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Full Research Article

Agricultural and oil commodities: price transmission and market integration between US and Italy

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Abstract. Purpose of this article it to get some evidences of market interaction between United States and Italy using the time series analysis of spot prices spanning from January 1999 to May 2012 for crude oil and three ag-commodities: wheat, corn and soybean. These crops have been selected for their relevance in ag-commodity exchanges between US and Italy markets. The integration between US and Italy agricultural markets is hypothesized for the consistent volume of crop traded between these two countries while the price transmission is related to the leading price signals of the CBT (Chicago Board of Trade). The integration between oil and ag-commodity markets is suggested both by the large use of energy intensive inputs, (fertilizer, seed, machinery) in production of these ag-commodities, and their use in biofuel production. The results suggest: a) for US market the evidence of market integration between crude oil and US ag-commodities; b) for Italy the integration with US ag-commodity markets and less evidence of integration with the oil market. These results are valuable information both for the agents and policy makers contributing to improve the information accuracy to predict the price movements used by marketing operators for their strategies and policy makers to set up policies to re-establish conditions of market efficiency and allocate these ag-commodities in alternative market channels.

Keywords. Agricultural commodity prices, time series analysis, cointegration, price transmission, market integration.

JEL Codes. C22, Q11, Q13

1. Introduction

Since 2006, the biofuel market in the United States has established a link between the prices of crude oil and grains such as corn characterized by the co-evolution of ag-commodity prices (Abbott *et al.*, 2008). The massive production of energy, mainly liquid fuels, from agricultural commodities has continued to strengthen these links between agricultural and energy markets and defined a dominant feature of current conditions in the agricultural sector. A resulting trend has been noted as a stronger dependency between crude oil, gasoline, and ag-commodity prices (Tyner and Taheripour, 2008). Brazil pro-

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vides a useful example of a well-integrated, long-term agro-energy market with the oil and cane-bioenergy (ethanol and electricity) market integration to define an energy market in which oil and sugar cane prices exhibit strong comovement. There are other countries where these price links have become increasingly strong: the prices of wood pellets and, to a lesser extent, the wood chips in Austria have been following with a growing degree of correlation the prices for heating fuel in 2006 and 2007 (Schmidhuber, 2007).

Prices of oil and agricultural commodities sharply rose in 2007, peaking in the second half of this year for some products and in the first half of 2008 for others. The causes of price spikes during 2007-2008 include some which were exogenous to the agriculture (macroeconomic), growth of food demand by the BRIC countries, and speculation on the oil prices and other factors (Piot-Lepetit and M'Barek, 2011; OECD, 2008). Others causes are due to physiological changes in market conditions from period to period, natural shocks such as weather, pests or regulatory restrictions in domestic markets (FAO *et al.,* 2011). These events raise new questions concerning ag-commodity price movement. First, have new trends been established for ag-commodity prices? Second, to what extent have recent price shock been temporary or permanent? Third, how have these shocks and possibly persistent trends affected comovement of commodity prices? Fourth, how has the nature of unanticipated shocks changed? Price transmission depends on the market efficiency and it may have limited by a number of causes, at least in the short-run. For example, these conditions may have included supply availability to the final consumer, demand circumscribed by bottlenecks in the distribution, logistic problems in transportation, blending systems (E10), spatial arbitrage, and political constraints like the border measures and subsidies affecting the exchanges. Thus, the purpose of this paper is to contrinbute to the understanding of the transmission of oil prices to agricultural markets. We expect that our results will contribute to better understand the nexus of agricultural and energy markets and its consequences for trading relations between US and Italy. We choose to study Italy and US as settngs that provide the basis for a test of the hypothesis of market integration and price transmission (Yang and Leatham, 1999). During the 2009-2010 commercial campaign, Italy imported 60% of soft wheat, 87% of soybeans and 20% of maize from USA (Associazione Nazionale Cerealisti, 2011). The paper is organized as follows: section 2 reviews relevant literature; section 3 presents the time series methodology used; section 4 describes the price series used and provides a preliminary descriptive analysis of price correlations; section 5 presents results; and final section 6 provides conclusions and suggestions for future research.

2. Literature review

Market efficiency exists if price pass through between markets is complete such that they differ only by the transaction costs (Ardeni, 1989). Rational expectations and competitive storage theory supports the hypothesis that commodity stocks, expected prices and hauling costs are keydrives of commodity prices in equilibrium. Importantly, shortages can induce substantial price shocks, Helmberger and Weaver (1977) and Helmberger *et al.*(1982). It is widely acknowledged that the increased use of central commodity exchanges affects the extent and speed of transmission to market levels in response to leading price signals (Rapsomanikis *et al.*, 2006). Deaton and Laroque (1991) have used the storage model to show that prices are not normally distributed, because the stock-holding behaviour by risk-averse agents generates an autoregressive pattern which is much stronger than what can be explained by the storage activity of risk-neutral agents. This market behaviour induces shocks in supply and demand that are correlated over time. On the supply side, this correlation is also induced by correlated shocks while on the demand side persistence in demand for working stocks induces intertemporal correlations. However, prices jumps may be induced by speculative demand by producers in response to anticipated stock-outs, see Helmberger and Weaver (1982). The decline of the dollar value (Trostle, 2008) and the speculation stemming from increased futures market volume are further factors contributing to recent agricultural commodity price movement (Robles *et al.*, 2009). Further, supply side factors include relatively lower growth in agricultural production and yield, increases in energy prices that have induced increased farm production costs (Tyner and Taheripour, 2008; Sumner, 2009; von Braun *et al.*, 2008), and climatic events (Trostle, 2008). The indirect price transmission through energy feed stock substitutes (e.g. sugar) led to increased demand for land and other limited resources diverting them from other agricultural crops, reducing their supply and driving up their prices (Schmidhuber, 2007). The growth of the biofuels production is an important driver of recent corn and oilseed demand growth (Gilbert, 2010; Zhang *et al.*, 2010; Ciaian and d'Artis, 2011a,b). Biofuel policies, encouraging farmers to produce feed-stocks for biofuel, have increased the dependency between agricultural and energy prices (Yu *et al.* 2006; Campiche *et al.*, 2007; Zhang and Reed 2008; Gilbert, 2010; Gohin and Chantret, 2010; Nazlioglu, 2011).

Policy has also conditioned price transmission as trade restriction, import tariffs, export subsidies or taxes, and macroeconomic exchange rate policies have impacted the efficiency of arbitrage by insulating the domestic markets and hindering the transmission of price signals (Sarris, 2013). Esposti and Listorti (2013) investigate the role of the trade policy by analyzing the agricultural price transmission in presence of bubbles, using Italian and international weekly spot (cash) price over the years 2006–2010. They observe that the bubble has had only a slight impact on the price spreads and that the temporary trade policy measures, when effective, have limited this impact. These interventions are responsible for excess demand or supply schedules of domestic commodity markets possibly generating asymmetric price responses with nonlinear price adjustment (Quiroz and Soto, 1996; Sharma, 2002, Rapsomanikis *et al.*, 2006, 2011; Harri *et al.*, 2009; Gutierrez *et al.*, 2013). The market integration between oil and ag-commodities has been explored using econometrically estimated demand and supply models based on partial or computable general equilibrium models (Lapan and Moschini, 2012; de Gorter and Just, 2010; Hertel *et al.*, 2010). These models incorporate calibrated price elasticities and long-run assuptions to simulate dependence of agricultural commodity prices to oil shock prices. An alternative approach used to explore market efficiency is time series analysis to test market integration, price transmission, cointegration, asymmetric response, and causal nexus among markets (2011; Ciaian, 2011b; Gilbert, 2010; Gohin *et al.*, 2010; Goodwin, 1992; Granger and Lee, 1989; Harri *et al.*, 2009; Minot, 2011; Rapsomanikis *et al.*, 2006; Sarris, 2013; Zhang *et al.*, 2008). Here, we use time series analysis to test the hypothesis of market integration between US and Italy and in a broader sense to verify the efficiency of agricultural markets for some ag-commodities selected for their importance in trade

(Tomek and Myers, 1993; Rosa, 1999; Nazlioglu, 2011; Rosa and Vasciaveo, 2012; Esposti and Listorti, 2013; Gutierrez *et al.*, 2013).

Whether agricultural and food commodity prices are unjustifiably volatile and unrelated to the market fundamentals has been extensively considered by Balcombe (2009, 2013) and Gilbert (2010). Persistence of these effects on prices has been considered by Serra and Gil (2012), Algieri (2012), Chatellier (2011), Balcombe (2009), Listorti and Esposti (2012), Rosa and Vasciaveo (2012). In presence of excess volatility beyond that which can be accounted for by changes in market fundamentals, the prices may be driven by fad or speculative bubbles, and commodity prices may become inefficient signals for resource allocation (Gilbert and Morgan, 2010; Balcombe and Fraser, 2013). Time-varying volatility of commodity price series leads to autocorrelation patterns in the conditional variance of price innovations where the variance is conditional on an information set available at the time forecasts are being formed. Engle (1982) has termed this conditional heteroscedasticity and developed the autoregressive conditional changes in economic fundamentals. Time-varying volatility in commodity prices has the same general effect on statistical inference as any other form of heteroscedasticity causing a loss of efficiency and estimated standard errors may be biased (Engle, 1982). Excess kurtosis causes also problems whenever inference requires a particular distributional assumption on the disturbance terms. Although the normal distribution is typically chosen, the actual distribution of commodity prices appears to have fatter tails than the normal. This can be a particular problem in maximum likelihood estimation of commodity market models. (Myers, 1992).

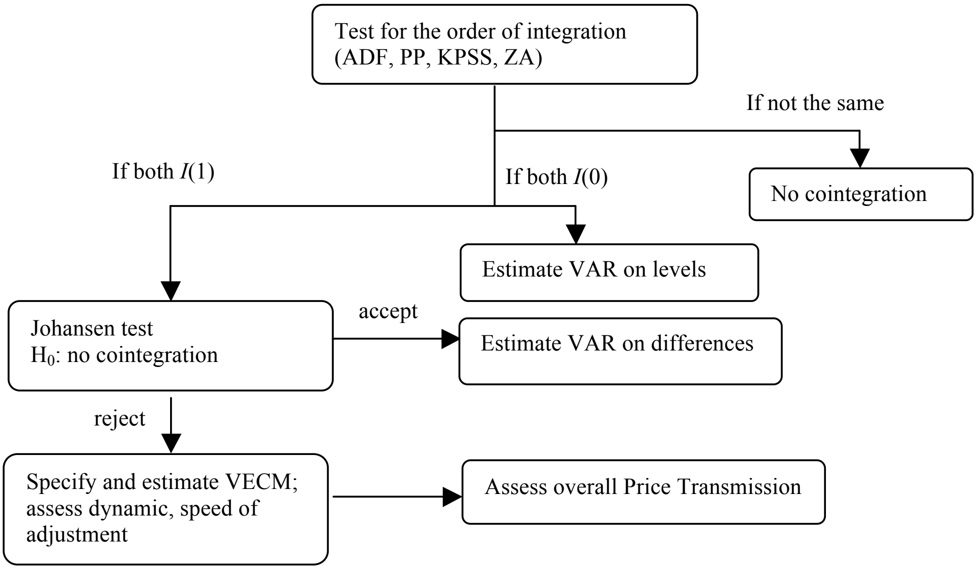
3. Methodology

Our analysis uses cointegration and vector error correction models to explore spatial market relationships and price transmission (Rapsomanikis *et al.*, 2006). The analysis is performed in the following three steps: 1) we determine whether univariate price series are nonstationary or $I(1)$ (if both price series are not $I(1)$, they cannot be cointegrated); 2) if they are both stationary or $I(0)$, we examine their dynamic interrelationship (Leucci *et al.*, 2013) with the vector autoregressive (VAR) model; 3) if the series are both $I(1)$, the null hypothesis that they are not cointegrated is tested with the Johansen procedure; 4) and if the results suggest evidence of long-run relationship between variables, we estimate vector error correction models (VECM). The scheme of this approach is reported in Figure 1.

3.1. Unit root test

The first step of the analysis is to test for stationarity and whether each series is integrated with the same order. We employ several unit root tests to consider robustness of our inference including: augmented Dickey-Fuller (1979) [ADF], Phillips-Perron [PP] (1988) and Kwaitkowski-Phillips-Schmidt-Shin [KPSS] (1992). These tests (except for KPSS) examine a null hypothesis of a unit root against the alternative of I(0) stationarity. Stationarity is the null hypothesis for KPSS. If the ADF statistic has a negative sign, as the absolute value increases the level of confidence for rejection of the hypothesis of unit root

Figure 1. Scheme of the price analysis to test the market integration and price transmission conditions.



Source: Own elaboration of the scheme proposed by Rapsomanikis *et al.* (2006).

is increased. The usefulness of ADF is limited in the presence of an explosive root (Balcombe and Fraser 2013). The ADF test follows from

$$y_t = \mu + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \quad y_{t-i} + \varepsilon_t \tag{1}$$

where μ is the constant, β is the coefficient on the time trend, k is the lag order of the autoregressive process, Dy_{t-i} is the lagged difference of y whose magnitude is measured by c and ε is the error. The unit root test is carried out under the null hypothesis $\alpha = 0$ against the alternative hypothesis of $\alpha < 0$; nonstationarity is rejected when α is significantly different from 1.

A common problem with conventional unit root tests is that they do not allow for any break in the data generation process. If a structural break is hypothesized, the conventional ADF test is biased toward the acceptance of the null resulting in a dramatic loss of power. Further, the ADF allows for higher-order autoregressive processes including lags of the order k that have to be pre-determined. Assuming the break time to be exogenous, Perron (1989) suggests that the power to reject a unit root decreases when the stationary alternative is true and the structural break is ignored. Following Perron's characterization of the form of structural break, Zivot and Andrews (ZA, 1992) formulate three different characterizations of the trend break: i) model A, "*the crash model*", allows the break in the intercept; ii) model B, "*the changing growth model*", allows for a one-time change in the slope of the trend function with the two segments joined at the break point; and iii) model C, "*the mixed model*", combines simultaneously the one-time changes in the level with the slope of the trend function of the series[2]. The aim of this procedure is to sequentially examine evidence of breakpoint candidates and select the one that gives most weight to the trend stationary alternative. Hence, to test for a unit root against the alternative of a one-time structural break, Zivot and Andrews propose the following regression equations (derived from equation 1) corresponding to the three cases noted above:

$$y_t = \mu + \beta t + \alpha y_{t-1} + \gamma DU_t + \sum_{i=1}^{k} c_i \quad y_{t-i} + \varepsilon_t \qquad \text{(Model A)}$$

$$y_t = \mu + \beta t + \alpha y_{t-1} + \theta DT_t + \sum_{i=1}^{k} c_i \quad y_{t-i} + \varepsilon_t \qquad \text{(Model B)}$$

$$y_t = \mu + \beta t + \alpha y_{t-1} + \theta DT_t + \gamma DU_t + \sum_{i=1}^{k} c_i \quad y_{t-i} + \varepsilon_t \qquad \text{(Model C)}$$

where DU_t is an indicator dummy variable for a mean shift occurring at each possible break-date while DT_t is the corresponding trend shift variable. The ZA unit root test is an endogenous structural break test with unknown timing in the individual series that uses the full sample and different dummy variables for each plausible break date. The break time is selected where the t-statistic from the ADF test of unit root is at a minimum (most negative), then a break date is chosen where the evidence is least favourable for the unit root null. The null hypothesis is that the series is integrated without an exogenous structural break against the alternative that the series can be represented by a trend-stationary process with only one break point occurring at some unknown time. The ZA test is a variation of PP's original test with the endogenous implementation of structural breaks in the analysis: the date of the break is determined with the t-statistics test of the unit root, with respect to the criteria of minimum values. The ZA test regards every point as a potential break-date and runs a regression for every possible break-date sequentially.

3.2. Cointegration analysis: the Johansen test

Cointegration analysis examines whether two series are linked to form an equilibrium relationship. The intuition of cointegration is that two price series cannot evolve in opposite directions for very long time if they are cointegrated. This condition is examined by estimation of the static regression between $I(1)$ variables:

$$y_t = \mu + \alpha x_t + \varepsilon_t \qquad (2)$$

[2] For the three models, Zivot and Andrews estimate the testing equation by allowing the break to take place beginning successively in the second, third, fourth, and so on, observation, up to observation T - l, where T stands for the total sample size used in the estimation and l are the lags. The alternative specifications are estimated by OLS, and the length of the lag (k) for the difference terms is determined by starting at k = 8, and working backwards until significant values are identified. The estimate of the breakpoint is that particular observation corresponding to the minimum t-value for the one period lagged term, for each model A, B, and C. In order to test the unit root hypothesis, this minimum t-value is compared with a set of asymptotic critical values from the work of Zivot and Andrews (1992).

where x_t is a vector of independent variables. The system is cointegrated if the errors ε_t are $I(0)$. In this case, equation (2) may be interpreted as a long-run equilibrium condition of the process $y(t)$. The Johansen cointegration test uses the vector autoregressive (VAR) model with k lags assuming the variables are I (1) written in error-correction form (Johansen, 1995). To determine the presence of cointegration between variables, the lag length (k) is determined with the Schwarz Bayesian Criterion (SBC or BIC test, Schwarz, 1978) and then the cointegration rank (r) is estimated.

3.3. Gregory Hansen test

Gregory and Hansen (1996) propose cointegration tests which are an extension of the Zivot and Andrew (1992) unit root tests to incorporate a single structural break in the underlying cointegrating relationship. The GH test extends the ADF^*, Z_t^* and Z_α^* type tests designed to test the null of no cointegration against the alternative of cointegration in presence of a single structural break. These authors consider three variations of equation (2) that includes dummies for the structural change:

Model C: Level Shift:

$$y_t = \mu + \theta DU_t + \alpha x_t + \varepsilon_t \quad (3a)$$

Model C/T: Level Shift with Trend

$$y_t = \mu + \theta DU_t + \beta t + \alpha x_t + \varepsilon_t \quad (3b)$$

Model C/S: Regime Shift (Intercept and Slope coefficients change)

$$y_t = \mu + \theta DU_t + \alpha_1 x_t + \alpha_2 DU_t x_t + \varepsilon_t \quad (3c)$$

where t is time subscript, ε is an error term and DU is a dummy variable.

Model C entails a level shift in the equilibrium relationship, model C/T adds a trend component to the previous model whilst model C/S deals with the regime shift by adding a change in the slope coefficients. The structural change is endogenously determined by the smallest value (the largest negative value) of the cointegration test statistics across all possible break points.

4. Data and descriptive analysis

To perform the empirical analysis, we use weekly spot prices[3] of three ag-commodities and the oil prices for the period spanning from 1999 to 2012; this frequency has been used to capture more accurately the price movements and linkages (Nazlioglu, 2011). Soft wheat, maize and soybeans are selected for their importance in ag-commodity trade between US and Italy: wheat is highly energy intensive and is a key product for human nutrition while corn and soybean are the most important ag-commodities for animal feeding and biofuel. Table 1 reports the list of variables used in the analysis.

[3] Spot prices are used because most of the transactions in Italy are made in these markets. For more details see Rosa and Vasciaveo (2012)

Table 1. Description of ag-commodity price series.

Variable		Description	Source
Italian	corn price (cit)	Weekly average of spot prices in €/ton of national hybrid corn-market at the origin (cit)	DATIMA provided by ISMEA[1]
	soybean price (sit)	Weekly average of spot prices in €/ton of soybeans with 14% of moisture--market at the origin (sit)	DATIMA provided by ISMEA
	wheat price (wit)	Weekly average of spot prices in €/ton of good mercantile wheat--market at the origin (wit)	DATIMA provided by ISMEA
US	corn price (cus)	Weekly average of spot prices converted in €/ton of US yellow no. 2 corn at the Gulf of Mexico (cus)	FAO International Commodity Price Database
	soybean price (sus)	Weekly average of spot prices converted in €/ton of US no. 1 yellow soybean at the Gulf of Mexico (sus)	FAO International Commodity Price Database
	wheat price (wus)	Weekly average of spot prices converted in €/ton of US no. 2 soft red winter wheat at the Gulf of Mexico (wus)	FAO International Commodity Price Database
Oil price	oil	Weekly spot prices of Brent crude oil converted in €/barrel (oil)	US Energy Information Administration (EIA, 2012)

DATIMA is a collection of statistical databases including Italian agricultural market data and foreign trade; ISMEA is the Italian agri-food market Institute

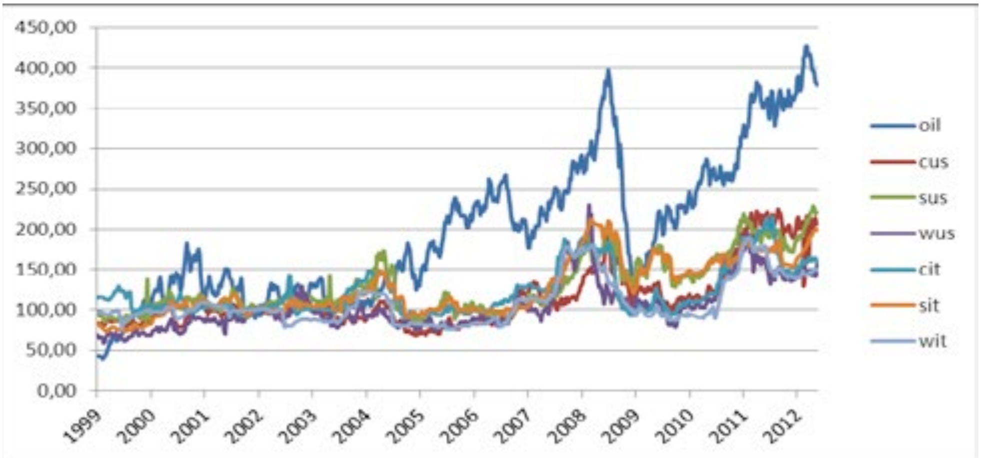
To be comparable, the US agricultural and oil commodities price series quoted in $ are converted into euro currency, using the official $/€ exchange rate[4] and converted to natural logarithms. Visual inspection of the price series reported in Figure 2 suggests a nonlinear trend component exists for each of the series. Figure 2 also suggests a relatively steady price period existed during 2005, followed by wider fluctuations to the end of 2008, and wider fluctuations in the final stage for all commodities prices. The wider oil price variability does not seem to affect the fluctuation of the ag-commodity prices. These observations motivate the need to examine the existence of structured breaks that define sub-samples to examine better the effect of volatility.

4.1. Testing for presence of bubbles

Figure 2 also suggests the possible presence of bubbles. During the period 2006-10, sub-periods of explosive price are apparent as also noted by Huchet-Bourdon (2011). Bubbles have been noted as occurring in 2006 when levels of agricultural and food prices increased sharply followed by a collapse, as well as between 2006 and 2008, 2008-2010 and more recently in autumn 2012. (Phillips, P.C.B., Shi S., Yu J., 2012). A number of

[4] Available at http://www.statistics.dnb.nl/index.cgi?lang=uk&todo=Koersen

Figure 2. Index of current prices of some agri commodities and oil prices



Source: own elaborations. cit, wit, sit, cus, wus, sus: €/ton; for oil: €/barrel; Jan 04, 2002= 100.

tests have been used to identify the sharp increases and declines in prices also known as explosive bubbles. We followed Phillips, Shi and Yu (2012), PSY hereafter, who developed a method to test for explosive behavior and date the origin and collapse of bubbles. This method is used to check for presence of multiple bubbles of the PCB[5] type in a sample data. (Phillips, Shi and Yu, 2012; Gilbert and Morgan, 2010). We apply the more recent generalized sup-augmented Dickey-Fuller (GSADF) test proposed by PSY, for explosive bubbles with variable windows widths in the recursive regression:

$$Y_t = \alpha r_1, r_2 + \beta r_1, r_2 Y_{t-1} + \sum_{i=1} \Phi r_1 r_2 \quad Y_{t-1} + \varepsilon_t \qquad e_\tau \sim N\ (0,s^2) \tag{4}$$

Here the null hypothesis of nonstationarity (H_0: $b_{r1,r2} = 0$) is tested against the alternative hypothesis H*1*: $b_{r1,r2} > 0$ which implies explosive behaviour. Our results reveal evidence of bubble behaviour for wheat, rice soybean oil and rapeseed oil price series during the first month of 2008. Beyond fundamentals, the GSDAF test does not provide sufficient evidence to infer whether these bubbles are the result of a trend and may persist in the ag-commodity market. We have tested with the PSY test the series used for this analysis and results are reported in Table 2 for different length of time series and window widths.

The analysis provides insights to price behaviour during the examined periods and their consequences for the analysis. The results of Table 2 do not support the hypothesis that bubbles occurred during the sample period with an exception for *sit* with window width 0.1, however for window width of 0.4 the values are substanitially below the critical threshold at 99 and 95% critical values. The period 2006:1-2008:52 is also examined for bubbles as past literature reports more evidence of price volatility during this period (Gil-

[5] PCB is the acronym for price collapsing speculative bubbles that are nonlinear processes (Evans,1991 explosive during the phase of bubble eruption, but they may be stationary over the whole sample period.

Table 2. Results of the GSADF recursive test with one lag.

Series	Lag	Window	Nr of observations	Test statistics GSADF	Critical values		
					99	95	90
cus	1	0.1	168	1.1787	2.9822	2.2381	2.0277
cus	1	0.4	168	1.1788	1.9957	1.4001	1.1322
cit	1	0.1	168	2.0310	2.9822	2.2381	2.0277
cit	1	0.4	168	-0.0909	1.9957	1.4001	1.1322
sit	1	0.1	168	3.5970	2.9822	2.2328	2.0278
sit	1	0.4	168	0.2132	1.9975	1.4001	1.1322
sus	1	0.1	168	1.8955	2.9822	2.2328	2.0278
sus	1	0.4	168	-0.1245	1.9957	1.4002	1.1392
wit	1	0.1	168	2.1124	2.9822	2.2328	2.0278
wit	1	0.4	168	-0.5607	1.9957	1.4001	1.1392
wus	1	0.1	168	0.9183	2.9822	2.2328	2.0277
wus	1	0.4	168	0.0573	1.9957	1.4002	1.1392
oil	1	0.1	168	1.2566	2.9822	2.2328	2.0278
oil	1	0.4	168	0.1011	1.9976	1.4002	1.1392

Time series 1999:1-2012:52.
Shaded values are above the critical values.

Table 3. Results of GSADF recursive test with one lag.

Series	Lag	Window	Nr of observations	Test statistics GSADF	Critical values		
					99	95	90
cus	1	0.1	156	2.5989	3.0681	2.2972	1.9947
cus	1	0.4	156	2.5989	2.0938	1.4633	1.1652
cit	1	0.1	156	2.1584	3.0681	2.2972	1.9947
cit	1	0.4	156	3.4991	2.1114	1.4575	1.1524
sit	1	0.1	156	2.3889	3.1413	2.2660	1.9882
sit	1	0.4	156	2.3889	2.0939	1.4633	1.1652
sus	1	0.1	156	1.9957	3.1413	2.2660	1.9882
sus	1	0.4	156	2.1787	3.0682	2.2972	1.9947
wit	1	0.1	156	4.7593	3.1413	2.2660	1.9882
wit	1	0.4	156	4.7593	2.0939	1.4633	1.1652
wus	1	0.1	156	3.1313	3.1413	2.2660	1.9882
wus	1	0.4	156	2.1584	2.0938	1.4633	1.1652
oil	1	0.1	156	3.0871	3.1413	2.2660	1.9882
oil	1	0.4	156	3.0871	3.0682	2.2972	1.9947

Time series 2006:1-2008:52.
Shaded values are above the critical values.

bert, 2010, Rosa and Vasciaveo, 2012). Results are reported in Table 3. For the series cus, cit, sit the test values are above the critical values, using the windows width 0.4 but below critical values using the windows width 0.1; a possible explanation is that the smaller window width includes price values less volatile compared to the larger window. For wheat, the results are above the critical values for both window widths. These results are more difficult to explain because in contrast with wus and other ag-commodity in Italy (Areal *et al.*, 2013), the 2007-2008 US wheat market experienced reduced stock levels. Reduced production levels, in conjunction with very low carryover stocks, resulted in an extremely tight global market and is likely to have affected the expectations of market operators in Italy. Another possible explanation is the interaction between spot and future markets. Given an high share of wheat open interest held by noncommercial traders in an already tight market, the demand for long-term wheat future contracts may have affected spot prices and generated bubbles due to strengthening inventory demand. The tests performed by Areal *et al.* (2013) have revealed weak presence of multiple bubbles in the food prices finding that when present, bubbles have been quite short, continuing between two and fourth months before collapsing.

4.2. Stationary and structural break tests

We next consider the order of integration and testing the stationary condition[6] with the unit root test for levels and first differences. A number of tests are used with results reported in Table 4. We find all variables are integrated of first order $I(1)$[7]

Table 5 reports the results of the Zivot-Andrews test with one break. Minimum ZA statistics for the levels of the variables reject the hypothesis of structured breaks implying the evidence of the unit root tests may be accepted with the exception of oil and sus. Allowing for the identified breaks and a deterministic trend for these products, the null hypothesis of unit root process is rejected. The test is performed in three versions reported in previous section 3. A structural break is found in the US soybeans series, the estimated date is July 2004 (week 29) fitted with drift (model A) and a change in the trend slope and drift (model C). The oil series is stationary with a break in October 2008 (week 40) and change in trend slope and drift. A possible explanation of the 2004 structural break is the massive growth in biofuel production in the US starting with 2004. The successive 2008 break in the oil series corresponds to the oil price peak.

The other price series are found to be $I(1)$ confirming the results of traditional unit root tests. While sus and oil price series are stationary in model C, this condition is not so evident in model A (only for oil) or in model B; for this reason it is conservatively assumed that all the variables are integrated of order one I*(1)*.

4.3. Preliminary evidence of comovement among Italian and US ag-commodity prices

The graphic evidence of comovement in levels for the historical price series (Figure 2) suggests strong comovement with a moderate deviations from cyclical long-run move-

[6] This condition implies that the mean, variance and autocorrelation of the series do not change over time;

[7] The differences of the alternative tests used in this analysis are not contradictory about stationary condition.

Table 4. Unit root test results.

		Levels			First differences		
		ADF	PP	KPSS	ADF	PP	KPSS
Intercept	cus	-0.31	-1.28	1.93*	-13.84*	-33.22*	0.14
	sus	-1.42	-1.35	2.20*	-30.21*	-30.99*	0.08
	wus	-2.23	-2.15	1.79*	-28.81*	-28.75*	0.03
	cit	-1.77	-2.08	1.18*	-16.35*	-16.41*	0.05
	sit	-1.28	-1.02	2.22*	-14.74*	-21.10*	0.06
	wit	-1.48	-1.77	1.03*	-18.51*	-19.42*	0.06
	oil	-1.23	-1.17	2.48*	-21.42*	-21.42*	0.04
Trend & intercept	cus	-1.59	-2.91	0.44*	-13.88*	-33.66*	0.05
	sus	-3.06	-3.08	0.36*	-30.20*	-31.02*	0.03
	wus	-3.36	-3.29^	0.12^	-28.80*	-28.73*	0.02
	cit	-2.47	-2.76	0.12^	-16.35*	-16.37*	0.03
	sit	-2.73	-2.46	0.17°	-14.73*	-21.09*	0.04
	wit	-2.07	-2.36	0.16°	-18.50*	-19.41*	0.04
	oil	-2.78	-2.71	0.14^	-21.41*	-21.41*	0.04

Schwarz Information Criterion to determine the optimal lags for ADF test; the bandwidth for PP and KPSS tests is selected with Newey-West using Bartlett kernel (by default). */°/^ denote statistical significance at 1, 5 and 10% respectively.

Table 5. Zivot Andrews one break test.

	Model A Change in drift	Model B Change in trend	Model C Change in drift and trend
cus	-3.63	-3.10	-3.42
sus	-4.98** (2004: w29)	-4.15	-5.48*** (2004: w29)
wus	-3.78	-3.50	-3.87
cit	-3.20	-2.83	-3.82
sit	-4.03	-3.02	-4.03
wit	-3.18	-2.81	-3.26
oil	-3.27	-3.30	-5.16** (2008: w40)

	Critical value		
	1%	5%	10%
Model A	-5.34	-4.80	-4.58
Model B	-4.93	-4.42	-4.11
Model C	-5.57	-5.08	-4.82

The asymptotic critical value for Zivot and Andrews (1992) test at different levels of significance

***/** denote statistical significance at 1% and 5% respectively; break date in brackets.

ments, non-linear trend components, and wider range fluctuation in latest periods. Oil price patterns could have affected the agricultural markets during the period (1999-2012) and changed the price dynamics generated by market fundamentals (Headey and Fan, 2008; OECD, 2008). The price correlation between two variables, here given by the prices of ag-commodities is measured linearly with the Pearson correlation coefficient (r). Given nonstationarity of the underlying series of price levels, we examine correlation of stationary first differences.[8] The correlation coefficients between the oil price and each ag-commodity price are computed over the whole sample and for sub- samples with critical 5% values; the results are reported in Table 6.

For the entire period (1999-2012), we find the Pearson correlation in price differences are small with values that vary in the range between 0.06 (dwit- dsus) and 0.37 (dcus-dwus) with two correlations below the critical 5% value: dsus-dcit and dsus-dwit. These results are consistent with the hypothesis of innovations in prices commove, though such comovement is small in magnitude. We also find that the innovations in Italian ag-commodity prices are not influenced by those of the oil prices, though we find evidence of comovement in innovations in oil and US ag-commodity prices. For the subsample period 1999-2004, we find the Pearson correlation values vary in the range between -0.03 (dwus and dcit) and 0.53 (dwus-dcus). However, compared to the full sample period we find more coefficients (9) are below the critical value indicating no comovement. Across countries, we find evidence of comovement between dwit and dcit and the US ag commodity prices, as well as with oil prices. For Italian prices, we find that only for the (dcit, dwit) pair can we reject the hypothesis of correlation. For the period 2004-2008, the estimated Pearson correlation values vary in the range between -0.02 (dcus-dwit) and 0.40 (dsus-dcus). Critical values indicate there are seven correlation coefficients below the critical value, in each case for pairs of Italian and US prices (dcit and dwit with US ag-commodity prices). For the period 2008-2012, we find the Pearson correlation values vary in the range between 0.11 (dsus and dwit) and 0.40 (dwit-dcit; dwus-dcus; doil-dsit). With respect to evidence of comovement across oil prices and commodity prices for entire period and sub-periods we find evidence of weak comovement of oil and ag-commodity prices that is weakest during the sub-period 1999-2004 and stronger in the later sub-periods; correlation is smaller with each of the Italian ag commodities compared to US. The analysis suggests that for sub-periods energy and agricultural market price comovement seems to become stronger in more recent periods. The general conclusion is that the US ag commodity price innovations appear to be more correlated with those of oil prices while little evidence of such correlation is found between Italian ag commodity price innovations and oil price innovations.

4.4 Market integration

The cointegration analysis is used to examine the comovements between oil and agri-commodity prices (Johansen and Juselius, 1990). Some of the series checked for unit roots are found to be stationary with a breaking trend, then the Johansen and Gregory- Hansen (GH) tests are used to check for the presence of cointegration for all pairwise price series that accounts for a break in the cointegration relationship.

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[8] In a bivariate time series characterized by nonstationarity, correlation of in nonstationary levels is meaningless as by definition the series are not generated by population data generation processes that are invariant with respect to time. In the absence of a population counterpart, correlation would result in spurious inference (Johansen, 1989).

Table 6. Pearson correlation coefficients for first difference series of ag-commodity prices.

Series 99/01/08 - 12/05/25; two tail critical value 5% = 0,0742*; n = 699

d_cit	d_wit	d_sit	d_cus	d_wus	d_sus	d_oil	
1.0000	0.2597	0.2037	0.1153	0.1420	0.0706	0.1130	d_cit
	1.0000	0.1544	0.0827	0.1907	0.0648	0.0751	d_wit
		1.0000	0.2329	0.1603	0.2513	0.2096	d_sit
			1.0000	0.3655	0.2150	0.2135	d_cus
				1.0000	0.2186	0.2149	d_wus
					1.0000	0.2151	d_sus
						1.0000	d_oil

Series 99/01/08 - 04/07/16; two tail critical value 5% = 0,1154*; n = 289

d_cit	d_wit	d_sit	d_cus	d_wus	d_sus	d_oil	
1.0000	-0.0451	0.2196	-0.0258	-0.0333	0.0346	-0.0831	d_cit
	1.0000	0.1664	0.0463	0.0921	0.0380	0.0411	d_wit
		1.0000	0.1564	0.1312	0.1860	0.0508	d_sit
			1.0000	0.5333	0.3176	0.1840	d_cus
				1.0000	0.2001	0.2180	d_wus
					1.0000	0.1676	d_sus
						1.0000	d_oil

Series 04/07/23 - 08/10/03; two tail Critical value 5% = 0,1323*; n = 220.

d_cit	d_wit	d_sit	d_cus	d_wus	d_sus	d_oil	
1.0000	0.4456	0.1257	-0.0064	0.1300	0.0769	0.1330	d_cit
	1.0000	0.0302	-0.0227	0.1644	0.0577	-0.0529	d_wit
		1.0000	0.2964	0.1082	0.2573	0.1080	d_sit
			1.0000	0.3518	0.3994	0.2014	d_cus
				1.0000	0.3007	0.1356	d_wus
					1.0000	0.2280	d_sus
						1.0000	d_oil

Series 08/10/10 - 12/05/25; two tail Critical value 5% = 0,1424*; n = 190.

d_cit	d_wit	d_sit	d_cus	d_wus	d_sus	d_oil	
1.0000	0.3850	0.2688	0.2402	0.3013	0.1207	0.2538	d_cit
	1.0000	0.2399	0.1363	0.2978	0.1076	0.1949	d_wit
		1.0000	0.2550	0.2356	0.3585	0.3905	d_sit
			1.0000	0.3943	0.1662	0.2519	d_cus
				1.0000	0.1970	0.3013	d_wus
					1.0000	0.2893	d_sus
						1.0000	d_oil

* Shaded values are below the critical values.
Source: Author computation.

Table 7. Johansen trace test for cointegration in price level: 1999:w1-2012:w21.

	cus	sus	wus	cit	sit	wit
cus						
sus	18.87					
wus	27.79**	26.44**				
cit	26.60**	27.15**	30.74**			
sit	12.74	25.91**	32.52***	29.56**		
wit	23.93*	29.33**	42.17***	31.24***	28.69**	
oil	15.36	17.73	21.16	23.11	16.38	19.50

The critical values are 31.15, 25.87 and 23.34 for 1%, 5% and 10% respectively (MacKinnon-Haug-Michelis, 1999, critical values). ***, ** and *denote statistical significance at 1%, 5% and 10% level of significance, respectively. Null is no cointegration.

The results of the trace test reported on Table 7 indicate that all ag-commodity prices are pair wise cointegrated with the exception of sus and cus and sit and cus. These results suggest that Italian and US agricultural markets are integrated while there is not statistical evidence that oil market affects the ag-commodity markets in US or in Italy. An absence of cointegration is found for cus and sus. This is consistent with a dominance of ethanol demand for corn that drives a wedge between the two markets. Table 8 reports the results of the cointegration test for the first sub-period and shows the null hypothesis of no cointegration is rejected for US corn and wheat with Italian commodity prices. US soybeans is cointegrated with the Italian ag-commodities, confirming the results obtained during the entire period of observation and is also cointegrated with the oil prices; US corn is cointegrated with US soybean and wheat. The Italian ag-commodity markets appear to be cointegrated with the US soybean market, and a significant cointegration exists between IT soybeans and oil prices.

Table 8. Johansen trace test for cointegration: period 1999:w1-2004:w29.

	cus	sus	wus	cit	sit	wit
cus						
sus	26.10**					
wus	30.80**	25.95**				
cit	21.47	34.33***	20.40			
sit	17.32	39.44***	15.48	18.18		
wit	17.24	34.24***	17.60	23.18	17.75	
oil	21.37	30.75**	18.89	22.44	24.10*	17.03

***, ** and *denote statistical significance at 1%, 5% and 10% level of significance, respectively.

The results reported in Table 9 for the sub-period (2004-2008) suggest a different market condition: Italian corn, soybeans and wheat prices are cointegrated with their corresponding Italian prices though not in all cases with US prices and never with the oil prices. The pairwise cointegrations between soybean and corn in the Italian markets suggest also that the price movements of these two major ag-commodity markets are moving together. The first observation inherent in this period is that US and Italian ag-commodity prices move together and US commodity prices are not in general cointegrated with Italian ag-commodity prices. We also find an absence of cointegration between US commodity price pairs.

Table 9. Johansen trace test for cointegration: period 2004:w30-2008:w40.

	cus	sus	wus	cit	sit	wit
cus						
sus	15.60					
wus	19.70	12.93				
cit	21.70*	23.18	12.69			
sit	13.71	24.74*	34.77***	31.06**		
wit	23.95*	24.37*	30.09	16.53	30.44**	
oil	15.32	19.03	12.66	10.85	15.24	17.36

***, ** and *denote statistical significance at 1%, 5% and 10% level of significance, respectively.

For the last sub-period presented in Table 10, the situation is consistently different. The null hypothesis of absence of cointegration with oil is rejected for all products except cit. These findings are consistent with the increasing use of ag-commodities in biofuel production that has generated more interdependence between oil and ag-commodity markets. We also find US prices to be cointegrated except between soy and wheat. US corn is cointegrated with Italian corn and with wheat while Italian soy appears not to be cointegrated with US corn prices. For each product, US and Italian markets appear cointegrated. No evidence of cointegration across Italian product prices is found.

Table 10. Johansen trace test for cointegration: 2008:w41-2012:w21.

	cus	sus	wus	cit	sit	wit
Cus						
Sus	27.75**					
Wus	29.53**	16.70				
Cit	24.50*	9.51	14.86			
Sit	18.05	47.26***	11.59	9.20		
Wit	42.80***	19.27	62.59***	13.35	9.62	
Oil	28.80**	36.68***	36.22***	20.46	32.39***	35.91***

***, ** and *denote statistical significance at 1%, 5% and 10% level of significance, respectively.

These results are supported by the observations of other authors. Campiche *et al.* (2007) found that while there is no evidence of cointegration among the variables for the period 2003-2005, corn and soybean prices are cointegrated with crude oil prices in the next period 2006-2007. Harri *et al.* (2009) found robust evidence of cointegration between crude oil and corn, soybeans starting in April 2006. Nazlioglu (2011) examined the cointegration between oil and three key ag-commodity prices and found evidence of corn and soybean price cointegration with the oil prices during the period 2008-2010. Structural break timing has been determined *a priori* in the previous papers. Here, we examine evidence of breaks within the context of our cointegration models, using Gregory-Hansen tests based on equations (3a – 3c) that allow for identification of structural break for the entire period 1999-2012. Results are reported in Table 11.

Table 11. G-H cointegration test with one structural break[9] for US ag-commodities and oil: 1999:w1-2012:w21.

		cus-oil	sus-oil	wus-oil
ADF^*	C	-3.45	-4.21	-4.23
	C/T	-3.84	-5.38** (2004: w34)	-4.26
	C/S	-4.06	-4.94* (2008: w10)	-4.65
Z_t^*	C	-4.69** (2010: w19)	-4.44* (2007: w39)	-3.81
	C/T	-5.52*** (2004: w22)	-5.72*** (2004: w33)	-3.85
	C/S	-5.72*** (2004: w37)	-5.20** (2007: w39)	-4.03
Z_α^*	C	-42.20** (2010: w19)	-40.26** (2007: w39)	-28.61
	C/T	-56.96** (2004: w22)	-61.15*** (2004: w33)	-28.91
	C/S	-62.39*** (2004: w37)	-52.52** (2007: w39)	-31.49

***/**/* statistical significance at 1%, 5% and 10% level of significance, respectively; break dates in brackets.

For oil and cus price, the ADF^* test did not reject the null hypothesis of no cointegration with model in the versions C, C/T and C/S whereas Z_t^* and Z_α^* type test results suggest the rejection of the null for each of the three models. Significance of structural breaks were found for May and September 2004 (week 22 and 37) and May 2010 (week 19). For soybeans and oil, the three tests do not support the rejection of the null hypothesis of cointegration and structural break evidence is found for August 2004 (week 33); besides, Z_t^* and Z_α^* fail to reject the null in the regime shift model with a break in July 2007 (week 39). For the long-run relationship between wheat and brent prices, no evidence was found

[9] Model C: Level shift, Model C/T: level shift with trend, Model C/S: Regime shift. Null hypothesis: no cointegration. For ADF^* and Z_t^* tests, critical values in Model C are: -5.13 at 1%, -4.61 at 5% and -4.34 at 10%; in Model C/T:-5.45 at 1%, -4.99 at 5% and -4.72 at 10%; in Model C/S: -5.47 at 1%, -4.95 at 5% and -4.68 at 10%. Critical values for Z_α^* test are -50.07, 40.48, -36.19 respectively at 1, 5 and 10% in Model C; -57.28, -47.96 and –43.22 at 1, 5 and 10% in Model C/T; -57.17, -47.04 and -41.85 at 1, 5 and 10% in Model C/S. The optimal lag length for ADF^* test was selected by Akaike information criterion.

for cointegration; a possible explanation is the wheat prices are less dependent on energy prices because only a limited quantity is used in the ethanol production.

Table 12. G-H cointegration test with one structural break for IT ag-commodities and oil.

		cit-oil	sit-oil	wit-oil
ADF*	C	-3.77	-4.11	-3.75
	C/T	-4.01	-4.13	-3.84
	C/S	-4.27	-4.20	-4.27
Zt*	C	-3.58	-3.69	-3.36
	C/T	-3.58	-3.71	-3.33
	C/S	-3.73	-3.84	-3.59
Zα*	C	-24.72	-27.95	-21.93
	C/T	-24.63	-28.21	-22.15
	C/S	-28.14	-29.42	-25.67

***/**/* statistical significance at 1%, 5% and 10% level of significance, respectively; break dates in brackets.

The results of Gregory Hansen tests reported in Table 12 do not in any case support the inference of cointegration between the crude oil and the Italian ag-commodity prices. The results of Table 13 suggest the cointegration between the Italy and US ag-commodity markets. These findings are consistent with those obtained by running the cointegration test without structural breaks. Results appear to be more robust for wheat and soybean commodities (confirmed by all the three tests). For corn, evidence of cointegration follows only from the Z_t^* and Z_α^* tests.

Table 13. Cointegration test with one structural break between Italian and US ag-commodities.

		cit-cus	sit-sus	wit-wus
ADF*	C	-3.89	-4.83** (2010: w19)	-5.29** (2001: w34)
	C/T	-4.21	-4.96* (2010: w19)	-5.26** (2001: w14)
	C/S	-4.43	-6.11*** (2008: w29)	-5.56*** (2004: w28)
Z_t^*	C	-4.81** (2008: w31)	-7.11*** (2010: w19)	-5.70*** (2001: w19)
	C/T	-5.10** (2003: w27)	-7.04*** (2010: w19)	-5.71*** (2001: w19)
	C/S	-4.86** (2008: w26)	-7.63*** (2010: w19)	-5.98*** (2004: w29)
Z_α^*	C	-45.21** (2008: w31)	-90.15*** (2010: w19)	-61.46*** (2001: w19)
	C/T	-50.52** (2003: w27)	-88.94*** (2010: w19)	-61.49*** (2001: w19)
	C/S	-46.00* (2008: w26)	-103.52*** (2010: w19)	-67.48*** (2004: w29)

***/**/* denote statistical significance at 1%, 5% and 10% level of significance, respectively. Break dates in brackets.

5. Price transmission

Market imperfections may interfere with the price adjustment process in many ways: asymmetric response, speed adjustment, biased information, decisions of storage and inventory holding, policy intervention and others (Granger *et al.*, 1989). Market conditions determine price transmission. If the condition of market efficiency holds, the price change in one market is instantaneously and completely transmitted to the related market and the price difference will reflect only the transfer cost (Fama, 1970; Goodwin, 1992). The cointegration condition of long-run equilibrium requires that the integrated pair-wise series comove together. Price transmission is tested with a cointegration error correction model (Rapsomanikis *et al.*, 2006; Minot, 2011). The hypothesis of price transmission from US (here assumed to be the leading market) versus Italy (domestic market) is empirically justified by the large volume of unidirectional commodity flow that has created a strong dependency of Italy on US exports of these ag-commodities: more than the 90% of the entire volume of Italy's ag-commodity trade is with US. The price transmission analysis is performed by using the vector error correction model VECM in the following general form:

$$p_t = \alpha + \Pi p_{t-1} + \sum_{k=1}^{q} \Gamma_k \ \ p_{t-k} + \varepsilon_t \tag{5}$$

where p_t is a n x 1 vector of n price variables; Δ is the difference operator, $\Delta p_t = p_t - p_{t-1}$; $\boldsymbol{\varepsilon}_t$ is a n x 1 vector of error terms; α is a n x 1 vector of estimated parameters that describe the trend component; Π is a n x n matrix of estimated parameters for the long-term relationship and the error correction adjustment; and $\boldsymbol{\Gamma}_k$ is a set of n x n matrices of estimated parameters for the short-run relationship between prices, one for each of q lags of the model. The VECM provides a basis for evaluation of relationships across cointegrated series given that cointegration implies that the two prices move closely in the long-run, though in the short-term the two series could drift apart. This approach is appropriate if the following two conditions are held:

i) all variables are nonstationary and integrated of order one $I(1)$, following a random walk;
ii) the variables are cointegrated in a linear combination that satisfies the stationary condition.

The cointegration equation is:

$$P^d = \alpha + \beta \ P^w + \varepsilon^{10} \tag{6}$$

P^d and P^w are the prices representative of two spatially separated markets integrated of the same order and the error term ε is stationary, β, the cointegrating vector is the price response of dominated market to price changes of the leading market in the long-run. Since prices are expressed in logarithms, β is the long-run elasticity of the domestic price with respect to the US price or the long-run elasticity of price transmission. The expected value for imported commodities is a β value ranging between $0 < \beta < 1$; for $\beta = 0$ the

[10] this equation is comparable to eq (2) of the previous section (Goodwin, 1992).

US market has influence on the Italian markets, for β=1 the price change in US market is entirely passed to the Italian market (Ravallion, 1986); considering the lagged effect i.e. β = 0.5 the 50% of the change in US price will be transmitted to the Italian price in the long-run (Minot, 2011). The regression equation form for VECM model is:

$$p_t^d = \alpha + \theta\left(p_{t-1}^d - \beta p_{t-1}^w\right) + \delta \ p_{t-1}^w + \rho \ p_{t-1}^d + \varepsilon_t \tag{7}$$

where p_t^d is the natural logarithm of the Italian (domestic) price of corn, soybeans and wheat respectively, p_t^w is the natural logarithm of the US (world) price of the same Italian commodities, α, θ, β, δ, and ρ are parameters to be estimated and ε_t is the error term, the expression in parenthesis (p^d_{t-1}- βp^w_{t-1}) is the deviation from the long-run equilibrium. The following two terms measure the short term impact of the lagged increments (Δ) of the natural logarithm of international and domestic prices (Conforti, 2004). The error correction coefficient (θ) measures the speed of adjustment, expected to fluctuate in the range between $-1<\theta<0$. If the lagged error correction term (the term in parentheses) is positive, the domestic price is too high given the long-term relationship, then the negative value of θ "corrects" the error by making it more likely that the Δp_t^d is negative. The larger θ is in absolute value (closer to 1), the more quickly the domestic price (p^d) will adjust to the value consistent with its long-run relationship to the world price (p^w). The coefficient of change in the world price (δ) is the short-run elasticity of the Italian price relative to the US price and represents the percentage adjustment of domestic price one period after a one percent shock in international price. The expected value is $0<\delta<\beta$ (Minot, 2011).

The coefficient of the lagged change in the domestic price (ρ) is the autoregressive term, indicating the change of the Italian price caused by the change of the corresponding price in the next period, the expected value ranges between $-1<\rho<1$. Table 14 reports the results for the transmission of US prices to the corresponding Italian prices. The unit root tests reported above suggest that each domestic price is nonstationary, and the Johansen cointegration test is used to test for a long-run relationship between the Italian and the US prices. The results suggest that all the domestic prices have a long-run relationship with the US prices for the corresponding commodity. The long-run elasticity of price transmission is statistically significant for all the commodities and very high for soybeans (0.96) and wheat (0.74) meaning that a high percent change of the US price is passed through to the Italian price in the long-run. The speed of adjustment coefficient (θ) is negative as expected for sit and wit and statistically significant at 1% level while for corn there is a slightly positive θ. The value of short run adjustment coefficient (δ) is in the expected range but is not significant for all pairs of commodities. The auto-regressive term is statistically significant for all the variables and is higher for corn.

Summarizing the results obtained from the transmission model, for each commodity we find the long-run relationship values (β coefficient) are larger than those that indicate short-run transmission (δ). An important role is performed by the autoregressive term meaning that for the corn market in Italy, the 42% of the change of corn price in period t, is transmitted to period t+1; for the soybean market the value of LR adjustment increases to 0.96 and the short run autoregressive value declines to 0.013; for the wheat market, the LR adjustment value is 0.74 and the autoregressive value is 0.026.

Table 14. Transmission of the US food prices (world) versus the Italian food prices (domestic).

Commodity	Unit root in Italian prices ADF PP KPSS ZA	Long-run relationship Johansen test	Error correction model			
			Long-run adjustment β	Speed of adjustment θ	Short-run adjustment δ	Auto-regressive term ρ
cit	yes	yes	0.569*	0.002	0.015	0.416*
sit	yes	yes	0.963*	-0.038*	0.013	0.202*
wit	yes	yes	0.738*	-0.017*	0.026	0.296*

* statistically significant at 1% level.

6. Conclusion

A number of studies have presented results that support the hypothesis of integration among the ag-commodity markets; more recently many researchers have demonstrated the growing interaction between oil and the ag-commodity price, and more difficulties in predicting the price changes of ag-commodities. Since the price comovements are becoming increasingly complex, this research has been dedicated to test the hypothesis of market integration and price transmission between US and Italy for oil and some relevant ag-commodities. Our results offer traders and policy makers more reliable information to improve their decisions in market trade and policy formation. Time series analysis has been used for testing the initial hypothesis that the oil price is an exogenous signal driving the ag-commodity prices. This is intuitively justified by the large amount of energy inputs used for the ag-commodity production (e.g. fuel, fertilizer, seed, machinery) and the growing quantity of ag-commodities used for biofuel production. The link between US and Italy ag-commodity markets is grounded on the large volume of ag-commodities flowing from US to Italy and the recognized leadership of US prices settled at the CBT. However, price transmission is a more complex phenomenon embedding the comovement, completeness, speed of adjustment asymmetries, in a contest of rapid market change (Engle and Granger, 1987; Johansen and Juselius, 1990).

In short our results highlight the importance of identifying sample breaks in time series. As clearly illustrated in Table 9 and 10, results and inferences are not robust across sample periods. Intuitively, this highlights the need to empirically determine the time location of structural breaks. By definition, the presence of such a break implies a change in the underlying data generating mechanism and, therefore, a change in parameter values and perhaps functional form of any relationships. Based on identified structural breaks, we find cointegration to vary across sample periods for both cointegration between US and Italian markets within US or Italian domestic settings. These results also highlight the sample dependence of structural estimates, an intuitive statement that is often overlooked when results are compared across papers using different sample periods.

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