

# Mean-Reversion in Unprocessed Food Prices

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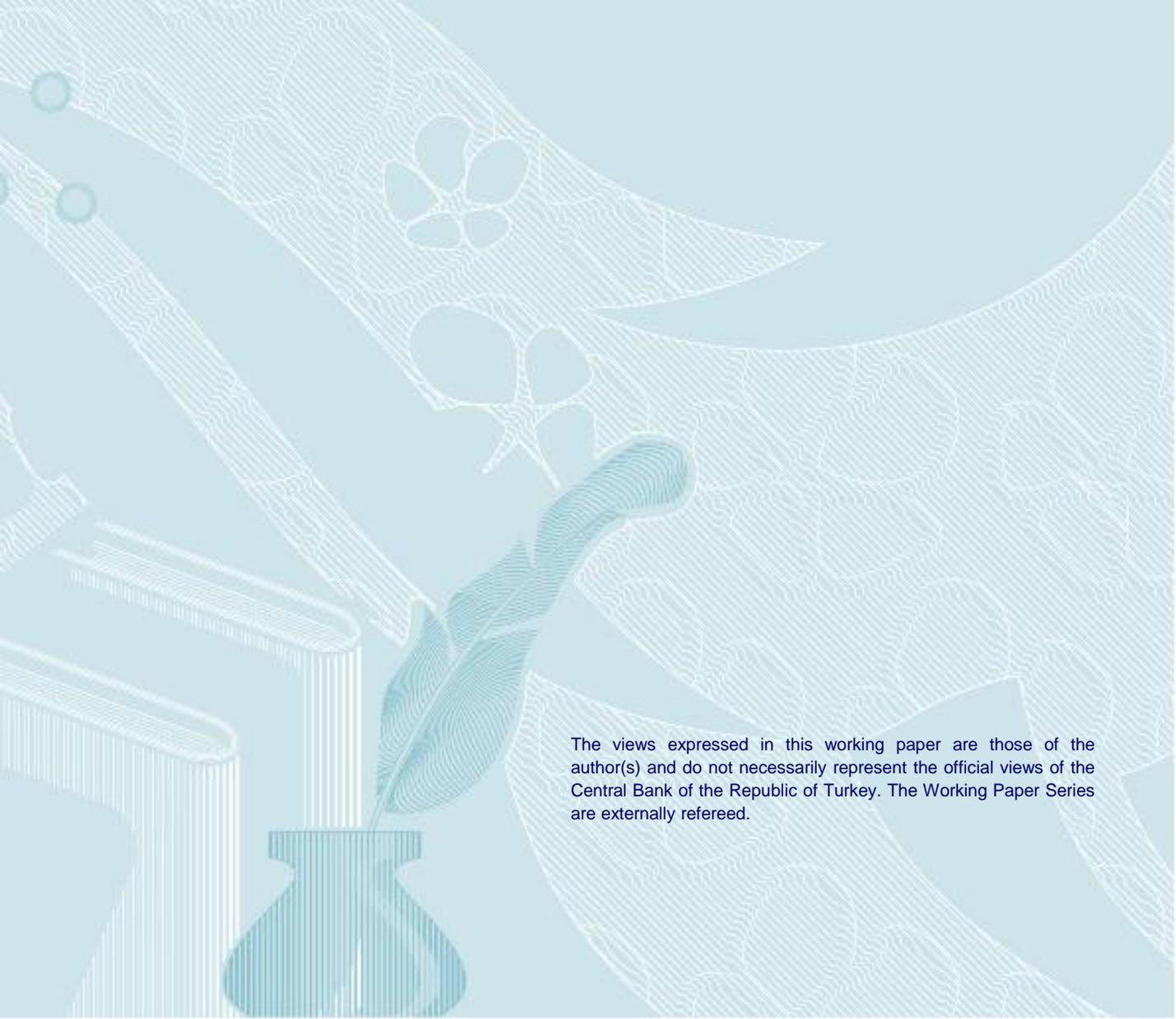
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# Mean-Reversion in Unprocessed Food Prices

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## **Abstract:**

The high volatility in food prices in the last decade is a major concern for policymakers across the globe. This study tests if there is mean-reverting behaviour in the unprocessed food prices towards a long-run trend, for twenty-four European countries, using linear and nonlinear unit root tests. The results indicate linear or non-linear mean reversion for more than one-third of the group. Non-linear models are useful in detecting asymmetric correction behaviour, depending on the size and sign of the deviation from the mean. However, out-of-sample forecasting performances of these models are poor relative to a simple autoregressive benchmark.

**JEL Classification:** Q11, C32, C53

**Keywords:** Food prices, unit root test, nonlinearity, forecasting

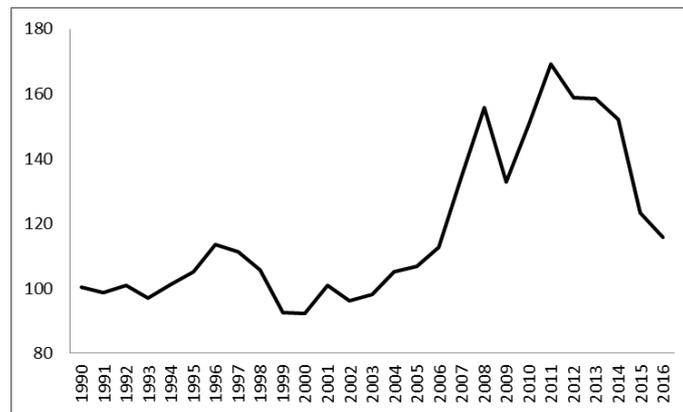
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## I. Introduction

The last decade displayed high volatility in global food prices (Figure 1). Whether this volatility, which is in sharp contrast with the secular downward trend over the last fifty years, is a “new normal” or would the prices go back to their long-term trend is the subject of an ongoing debate among policymakers<sup>2</sup>. Hence, a good understanding of the time-series characteristics of food price series is an important step in developing policies aimed to stabilize inflation and establishing models to forecast inflation. From this standpoint, this paper examines whether the unprocessed food prices (the most volatile component of the food prices) has a tendency to revert to a long-term mean level (or a trend); and if so, whether this adjustment behaviour towards this attractor differs depending on the distance and sign of the deviation from it.

**Figure 1: Annual Real Food Price Indices**



Source: Food and Agricultural Organization (FAO). 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”.

Fluctuations in food prices have diverse impacts on the well-being of economic agents<sup>3</sup>. For instance, higher food prices could benefit small-scale farmers while endangering the access to food by low-income households in many underdeveloped countries. Frequent changes in prices would lead to sub-optimal investment and consumption decisions in agriculture, with possible repercussions at the national and global level. Moreover, changes in food prices are one of the main determinants of the consumer inflation for many countries. Hence, volatility in food prices is a major concern for governments that would aim for food security and relatively stable inflation, at times necessitating coordinated action around the world<sup>4</sup>.

<sup>2</sup> One statement by Food and Agricultural Organization (FAO, 2013) emphasizes that the volatility in food price might continue over the following years and policies might “need to adapt this *new normal*”. A more recent policy note by FAO (2016) takes a slightly different stance: “The conventional wisdom is that, in the long-term, real commodity prices would follow a declining trend interrupted by periodic and sudden surges”. Also, See FAO (2012a) for outlook of food price developments of the period.

<sup>3</sup> 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 Dilon and Barrett (2015) for the impact of oil prices on food prices.

<sup>4</sup> See G20 statement in FAO (2012b) for suggestions for coordinated policy actions against rising volatility in global food prices.

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. Regarding our focus on food prices; existence of a long-run mean level, or a trend, is important for a couple of reasons. First, if there exists such a level of equilibrium that the prices revert; then we have reasons to believe that *fundamentals* are the main factors that drive the behaviour of food prices in the long-run. As the argument goes, food prices might display short-term deviations due to many reasons such as sudden changes in climate or speculative reasons; yet, they convert to an equilibrium level in the long-run. Otherwise, if there is no such convergence, then it means that short-term shocks to the prices are persistent and the prices can easily settle to a new equilibrium level after relatively short-term disturbances

This aforementioned behaviour in prices is mainly analysed by linear and nonlinear unit root analysis in the literature. Stationarity in prices ensures that prices converge to a long-run equilibrium level. Linear unit root analysis test whether the short-term deviations are corrected in a linear way. This means that any small deviation from the long-run equilibrium level (or trend if we assume a trend-stationary process) has a tendency to be corrected in the long-run. On the other hand, nonlinear unit root tests assume asymmetric correction behaviour. In such models, the strength of the correction towards the long-run equilibrium level depends on the distance and/or sign of the deviation from this equilibrium level.

In this paper we test for the stationarity of unprocessed food prices for 24 European countries, using two different nonlinear models of smooth transition - Exponential Smooth Transition Autoregressive (ESTAR) and Asymmetric Exponential Smooth Transition Autoregressive (AESTAR). Both models assume that the correction behaviour is stronger when the series get far away from the long-run level. Yet, the sign of the deviation does not matter in ESTAR model while it is important in AESTAR model.

A number of reasons are discussed in the literature for the asymmetric behaviour of food prices. First, 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).

A second explanation is provided by Holt and Craig (2006) study which shows that the US hog-corn cycle could be characterized by a nonlinear form, in particular a time-varying ESTAR model. They argue that farmers might give overnight decisions to slaughter their animals in response to changes in expectations of the market price for corn. 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<sup>5</sup>. Later on Balagtas and Holt (2009) also display the ability of STAR models to reveal the nonlinear behaviour of selected primary commodity prices<sup>6</sup>. Third, Thirlwall and Bergevin (1985) argued 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.

Our paper contributes to this literature in three important ways. First, similar to the previous literature we conduct ESTAR test suggested by Kapetanios, Shin and Snell (2003), taking cognisance of asymmetric behaviour around the long-run mean which depends on the *size* of the deviations. We further take into account the asymmetric correction behaviour in food prices depending on the size as well as the *sign* of the deviations from the equilibrium, using AESTAR test suggested by Sollis (2009). This type of correction behaviour could indicate that positive deviations of food prices from the equilibrium level could be more persistent compared to negative deviations, or vice versa. In addition to the aforementioned motives in the previous paragraph, this idea is in line with the recent literature documenting different strength of correction behaviour for positive or negative shocks to inflation<sup>7</sup>. We take into account the possibility of a trend in series and conduct the exercise with and without assuming trends. Our analysis that covers the unprocessed food price series of 24 European countries indicates nonlinear adjustment process for one-third of the countries.

A second contribution of our paper is our consideration of nonlinearity in the presence of structural breaks. Certain shocks, like technological innovations, could have permanent effects on the price process. Previous literature includes many papers that focus on the structural break and multiple equilibria while testing for trends and stationarity in commodity prices<sup>8</sup>. We take into account possible structural tests in nonlinearity testing using Christopoulos and Leon-Ledesma (2010) test.

Third, we conduct a forecasting exercise to examine the power of our proposed nonlinear models in predicting the unprocessed food prices. It is shown that these models do not have a significant contribution above a simple benchmark autocorrelation model. We believe that our results will benefit researchers that aim to understand the time series behaviour of one of the most volatile price series, the unprocessed food prices.

The remainder of the paper is organized as follows. Second section describes the econometric methodology. Third section demonstrates our data, documents the results of the linear and nonlinear unit root tests; estimation of nonlinear models as well as the results of the out-of-sample forecasting exercise. The fourth section concludes.

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<sup>5</sup> Similarly, regime-switching and smooth transition frameworks are used to analyse asymmetries in fish meal and soybean meal by Asche et al. (2013); and US soybean-to-corn price ratio by Ubilava (2012), respectively.

<sup>6</sup> Beckmann and Czudaj (2014) explore the nonlinearities in agricultural future prices using a smooth transition framework.

<sup>7</sup> 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. Also see Tillmann and Wolters (2012) and Manzan and Zerom (2014) for nonlinearities in persistence of shocks to US inflation.

<sup>8</sup> 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).

## II. Methodology

The literature widely reports the low power of conventional unit root tests in rejecting the null of nonstationarity, in the presence of nonlinearities in time series. This issue is addressed by means of nonlinear unit root tests in numerous studies. From this standpoint, this study first conducts three linear unit root tests that are extensively used in the literature: Augmented Dickey Fuller (ADF), Elliot-Rottenberg-Stock (ERS) and Phillips-Perron tests. Second, nonlinear unit root tests of Kapetanios, Shin and Snell (2003), Sollis (2009) and Christopoulos and Leon-Ledesma (2010) are performed. Third, nonlinear models are estimated and an out-of sample forecasting exercise of these models is implemented. However, the nonlinear models estimation is only conducted for countries for which the nonlinear unit root tests reject the null of stationarity, following the suggestion of Teräsvirta (2006) who argues that nonlinear models nest a linear regression model which cannot be identified once the data generating process is linear. The remaining of this section provides a brief review of nonlinear modelling and demonstrates three models that are used in the analysis.

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. As argued in the previous section, this specification provides us a convenient framework to capture the asymmetric correction behaviour in food prices.

Kapetanios et.al (2003) suggests the model:

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

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  $\theta$ . 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:

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

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:

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

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

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

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

where

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

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

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  $\infty$  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  $\theta_1$ . The asymmetric adjustment is maintained by  $a_1 \neq a_2$ . The generalized model including serially controlled errors reads:

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

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  $\theta_1$  followed by another one around  $\theta_2$ . 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).

### III. Data and Results

Our data consists of monthly unprocessed food price series of 24 European countries (as in Table 1) taken from Eurostat database. The data spans January 1996 to July 2016 period, adding up to 246 data points. The series are indexed as (2005=100) and are seasonally adjusted using Census X-12 method. Figure 2 provides a graph of these indices for our sample countries along with the EU average. The unprocessed food prices (of EU countries) follow an increasing trend after 2000, partly in line with the food prices series (for world) given in Figure 1. Yet, the countries are dispersed in a wide range along the average, indicating differences between the food price behaviour among EU countries over the sample period.

#### a. Unit Root Tests

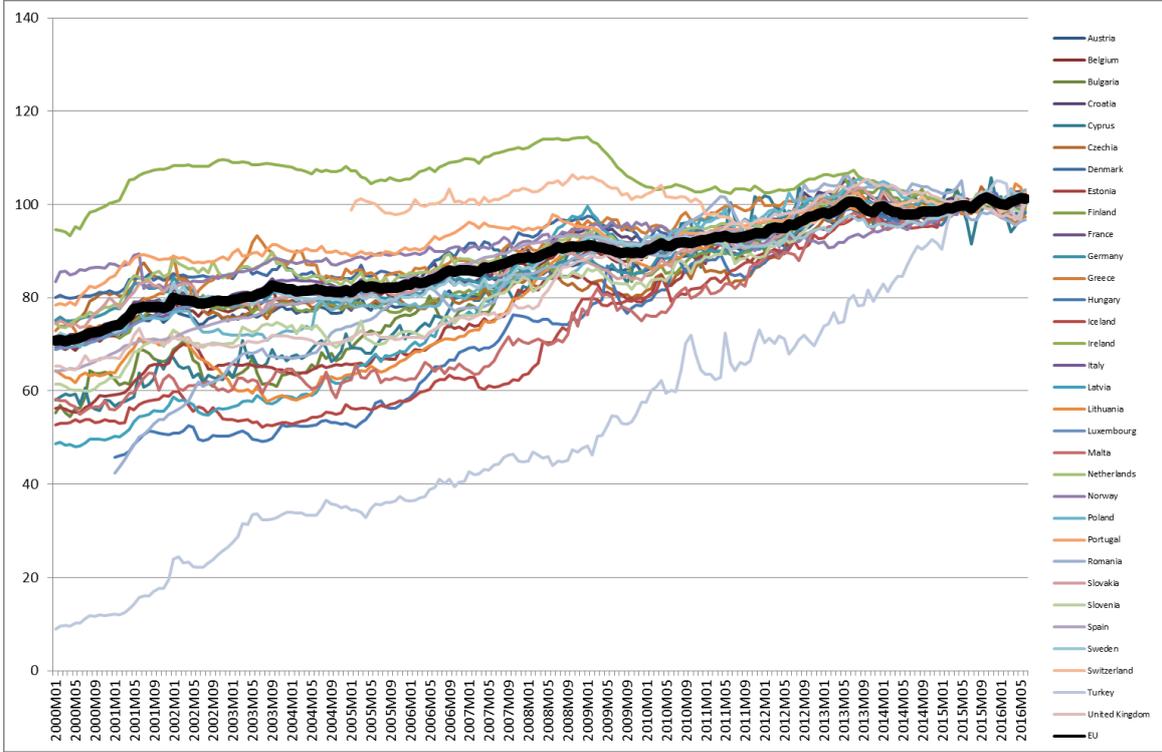
Both Kapetanios, Shin and Snell (2003) and Sollis (2009) provide critical values for unit root tests that are estimated with and without time trends. Both theoretically and empirically, the idea of a time trend in food price index is plausible and tested in many studies in previous literature<sup>9</sup>. Hence, we conducted both linear and non-linear tests with and without trends. Table 1 and Table 2 present the results with no trend and with trend case, respectively.

In both tables, the first three columns 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, Shin and Snell (2003); and AESTAR F-statistics ( $F_{AE,\mu}$ ) for Sollis (2009). The

<sup>9</sup> See Baffes and Etienne (2016) for a review of this literature and tests on Prebisch-Singer hypothesis.

Christopoulos and Leon-Ledesma (2010) test is designed with trend case and is given by  $(t_{nl,II})$  in Table 3 with the optimal lag value  $(k_{II})$ .

**Figure 2: Unprocessed Food Price Indices**  
*(Sample countries, seasonally adjusted, thick line indicates EU average)*



Source: Eurostat

A first look at Tables 1 and 2 suggest that for both linear and nonlinear unit root tests, adding a trend to the estimation results in a higher rejection of the null of unit root. In no trend case (Table 1), the linear unit root tests only suggests signs of stationarity for France whereas AESTAR test points out stationarity only for four countries. However, as documented in the table, linear unit root tests reject the null of unit root for eight countries (Belgium, Cyprus, Denmark, France, Luxemburg, Netherlands, Slovakia and Sweden) when we include time trend to the estimation.

Regarding nonlinear unit root tests with the trend case, both ESTAR and AESTAR tests suggest stationarity for Cyprus. For Iceland, Latvia, Lithuania and United Kingdom only ESTAR test reject the null of unit root. Note that for none of these countries linear unit root test could reject the null of unit root. These results highlight the importance of taking into account possible nonlinear behaviour in testing for unit root for food prices. Similarly, for France, Norway and Poland only AESTAR test could reject the null of unit root whereas for Belgium, Luxemburg, Netherlands and Sweden both linear unit root tests and AESTAR test could reject the null of unit root.

**Table 1**  
**Linear and Non-Linear Unit Root Tests (constant, no trend)**

	ADF	ERS	PP	$t_{nl,kss}$	$F_{AE,\mu}$
Austria	0.09	178.99	0.46	-0.58	0.24
Belgium	-0.12	244.78	-0.11	-0.27	0.63
Cyprus	-1.12	42.29	-0.77	-1.84	7.56 ***
Denmark	-0.77	93.51	-0.68	-0.52	0.76
Finland	-0.50	126.08	-2.50	-1.31	1.64
France	-3.09	5.12 **	-3.22 *	-0.51	1.10
Germany	-1.94	13.51	-1.94	-0.39	1.28
Greece	-2.43	17.24	-2.22	-2.05	2.52
Iceland	0.59	264.55	0.53	-0.52	1.07
Ireland	-2.13	65.42	-2.16	-1.87	2.52
Italy	-0.59	414.47	-0.63	-1.10	5.74 **
Latvia	-0.55	332.99	-0.56	-1.26	17.23 ***
Lithuania	-0.39	95.89	-0.57	-1.58	2.99
Luxembourg	-0.38	365.60	-0.44	-0.81	1.55
Malta	0.35	104.01	0.59	-0.08	0.04
Netherlands	-0.87	115.81	-0.81	-0.50	0.20
Norway	-0.99	134.97	-0.88	-1.27	1.87
Poland	-1.78	323.03	-1.75	-1.89	1.90
Portugal	-1.85	231.47	-1.79	-1.66	1.88
Slovakia	-1.58	75.33	-1.35	-0.97	0.57
Spain	-0.99	1007.86	-0.96	-2.09	11.23 ***
Sweden	-0.37	118.05	-0.27	-0.56	0.47
Turkey	0.68	274.72	1.40	0.05	1.07
United Kingdom	-0.20	200.40	-0.31	-0.96	3.29

Note: \*, \*\* and \*\*\* stand for significance levels at 10%, 5% and 1%, 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 results of the Christopoulos and Leon-Ledesma (2010) test in Table 2 suggest rejection of the null of unit root only for Spain. Note that for the case with trend this is the only test that would imply stationarity for Spain in trend case. This underscores the importance of taking into account structural breaks while testing for nonlinearities.

These results imply that the unprocessed food price series might imply nonlinear behaviour due to frictions and cyclical asymmetries discussed in the introductory part; and hence unit root tests considering asymmetric behaviour is of use in testing for the stationarity of these series. Next subsections explore how well our proposed nonlinear models of smooth transition fit the data and perform in forecasting exercises.

**Table 2**  
**Linear and Non-Linear Unit Root Tests (constant, with trend)**

	ADF		ERS		PP		$k_{II}$	$t_{n,II}$		$t_{n,ESS}$		$F_{AE,\mu}$
Austria	-3.11		9.28		-2.81		1	0.23		1.83		5.55
Belgium	-4.49 ***		3.26 ***		-4.27 ***		2	0.93		1.54		9.76 ***
Cyprus	-4.58 ***		3.48 ***		-4.19 ***		5	-1.56		2.88 *		10.71 ***
Denmark	-3.00		5.32 **		-2.87		1	0.03		1.94		1.78
Finland	-2.16		11.81		-2.50		3	-2.01		1.56		4.71
France	-3.09		5.12 *		-3.22 *		5	-0.35		1.11		6.57 *
Germany	-1.94		13.51		-1.94		1	-0.19		2.17		4.79
Greece	-2.43		17.24		-2.22		1	-2.59		0.30		3.84
Iceland	-1.60		39.44		-1.67		5	-0.63		2.84 *		2.60
Ireland	-1.14		68.24		-1.08		3	-2.30		-0.36		0.45
Italy	-2.21		9.54		-1.90		5	-0.85		1.00		0.59
Latvia	-1.05		38.67		-1.19		5	0.19		3.39 **		0.95
Lithuania	-1.51		22.43		-1.46		5	-2.08		3.41 **		1.78
Luxembourg	-3.43 *		3.22 ***		-2.79		5	0.52		1.13		6.77 **
Malta	-2.98		18.30		-1.66		1	0.30		2.05		2.56
Netherlands	-2.98		5.57 **		-2.83		1	0.01		0.78		6.59 *
Norway	-2.46		15.20		-2.62		2	-0.78		0.67		26.73 ***
Poland	-2.75		20.24		-2.75		2	-0.98		2.35		6.21 *
Portugal	-2.14		31.42		-2.30		4	-0.13		0.50		4.30
Slovakia	-3.03		5.57 **		-2.69		1	-2.46		-1.31		0.05
Spain	-0.84		32.29		-0.98		1	-6.40 ***		0.05		0.09
Sweden	-3.25 *		5.80 *		-3.10		1	-0.20		1.24		9.05 ***
Turkey	-2.71		16.20		-2.51		5	-1.25		1.75		2.61
United Kingdom	-1.44		52.60		-1.62		2	-1.10		2.93 *		1.57

Note: \*, \*\* and \*\*\* stand for significance levels at 10%, 5% and 1%, 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

### b. Model Estimation

As mentioned in the introductory section, nonlinear models are estimated only for the countries for which the nonlinear unit root tests suggest nonlinear behaviour. Considering our results in Table 2 we estimate ESTAR model for Cyprus, Iceland, Latvia, Lithuania and UK; AESTAR model for Belgium, Cyprus, France, Luxemburg, Netherlands, Norway, Poland and Sweden.

The predictive power of the AESTAR model is examined in a limited number of studies so far, as will be documented in the next subsection. Therefore, only the estimation results for AESTAR model is presented here, while the ESTAR estimation results are left out due to space considerations<sup>10</sup>. However, next section presents the results of out-of-sample forecasting exercise for both ESTAR and AESTAR models.

<sup>10</sup> ESTAR estimation results are available upon request.

Estimation of the raw form of the AESTAR model in Equation 4 is conducted with the restrictions  $\theta_1, \theta_2 > 0$  and  $a_1, a_2 < 0$ . The resulting  $\{\theta_1, \theta_2, a_1, a_2\}$  values are presented in the first four columns of the Table 3, where the standard errors are displayed in parenthesis below the figures<sup>11</sup>.

**Table 3**  
**AESTAR Model Estimation**

	$\theta_1$	$\theta_2$	$a_1$	$a_2$	$a_1 - a_2$
Belgium	0.26	70.51	-0.01	-1.44	1.43
	(162.26)	(26326.38)	(0.00)	(880.75)	(274.84)
Cyprus	0.01	0.75	-34.34	-0.19	-34.15
	(0.00)	(0.00)	(0.00)	(0.00)	(79878.18)
France	0.01	0.54	-0.01	-0.24	0.23
	(0.00)	(0.33)	(0.00)	(0.00)	(2.13)
Luxemburg	0.03	32.95	-0.01	-5.20	5.19
	(0.00)	(3.07)	(1.13)	(0.85)	(7026.81)
Netherlands	0.11	115.59	-0.01	-1.97	1.96
	(0.00)	(0.00)	(0.00)	(0.00)	(1967.17)
Norway	0.03	48.67	-0.01	-2.33	2.32
	(0.00)	(0.00)	(0.00)	(0.00)	(2709.22)
Poland	0.09	73.31	-0.01	-4.21	4.20
	(0.00)	(0.00)	(0.00)	(0.00)	(3149.72)
Sweden	0.03	14.94	-374.75	-0.10	-374.65
	(0.00)	(0.00)	(0.02)	(0.00)	(5655322.00)

Note that, in Equation 4, the asymmetry in AESTAR model is maintained if  $a_1 \neq a_2$  otherwise the system would be equal to an ESTAR model. Hence, the difference  $(a_1 - a_2)$  determines the degree of asymmetry. Accordingly, Akdogan (2016) develops a Wald test with the null hypothesis  $H_0 = a_1 - a_2 = 0$ , which could be used in addition to F-test suggested in Sollis (2009). The test statistics is:

$$F = (R\hat{\beta} - r)'[\hat{\sigma}^2 R\{\sum_t x_t x_t'\}^{-1} R']^{-1} (R\hat{\beta} - r) / m \quad (9)$$

where R is a 2 x 2 identity matrix;  $\hat{\beta} = [\hat{a}_1 - \hat{a}_2]'$ ;  $\hat{a}_1$  and  $\hat{a}_2$  are the least square estimates of  $a_1$  and  $a_2$  respectively.  $r = [0, 0]'$  and  $\hat{\sigma}^2$  is the least square estimate of  $\sigma^2$ .  $x_t = [u_{t-1}^3, u_{t-1}^4]'$  as in equation 8 and  $m=2$ . This test statistics is presented in the parenthesis below  $(a_1 - a_2)$  in last column of Table 4. The statistics is significant for all the countries except France. Hence, even though unit root test of Sollis (2009) would point out a nonlinear behaviour for France, the result of the Wald test that we suggest could not reject the null hypothesis that  $a_1$  is not different than  $a_2$ , denying an AESTAR type nonlinearity. Hence, we suggest the use of that Wald test as a further examination of the sign

<sup>11</sup> The sequential quadratic programming method of Gauss 14 is used to solve the nonlinear problem. For some parameters, estimations return the minimum value to comply with the given restrictions. Standard errors are close to zero for these cases.

asymmetry in AESTAR model. That being said we nevertheless present the results of the out-of-sample forecasting exercise results for France, in the next subsection.

The sign of the difference between parameters (a1-a2) provides us information about the characteristic of asymmetric correction. For example, the results for Belgium in Table 3 states that when food prices are below the mean the combined function in equation 4:

$$G(0.26, u^*_{t-1}) [S(70.51, u^*_{t-1})(-0.01) + \{1 - S(70.51, u^*_{t-1})\}(-1.44)] u^*_{t-1}$$

would change between -1.44 and 0. Otherwise, if the food price is above the mean, then the combined function changes between -0.01 and 0. Hence, a positive (a1-a2) would indicate that the correction towards the mean is stronger when the food prices are below the band as compared with the case when the food prices are above the band. The reason for this could be strong persistence in food price inflation. That is the case for all countries with a positive (a1-a2) in Table 3 (Belgium, Luxemburg, Netherlands, Norway and Poland). For Cyprus and Sweden the opposite case holds where the mean-reversion is stronger once the food prices are above the band compared to the case where the index falls below it. This result suggests that upwards deviations are relatively short-lived for these countries.

As a final remark on Table 3, since higher  $\theta_1$  implies a higher speed of transition, we can conclude that for Belgium, Netherlands and Poland the mean-reversion is relatively faster compared to the rest of the group.

### c. Out-of-Sample Forecasting Analysis

The next step in our analysis after estimation is testing the power of our nonlinear models in predicting food prices. We first divide the sample in two parts. Our training sample starts with the initial point of the series (January 1996) and ends at December 2014. Hence the forecasting sample covers 18 months between January 2015 and June 2016. Using our ESTAR and AESTAR estimations, we derive one, three, six, nine and twelve months-ahead forecasts. We carry out his task by extending the estimation period one month at a time until the end of the pseudo out-of-sample period.

The forecasting power of our models is compared with a simple benchmark model. In practice, usually a naïve random walk or a simple autoregressive (AR) model is used as a benchmark. Since our nonlinear models have an AR component, we chose a simple linear AR model as our benchmark. Table 4 displays the relative root mean square errors (RRMSE's) for ESTAR and AESTAR against the linear AR benchmark. The predictive power of the models for each forecast horizon is displayed in consecutive columns. A figure below 1 denotes improvement in prediction over AR model.

**Table 4****a) Out-of-Sample Exercise: ESTAR model RRMSE's**

	h=1	h=3	h=6	h=9	h=12
Cyprus	1.11	1.08	1.07	1.05	0.92
Iceland	1.05	0.99	0.90	0.90	0.87
Latvia	1.01	1.02	0.96	1.08	0.85
Lithuania	1.11	1.11	1.18	1.25	1.01
United Kingdom	0.99	1.00	2.32	0.74	1.02
<b>average</b>	<b>1.05</b>	<b>1.04</b>	<b>1.29</b>	<b>1.00</b>	<b>0.93</b>

**b) Out-of-Sample Exercise: AESTAR model RRMSE's**

	h=1	h=3	h=6	h=9	h=12
Belgium	4.41	5.93	3.99	3.67	3.25
Cyprus	1.45	1.40	1.35	1.33	1.17
France	6.27	5.95	6.37	6.89	4.69
Luxemburg	7.52	8.17	10.00	8.46	7.02
Netherlands	2.94	3.10	3.10	3.23	3.61
Norway	3.07	2.57	2.53	2.23	1.71
Poland	4.60	4.66	4.33	4.33	5.75
Sweden	4.14	4.38	3.94	3.97	3.00
<b>average</b>	<b>4.30</b>	<b>4.52</b>	<b>4.45</b>	<b>4.26</b>	<b>3.77</b>

A general look at the results indicated that the forecasting power of both models relative to a simple linear AR model is poor. While ESTAR model shows slight improvements in longer horizons for some countries (e.g. UK for nine months or Latvia for twelve months), the predictive power of AESTAR model compared to the linear AR models is very poor.

The previous literature provides mixed results on the issue of forecasting power of nonlinear models. Teräsvirta and Anderson (1992) argue that forecasting power of STAR models is not strong relative to the linear alternatives. Some studies argue that the predictive performance of nonlinear models depend on factors such as estimation and forecasting period (Clements and Smith, 2001) as well as the choice of the model or macroeconomic variables (Ferrara et al., 2015). McMillan and Wohar (2010) shows that the forecasting power of AESTAR model is better than linear models and ESTAR model for dividend–price ratio of the stock returns. Akdoğan (2015) shows that the ESTAR and AESTAR models show slight improvements over a naive random walk benchmark for inflation in the longer horizon for some inflation targeting countries while Akdoğan (2016) argues that the predictive performance of these models are poor for unemployment, once we chose the benchmark as a simple AR model. This paper also report poor performance of such models for predicting food prices. Hence, according to our results, the specification of the benchmark is also important in determining the forecasting power of nonlinear models. On the other hand, literature provides evidence of

improvements in forecasting power for nonlinear models over their linear alternatives for exchange rate (Killian and Taylor, 2003; Altavilla and De Grauwe, 2010)

#### **IV. Conclusion**

Stability in food prices is essential for many reasons such as food security or price stability in general. The high volatility 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 food price series. Our interest lies in assessing the linear or non-linear mean-reverting behaviour in unprocessed food prices for 24 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 for some countries the unprocessed food price series display nonlinear behaviour, in alternative smooth transition forms including asymmetries depending on the sign and size of the deviation from the long-run mean. We further perform a forecasting exercise using ESTAR and AESTAR models. Nonetheless, these models hardly display improvements over a simple AR model.

A number of research questions emerge for future analysis. First, what constitutes excessive food price volatility 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 volatility in 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 volatility 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, a further detailed exploration might study the main components of these aggregate indices; such as meat, dairy, sugar or grain. Such disaggregation would help the researchers in taking cognizance of the disproportionate impact of price volatility 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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