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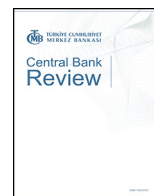
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Mean-reversion and structural change in European food prices

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ABSTRACT

Mean-reversion in unprocessed food prices and beef prices towards the long-run trend is examined for twenty-two European countries, using linear and nonlinear unit root tests. As the argument goes, food prices might display short-term deviations from their long-run values due to disturbances such as changes in climate or speculation; yet, once the impact of these short-lived shocks fade away, the prices convert to the long-run equilibrium level determined by fundamentals. The nonlinear smooth transition framework suggest that the speed of this adjustment might depend on the size and sign of the deviation of prices from their long-run values. The results carry important policy implications regarding the benefits of short-term demand management policies along with structural policies.

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1. Introduction

Global food prices revealed sizable fluctuations in the last two decades (Fig. 1). This raises concerns for policymakers since unstable food prices have diverse impacts on the well-being of economic agents. On the one hand, higher food prices could endanger the access to food by low-income households in many underdeveloped countries, albeit benefiting small-scale farmers. Lower food prices, on the other hand, could result in insufficient investment in agricultural sector -hence forming a threat to food security- as well as indicate lower returns for small farmers. Consequently, frequent changes in prices would lead to sub-optimal investment and consumption decisions in agriculture, with possible repercussions at both national and global levels².

This paper focuses on the time-series characteristics of two selected food price indices: unprocessed food prices and beef prices

for twenty-two European countries. We focus on whether fluctuations in these series have long-lasting effects or would the prices go back to their long-term equilibrium value (or trend) after a certain amount of time. In more technical terms, we ask whether there exists a mean-reverting long-run equilibrium in these series, using linear and nonlinear unit root tests. The particular focus of our study is on the sign and size asymmetries in mean-reversion as well as possible structural breaks (shifts in the intercept or the trend) in the series while testing for unit roots, using two types of nonlinear models: Exponential Smooth Transition Autoregressive (ESTAR) and Asymmetric Exponential Smooth Transition Autoregressive (AESTAR) models.

The presence of a long-run equilibrium in food prices indicates the dominance of *fundamentals*, such as growth in population and income or changes in tastes and technology, in driving the behaviour of food prices. However, food prices could also be influenced by temporary factors such as changes in climate, price speculation or hoarding behaviour against temporary shocks⁴. As the argument

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¹ The views and opinions presented in this study belong to the author and do not necessarily represent those of the CBRT or their staff. We would like to thank to the anonymous referees for their valuable comments.

² Instability in food prices is not only a major concern for governments that would aim for food security, but also a global problem necessitating coordinated action around the world. See FAO (2012a) and FAO (2012b) for collaborative policy reports.

³ The index is defined as "... a measure of the monthly change in international prices of a basket of food commodities. It consists of the average of five commodity group price indices (meat, dairy, cereals, oil and sugar, weighted with the average export shares of each of the groups for 2002–2004".

⁴ See Timmer (2012) and Robles et al. (2010) for the impact of speculation on food prices such as financial speculation in futures and options markets or commodity index funds.

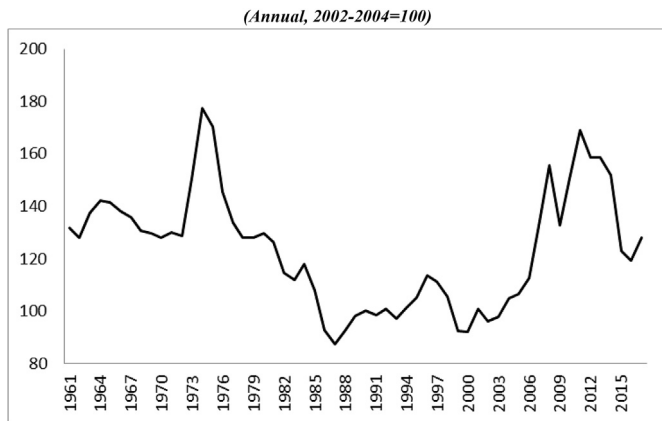


Fig. 1. World real food price index.³

goes, once the impact of these short-lived shocks fade away, the food prices convert to their equilibrium level in the long-run. Absence of such convergence means that short-term shocks to the prices are persistent and the prices would settle to a new equilibrium level.

The characteristic of the correction behaviour is important from the policymaking perspective since the policy responses to food price fluctuations would be different against temporary or permanent shocks. If the policymaker believes that the short-lived shocks have only temporary effects on prices then she might make use of demand-management policies that would provide short-term relief, especially for the most vulnerable parts of the sector. Some examples of such policies could be buffer stock policies, cash transfers or other safety net measures. Otherwise, if the fluctuations are believed to be persistent, then the policies should include more structural components, rather than a short-run demand-management framework, such as infrastructure investment, education or R&D incentives on agriculture. Obviously, most of the times the policymakers prefer a synthesis of these short-run and long-run policies but the degree of urgency for short-run policies depend on the policymakers view on the duration of the deviation from the trend.

The first series of our interest, unprocessed food price index consists of four sub-components of food items: meat, fish, fruit and vegetables. These components might reveal deviations from their equilibrium (or trend) value due to two intrinsic characteristics of them. First, their production is vulnerable against climate conditions. Second, they have relatively lower price elasticity due to their prominence in the consumer basket; hence changes in their input prices or short-term supply problems could easily be reflected in their final prices (Öğünç, 2010).⁵ The second series of our interest is the beef prices which has a significant impact on the unprocessed food prices. ECB (2013) states that meat prices consists of more than half of the contributions to the changes in the unprocessed food price index. Last decade witnessed increasing demand for meat in developing areas of the world which led to substantial increases in livestock production. Moreover, livestock production still constitutes a considerable part of rural employment around the world, catering food security for households.

Fig. 2 and Fig. 3 depicts respectively the monthly unprocessed food price and beef price series for 22 European countries taken from Eurostat database. The unprocessed food price series covers

January 1996 to June 2016 period while the beef price series covers January 2005 to March 2016 period. All series are seasonally adjusted using Census X-12 method. The countries are selected to ensure that their data coverage allows comparison among them.

A first general look at Fig. 2 suggests a steady upwards pattern in unprocessed food prices in most countries. Moreover, there are three periods that shows signs of sudden increases: early 2000s, 2007–2008 and post-2010 period.⁶ Fig. 3 suggest a change in trend in beef prices for most countries around 2010. Hence, in a technical manner, it is important to take into account possible structural breaks and trend-shifts while testing for mean reversion, as will be detailed in the third section.

Our conjecture for the mean reversion taking a nonlinear form could stem from a couple of reasons. First one could be the cyclical behaviour in livestock production. Holt and Craig (2006) argue that farmers might give overnight decisions to slaughter their animals in response to changes in expectations of the market price for corn.⁷ Yet, once the price of corn goes up, it takes considerable time to reproduce livestock, which would result in nonlinearities in the hog as well as corn prices. They show that the US hog-corn cycle could be characterized by a nonlinear form, in particular a time-varying STAR model.⁸ Later on, the smooth transition framework is also employed to examine the nonlinear behaviour of selected primary commodity prices by Balagtas and Holt (2009); the US ethanol, corn, oil, and gasoline prices by Serra et al. (2011); the EU pork market by Emmanouilides and Fousekis (2012); US soybean-to-corn price ratio by Ubilava (2012); and agricultural future prices by Beckmann and Czudaj (2014).⁹

There are alternative motivations for asymmetries in beef prices, some of which could be generalized to the food prices, or commodity prices in general. One of them is the positive storage. An important driver of the food prices is the stock level of the product as a percentage of the total demand (Baffes and Dennis, 2013). By definition, storage should always be greater than or equal to zero (negative levels are not possible!) which brings in an inherent nonlinearity to the price process (Deaton and Laroque, 1995).

The nonlinear mean-reverting behaviour in prices is mainly analysed by unit root tests in the literature. This paper employs two

⁶ ECB (2013) report argues that the surge in unprocessed food prices in the latest period is mainly due to contributions from fruit and vegetable prices while FAO (2009) report blames particularly surging energy prices and low harvest as a result of adverse weather conditions for the surge in 2007–2008 period. On the other hand, ECB (2013) argues that the sharp increase at the beginning of the century is mainly due to an upward pressure in beef prices as a result of health concerns associated with animal diseases that spread towards other substitute meat products. Baffes and Dennis (2013) shows that the main drivers of the food prices are the stocks of the product as a percentage of the total demand, crude oil prices and exchange rates over a fifty year period. Recently, Baffes and Etienne (2016) argue that income growth in developing countries has a negative impact on real food commodity prices in line with Engle's law. Also see Dillon and Barrett (2015) for the impact of oil prices on food prices.

⁷ OECD-FAO Agricultural Outlook (2015) states that "Nominal prices for beef will remain high in the short term, as herds are being rebuilt in several meat producing countries. Over the medium term, prices will ease due to rising production levels.

⁸ This story is somewhat similar to the hysteresis idea in some economic aggregates. A line of the recent literature states that short-run unemployment shocks might have protracted effects due to rigidities in labour markets (Blanchard and Summers, 1986). For instance, a negative stigma on the unemployed workers could lower the possibility of them to find new jobs (Blanchard and Diamond, 1994). Hence, shocks that are regarded as cyclical at first could lead to structural problems in the long-run. In the example above, a sudden change in meat (or corn) prices might lead the farmers to slaughter their animals and quit production at all. Hence, a shock that shows short-term characteristics initially might turn up having long-run structural effects which could be an example of hysteresis in meat prices.

⁹ The recent literature also includes studies using regime-switching models to analyse asymmetries in food price markets such as Taiwanese wholesale hog prices by Chen and Lee (2008) and imported rice prices in Asian countries by Valera and Lee (2016).

⁵ For these reasons, they are usually the hardest components in the inflation basket to forecast as discussed by Bruneau et al. (2013).

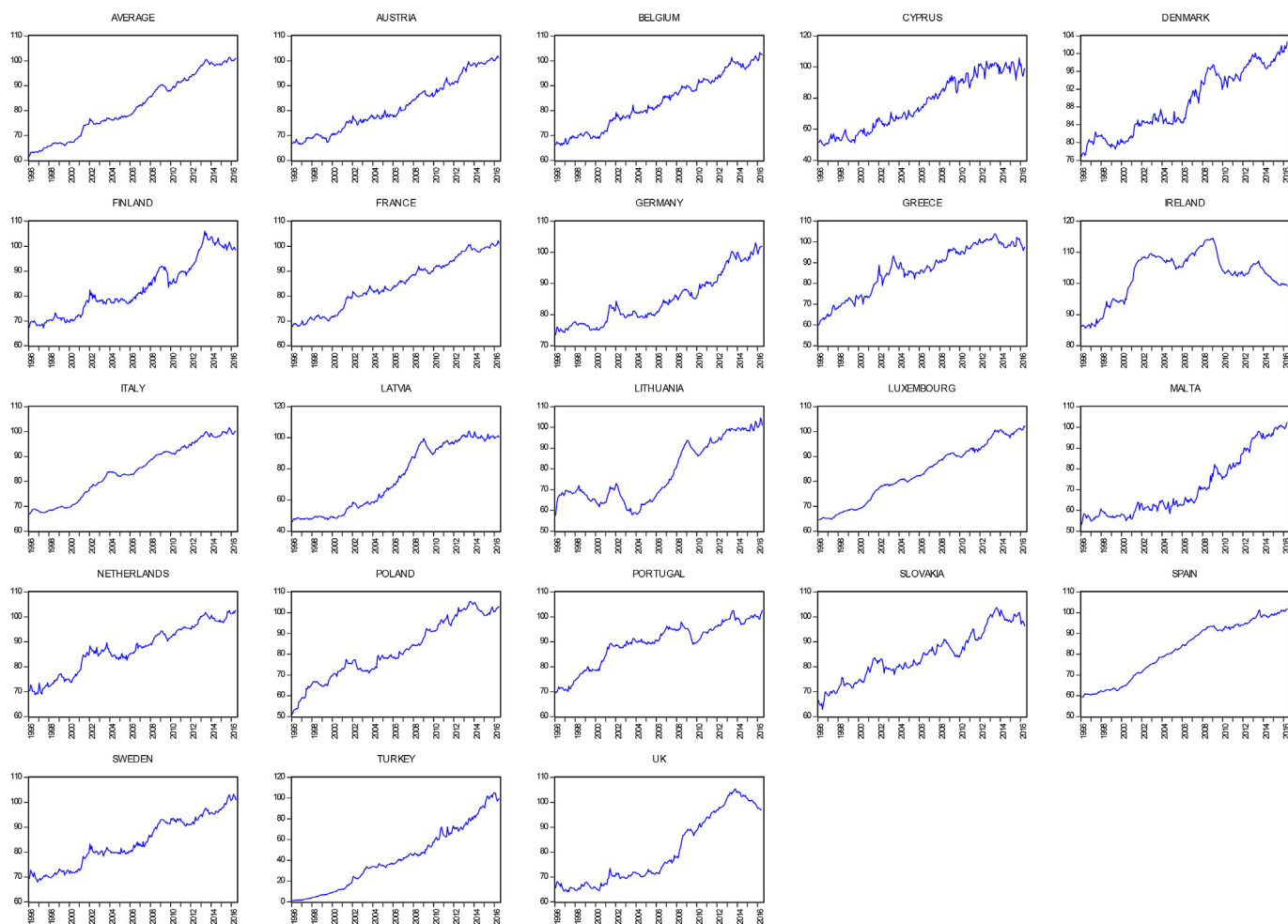


Fig. 2. Unprocessed Food Price Indices. (Seasonally adjusted, 2015 = 100, the first graph is the average for all countries).

different nonlinear smooth transition models in unit root testing. The first one is the ESTAR test suggested by [Kapetanios et al. \(2003\)](#), taking cognisance of asymmetric behaviour around the long-run mean which depends on the *size* of the deviations for the two series of our interest. As the argument goes, there is no correction within a threshold band around the equilibrium level (hence the system shows unit root behaviour in this inaction band) while above or below this band there is adjustment towards the long-run value (hence the system displays stationary behaviour). In particular, the model assumes that the further the prices from the long-run equilibrium, the faster the adjustment towards the long-run level.

An important characteristic of the aforementioned ESTAR framework is its conjecture of a *symmetric* threshold band. It is assumed that the market adjustment against a positive deviation of prices from the equilibrium level has the same strength with a negative deviation. However, the inherent nonlinearity in the series could not only result from size of the deviation from trend but could also be influenced by the *sign* of the deviation, as suggested by the AESTAR test of [Sollis \(2009\)](#). As the argument goes, positive deviations of food prices in general (and beef prices in our example) from the equilibrium level could be more persistent compared to negative deviations, or vice versa. This sign asymmetry could be motivated with the conjecture that the firms could be more reluctant to lower prices in response to lower input prices, while

they are quick to raise prices against an increase in input prices ([Peltzman, 2000](#)). Also, [Thirlwall and Bergevin \(1985\)](#) argue that the fall in the primary commodity prices during cyclical downturns could be greater than the rise during cyclical upturns as an explanation of Prebisch-Singer hypothesis,^{10, 11}. To the best of our knowledge, our paper is the first study which applies AESTAR type of unit root test for the food prices.

Another contribution of our paper is our consideration of nonlinearity in the presence of structural breaks. As discussed before, certain shocks might have permanent effects on the price process. Previous literature includes many papers that focus on the structural break in linear unit root tests and multiple equilibria

¹⁰ [Prebisch \(1950\)](#) and [Singer \(1950\)](#) argued that primary commodity prices display a long-term declining trend relative to manufactured good prices (Prebisch-Singer Hypothesis). Previous literature documents numerous studies that explored the existence of such behaviour in primary commodities.

¹¹ This idea is also in line with the recent literature documenting different strength of correction behaviour for positive or negative shocks to inflation. [Tsong and Lee \(2011\)](#) shows that mean-reversion in inflation is stronger for negative shocks relative to the positive ones for 12 OECD countries. [Koenker and Xiao \(2004, 2006\)](#) documents asymmetries in the mean-reversion in inflation process depending on the size and the sign of the shocks, using quantile-regression methods. [Akdoğan \(2015\)](#) employs an ESTAR and AESTAR framework to analyse asymmetric behaviour of inflation around the inflation-target for inflation-targeting countries. Also see [Tillmann and Wolters \(2012\)](#) and [Manzan and Zerom \(2014\)](#) for nonlinearities in persistence of shocks to US inflation.

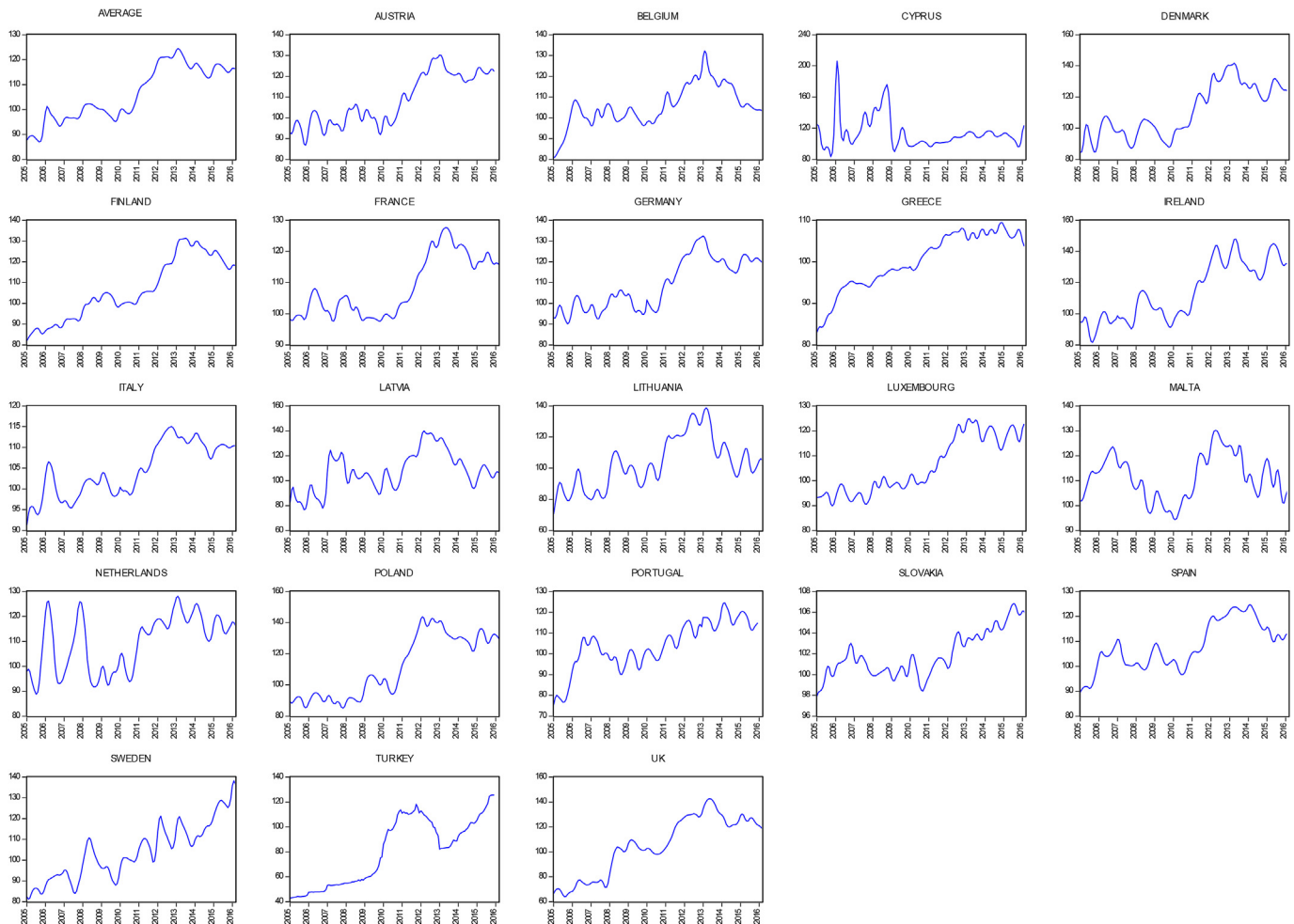


Fig. 3. Beef Price Indices.
(Seasonally adjusted, 2010 = 100, the first graph is the average for all countries).

while testing for trends and stationarity in commodity prices.¹² In line with these studies, we take into account possible structural tests in the linear case using Lee and Strazicich (2003, 2004). However, we further take into account the structural breaks in nonlinear unit root testing for food prices, using Christopoulos and Leon-Ledesma (2010) test. Again, to the best of our knowledge, our analysis is unique in applying nonlinear unit root test considering structural breaks (Christopoulos and Leon-Ledesma, 2010) for food prices.

Overall, our results imply the importance of considering possible asymmetric behaviour in mean-reversion in unprocessed food and beef prices. Furthermore, our results suggest that structural breaks should be taken into account in both linear and nonlinear tests. These results have implications for the policy. Existence of mean-reversion could imply that short-term deviations are temporary and demand-management policies could help to overcome the impact of the short-lived shocks. However, absence of mean-reversion would call for stronger, and possibly structural measures, as will further be discussed in our analysis on country examples.

¹² This literature starts with Sapsford (1985) and Cuddington and Urzua (1989) papers. For an extensive review of the following literature see Baffes and Etienne (2016).

The remainder of the paper is organized as follows. Second section describes the econometric methodology. Third section documents the results of the linear and nonlinear unit root tests. The fourth section concludes.

2. Methodology

Our exercise includes linear and nonlinear unit root tests, accounting for possible structural breaks for each series. The three linear unit root tests that are employed are extensively used in the literature: Augmented Dickey Fuller (ADF), Elliot-Rottenberg-Stock (ERS) and Phillips-Perron tests. However, the literature widely reports the low power of conventional unit root tests in rejecting the null of nonstationarity, under the presence of nonlinearities in time series. This issue is addressed by means of nonlinear unit root tests in numerous studies. In this direction, we employ nonlinear unit root tests of Kapetanios et al. (2003) and Sollis (2009). Moreover, we take into account possible structural break problems in unit roots under linear case by Lee and Strazicich (2003) and Strazicich et al. (2004) papers; and under the presence of nonlinearities with while we focus on nonlinearities and structural breaks Christopoulos and Leon-Ledesma (2010). Accordingly, the structural break problem in unit root testing is described in the first subsection below. The second subsection describes our nonlinear unit root testing methodology.

2.1. Unit root with structural breaks

The literature suggests caution on the presence of a structural break in unit root testing. Perron (1989) suggests that unit-root behaviour might be the result of disregarding a structural break in the deterministic trend. His modified Dickey-Fuller test including dummies assumes an exogenous break date. However, choosing an incorrect break date might lead to distortions in size and loss in power (Hecq and Urbain, 1993). Hence, subsequent literature incorporates unknown breaks which are endogenously determined through data. In line with this framework, Zivot and Andrews (1992) employs a single break that is estimated by minimizing the Dickey-Fuller t-statistic for the unit root null. Later on, Lumsdaine and Papell (1997) extend this test for a two-break alternative.

The consequent literature also underlines the impact of the specification of breaks under the null and alternative hypothesis on the power of tests. Among the aforementioned tests, Perron (1989) considers breaks for both under the null of and alternative hypothesis while Zivot and Andrews (1992) and Lumsdaine and Papell (1997) tests assume for structural breaks only under the alternative specification. However, Kim and Perron (2009) shows that the power of this latter group of tests which neglect the breaks under the null would be lower since these tests are not invariant to the break parameters. In line with this thought, Lee and Strazicich (2003) develop an endogenous two-break Lagrange multiplier (LM) tests which allow for structural breaks under both the null and alternative hypothesis. The alternative hypothesis unambiguously implies trend stationarity in their test.

The unit root analysis that would allow for structural breaks is conducted in three steps in our analysis. First, assuming a linear data generating process, the endogenous two-break LM test of the Lee and Strazicich (2003) test is applied. If this test suggests only one-break for a particular country, then, in the second stage of the exercise, we repeat the exercise with Lee and Strazicich (2004) test that would assume one break. The third and the last step consists of determining the source of the parameter instability in nonlinear estimations using Christopoulos and León-Ledesma (2010) test. The details of this test which simultaneously allow for an ESTAR type nonlinearity and structural breaks are provided in the end of the following subsection.

2.2. Nonlinear unit root tests

The recent literature provides many studies that employ nonlinear models to capture the asymmetries in macroeconomic time-series. A rough categorization of these models rests on the presumed regime-switching behaviour of the series. The first group consists of Markov-switching type models which assume that the regime change depends on an unobservable variable. Alternatively, threshold models presume that the regime-switch is controlled by an observable variable. In particular, self-exciting threshold models posit that the change in the regime is determined by the past values of the time-series itself. One popular specification of the self-exciting threshold models is the smooth transition autoregressive (STAR) type models (Granger and Teräsvirta, 1993) which allows for a gradual adjustment towards a long-run equilibrium. In this paper, we employ two extensions of the STAR framework as will be demonstrated below.

The first of our nonlinear models is the ESTAR model which assumes a symmetric and gradual adjustment towards a long-run equilibrium. This specification provides us a convenient framework to capture the asymmetric correction behaviour in food prices.

Kapetanios et al. (2003) suggests the model:

$$\Delta f_t = a_1 f_{t-1} + a_2 f_{t-1} \left[1 - \exp\left(-\theta(f_{t-d} - \lambda)^2\right) \right] + \varepsilon_t \tag{1}$$

where f stands for food prices. The speed of adjustment is determined by the transition function inside the hard brackets. The speed of adjustment between different regimes is determined by θ . A mean-zero stochastic process is imposed by choosing $\lambda=0$ and further $a_1=0$, similar to Kapetanios et al. (2003). This way, we posit that the series has unit root behaviour when it is close to the long-run equilibrium, but displays mean-reversion when it is far away from this value. We select the delay parameter $d=1$. Then, the equation turns into:

$$\Delta f_t = a_2 f_{t-1} \left[1 - \exp\left(-\theta f_{t-1}^2\right) \right] + \varepsilon_t \tag{2}$$

In this equation, we would like to test the null hypothesis is $H_0: \theta=0$ against the alternative $H_1: \theta > 0$. Nonetheless, a well-known problem in these types of specifications is that the parameter a_2 is unidentified under the null. To deal with this problem, Kapetanios et al. (2003) employs an auxiliary regression, making use of a first order Taylor series approximation. Then, the general model with serially correlated is:

$$\Delta f_t = \sum_{j=1}^p p_j \Delta f_{t-j} + \gamma f_{t-1}^3 + error \tag{3}$$

Kapetanios et al. (2003) tabulates the asymptotic critical values for the t-statistics by employing the OLS estimation of $\gamma(\hat{\gamma})$.

AESTAR model (Sollis, 2009), provides a more specific form of ESTAR model where the speed of adjustment is allowed to be different below or above the threshold band. The formal model is:

$$\Delta f_t = G(\theta_1, f_{t-1}) [S(\theta_2, f_{t-1}) a_1 + \{1 - S(\theta_2, f_{t-1})\} a_2] f_{t-1} + \varepsilon_t \tag{4}$$

where

$$G(\theta_1, f_{t-d}) = 1 - \exp\left(-\theta_1 f_{t-1}^2\right) \theta_1 > 0 \tag{5}$$

$$S(\theta_2, f_{t-d}) = [1 + \exp(-\theta_2 f_{t-1}^2)]^{-1}, \theta_2 > 0 \tag{6}$$

Equation (4) tells that, without loss of generality, assuming $\theta_1 > 0$ and $\theta_2 \rightarrow \infty$, if f_{t-1} moves from 0 to $-\infty$ then $S(\theta_2, f_{t-d}) \rightarrow 0$ and ESTAR transition take place between the central regime model $\Delta \pi_t = \varepsilon_t$ and the outer regime model $\Delta f_t = a_2 f_{t-1} + \varepsilon_t$. Correspondingly, if f_{t-1} changes from 0 to ∞ then $S(\theta_2, f_{t-d}) \rightarrow 1$ and ESTAR transition takes place between the central regime model $\Delta f_t = \varepsilon_t$ and the outer regime model $\Delta f_t = a_1 f_{t-1} + \varepsilon_t$. In both cases, the speed of transition is determined by θ_1 . The asymmetric adjustment is maintained by $a_1 \neq a_2$. The generalized model including serially controlled errors reads:

$$\Delta f_t = G(\theta_1, f_{t-1}) [S(\theta_2, f_{t-1}) a_1 + \{1 - S(\theta_2, f_{t-1})\} a_2] f_{t-1} + \sum_{i=1}^k \kappa_i \Delta f_{t-i} + \varepsilon_t \tag{7}$$

Similar to the ESTAR case, the unit root testing procedure confronts the identification problem under the null. Sollis (2009) adopts a two-step Taylor series expansion; around θ_1 followed by another one around θ_2 . Then the equation yields:

$$\Delta f_t = \phi_1 (f_{t-1})^3 + \phi_2 (f_{t-1})^4 + \sum_{i=1}^k \kappa_i \Delta f_{t-i} + \mu_t \quad (8)$$

where $\phi_1 = a_2^* \theta_1$ and $\phi_2 = c(a_2^* - a_1^*) \theta_1 \theta_2$ with $c=0.25$. Here, a_1^* and a_2^* are functions of a_1 and a_2 as described in Sollis (2009). The null hypothesis is:

$H_0: \phi_1 = \phi_2 = 0$. In the auxiliary equation (8), Sollis (2009) provides asymptotic distribution of an F-test and critical values for the cases with zero mean, non-zero mean and deterministic trend.

The last test, Christopoulos and Leon-Ledesma (2010) examines the joint presence of structural breaks and nonlinearities in the food price series. The series could display long-run mean reversion along with temporary breaks. This test is basically a modified version of the aforementioned Kapetanios et al. (2003) ESTAR test, allowing for infrequently smooth mean changes. For the sake of space limitations, we do not provide the full details of the test here, referring the reader to three-step procedure described in Christopoulos and Leon-Ledesma (2010, p1082).

3. Results

Table 1 and Table 2 documents the results for unprocessed food price index and for beef prices, respectively, for twenty-two European countries. The left parts of Tables 1 and 2 display the results for the endogenous two-break LM test of Lee and Strazicich (2003) and endogenous one-break LM test of Lee and Strazicich (2004) for unprocessed food prices and beef prices, respectively. The right parts of the same tables document the linear and nonlinear unit root tests for two series.

In both tables, the first column stands for the type of structural break model of Perron (1989). Model A assumes a one-time change in level and model C allows for a change in both the level and trend. The second column denotes whether the estimation includes a time-trend ($t = 1$) or not ($t = 0$). A visual analysis of Figs. 2 and 3 suggests a continuous trend for most countries for both aggregates. Both theoretically and empirically, the idea of a time trend in food price index and beef prices is plausible and tested in many studies in previous literature.¹³ In our analysis, the autocorrelation functions of the series are used to check for the existence of a trend in series. Accordingly, with the exception of beef price series in Cyprus and Netherlands, all estimations contain a trend. Consequently, Model C is chosen for all countries except these two.

Third column provides the optimal lag length k which is derived using a general-to-specific procedure as described in Lee and Strazicich (2003). Fourth and ninth columns, LM_{T2} and LM_{T1} report the LM test statistic where the subscript stands for the number of breaks considered. Fifth and tenth columns, TB_2 and TB_1 denote the structural break dates suggested by two-break and one-break tests, respectively. The critical values depend on the location of the breaks ($\lambda_i = TB_i/T$), given on the sixth, seventh and eleventh columns, with T being the number of observations.

In both tables, the first three columns of the right part document the results of the test statistics for the linear unit root tests, respectively Augmented Dickey Fuller (ADF), Elliot-Rottenberg-Stock (ERS) and Phillips-Perron tests. The following two columns document the ESTAR t-statistics ($t_{nl, kss}$) for the Kapetanios et al. (2003); and AESTAR F-statistics ($F_{AE, \mu}$) for Sollis (2009). The Christopoulos and Leon-Ledesma (2010) test is designed with trend case and is provided by ($t_{nl, \mu}$) in Table 3 (page 1084) of that study

with the optimal lag value (k_{ll}).

3.1. Results for unprocessed food prices

A joint look at the results in Table 1 suggest that for five countries out of twenty-two, neither linear nor nonlinear tests reject the null of unit root: Finland, Greece, Italy, Malta and Turkey. This initial result deserves a more stringent examination because a common characteristic of these countries except Italy is the high volatility of their food prices. Akcelik and Yucel (2016) study suggests that these four countries, Malta, Turkey, Finland and Greece, are in top five out of eighteen countries in terms of food price volatility among European countries. Hence, our results suggest that for these countries for which the magnitude of the fluctuations is higher than the other European countries, the impact of these fluctuations are far from being transitory but result in permanent shifts in unprocessed food price series.

For the rest of the group, for seventeen countries, the results suggest mean-reversion in linear or non-linear way, indicating that shocks are relatively transitory and the series would converge back to the trend level at some point in the future after the impact of the shocks fade away. However, for the five countries with no sign of stationarity, it can be suggested that the cyclical shocks have a tendency to turn into more structural problems. Hence, the policies against food price fluctuations should include more structural components for these countries. A further analysis of the evolution of food prices in countries with nonstationary prices in the following paragraphs provides us with interesting insights.

Sammur (2015) points out that during the surge in food prices around 2007–2011 period imports of certain food items (such as wheat, maize, rice and sugar) have also increased in Malta, which might suggests expectations of further price increase. He explains this importing and further hoarding behaviour phenomenon with the “entrepreneurial spirit in seeking out quick profits” in this small island state. This speculative behaviour through Maltese export-import companies combined with our results of nonstationarity in food prices could support the idea that the impact of speculative motives could linger and dominate that of fundamentals, preventing the prices to revert to the trend level after a certain period of time. Sammur (2015) also underlines that in addition to the primary market import-export behaviour, the speculative behaviour through the index-linked derivatives in food market further exacerbates the long-term trend in higher food prices.

A recent OECD (2014) study based on a research by Finnish authorities admits that consumer prices of food items are relatively higher in Finland in an internationally comparative manner. They suggest four reasons for that. First one is the high value added tax on food items. Second, the competitiveness in agricultural market in Finland is low which is due to unfavourable climate conditions of the country and small average farm size. Third, market concentration in retail sector is much higher in Finland compared to other European countries. The same report tells that the share of wholesale and retail trade in food chains is higher in Finland than other EU countries mainly due to high costs of transportation between long distances.

Apergis and Rezitis (2011) examine the evaluation of food prices in Greece. They suggest high competition from EU countries and Common Agricultural Policy (CAP) reforms as determinants of the fluctuations of food prices in addition to the surges in input prices. In particular, they argue that CAP reforms resulted in lower intervention prices and higher direct payments to producers as a part of compensation policy regardless of the production level which, in turn, increased price fluctuations. They further argue that General Agreement on Tariffs and Trade (GATT) changes have also added to instability in food process since they increased international

¹³ See Baffes and Etienne (2016) for a review of this literature and tests on Prebisch-Singer hypothesis.

Table 1
Unprocessed food prices.

	Minimum LM unit root test with structural breaks										Linear and nonlinear unit root tests											
	Two-break test					One-break test					Linear tests			Nonlinear tests								
	model	t	k	LM _{T2}	T _{B2}	λ_1	λ_2	k	LM _{T1}	T _{B1}	λ	ADF	ERS	PP	t _{nl,il}	t _{nl,kss}	F _{AE,il}					
Austria	C	1	12	-5.48	*	2005m02, 2013m02	0	0.8				-3.11	9.28	-2.81	0.23	1.83	5.55					
Belgium	C	1	11	-6.21	***	2001m01, 2003m11	0	0.4				-4.49	3.26	-4.27	***	0.93	1.54	9.76	***			
Cyprus	C	1	12	-4.86		2006m04, 2013m09	1	0.9				-4.58	***	3.48	***	-4.19	***	-1.56	2.88	*	10.71	***
Denmark	C	1	10	-5.82	**	2005m02, 2007m11	0	0.6				-3.00	5.32	**	-2.87	0.03	1.94	1.78				
Finland	C	1	9	-5.14		2012m06	1		9	-1.37	1998m04	0.1	-2.16	11.81	-2.50	-2.01	1.56	4.71				
France	C	1	9	-5.01		2001m01, 2004m08	0	0.4				-3.09	5.12	*	-3.22	*	-0.35	1.11	6.57	*		
Germany	C	1	9	-5.52	*	2003m10, 2012m08	0	0.8				-1.94	13.51	-1.94	-0.19	2.17	4.79					
Greece	C	1	10	-4.79		2004m05, 2011m12	0	0.8				-2.43	17.24	-2.22	-2.59	0.30	3.84					
Ireland	C	1	9	-5.94	**	2000m04, 2010m06	0	0.7				-1.14	68.24	-1.08	-2.30	-0.36	0.45					
Italy	C	1	7	-4.65		2001m09, 2013m11	0	0.5				-2.21	9.54	-1.90	-0.85	1.00	0.59					
Latvia	C	1	9	-5.19		2005m01, 2007m12	0	0.6				-1.05	38.67	-1.19	0.19	3.39	**	0.95				
Lithuania	C	1	11	-5.49	*	2003m06, 2007m10	0	0.6				-1.51	22.43	-1.46	-2.08	3.41	**	1.78				
Luxembourg	C	1	8	-5.51	*	2001m03, 2011m04	0	0.7				-3.43	*	3.22	***	-2.79	0.52	1.13	6.77	**		
Malta	C	1	12	-4.25		2005m01, 2012m02	0	0.8				-2.98	18.30	-1.66	0.30	2.05	2.56					
Netherlands	C	1	10	-6.27	***	2001m01, 2004m08	0	0.4				-2.98	5.57	**	-2.83	0.01	0.78	6.59	*			
Poland	C	1	9	-4.51		2003m03, 2013m12	0	0.9				-2.75	20.24	-2.75	-0.98	2.35	6.21	*				
Portugal	C	1	8	-5.84	**	2000m10, 2009m04	0	0.7				-2.14	31.42	-2.30	-0.13	0.50	4.30					
Slovakia	C	1	10	-4.01								-3.03	5.57	**	-2.69	-2.46	-1.31	0.05				
Spain	C	1	12	-4.86		2000m07, 2009m03	0	0.6				-0.84	32.29	-0.98	-6.40	***	0.05	0.09				
Sweden	C	1	7	-4.38		2008m01, 2011m07	1	0.8				-3.25	*	5.80	*	-3.10	-0.20	1.24	9.05	***		
Turkey	C	1	12	-3.79		2002m10	0		12	-3.77	2005m02	0.4	-2.71	16.20	-2.51	-1.25	1.75	2.61				
UK	C	1	12	-4.84		2008m02, 2014m05	1	0.9				-1.44	52.60	-1.62	-1.10	2.93	*	1.57				

Note: *, ** and *** stand for significance levels at 10%, 5% and 1%, respectively. The critical values of the Lee and Strazicich (2003, 2004) test on the left depend on the location of the structural breaks where ($\lambda_i = T_{Bi}/T$) and T is the number of observations. Critical values for LM_{T2} and LM_{T1} are tabulated in Table 2 of Lee and Strazicich (2003) and Table 1 of Lee and Strazicich (2004), respectively. Critical values for 10%, 5% and 1% are -2.66, -2.93 and -3.48 for ESTAR test; 4.17, 4.97 and 6.81 for AESTAR test, respectively. The critical values of the Christopoulos and Leon-Ledesma (2010) test depend on the optimal value of lag and are tabulated in Table 3 (page 1084) of this article.

Table 2
Beef prices.

	Minimum LM unit root test with structural breaks											Linear and nonlinear unit root tests										
	Two-break test					One-break test						Linear tests			Nonlinear tests							
	model	t	k	LM _{T2}	T _{B2}	λ_1	λ_2	k	LM _{T1}	T _{B1}	λ	ADF	ERS	PP	$t_{nl,II}$	$t_{nl,kss}$	F _{AE,μ}					
Austria	C	1	10	-4.87	2013M01	1	0	-3.00	2014M11	0.9	-2.48	1.14	***	-2.17	-3.80	***	-1.02	0.03				
Belgium	C	1	4	-5.03	2008M02	0	4	-4.98	**	2007M06	0.9	-1.75	26.55	-1.44	-2.80	*	0.52	0.75				
Cyprus	A	0	4	-6.28	***	2008M12, 2009M03	0	0.4				-2.72	1.48	-2.60	-6.10	***	-6.10	***	10.20	***		
Denmark	C	1	12	-5.42	*	2008M10	0	8	-4.57	**	2007M12	0.3	-1.78	6.44	*	-2.09	-3.10	*	0.97	0.40		
Finland	C	1	10	-7.00	***	2009M03, 2013M03	0	0.7				-3.35	*	5.65	*	-1.08	-3.60	***	1.32	5.07		
France	C	1	10	-4.99		2012M07	1	11	-3.03		2007M04	0.2	-1.93	6.05	*	-1.53	-2.39		0.86	0.67		
Germany	C	1	12	-6.46	***	2009M02, 2011M12	0	0.6				-2.17	7.92	-1.92	-3.30	**	0.81	1.09				
Greece	C	1	9	-5.16		2007M02, 2011M07	0	0.6				-1.55	32.50	-1.12	-1.47		0.66	6.04	*			
Ireland	C	1	12	-6.70	***	2011M01, 2013M07	1	0.8				-2.76	7.23	-2.64	-3.30	***	2.32	0.12				
Italy	C	1	12	-6.96	***	2006M05, 2011M12	0	0.6				-2.59	8.79	-2.52	-2.90	*	2.50	1.21				
Latvia	C	1	4	-4.64		2007M11	0	10	-3.76		2008M01	0.3	-1.78	9.64	-2.30	-3.70	***	-0.21	0.96			
Lithuania	C	1	12	-6.12	**	2009M07, 2013M04	0	0.7				-1.41	10.85	-2.36	-2.38		-0.72	0.33				
Luxembourg	C	1	10	-3.65		2014M01	1	10	-2.56		2007M09	0.2	-1.85	6.50	*	-2.05	-1.63	2.23	0.88			
Malta	C	1	9	-3.96		2008M06	0	7	-3.67		2007M10	0.3	-2.72	0.89	***	-1.87	-1.15	0.39	0.89			
Netherlands	A	0	1	-7.44	***	2008M03	0	10	-4.01		2010M11	0.5	-3.55	***	3.28	*	-2.48	-3.50	**	-3.30	**	0.89
Poland	C	1	11	-6.34	**	2008M09, 2013M03	0	0.7				-2.08	2.54	***	-1.71	-3.70	***	0.94	0.34			
Portugal	C	1	1	-4.90		2007M07	0	1	-3.75		2014M11	0.9	-3.20	*	9.50	-2.19	-0.20	-1.86	0.07			
Slovakia	C	1	4	-5.23		2006M12, 2011M01	0	0.5				-1.86	11.11	-2.04	-0.70		1.81	1.85				
Spain	C	1	10	-8.18	***	2009M01, 2012M08	0	0.7				-3.04	0.05	***	-1.93	-2.90	**	1.30	0.15			
Sweden	C	1	11	-6.54	***	2009M05, 2013M05	0	0.7				-3.13	0.08	***	-2.59	-1.70		1.48	0.86			
Turkey	C	1	2	-4.33								-1.84	12.07	-1.66	-7.10	***	1.28	0.04				
UK	C	1	12	-5.65	**	2008M01, 2013M11	0	0.8				-1.04	14.27	-1.21	-2.31		0.44	3.74				

Note: *, ** and *** stand for significance levels at 10%, 5% and 1%, respectively. The critical values of the Lee and Strazicich (2003, 2004) tests on the left depend on the location of the structural breaks where ($\lambda_i = T_{Bi}/T$) and T is the number of observations. Critical values for LM_{T2} and LM_{T1} are tabulated in Table 2 of Lee and Strazicich (2003) and Table 1 of Lee and Strazicich (2004), respectively. Critical values for 10%, 5% and 1% are -2.66, -2.93 and -3.48 for ADF test; 4.17, 4.97 and 6.81 for AESTAR test, respectively. The critical values of the Christophoulos and Leon-Ledesma (2010) test depend on the optimal value of lag and are tabulated in Table 3 (page 1084) of this article.

competition through limiting agriculture support programs. Our nonstationary result confirms with these in the sense that these structural policy changes might prevent small shocks to be contained. As the argument goes, due to the problems in the adjustment to agricultural liberalization in a relatively small economy, the impact of cyclical shocks could magnify and turn into structural problems.

In Turkey, the highest contribution to consumer inflation is provided by unprocessed food prices and exchange rate in the last decade (Kara et al., 2017). Furthermore, Ganioglu (2017) and Bastan and Chadwick (2017) emphasizes the key role of high food price volatility in deterioration of inflation expectations. Different studies report inefficiencies in intermediation due to long-supply chains engaging in transport, packaging and classification of products (Songul, 2017); infrastructure problems in storage (Tumen and Songul, 2017) or transport highways (Saygili and Tumen, 2017); lack of a quality standard in exported food items (Sevinc, 2017) or agricultural credit policies (Seven, 2016) and climate change (Dudu and Çakmak, 2018) as possible determinants of food price fluctuations in Turkey.

A second observation from Table 1 is about the structural break dates. While the structural break dates show dispersion over the years within countries, the period between 2001–2005 and 2011–2013 constitute the break dates for many countries. These results make sense considering sudden surges in unprocessed food prices in these periods as discussed in our introductory section.

Third, our results confirm with the conjecture that while testing for unit root in under linear alternatives, it is important to consider possible structural breaks. To notice this, we compare the test results of Lee and Strazicich (2003, 2004) tests with the other linear unit root tests (ADF, ERS and PP) in Table 1. The two break LM unit root test of Lee and Strazicich (2003) rejects the null of unit root for 9 countries: Austria, Belgium, Denmark, Germany, Ireland, Lithuania, Luxemburg, Netherlands and Portugal. However, for five countries among these nine (Austria, Germany, Ireland, Lithuania and Portugal) none of the linear tests reject the unit root null.

Our fourth conclusion is in regard to the importance of considering nonlinearities in unit root testing. A joint look at both sides of the table documents six countries for which only linear tests reject the null of unit root: Austria, Denmark, Germany, Ireland, Portugal and Slovakia. These results suggest that these countries reveal mean-reversion in the long-term that could be characterized by a linear process. However, for Latvia, Poland, Spain, Sweden and UK; while linear unit root tests do not reject the null of unit root, nonlinear unit root tests do reject the null. This means that the mean-reversion displays nonlinear characteristics for these countries. Among these countries, Latvia and UK display ESTAR type of mean-reversion. Hence, for these two countries, the series show nonstationary behaviour inside a symmetric threshold band while it has a tendency to revert back to the trend above or below this band. Poland on the other hand displays an AESTAR type mean reversion where the band inside which the series show nonstationary behaviour is not symmetric. For the rest of the group (Belgium, Cyprus, France, Lithuania, Luxembourg, and Netherlands) both linear and nonlinear tests reject the null of unit root.

Fifth, Christopoulos and Leon-Ledesma (2010) test results indicate that ESTAR type mean reversion could only be detected once we consider structural breaks for Spain. This confirms with the conjecture that, it is important to consider the structural breaks while testing for unit root under nonlinear alternatives.

3.2. Results for beef prices

A first joint look at the linear and non-linear unit root test results in Table 2 tells that the null of unit root is rejected for all

countries except Slovakia. Compared to the unprocessed food price results, this higher rate of mean-reversion could be result of two important characteristics of beef prices. First, as pointed out before, meat sector constitute a significant part of unprocessed foods and hence it takes significant aid by EU for a long time through different and changing methods. Nielsen and Jeppesen (2001) indicate meat producers are compensated for their losses due to falling prices in the form of higher headage premiums while a recent OECD-FAO (2015) report underlines the use of coupled payments. These policies could smooth the impact of short-term shocks and help the market prices come back to equilibrium levels in a shorter-period. Second, as mentioned in the motivational part, speculation could be a reason for certain agricultural prices to reveal longer deviations from their equilibrium prices. Since it is harder to stock beef (or livestock) than other agricultural products (such as wheat or rice), beef prices could be comparably less subject to speculations than other products.

Table 2 document that linear and nonlinear tests reject the null of unit root in ten countries: Austria, Belgium, Cyprus, Denmark, Finland, Ireland, Italy, Netherlands, Poland and Spain. For eight countries, only linear tests reject the null of unit root: France, Germany, Lithuania, Luxembourg, Malta, Portugal, Sweden and UK. For three countries, only non-linear unit root tests reject the null of unit root: Greece, Latvia and Turkey. Similar to the previous section, these results suggest that using non-linear alternatives is important in unit-root testing. Among this last group Latvia and Turkey shows ESTAR type of mean-reversion whereas Greece displays AESTAR type of mean reversion.

On the subjects of the structural breaks and nonlinearities, three points are noteworthy. First, the period after 2011 consists almost half of the breaks in as illustrated in under T_{B2} column of Table 2. This result is in line with that of the same column in Table 1 for unprocessed food prices since this period marks sudden changes in beef prices as well as other agricultural prices. Second, both linear (Lee and Strazicich, 2003, 2004) and nonlinear (Christopoulos and Leon-Ledesma, 2010) tests that takes structural breaks into account provide higher rejection rates for beef prices compared to unprocessed food prices. In a similar manner, the rejection rate for the null of unit root by nonlinear tests is higher in beef prices compared to unprocessed food prices. These suggest that structural breaks and nonlinear characteristics inherent to the production process for beef prices discussed in the introduction part are less influential in the mean-reversion process compared to other unprocessed food products. This tells that, while the beef prices could reveal some fluctuations around its equilibrium for many countries; mean-reversion exists. However, this reversion might reveal nonlinear characteristics due to the intrinsic nonlinearities in beef production or price formation.

4. Conclusion and discussion

Stability in food prices is essential for many reasons such as food security or price stability in general. The high fluctuations in food prices in the last decade urges policy makers to design policies to smooth the potential negative impacts. From this standpoint, this paper contributes to the literature that explores the time-series characteristics of unprocessed food prices and beef prices. Our interest lies in assessing the linear or non-linear mean-reverting behaviour in these series for twenty-two European countries. We employ linear and nonlinear unit root tests, taking into account possible trends and structural breaks in the series. Our results indicate that mean-reversion does not exist for some countries, especially in the general index of unprocessed food prices. Moreover, it is important to take into account structural breaks or nonlinearities in reverting behaviour. For some countries the mean

reversion in these series could take alternative smooth transition forms including asymmetries depending on the sign and size of the deviation from the long-run mean. These results have policy implications since the benefits of demand-management policies depend on the persistence of shocks as discussed in the text.

A number of research questions emerge for future analysis. First, what are the determinants of the excessive fluctuations in food prices for an economy? While our models argue that for some countries there is a threshold level beyond which strong correction behaviour towards the equilibrium is observable; the impact of policies prescribed against extreme deviations along these mean-reverting behaviour could not be singled out straightforwardly. There are many other forces that would affect the strength and the speed of the drive towards the mean, including the domestic market structure, the extent of the price transmission from international prices to domestic market or the degree of market openness. Hence, assessing the appropriate timing of the policy response requires timely and detailed information on the price dynamics at the national and the global level. From this standpoint, a further avenue of research would explore the diverse impacts of food price fluctuations on food importing countries and the countries which mainly depend on export revenues from agricultural sector.

Second, while our focus on this analysis is on the national food price index and beef prices in particular, a further detailed exploration might study other components of these aggregate indices; such as dairy, sugar or grain. Such disaggregation would help the researchers in taking cognizance of the disproportionate impact of price fluctuations on key exporters of agricultural commodities compared to the others. Moreover, the weights of these components in the consumer price indices would differ among countries which would also affect the pace and the of the price adjustment.

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