dickey-Fuller (1979) test equation for cointegration, which is given by the following relationship:

$$\Delta\hat{u}_t = \rho\hat{u}_{t-1} + v_t \tag{2}$$

is augmented, incorporating a functional importance indicator I_t, according to which the residuals of the long-run relationship (1) are split into two components with a time lag and which is given by the following relationship:

$$\Delta\hat{u}_t = I_t\rho_1^{up}\hat{u}_{t-1}^{up} + (1 - I_t)\rho_2^{down}\hat{u}_{t-1}^{down} + v_t \tag{3}$$

If after the estimation of relationship (3) some form of autocorrelation is observed, then the appropriate number of time lags is inserted $\Sigma\Delta\hat{u}_{t-j}$.

The functional importance indicator I_t is given by relationship (4):

$$\mathrm{I}_t = \begin{cases} 1 & \text{if } \hat{\mathrm{u}}_{t-1}^{up} \geq \hat{\tau} \\ 0 & \text{if } \hat{\mathrm{u}}_{t-1}^{down} < \hat{\tau} \end{cases} \tag{4}$$

where the value of the point (or threshold) is stochastic and denoted with the letter $\hat{\tau}$. $\{v_t\}$ denotes the residual series of relationship (3) with a mean of zero and constant variance, so that the vector $\{v_t\}$ is independent of $\hat{u}_t$. The combination of relationships (3) and (4) is the threshold autoregressive (TAR) model. Allowing the parameters ϱ_1^{up} and ϱ_1^{down} to take different values, the model acknowledges that positive and negative deviations from equilibrium can be corrected at different adjustment speeds.

The cointegration test can be performed using the critical values of the distributions t-Max and Φ^*, as suggested by Enders and Siklos (2001). Furthermore, Tong (1982; 1983; 1990) showed that in the event of a threshold autoregressive cointegration, i.e. if the hypothesis H_0: $\varrho_1^{up}=\varrho_1^{down}=0$ is rejected, the parameters ϱ_1^{up} and ϱ_1^{down} of relationship (3) that were estimated using the least squares method follow an asymptotic multi-variate normal distribution, which allows to test for the hypothesis H_0: $\varrho_1^{up}=\varrho_1^{down}=0$ using a typical F test (e.g. Wald test), to check whether the adjustment is symmetric. In reality, the true value of the threshold is unknown and must be estimated. To determine the threshold value $\hat{\tau}$ where a statistically significant asymmetry is likely to emerge, Enders and Siklos (2001) propose the application of a search method that is widely known in the literature as "Chan's approach". Chan (1993) showed that under certain typical circumstances the OLS method yields a super-consistent estimator of the true threshold value. More specifically, according to this grid search procedure, the vector values of the residuals $\hat{u}_t$ resulting from the cointegration relationship (1) are identified as potential threshold values. The residual vector $\{\hat{u}_t\}$ is sorted in ascending order as follows, $\left\{\hat{u}_1^\tau < \hat{u}_2^\tau < \ldots < \hat{u}_T^\tau\right\}$ where T denotes the number of usable observations. From the distribution of threshold values sorted in ascending order, the highest and the lowest 15% of values (outliers) are discarded and the remaining (central) 70% of residual values are considered possible thresholds. For each one of those possible thresholds belonging to 70% of the values, relationship (3) is successively estimated using the specification of the importance indicator in relationship (4). The value of the residual $\hat{u}_t$ that was estimated using relationship (3) and which yields the lowest residual sum of squares (RSS) is deemed to be the appropriate threshold $\hat{\tau}$.

If a TAR cointegration is verified, an error correction model with a long-run asymmetric equilibrium can then be estimated, which in its generic form is described by the following relationship (5):

$$\Delta r_t^r = \mu_0 + \sum_{i=1}^{k_1} \beta_{1,i} \Delta r_{t-i}^r + \sum_{i=0}^{k_2} \beta_{2,i} \Delta r_{t-i}^w \cdot + a_1^{up} \hat{u}_{t-1}^{up} + a_2^{down} \hat{u}_{t-1}^{down} + e_t \tag{5}$$

where $\hat{u}_{t-1}^{up} = I_t \hat{u}_{t-1}$ and $\hat{u}_{t-1}^{down} = (1-I_t) \hat{u}_{t-1}$.

On the basis of the specification of relationship (5), the adjustment speed of retail prices is a_1^{up}, if $\hat{u}_{t-1}$ is above the long-run equilibrium, and a_2^{down}, if $\hat{u}_{t-1}$ is below the long-run equilibrium. To test for long-run symmetry, i.e. whether the coefficients a_1^{up} and a_2^{down} are statistically equal, the H_0: $a_1^{up}=a_2^{down}$ hypothesis of model (5) is controlled. The asymmetric mean adjustment lag (MAL) is calculated using the following formulas (Hendry 1995):

$$\text{MAL}^{up} = (\sum_i^{k_1} \beta_{2i} - 1) / \alpha_1^{up} \tag{6}$$

and

$$\text{MAL}^{down} = (\sum_i^{k_2} \beta_{2i} - 1) / \alpha_2^{down} \tag{7}$$

Relationship (6) measures the mean adjustment lag MAL^{up} in months during which food retail prices are above their equilibrium level. Respectively, relationship (7) measures the mean adjustment lag MAL^{down} in months during which food retail prices are below their equilibrium level.

7 EMPIRICAL RESULTS

Prior to the cointegration tests, the presence of unit root *I(1)* in food price variables is examined performing the Augmented Dickey-Fuller (ADF) (1979; 1981), Phillips-Perron (PP) (1988) and DF-GLS (Elliot et al. 1996) tests. The series are log-transformed for all variables. Unit root tests show that, while the unit root hypothesis cannot be rejected as to the levels of the series, their first differences were found to be stationary. Thus, the series of both food retail prices and producer prices are integrated of order one, *I(1)*.[5]

The estimation of relationship (1) is carried out using the FMOLS (Fully Modified OLS) method.[6] The cointegration tests reject the null hypothesis of no cointegration (see Table 5).

If the reviewed food industries did operate under perfect competition, the degree of adjustment should be expected to be 100%, i.e. elasticity would be $\beta_0=1$. This would mean that producer price movements would be fully passed through to food retail prices. The estimations of the long-run relationships suggest that the values of the coefficient β_0 across the reviewed food categories are statistically less than unity, with the exception of the dairy products category where the coefficient is inferred to be equal to unity. The finding that the estimated values of the degree of adjustment are less than unity, i.e. indicating that imperfect competition conditions prevail in the reviewed industries, does not necessarily imply a long-run asymmetric adjustment of retail prices as well.

The results of the TAR model estimation, which are presented in Table 6, suggest that the consistent value of the threshold $\hat{\tau}$ for each individual food category is different. A positive threshold value implies that the current food retail price r_t^r is above the long-run equilibrium, as determined by the cointegration relationship (1), which in turn indicates an environment of expanded profit margins. By contrast, a negative threshold value points to an environment of compressed profit margins. In other words, the critical point beyond which the residuals respond and change their speed

5 The results of unit root tests on individual series are not presented in this article for the sake of brevity, but are available upon request.
6 The FMOLS method yields super-consistent estimators and the ensuing residuals denote stochastic (unpredictable) shocks. The FMOLS method is preferable to the OLS method because it increases effectiveness and reduces small sample bias compared with the OLS estimate (Phillips and Hansen 1990). Moreover, the FMOLS method yields an asymptotic unbiased and effective estimator, thereby allowing reliable statistical testing.

Table 5 Estimation of the long-run relationship for food price equations and cointegration testing

Long-run relationships	Price equations									
	Cereals		Meat		Dairy products		Fruit		Vegatables	
Adjusted sample size	January 2002 - June 2016									
Total observations	174		174		150		138		126	
α_0	*1.098***	*(5.123)*	*0.386***	*(2.681)*	*0.159*	*(1.288)*	*1.647***	*(3.570)*	*2.251***	*(4.079)*
β_1	*0.734***	*(13.971)*	*0.917***	*(28.800)*	*0.965***	*(28.936)*	*0.650***	*(19.328)*	*0.505***	*(4.158)*
Adjusted R^2	0.987		0.968		0.985		0.633		0.285	
Standard error of regression	0.012		0.012		0.006		0.043		0.041	
Engle-Granger cointergation tests										
Null hypothesis of cointegration H_0										
Engle-Granger tau-statistic	-4.607**	[0.018]	-5.604**	[0.000]	-4.956**	[0.001]	-5.300**	[0.000]	-4.878**	[0.000]
Engle-Granger z-statistic	-37.657**	[0.015]	-44.948**	[0.000]	-43.313**	[0.000]	-44.365**	[0.000]	-40.278**	[0.000]
Conclusion: the series are cointegrated										

Source: Author's econometric estimations.
Notes: t statistics in parentheses and P-values in square brackets. ** denotes statistical significance at 5%. Critical values for the cointegration test are provided by MacKinnon (1996). The lags of the Engle-Granger cointegration tests were obtained using the Schwarz information criterion (SBC).

of adjustment to the long-run equilibrium is different for each food category, yet very close to zero.

According to the estimations of the TAR model, the residual convergence coefficients of the long-run relationship ϱ_1^{up} and ϱ_2^{down} are statistically significant on the basis of the critical values of t-Max* statistics and have the expected negative sign across all food categories. It should be noted that in all food categories the absolute value of the coefficient ϱ_2^{down} is higher than that of ϱ_1^{up}, implying a faster adjustment of deviations when residuals are below the consistent threshold $\hat{\tau}$ (profit margin compression) than when they are above it (profit margin expansion).

The test for the hypothesis of a symmetric adjustment in deviations towards their mean $\varrho_1^{up}=\varrho_2^{down}$ for all food categories, except for meat, shows that residuals symmetrically adjust to their attractor.

The existence of threshold autoregressive cointegration allows the estimation of the ECM with TAR cointegration, which in its generic form is described by relationship (5). In this way, the nature of the long-run adjustment is explored, i.e. whether food retail prices exhibit an asymmetric response to their long-run equilibrium. The results of these estimations are presented in Table 7.

In all food categories, the estimations of the coefficients measuring asymmetric adjustment to the long-run equilibrium appear to vary.

The values of the coefficients measuring the speed of adjustment when retail prices are below their equilibrium (compressed profit margins) are higher in all food categories than what is the case when retail prices are above their equilibrium (expanded profit margins). Against this backdrop, it can be assumed that, when profit margins are compressed, retailers exhibit a higher tendency to raise food prices after a rise in producer prices. Conversely, when profit margins are expanded, retailers are less inclined to lower prices after a drop in producer prices.

Table 6 TAR tests for threshold autoregressive cointegration

Adjusted sample size			Feb. 2002-June 2016 (n=173)			Feb. 2002-June 2016 (n=173)			Feb. 2002-June 2016 (n=149)			Feb. 2002-June 2016 (n=137)			Feb. 2002-June 2016 (n=128)		
Estimation method: OLS			Cereals			Meat			Dairy products			Fruit			Vegetables		
Threshold			$\tau = -0.012288$			$\tau = 0.008340$			$\tau = 0.006592$			$\tau = -0.013076$			$\tau = -0.023808$		
					P-value			P-value			P-value			P-value			P-value
	Enders-Siklos, t-MAX*	ρ_1^{up}	-0.148**	(-2.496)	[0.013]	-0.166**	(-2.176)	[0.030]	-0.231**	(-2.120)	[0.035]	-0.292**	(-4.299)	[0.000]	-0.255**	(-3.101)	[0.002]
	Enders-Siklos, t-MAX*	ρ_2^{down}	-0.238**	(-3.657)	[0.000]	-0.402**	(-5.678)	[0.000]	-0.341**	(-4.638)	[0.000]	-0.351**	(-3.162)	[0.001]	-0.397**	(-3.776)	[0.000]
Testing for TAR cointergation	Statistic Enders-Siklos Φ^*	H_0: $\rho_1^{up}=\rho_2^{down}=0$		9.805**	[0.000]		17.874**	[0.000]		13.008**	[0.000]		14.244**	[0.000]		11.942**	[0.000]
Testing for symmetry	Wald F-test	H_0: $\rho_1^{up}=\rho_2^{down}$		1.048**	[0.307]		5.441	[0.020]		0.693**	[0.406]		0.227**	[0.634]		1.122**	[0.291]
Conclusions			Cointegration Symmetric adjustment of residuals			Cointegration Asymmetric adjustment of residuals			Cointegration Symmetric adjustment of residuals			Cointegration Symmetric adjustment of residuals			Cointegration Symmetric adjustment of residuals		

Source: Author's econometric estimations.
Notes: t statistics in parentheses and P-values in square brackets. ** denotes statistical significance at 5%. The maximum lag length is obtained using the AIC (1973) and SBC (1978) information criteria. Critical values for TAR cointegration testing are provided by Enders and Siklos (2001) and Wane (2004). The critical values of Φ^* statistics (with threshold) without lags are 6.95 (with 100 observations) and 6.93 (with 250 observations) at a 5% level of significance. The critical values of t-Max* statistics without lags are -1.85 (with 100 observations) and -1.84 (with 250 observations) at a 5% level of significance. To test the hypothesis of symmetry, F values are used, provided that there is TAR cointegration.

Table 7 Error correction model (ECM) with TAR cointegration

Estimation method: OLS (HAC standard errors & covariance)		ECM with TAR cointegration									
Adjusted sample size (n=number of observations)		Feb. 2002-June 2016 (n=173)		Feb. 2002-June 2016 (n=173)		Feb. 2002-June 2016 (n=149)		Feb. 2002-June 2016 (n=137)		Feb. 2002-June 2016 (n=128)	
Error correction models	Coefficients	Cereals	p-value	Meat	p-value	Dairy products	p-value	Fruit	p-value	Vegetables	p-value
Fixed term	μ_0	0.000 (0.800)	[0.424]	0.000 (1.448)	[0.149]	0.001 (0.291)	[0.771]	-0.004 (-1.182)	[0.239]	-0.001 (-0.245)	[0.806]
$\Delta(r^r)_{t-1}$	$\beta_{1.1}$	0.266** (7.615)	[0.000]			0.224** (2.439)	[0.016]			0.084 (1.043)	[0.298]
$\Delta(r^r)_{t-2}$	$\beta_{1.2}$					0.098** (1.414)	[0.159]				
$\Delta(r^r)_{t-3}$	$\beta_{1.3}$										
$\Delta(r^r)_{t-5}$	$\beta_{1.5}$					0.137** (1.906)	[0.058]	0.215** (2.772)	[0.006]		
$\Delta(r^w)_t$	$\beta_{2.0}$					0.328** (2.264)	[0.025]	0.651* (1.881)	[0.062]		
$\Delta(r^w)_{t-1}$	$\beta_{2.1}$	0.153** (3.271)	[0.001]	0.203** (5.253)	[0.000]						
$\Delta(r^w)_{t-2}$	$\beta_{2.2}$			0.035** (1.781)	[0.076]					0.600** (1.993)	[0.048]
$\hat{u}_{t-1}^{up}$	α_1^{up}	-0.017 (-0.546)	[0.585]	-0.047 (-1.058)	[0.291]	-0.151** (-2.111)	[0.036]	-0.268** (-4.175)	[0.000]	-0.225** (-2.207)	[0.029]
$\hat{u}_{t-1}^{down}$	α_2^{down}	-0.143** (-4.392)	[0.004]	-0.133** (-2.808)	[0.005]	-0.260** (-3.025)	[0.003]	-0.320** (-2.902)	[0.004]	-0.424** (-3.362)	[0.001]
Wald F-statistic	$\alpha_1^{up}=\alpha_2^{down}$	7.206 df(1.165)	[0.008]	5.472 df(1.164)	[0.026]	0.934** df(1.139)	[0.335]	0.125** df(1.130)	[0.723]	1.136** df(1.118)	[0.288]
Main conclusion		**Rejection of symmetry**		**Rejection of symmetry**		**Verification of symmetry**		**Verification of symmetry**		**Verification of symmetry**	
Adjusted R^2		0.810		0.360		0.421		0.330		0.244	
DW statistic		2.16		2.000		2.010		2.075		1.976	
LM(6), Breusch-Godfrey Serial Correlation Test		1.966 Prob. F-stat (6.159)	[0.073]	0.928 Prob. F-stat (6.158)	[0.476]	1.002 Prob. F-stat (6.133)	[0.426]	1.220 Prob. F-stat (6.124)	[0.300]	0.615 Prob. F-stat (6.112)	[0.717]

Source: Author's econometric estimations.
Notes: t statistics in parentheses and P-values in square brackets. ** and * denote statistical significance at 5% and 10%, respectively. The maximum lag length was obtained using the AIC (1973) and SBC (1978) information criteria.

Table 8 Mean adjustment lag

Mean adjustment lag		Cereals	Meat	Dairy products	Fruit	Vegatables
MAL^{up}	Months	49.8	16.2	4.5	1.3	1.8
MAL^{down}	Months	5.9	5.7	2.6	1.1	0.9

Source: Author's econometric estimations.

With regard to the cereals category, the coefficient measuring adjustment when the retail prices of cereal products are above their equilibrium (thereby suggesting a decreasing trend) takes the value ϱ_1^{up}= -0.017, but is statistically insignificant. In other words, it is presumed to be equal to zero, even though it bears the expected sign. On the basis of the available sample of observations, -0.017 is considered to be the most likely value for this parameter. If the value of the term ϱ_1^{up} is not in reality other than zero, this implies that expanded profit margins are not corrected and that the values r_t^w and r_t^r are not cointegrated. Nevertheless, the tests performed suggest a cointegration between both series. Studies with similar findings, such as von Cramon-Taubadel (1998) and Reziti (2014), rested upon the same assumption, i.e. that the estimated value ϱ_1^{up} = -0.017, although it is assumed to be zero, is the most likely value for this parameter on the basis of the specific sample of observations. In other words, it can be deduced that either expanded profit margins are not corrected or that the adjustment is very slow. The coefficient measuring adjustment when retail prices are below their equilibrium (thereby suggesting a rising trend) takes the value ϱ_2^{down}= -0.143 and is statistically significant. That is, when deviations from long-run equilibrium are negative, the retail prices of cereal products move upwards to converge towards their equilibrium, whereas positive deviations from long-run equilibrium do not affect the retail prices of cereals, which remain sticky. A typical Wald test (F(1.165) = 7.206, P-value = 0.008) rejects the equality of the adjustment coefficients $\varrho_1^{up}=\varrho_2^{down}$, at a 5% level of statistical significance. Asymmetric speeds of adjustment to long-run equilibrium also determine the mean adjustment lag (MAL). As shown in Table 8, the mean adjustment lag, when the retail prices of cereals are above equilibrium, is MAL^{up} = 49.8 months, implying a marginally zero adjustment and therefore the adjustment of the deviation from the original shock is very slow. By contrast, when retail prices are below equilibrium, MAL^{down} = 5.9 months. The empirical findings lead to the conclusion that there is a positive asymmetric adjustment of the retail prices of cereals to their long-run equilibrium, at the expense of consumers.

For the meat category, the findings are similar with those for the cereals category. That is, the coefficient measuring adjustment when meat retail prices are above equilibrium takes the value ϱ_1^{up} = -0.047 and is statistically insignificant, whereas the respective coefficient measuring adjustment when prices are below equilibrium takes the value ϱ_2^{down}= -0.133 and is statistically significant. MAL^{up} is 16.2 months (marginally zero adjustment) and therefore the deviation from the original shock corrects at a very slow speed, taking too long to fully adjust. By contrast, when retail prices are below equilibrium, MAL^{down}= 5.7 months.

The empirical findings provide evidence that there is an asymmetric adjustment of meat retail prices to their long-run equilibrium.

In the cases of dairy products, fruit, and vegetables, the coefficients measuring the speeds of adjustment are statistically significant under both premises and, although they are seemingly divergent, the tests performed support the equality of coefficients.

More specifically, with respect to the category of dairy products and eggs, there appears to be a divergence of $\varrho_1^{up} = |-0.151| < \varrho_2^{down} = |-0.260|$ between the coefficients measuring the speed of long-run adjustment, probably suggesting asymmetries in the speed of price adjustment. The adjustment coefficients are statistically significant and a typical Wald test (F(1.139) = 0.934, P-value = 0.335) for $\varrho_1^{up} = \varrho_2^{down}$ showed that the equality hypothesis cannot be rejected at a 5% level of statistical significance. When the retail prices of dairy products are above equilibrium, MAL^{up} is 4.5 months, while MAL^{down} is 2.6 months when prices are below equilibrium (see Table 8).

For the fruit category, the coefficients measuring the speed of long-run adjustment also seem to be divergent, as $\varrho_1^{up} = |-0.268| <= \varrho_2^{down} |-0.320|$.

The adjustment coefficients are statistically significant and a Wald test (F(1.130) = 0.125, P-value = 0.723) for $\varrho_1^{up} = \varrho_2^{down}$ showed that the equality hypothesis cannot be discarded at a 5% level of significance. When fruit retail prices are above equilibrium, MAL^{up} is 1.3 month, while MAL^{down} is 1.1 month when prices are below equilibrium.

For the category of vegetables, price adjustment is broadly the same as for the category of fruit. The coefficients measuring the speed of long-run adjustment are seen to be different, since $\varrho_1^{up} = |-0.225| <= \varrho_2^{down} |-0.424|$. The adjustment coefficients are statistically significant and a Wald test (F(1.118) = 1.136, P-value = 0.288) showed that the hypothesis of coefficients equality $\varrho_1^{up} = \varrho_2^{down}$ cannot be rejected at a 5% level of significance. When vegetable retail prices are above equilibrium, MAL^{up} is 1.8 month, while MAL^{down} is 0.9 month when prices are below equilibrium.

8 CONCLUSIONS

Asymmetric price transmission is of particular relevance when it refers to food commodities, as it has an incommensurate impact on low-income consumers, who tend to spend a higher share of their income on basic food items.

The main objective of this paper was to empirically investigate the existence of asymmetries in the speed of retail price adjustment to long-run equilibrium, focusing on a group of basic food categories. The food categories examined were cereals, meat and meat products, dairy products, fruit, and vegetables.

The key factor behind food price adjustment asymmetries that is cited in the literature is the monopoly power of firms or the degree of concentration in a sector, which enables firms to engage in a tacit collusion (cartels) to systematically maintain high profit margins. Measuring concentration in the sectors reviewed on the basis of the Herfindahl-Hirschman Index (HHI) for 2014 attested to high market concentration in all sectors, and particularly so in those of cereals and meat. Other reasons mentioned in the literature are the perishability of some food items, re-pricing costs, government-mandated minimum prices, and information asymmetries.

To empirically investigate whether there is an asymmetric adjustment of food retail prices relative to producer price movements, an error correction model (ECM) with threshold autoregressive (TAR) cointegration was employed. The empirical findings support a TAR cointegration for each of the reviewed food categories.

The empirical investigation into a possible asymmetric behaviour of retail prices using the estimations of the ECM-TAR models provides strong evidence that there are in fact such asymmetries in the categories of cereals and meat. The empirical findings of retail price asymmetric adjustment in the categories of cereals and meat may be associated with the oligopolistic structure of those sectors, where firms tend to keep prices relatively unchanged (rigid) in times of declining producer prices, while rushing to raise them in times of rising producer prices. Price adjustment asymmetries for the categories of cereals-flour products and meat are consistent with the high concentration ratio of those sectors on the basis of the Herfindahl-Hirschman Index (HHI).

Conversely, turning to the categories of dairy products, fresh fruit, and vegetables, the divergences in the coefficients of long-run adjustment cannot justify a rejection of symmetrical adjustment. The recommendations contained in the OECD Toolkit (OECD 2014a; 2014b; 2016), which were to a great extent adopted by the Greek government with Laws 4254/2014, 4336/2015 and 4441/2016, contributed to a more effective functioning of those specific markets.[7]

7 The most important measure in this direction was to extend the shelf life of fresh milk as well as to broaden the scope of its definition. Under the previous framework, the maximum lifetime of fresh milk was limited to 5 days. This restriction hampered imports, undermined competition and raised return costs. Under Law 4336/2015 the definition “pasteurised milk” was instead adopted and the five-day restriction was abolished, but both the pasteurisation and the expiry dates must be clearly indicated.

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