Agricultural market integration: price transmission and policy intervention

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Summary

The increasing co-movements between world oil and food prices in the 2000s has prompted interest in the transmission mechanism among markets. This research investigates integration and price transmission of some important agricultural commodities traded in market area that includes United States and Italy for a period spanning from January 1999 to May 2012. The hypothesis of market integration is verified for crude oil and three agri-commodities wheat, corn and soybean in Italy and US. They are selected for their market relevance due to growingly demand diversified in food, feed and fuel; wheat for its higher accounting for much of the world food consumption. It is hypothesized that US and Italy agricultural markets are integrated by a consistent volume of trading and by the recognized role of the CBT price signals transmitted to the Italian agri-commodity markets. This study extends the knowledge about the oil–agricultural commodity price transmission dynamics from international (US) to domestic market (Italy).The time series analysis is used to test the structural breaks, the co-integration and price transmission and the causality. Results suggest: i)t for the US markets the evidence of market integration between crude oil and US agri-commodity prices with non linear causality direction going from oil to agri-commodity prices; ii) the integration between US and Italian agricultural markets, with no clear evidence of causality between oil and Italian agri-commodities, while there is the evidence of linear causality going from US to Italian agricultural markets. The conclusion is a presence of causality going from Oil to US agri-markets and from US agri-markets to the Italian ones These information can be useful both for investors and policy makers interested in the knowledge about the nature of price movement in the international arena the close market integration and price transmissions with consequence for co-movement ,inherent dynamic market relationship, speed of adjustment, consequences of t price support policies. The agricultural policies in different countries may be organized to countervail the destabilizing effect of the oil price movement in disrupting the world market equilibrium by arbitraging the market condition to return to a situation of competitive pricing behaviour.

Keywords: price transmission, market integration, agricultural commodity prices, oil price, cointegration, causality

JEL Classification codes: C22, Q11, Q13

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1. INTRODUCTION

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 '07-'08 price spike seems to have been caused by different factors: a macroeconomic instability influencing the world commodity markets as the rapid growth of food demand by the BRIC countries, the international financial crisis, and the growing influence of the oil price on the other commodity markets (Piot-Lepetit and M'Barek, 2011). The agro-fuel commodities deserve an additional so-called knock-on effect due to the expanding U.S. corn production for ethanol use, reducing the oilseed acreage, such that the oilseed prices tended to increase for the expected tightening supplies. The upward price trend is enhanced by the rising demand for meals being the cereal feedstock substitutes and for vegetable oils used for bio-diesel production (OECD, 2008). Most agricultural commodity markets seems to manifest in recent times a higher variability; however, a physiological price fluctuation is accepted for the changes of agricultural output from period to period caused by natural shocks such as weather ad pests. Another reason is the rigidity of demand due to the length of time for the production to adjust to market changes.

It is widely acknowledged the growing exchanges and integration among the world market with price transmission affecting the condition of market efficiency and speed of adjustment to market level in response to leading price signals (Rapsomanikis *et al.*, 2006). The market theory, suggests that the spatial price transmission is an essential condition for the existence of the market efficiency condition based on the "Law of one price"[1]: the price transmission is complete if the prices generated in two competitive markets, converted in a unique currency, will differ only by the transaction costs. The spatial arbitrage will reduce these price differences to the level of transport cost (Ardeni, 1989). The rational expectation based on the competitive storage theory support this condition: the commodity stocks, expected prices and hauling costs drive the commodity prices to an equilibrium prices for the presence of competitive speculators who trade their stocks according with price expectation and carrying costs. Their unwillingness to hold negative inventories generate asymmetry in storage causing non linear price components (Deaton and Laroque, 1995). The absence of market integration or of a complete pass-through of price changes from one to another market, has important implications for the welfare distribution (Sharma, 2002). Market integration and price transmission, both spatially and vertically, are supported by theories and quantitative techniques apt to test the degree of market efficiency and has highlighted several factors interfering with the complete pass-through of the price signals. An important cause of market inefficiency is the government action either in the

[1] The law of one price states that the price changes in one market caused by variation in demand and supply are instantaneously transmitted to other markets so that the price variation in related markets will uniquely reflect the local changes in market equilibrium. In this sense the markets are integrated and price changes are due to spatial arbitrage (Enke, 1951, Samuelson, 1952, Takayama and Judge, 1971). A distinction is needed between short and long run equilibrium: beside price transmissions can be incomplete in short run they are complete in long run because the efficient arbitrage will fade out the differences. In this case price changes are not passed instantaneously from one market to another and delay in price changes can be imputed to policy intervention, number of stages in marketing and corresponding contractual arrangement, transport condition, rapidity of the operators to adapt to new market conditions, inventory holdings, financial speculation. These causes of market inefficiencies will affect the price adjustment either in rapidity of change and asymmetric response.

form of policies at the border or as price support mechanisms, that alter the market equilibrium by weakening the flows of products between the international and domestic markets.

The biofuel policies, encouraging farmers to produce biofuel feed-stocks, have increased the dependency of the agricultural prices to the energy prices (Gohin and Chantret, 2010). Trade limitation as import tariffs, quotas and export subsidies or taxes, trade barriers, exchange rate policies, have caused inefficient arbitrage by insulating the domestic markets and hinder the full transmission of price signals. They are responsible of excess demand or supply schedules of domestic commodity markets possibly inducing asymmetric price response reflected in a non linear adjustment between prices (Quiroz and Soto, 1996; Sharma, 2002, Rapsomanikis *et al*, 2006). Over the past decades, the world-wide integration has led to a significant increase in global trade; since late 2002, the major grains and oilseeds world markets have experienced a period of tight supplies and severe contraction in world trade; in addition the production of biofuels (ethanol and biodiesel from agricultural feedstock) and growing use of chemical and petroleum derived inputs has experienced a remarkable growth over the last decade, causing a growing dependence among oil and agri-commodity prices (Harri *et al*., 2009)[2].

Three countries are the main integrated areas of agricultural commodities used for biofuel production today: United States, Europe and Brazil. The ethanol is the main biofuel for US and Brazil while biodiesel is the most produced biofuel in the EU[3]. Ethanol production is diffused in the Northern countries for the higher energy balance compared to biodiesel, and is now accounting for the three–fourth of the world biofuel output. Studies directed to test the market integration and influence of the oil on agri-commodities have moved in two directions: econometric studies based on traditional demand and supply models adapted to capture the demand diversification in the world relied upon partial and computable general equilibrium models (Lapan and Moschini, 2012; de Gorter and Just, 2010; Hertel *et al.*, 2010). These models could suffer from arbitrarily determined or calibrated price elasticity used in stimulating the long-run sensitivity of agricultural commodity prices to oil shock prices. The second approach is based on time series analysis addressed to capture the price transmission, cointegration between prices and causal nexus among markets.

This research follows the second approach by analyzing the agricultural markets linkages, with spatial integration and price transmission that have become the most influential effects on agri-market equilibrium induced by external factors as the oil price volatility and growing market integration. The above observations suggest to test the hypothesis that the oil price is an exogenous price signal, and the main responsible of transmitting volatility to the agricultural markets, due to the strengthening of spatial price relationship and integration between oil and agricultural markets. (McNew and Fackler, 1997).

The results of this study may have important implications for both policy makers and global investors who need to follow the price shocks and transmission mechanisms between alternative investment areas closely. Therefore, research results will be beneficial for forecasting prices, establishing strategies.

The remaining of this paper is organized as it follows: in the next section is discussed the relevant literature about the market integration, volatility and price transmission; the third section introduces the methodology, the fourth section presents the data organization and preliminary descriptive analysis, the fifth section discusses the tests and empirical findings, finally the last section draws the main conclusion with policy suggestions.

[2] The US Energy Information Agency indicate that the total world production of biofuel increased nearly six times over the 2000-2010 period from 315 thousand barrels/day to 1,856 barrels/day.
[3] Ethanol growth is impressive in US, with 57.1% of the world production in 2010, in future the 2nd generation ethanol produced from the most performing non-food cellulosic feedstock (*Arundo, Panicum, Sorghum*), is the most promising biofuel source estimated to reach in **optimal** conditions 10-15,000 liters per hectare (EIA, 2012).

2. LITERATURE REVIEW

The recent rise in agricultural commodity prices has increased the interest on the determinants of this price surge; the literature suggests that these price movements have been driven by a combination of different factors, including the supply and demand changes in the agri-commodity markets and other financial factors.

The *demand-side* is thought to be the main driving forces of increasing agricultural commodity prices. The rising of the world demand for agricultural commodities is justified by the increasing population, diet changes of households in the emerging economies driven by rapid economic growth and improvement in the standards of living, that have determined the rising per capita meat consumption (Headey and Fan, 2008). Another factor is the growth of the biofuels production that has been considered the main responsible of the raise in corn and oilseed commodity demand and prices and the acreage invested in biofuel crops (Gilbert, 2010; Mitchell, 2008; Zhang *et al*, 2010). The decline of the dollar value (Trostle, 2008); and the speculation stemming from increased activity in futures markets (Robles *et al.*, 2009) are also the demand-driven factors contributing to the agricultural commodity prices movement[4].

On the *supply side*, several factors have had an effect on prices: the slow growth in agricultural production, and increase in energy prices have determined the rise of farm production costs like transportation fertilizer, pesticide (Tyner and Taheripour, 2008; Sumner, 2009; von Braun *et al*., 2008), and climatic events (Trostle, 2008) are the more pronounced supply-side explanations.

According to the OECD (2008), oil prices and feedstock demand for biofuel production seem to keep their importance in determining the recent behaviour of agricultural commodity prices and appear to be permanent factors of demand for agricultural products and agricultural prices. Others authors have attributed to the link between agricultural and oil prices the responsibility of the rising production of biofuels (Zhang *et al.*, 2010; Ciaian and d'Artis, 2011a,b).

The existing literature involved in the interaction between energy and agricultural commodities is quite recent; a lot of these studies dated after 2008 food crises, although this topic was somehow highlighted before. Some studies have examined the relationship and the long-run relationship among soft commodity prices and in some cases among selected soft commodities prices and crude oil price.

Campiche *et al.* (2007) by examining the relationship between crude oil prices and corn, sorghum, sugar, soybeans, soybean oil, and palm oil prices during the 2003-2007 period, with a vector error correction model conclude that there wasn't evidence of cointegration during the 2003-2005 period whereas corn prices and soybean prices were cointegrated with oil prices during the 2006-2007 time period. The authors conclude from their analysis that soybean prices seems to be more correlated to crude oil prices than corn prices. Yu *et al.* (2006) investigate the long-run interdependence between major edible oil prices oils prices, including soybean, sunflower, rapeseed and palm oils, and the dynamic relationship between vegetable and crude oil prices. They conclude that shocks in crude oil prices do not have a significant influence on the variation of edible oil prices, which appears to confirm the results of Campiche (2007).

Zhang and Reed (2008) examined the impact of the crude oil price on China's agricultural commodity prices focalizing their attention on feed grain (corn and soybeans) and pork and suggest non significant crude oil price fluctuation over the study period (2000-2007). Similar results are found by Nazlioglu and Soytas (2011) for Turkey. Tyner and Taheripour (2008) emphasize the relation between rise in oil prices and the increase in corn prices. Gilbert (2010) suggests that all agricultural markets are affected by the change in oil prices; namely the oil prices influence the food prices either by increasing the production costs or by using

[4] The price surges caused by speculation could determinate turbulence to the global grain markets affecting the market's efficiency in responding to fundamental changes in supply, demand, and costs of production (Spears, 2011)

food an input for biofuel production. The author suggests also that the cost of food production were affected by the transport and fertilizer cost.

Nazlioglu (2011) by examining the relationships between oil and ag-commodities corn, soybeans, and wheat has found evidence of non linear causality between oil and agricultural commodity prices; in another work, Nazlioglu and Soytas (2012) have found strong evidence of the world oil price changes on agricultural commodity prices using the panel cointegration and Granger causality methods.

Moving to the relations oil – biofuel - crops, Serra *et al.* (2010a) find that the prices of oil, ethanol and corn for the US to be positively correlated, and the existence of a long term equilibrium relationship between these prices, with ethanol. In Brazil, using the sugar as feedstock Serra *et al*. (2010b) demonstrate that sugar and oil prices are exogenously determined; by focusing their attention on the response of ethanol prices to changes in these two exogenous drivers, these authors conclude that ethanol prices respond relatively quickly to sugar price changes, but more slowly to oil prices.

Serra and Gil (2012) suggest the price volatility of agricultural commodities is affected by the energy prices, corn stocks and global economic conditions. Their findings support evidence of price volatility transmission between ethanol and corn markets. While the impacts of stocks in the very short-run are higher compared to the energy prices and macroeconomic instability, in the long-run the ethanol price and interest rate volatility are found to have the strongest impacts.

Considering the role of trade policy intervention, Esposti and Listorti (2013) investigate about the national and international markets; trade policy regime has an important role in price transmission mechanisms and the trade policy intervention to mitigate the impact of price exuberance is considered. The authors analyze agricultural price transmission during the time of price bubbles, using Italian and international weekly spot (cash) price data over the years 2006–2010. Their conclude that the bubble has had only a slight impact on the price spread and the temporary trade-policy measures, when effective, have limited this impact.

As the agri-commodity markets are open to investors and speculators, similarly to the oil markets, the prices in both markets may be governed by similar dynamics due to their reciprocal influence. Food, oil and energy prices have been studied extensively in the literature. The innovation of this study is to consider the interaction among US and Italian agri-commodity markets using the corn, soybeans, wheat and oil prices to observe possible evidence of market integration and price transmission.

3. METHODOLOGY: TIME SERIES ANALYSIS

Time series analysis has been frequently used to study the agricultural markets (Tomek and Myers, 1993; Rosa, 1999; Esposti and Listorti, 2010; Carraro and Stefani, 2010; Nazlioglu, 2011).

Time series analysis includes methods for analysing time series data in order to extract meaningful statistics and relevant information from time series data. Most of these studies analyse the co-movement in prices using cointegration relationship and the error correction models, that is a common procedure for analyzing spatial market relationships and price transmission, replacing earlier empirical tools, such as the bivariate correlation coefficient or simpler regressions analyses. This approach is well suited to test market conditions such as completeness, speed and asymmetry of price relations and is appropriate to give evidences about market failures, direction, magnitude and distribution of welfare changes.

In our approach traditional unit root tests and the Zivot Andrews one break test are used to detect the stationary condition in data time series. Then, cointegration tests has been performed to determine whether there exists a long-run relationship among the series in the system and a vector error correction model has been specified to detect the price transmission. This is followed by causality tests with the linear (Granger causality test) and a non linear approach to examine the causal linkage among the variables.

3.1. Unit root test

The condition of stationary is checked because all series must be integrated of the same order. In order to have robust estimation results, identification of the stationarity of the data has an utmost importance. Stationarity properties of the variables are determined by various unit root tests. Since some tests can give contradictory results, a variety of tests are conducted to check reliability. Aforementioned tests are augmented Dickey-Fuller (1979) [ADF], Phillips-Perron (1988) [PP], Kwaitkowski-Phillips-Schmidt-Shin (1992) [KPSS]. The null hypothesis of the unit root tests, apart from KPSS, is the series must have a unit root against an alternative of stationarity; on the other hand, stationarity of the variable is the null hypothesis for KPSS. As a standard procedure to test the non stationarity of a series, the ADF test is based on the regression:

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

where μ is a constant, β the coefficient on a time trend and k the lag order of the autoregressive process. The unit root test is then carried out under the null hypothesis $\alpha = 0$ against the alternative hypothesis of $\alpha < 0$. Non stationarity is refused when the test suggests that α is different from 1.

However, if a structural break is present in the data generating process, the conventional ADF test is biased toward the acceptance of the null resulting in a dramatic loss of power.

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 exists in the series, the conventional unit root test may result in misleading inferences. Assuming the time of the break as an exogenous phenomenon, Perron (1989) has demonstrated that the power to reject a unit root decreases when the stationary alternative is true and a structural break is ignored. This is the reason why the results of conventional tests are compared with those obtained with the Zivot and Andrews (1992)[ZA] unit root test; it is an endogenous structural break test with unknown timing in the individual series using the full sample and different dummy variable for each possible 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 favorable 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 a once only break point occurring at some unknown time. ZA test is a variation of Perron's original test with the endogenous implementation of structural breaks in the analysis: the date of the break is determined on the basis of t-statistics test of the unit root, with respect to the criteria of minimum values.

Following Perron's characterization of the form of structural break, ZA formulate three different characterizations of the trend break to test for a unit root: i) model A, "*the crash model*", that allows the break in the intercept; ii) model B, "*the changing growth model*", which allows for a one-time change in the slope of the trend function with the two segments joined at the break point; iii) model C, "*the mixed model*", which combines simultaneously the one-time changes in the level with the slope of the trend function of the series[5]. The aim of this procedure is to sequentially test the breakpoints candidates and select the one that

[5] 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).

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 above three situations.

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

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

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

where DU_t is an indicator dummy variable for a mean shift occurring at each possible break-date while DT_t is corresponding trend shift variable. The null hypothesis in all the three models is $\alpha = 0$, which implies that the series y_t contains a unit root with a drift that excludes any structural break, while the alternative hypothesis $\alpha < 0$ implies that the series is a trend-stationary process with a one-time break occurring at an unknown point in time. The Zivot and Andrews test regards every point as a potential break-date and runs a regression for every possible break-date sequentially.

3.2. Cointegration analysis

The main goal of a co-integration test is to examine if two or more series are linked to form an equilibrium relationship. The concept of cointegration means that the two price series cannot evolve in opposite directions for very long time without converge to a mean distance.

Let us consider a static regression between $I(1)$ variables:

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

where x_t is a vector of independent variables. The system is cointegrated if the errors ε_t are $I(0)$. In this case the relation (2) may be interpreted as a long run equilibrium toward which the process *yt* tends.

Assuming x_t and y_t two integrated processes, if there exist a linear combination which is integrated of a lower order, both variables are cointegrated. Failure to reject the null of no cointegration implies that two price series drift apart in the long run driven by non proportional stochastic trends (Rapsomanikis *et al*., 2006). Johansen cointegration test starts considering a vector autoregressive (VAR) model with *k* lags under the consideration that variables are I (1): written in error-correction form:

$$\Delta p_t = \alpha + \Pi p_{t-1} + \sum_{k=1}^{t} \Gamma_k \Delta_{t-k} + \varepsilon_t \qquad (3)$$

where $\mathbf{p_t}$ is an n x 1 vector of n price variables, $\mathbf{\Delta}$ is the differencing operator such that $\Delta p_t = p_t - p_{t-1}$, $\boldsymbol{\alpha}$ is an n x 1 vector of estimated parameters that describe the trend component, $\mathbf{\Pi}$ is an n x n matrix of estimated parameters that describe the long-term relationship and the error correction adjustment, $\mathbf{\Gamma_k}$ is a set of n x n matrices of estimated parameters that describe the short-run relationship between prices, one for each of q lags included in the model, and $\boldsymbol{\varepsilon_t}$ is an n x 1 vector of error terms.

The rank of matrix Π is of interest with regard to the long-run cointegrating relationships between variables in the model. That is if Π = 0, all variables are non-stationary and model (3) reduces to a differenced vector time series model implying that no cointegration relationships exist among variables and there are no cointegrating vectors and for Π = 1, there is one cointegrating vector. If Π > 1, there is more than 1 cointegrating vectors. If the rank of Π equals zero, (Johansen and Juselius, 1990).

The procedure proposed by Johansen and Juselius (1990), is used to determine the absence or presence of cointegrating relationship among variables. Although there are other tests like Engle and Granger (1987), Johansen's cointegration test has demonstrated to be superior by considering all variables to be endogenous and its capability of testing more than one cointegrating relationship. The Johansen procedure uses Trace and Eingenvalue tests. The trace statistic reports the null hypothesis of r cointegrated relations against the alternative of k cointegrating relations, where k is the number of endogenous variables. The maximum eigenvalue test, on the other hand, tests the null hypothesis of r cointegrating vectors against the alternative hypothesis of r + 1 cointegrating vectors. The rank r is calculated with the eigenvalues of a matrix. If all the eigenvalues are significantly different from zero, all processes are stationary. On the contrary, if there is at least one eigenvalue equal to zero, the process x_t is integrated. On the other side, if none eigenvalue is significantly different from zero, not only the process x_t is non stationary but this is for all the linear combinations. In other words there is no evidence of cointegration.

It has been found that the trace test is the better test, since it appears to be more robust to skewness and excess kurtosis (Sjö, 2009). To determine if cointegration relationships exist between the variables, first the lag length (k) is determined and then cointegration rank (r) is determined. To determine the lag length the Schwarz Bayesian Criterion (SBC) (Schwarz, 1978) (also known as BIC) is used. Finally, Gregory and Hansen (1996) apply a test that is an extension of the Engel - Granger residual based cointegration analysis. This approach is an extension of the endogenous univariate test of Zivot and Andrew (1992) unit root tests with structural breaks: Gregory and Hansen propose the cointegration tests which accommodates a single endogenous break in an underlying cointegrating relationship.

The null hypothesis of cointegration is tested against the alternative of cointegration with a break in cointegrating relationship. The GH tests for cointegration allow the possibility of regime shift and develop ADF^*, Z_t^* and Z_α^* type tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible structural break. The authors consider three modified version of equation (2) that includes dummies for the structural change :

Model C: Level Shift

$$y_t = \mu + \theta DU_t + \alpha x_t + \varepsilon_t \tag{4a}$$

Model C/T: Level Shift with Trend

$$y_t = \mu + \theta DU_t + \beta t + \alpha x_t + \varepsilon_t \tag{4a}$$

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

$$y_t = \mu + \theta DU_t + \alpha_1 x_t + \alpha_2 DU_t x_t + \varepsilon_t \tag{4c}$$

where y is the dependent and x is the independent variable, 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 to the previous model whilst model C/S deals with 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 point.

Next step will be testing for Granger causality which plays an important part in many vector error correction models. The cointegration implies causality in the Granger sense defined in terms of predictions of future values of Y improved by using present and past values of X.

3.3. Granger causality

The Granger test (1969) is introduced to observe how much of the current y can be explained by a past value of x and then to see whether adding lagged value can improve the explanation. y is said to be Granger-caused by x if x helps in the prediction of y, or equivalently if the coefficients on the lagged x's are statistically significant.

In this step the examining of the relationship by the traditional Granger causality test is simply to give an indicator of the direction relationship. The Granger causality test has to be run on $I(0)$ series, and it is executed by a simple F-test. The causality relationship can be evaluated by estimating the following:

$$\Delta X_t = \sum_{j=1}^{m} \alpha_{1j} \Delta X_{t-j} + \sum_{j=1}^{m} \beta_{1j} \Delta Y_{t-j} + \varepsilon_{1t} \quad (5a)$$

and

$$\Delta Y_t = \sum_{j=1}^{m} \alpha_{2j} \Delta X_{t-j} + \sum_{j=1}^{m} \beta_{2j} \Delta Y_{t-j} + \varepsilon_{2t} \quad (5b)$$

where X_t and Y_t are the prices of time series to test for causality. In this specific case, the null hypothesis to be tested are:

- oil price does not Granger-cause US food commodity price and US food commodity price does not Granger-cause crude oil price i.e. $H_0{:}\beta_{1j}=0$, j=1,2,...m and $H_0{:}\alpha_{2j}=0$,j=1,2,...m.;
- oil price does not Granger-cause Italian food commodity price and Italian food commodity price does not Granger-cause crude oil price;
- US food commodity price does not Granger-cause Italian food commodity price and Italian food commodity price does not Granger-cause US food commodity price.

As a complementary analysis, nonparametric Granger causality tests are performed to uncover potential nonlinear dynamic relations between oil and agri-commodities. Traditional linear Granger causality tests have high power in identifying linear causal relations, but some authors argue that the linear Granger causality is ineffective in capturing nonlinear causal relations, and recommend to test for nonlinear Granger causality (Baek and Brock,1992; Hiemstra and Jones,1994).

Assuming that linear causality tests might overlook nonlinear dynamic relations between oil and agri-commodities, we verify the hypothesis of non linear Granger causality performing the nonparametric causality test proposed by Diks and Panchenko, (2006) [DP] which avoids the over-rejection observed in the test proposed by Hiemstra and Jones (1994).

Considering that the null hypothesis of Granger non-causality can be rephrased in terms of conditional independence of two vectors X and Z given a third vector Y, Diks and Panchenko (2006) show that the Hiemstra and Jones test is sensitive to variations in the conditional distributions of X and Z that may be present under the null hypothesis[6].

4. DATA CHARACTERISTICS

For the empirical analysis, weekly spot prices[7] of major agricultural and oil commodities have been examined. The high frequency of observation has been chosen to capture the dynamic market linkages and causal nexus among prices (Nazlioglu, 2011). Soft wheat, maize and soybeans have been selected because their high importance for food, feeds and fuel. Wheat is more energy intensive and the key product for food, corn and soybean are the most important biofuel feedstock; wheat, corn and soybean are still used in crop rotation for sustainable reasons. Table 1 presents the variables used in analyzing both markets.

[6] For details in the methodology refer to Diks and Panchenko (2006)
[7] Spot prices are used because most of the transactions in Italy are made in these markets.

Table 1. List of variables

Variable		Description	Source
Italian	corn price	Weekly average of spot prices in €/ton of national hybrid corn-market at the origin (cit)	DATIMA provided by ISMEA[8]
	soybean price	Weekly average of spot prices in €/ton of soybeans with 14% of moisture--market at the origin (sit)	DATIMA provided by ISMEA
	wheat price	Weekly average of spot prices in €/ton of good mercantile wheat--market at the origin (wit)	DATIMA provided by ISMEA
US	corn price	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	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	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		Weekly spot prices of Brent crude oil converted in €/barrel (oil)	US Energy Information Administration

Weekly price series have been monitored from January 1999 to May 2012, a total of 699 observations. The prices of US agri-commodities and oil, originally expressed in US dollar, are converted into euro by using the official $/€ exchange rate[9] considering the weekly averages of daily quotations.

Line graphs of spot market prices presented in Figure 2 indicate at glance non-stationary trend and suggests to divide the all sample in sub-samples with different characteristics: a relatively quiet period till 2004, followed by wider fluctuations in the following period (till the end of 2008), due to turmoil in financial and energy markets transmitted to the agricultural markets, and finally an instable period with relative increment of all commodity prices.

Figure 1. Index of current prices of some agri commodities and oil (Jan 04, 2002= 100)

Source: own elaborations. cit, wit, sit, cus, wus, sus: €/ton; for oil: €/barrel

4.1. Testing for stationarity and structural break

To check for the order of integration of the time series and identify the subdivision of the samples, stationary tests were firstly performed. A stationary process implies that the mean, variance and autocorrelation do not change over time. These tests are usually incorporated as first in all the time series econometric analysis since the stationary condition must be achieved before further proceeding with other analysis. Such condition has been studied even before the description of the price series characteristics to better define the split-periods of investigation. The results of the conventional unit root tests for levels and

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

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

first differences are presented in Table 2 and suggest that all the all variables are integrated of order 1, even though there are slight differences between the results of different tests.

Table 2. 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 is used 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 3 reports the results of Zivot-Andrews one break test. Minimum ZA t-statistics for the levels of the variables show similar results with those obtained from the unit root tests without accounting for structural breaks with the exception for oil and sus: once a break in the deterministic trend is allowed for, the null hypothesis of a unit root process is rejected. The test was run in the three versions illustrated in section 3. A structural break is found in the US soybeans series. The estimated date is July 2004 (2004: week 29) with a model fitted in the drift (model A) and a change in the trend slope and drift (model C). The oil series appears to be stationary with a break in October 2008 (2008: week 40) with change in trend slope and drift. As underlined by Piehl *et al*., (1999), the knowledge of break time is a central point in the accurate evaluation of any program intended to bring about structural changes; such as the climate shocks, market disruption, regime shifts and others.

Table 3. 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)

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

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

All the other price series are found to be *I*(1) confirming the traditional unit root tests. Even though sus and oil price series are stationary in model C, this condition does not appears with the same evidence in model A neither in model B. For this reason we conservatively assume that all the variables are integrated of order one I*(1)*.

4.2. Descriptive statistics

Table 4 reports the descriptive statistics of the price in each market. Skewness is particularly important for the investment theory: a positive value (long RHS tail) means frequent small price drops and few extreme price run up, while a negative value (long LHS tail) means frequent small gains and few extreme losses. Positive skewness implies larger price increase while negative skewness implies large price drops.

Table 4. Summary statistics (entire period)

	oil	cus	sus	wus	cit	sit	wit
Mean	44.15	116.77	246.83	139.64	156.75	283.85	168.15
Median	39.73	104.00	219.86	127.65	140.50	252.00	152.97
Maximum	96.30	229.35	434.00	307.81	273.80	475.67	292.13
Minimum	8.75	69.44	157.65	78.98	113.85	166.30	118.00
Std. Dev.	20.30	38.86	67.21	40.31	37.71	77.50	43.64
Skewness	0.62	1.40	0.83	1.18	1.15	0.62	1.20
Kurtosis	2.48	4.05	2.56	3.85	3.24	2.26	3.42
Jarque-Bera	52.77	261.23	85.00	183.94	155.69	60.74	171.94
p-value	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Source: own elaboration

All price series follow a non-normal distribution. *Jarque-Bera* test statistics are significant implying a deviation from normality[10]. According to Table 4, *kurtosis*[11], indicating the flatness of the curve, exceeds 3 pointing out the presence of fat tails in cus, wus, cit and wit; suggesting a leptokurtic distribution with value concentration around the mean. *Skewness* measures the symmetry of the curve: for perfect symmetric normal distribution, the skewness value is zero, the negative skewness indicates that the mean is inferior to the median, the asymmetric distribution will show a long left tail with the mass of the distribution concentrated on the right side of the figure. At these conditions the market activity is higher, because operators have positive expectation about price increase. The reverse RHS tail implies higher risk for operators and minor willingness for producers to enter the market. The commodities observed, show long right tail and leptokurtic distribution meaning that there is a lower frequency of values with positive small deviation and a higher risk of losses.

4.3. Price variability

Food price variability is observed over time: a rather intuitive measure is the coefficient of variation, CV = σ/μ where σ is the standard deviation of the variable of interest and μ is the mean value measured over a given time period and is a quite diffused measure to estimate price volatility (Minot, 2012) among commodities with different average prices. This measure refers to *ex post* observations of actual prices. The higher the coefficient of variation is, the larger the dispersion of series and greater the volatility. Clearly, some variability can be predicted (*e.g.* seasonal variation, business cycles, or other trending behavior) such that results from using the simple standard deviation may overstate the degree of volatility or uncertainty (for more discussion see Moledina *et al.*, 2004). Therefore, in order to have a better measure of the

[10] The **Jarque–Bera** test is a goodness-of-fit measure of departure from normality, based on the sample kurtosis and skewness. The test statistic *JB* is defined as

$$JB = \frac{n}{6}\left(S^2 + \frac{1}{4}K^2\right)$$

where *n* is the number of observations (or degrees of freedom in general); *S* is the measure of skewness (third moment), and *K* is the kurtosis (fourth moment) of the data. The statistic *JB* has an asymptotic chi-square distribution with two degrees of freedom and can be used to test the null hypothesis that the data are from a normal distribution.

[11] **Kurtosis** is a measure of whether the data are peaked or flat relative to a normal distribution. The value 3 signals a normal distribution, value above 3 signals a peaked curve near the mean declining rather rapidly and having heavy tails. With kurtosis value below 3 the curve tend to have a flat top near the mean; value 0 suggest that the curve is horizontal.

unpredictability or uncertainty faced by the market, it is common to take into account only movements of the series that cannot be predicted on the basis of its previous values.

Table 5. Coefficient of variation

	total sample	sub-samples		
	1999:w1-2012:w25	1999:w1-2004:w29	2004:w30-2008:w40	2008:w40-2012:w25
oil	0,460	0,229	0,247	0,287
cus	0,333	0,090	0,285	0,269
sus	0,272	0,151	0,263	0,151
wus	0,289	0,154	0,312	0,214
cit	0,241	0,127	0,254	0,255
sit	0,273	0,156	0,301	0,101
wit	0,260	0,106	0,313	0,254
# obs	699	289	220	190

Source: own elaboration

Table 5 reports the coefficient of variation for the whole period and the sub-periods as individuated by ZA one break test. Considering the total sample, January 1999 to May 2012, oil and US commodity markets generally experienced more volatility than Italian markets. Coefficient of variation increased both on the oil and agricultural commodity markets between 2004-2008 and 2008-2012, with the oil recording more dramatic increases. However, comparing the three sub-periods, dispersion of prices in 2004–08 measured by coefficient of variation is higher than the other two sub-period; note that 2004-08 time period includes price peaks, significantly shifting the means of the time series. In absolute terms the coefficient of variation remains higher on the US than on the IT markets during 2004-2008 and 2008-2012 for all products but wheat where the levels are comparable. During the last sub-period, however, comparing the agri-commodity prices, volatility of soybeans is relatively low.

4.4. Descriptive analysis of price correlation

This preliminary analysis is addressed to provide "ex ante" information about price dynamics but doesn't support any evidence about market integration or efficiency. A common view is that crude oil prices and agricultural product prices should be related through production costs for high energy intensive agriculture and more recently as a result of increasing use of agricultural feedstock (cereals, oilseeds and sugar crops) for biofuel production (Houchet-Bourdon, 2010).

The visual inspection of the historical price series (Figure 1) suggests a closer co-movement with moderate unvaried pattern, cyclical long run movements, non-linear trend components and random fluctuation in shorter period.

Oil price patterns could have affected the efficiency of agricultural market in the last period (1999-2012) with substitution of price signals generated by market fundamentals (demand-supply-stock) with other reference signals (Headey and Fan, 2008; OECD, 2008). The integration among agricultural, energy and financial markets is the relevant topic to frame the policies to prevent the agri-commodity destabilization in the long term (Tyner and Taheripour, 2008).

The most familiar measure of correlation is the Pearson correlation coefficient (r) used to measure the linear relationship between two variables represented by the prices of commodities X and Y.

The correlations between the oil price and each of the agricultural price series are computed over the whole sample and within the split samples and results are reported in Table 6. The Pearson correlation values for the whole period of observation (Table 6 a) suggest the following considerations: the pairwise correlation coefficient between oil and most of the commodities is generally high (>.70), attesting the close co-movements among the prices. Some values deserve more attention: the lower values of correlation between oil and respectively the corn and wheat prices in Italy suggest that the agri-markets are not strictly influenced by the oil price movements, while the higher correlation between sus and sit suggests a stronger link, probably due to the high quantity of soybean imported by Italy and quoted in US market.

Table 6. Pearson correlation coefficients

	1999:w1-2012:w25 (a)							**1999:w1-2004:w29 (b)**						
	oil	cus	sus	wus	cit	sit	wit	oil	cus	sus	wus	cit	sit	wit
oil	1.00	0.80	0.77	0.74	0.68	0.80	0.66	1.00	0.39	0.43	0.39	-0.17	0.57	0.10
cus		1.00	0.89	0.79	0.77	0.80	0.78		1.00	0.50	0.75	-0.08	0.51	0.11
sus			1.00	0.79	0.76	0.93	0.75			1.00	0.50	0.53	0.87	0.46
wus				1.00	0.81	0.81	0.88				1.00	-0.03	0.58	0.10
cit					1.00	0.75	0.92					1.00	0.43	0.50
sit						1.00	0.76						1.00	0.47
wit							1.00							1.00
# obs							699							289
	2004:w39-2008:w40 (c)							**200:w41-2012:w25 (d)**						
	oil	cus	sus	wus	cit	sit	wit	oil	cus	sus	wus	cit	sit	wit
oil	1.00	0.79	0.83	0.56	0.71	0.81	0.64	1.00	0.84	0.79	0.74	0.81	0.72	0.78
cus		1.00	0.89	0.67	0.81	0.86	0.77		1.00	0.85	0.85	0.89	0.70	0.91
sus			1.00	0.76	0.84	0.96	0.81			1.00	0.83	0.84	0.90	0.84
wus				1.00	0.89	0.80	0.95				1.00	0.89	0.69	0.94
cit					1.00	0.87	0.97					1.00	0.75	0.93
sit						1.00	0.86						1.00	0.71
wit							1.00							1.00
# obs							220							190

Source: own elaboration

The correlation coefficients for sub-periods reported in sections b, c and d of Table 6, indicates that energy and agricultural prices co-movement are stronger in the more recent years with the signs as they were expected. The period 99-04, characterized by lower price volatility, presents a pairwise correlation (oil/agri-commodities) with lower values and negative in one case; Figure 1 helps to explain this result by showing a great oil price variability not followed by the same fluctuation of the agricultural prices. Examining correlations among the food prices, values are in general positive but lower, meaning a low co-movement except for the soybean market that confirms the previous observation.

A comparison across periods, indicates that energy and agricultural markets became more and more interconnected in the recent period of observation with positive correlation for all markets. Even considering the results between international and Italian agri-commodities, we can notice that starting from 2004 there is a strong positive linear correlation. Anyhow, higher (or lower) values of linear correlation coefficients do not necessarily imply existence (or not existence) of causal nexus between two variables.

5. ECONOMETRIC ANALYSIS RESULTS

Usually, checking for stationary condition of price series is the first test in any econometric analysis. In order to have robust estimation results, identification of the stationarity of the data has an utmost importance. Anyhow, it has already been studied at the beginning of the data analysis to delineate the break point in order to divide the sample into sub-samples. From results of section 4.1., the time series under investigation are all integrated of first order so it is possible to continue with the cointegration and VEC analyses based on Johansen and Juselius (1990) procedure.

5.1. Cointegration analysis

The increasing co-movements between oil and agri-commodity prices during the recent years suggest to consider the cointegration relationship among the variables under investigation.

Although some of the series checked for unit roots were found to be stationary with a breaking trend, tests of cointegration were conservatively run for all possible couple of series: Johansen test and Gregory and Hansen (GH) test that accounts for a break in the cointegration relationship as described in section 3.

Table 7. Matrix of Cointegration Test: trace test results (1999:w1-2012:w25)

	cus	sus	wus	cit	sit	wit	oil	
r=0	NA	18,87	27,79	26,60	12,74	23,93	15,36	cus
r≤1		4,61	6,03	7,24	4,56	7,44	3,83	
r=0	18,87	NA	26,44	27,15	25,91	29,33	17,73	sus
r≤1	4,61		9,23	8,04	7,04	6,59	7,33	
r=0	27,79	26,44	NA	30,74	32,52	42,17	21,16	wus
r≤1	6,03	9,23		5,54	10,36	6,13	6,69	
r=0	26,60	27,15	30,74	NA	29,56	31,24	23,11	cit
r≤1	7,24	8,04	5,54		10,60	4,91	7,02	
r=0	12,74	25,91	32,52	29,56	NA	28,69	16,38	sit
r≤1	4,56	7,04	10,36	10,60		7,44	7,68	
r=0	23,93	29,33	42,17	31,24	28,69	NA	19,50	wit
r≤1	7,44	6,59	6,13	4,91	7,44		4,78	
r=0	15,36	17,73	21,16	23,11	16,38	19,50	NA	oil
r≤1	3,83	7,33	6,69	7,02	7,68	4,78		

	Critical value			
		1%	5%	10%
	r=0	31,15	25,87	23,34
	r≤1	16,55	12,52	10,67

MacKinnon-Haug-Michelis (1999) critical values

H_0 no cointegration. Red indicates rejection of Ho

The trace test of Table 7 indicates that all food prices are cointegrated in bivariate pairs with the exception of US and IT soybeans that have no cointegration with US corn. This support the idea that Italian markets are integrated with the US agri-markets and consequently the price changes in the two markets are quite interdependent. This hypothesis is supported by the Law of One Price

No agri-commodity prices are found to be cointegrated with Brent Blend price. Very weak evidence of cointegration was found for IT corn but oil and it cannot be rejected the null of no cointegration at 10% level.

Table 8. Matrix of Cointegration Test: trace test results (1999:w1-2004:w29)

	cus	sus	wus	cit	sit	wit	oil	
r=0	NA	26.10	30.80	21.47	17.32	17.24	21.37	cus
r≤1		8.73	9.70	8.12	6.97	4.97	7.84	
r=0	26.10	NA	25.95	34.33	39.44	34.24	30.75	sus
r≤1	8.73		7.82	8.85	6.88	6.60	8.95	
r=0	30.80	25.95	NA	20.40	15.48	17.60	18.89	wus
r≤1	9.70	7.82		7.43	6.21	6.79	7.67	
r=0	21.47	34.33	20.40	NA	18.18	23.18	22.44	cit
r≤1	8.12	8.85	7.43		7.83	5.89	11.02	
r=0	17.32	39.44	15.48	18.18	NA	17.75	24.10	sit
r≤1	6.97	6.88	6.21	7.83		5.59	5.63	
r=0	17.24	34.24	17.60	23.18	17.75	NA	17.03	wit
r≤1	4.97	6.60	6.79	5.89	5.59		6.59	
r=0	21.37	30.75	18.89	22.44	24.10	17.03	NA	oil
r≤1	7.84	8.95	7.67	11.02	5.63	6.59		

H_0 no cointegration. Red indicates rejection of Ho

Table 8 shows results of the cointegration test in the first sub-period among all the variables. The null hypothesis of no cointegration relation is rejected in few cases: there is a reduction of cointegration relationships among food prices. US soybeans is cointegrated with all the other agricultural commodities and with oil; besides, wheat and corn are cointegrated only in the US market. The Italian market doesn't experiment any cointegration relationship a part from soybeans cointegrated with the US corresponding commodity and with oil prices.

Table 9. Matrix of Cointegration Test: trace test results (2004:w30-2008:w40)

	cus	sus	wus	cit	sit	wit	oil	
r=0	NA	15.60	19.70	21.70	13.71	23.95	15.32	cus
r≤1		5.26	5.61	3.60	3.25	3.17	4.47	
r=0	15.60	NA	12.93	23.18	24.74	24.37	19.03	sus
r≤1	5.26		4.40	3.81	3.00	3.51	5.55	
r=0	19.70	12.93	NA	12.69	34.77	30.09	12.66	wus
r≤1	5.61	4.40		1.45	4.83	2.57	5.23	
r=0	21.70	23.18	12.69	NA	31.06	16.53	10.85	cit
r≤1	3.60	3.81	1.45		5.12	1.13	3.79	
r=0	13.71	24.74	34.77	31.06	NA	30.44	15.24	sit
r≤1	3.25	3.00	4.83	5.12		4.80	6.17	
r=0	23.95	24.37	30.09	16.53	30.44	NA	17.36	wit
r≤1	3.17	3.51	2.57	1.13	4.80		4.14	
r=0	15.32	19.03	12.66	10.85	15.24	17.36	NA	oil
r≤1	4.47	5.55	5.23	3.79	6.17	4.14		

H_0 no cointegration. Red indicates rejection of Ho

Next period (2004-2008) presents a different situation. Table 9 results show that there isn't cointegration between oil and food prices but Italian soybeans and wheat prices show a cointegration linkage with almost all the other agricultural prices.

In the last sub-period the situation is radically changed especially with regard to the relation to the linkage between oil and food prices. In this case the null hypothesis of no cointegration relation is rejected for all the combination of prices. This finding is in fact consistent with increasing importance of corn and soybeans as a consequence of the significant expansion of biofuels in the last years and the fact that wheat production process is becoming more and more energy intensive.

Table 10. Matrix of Cointegration Test: trace test results (2008:w41-2012:w25)

	cus	sus	wus	cit	sit	wit	oil	
r=0	NA	27.75	29.53	24.50	18.05	42.80	28.80	cus
r≤1		6.45	6.67	3.63	3.35	2.79	10.04	
r=0	27.75	NA	16.70	9.51	47.26	19.27	36.68	sus
r≤1	6.45		4.65	3.00	6.15	2.86	7.48	
r=0	29.53	16.70	NA	14.86	11.59	62.59	36.22	wus
r≤1	6.67	4.65		3.24	5.37	3.12	6.88	
r=0	24.50	9.51	14.86	NA	9.20	13.35	20.46	cit
r≤1	3.63	3.00	3.24		2.84	2.43	3.37	
r=0	18.05	47.26	11.59	9.20	NA	9.62	32.39	sit
r≤1	3.35	6.15	5.37	2.84		3.86	3.23	
r=0	42.80	19.27	62.59	13.35	9.62	NA	35.91	wit
r≤1	2.79	2.86	3.12	2.43	3.86		1.38	
r=0	28.80	36.68	36.22	32.39	32.39	35.91	NA	oil
r≤1	10.04	7.48	6.88	3.23	3.23	1.38		

H_0 no cointegration. Red indicates rejection of Ho

Quite similar results are reported in literature. Campiche *et al.* (2007) examine the co-movements between world crude oil prices and corn, sorghum, sugar, soybeans, soybean oil, and palm oil prices during the period 2003–2007 based on weekly data. The empirical analysis with the Johansen cointegration test shows that while there is no cointegrating relation among the variables in concern for the period 2003–2005, corn and soybean prices are cointegrated with crude oil prices during the period 2006–2007. Harri *et al.*

(2009) report a consistent cointegrating relationship between crude oil and corn, soybeans starting in April 2006. Nazlioglu (2011) considers the cointegration between oil and the three key agri-commodity prices (corn, soybeans and wheat) and reports that corn and soybeans are cointegrated with the oil prices during the period 2008-2010. If our results differ from those found in literature is mostly because samples length is different. To be sure to give a good interpretation of the results from Johansen's testing framework, since the structural break dates were determined *a priori* instead of finding them endogenously in the cointegration model, the relationship between brent and agri-commodity prices is also analyzed by running the Gregory-Hansen test with structural break.

Table 11. Cointegration test with structural break[12] between US agri commodities and oil

		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

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

Table 11 reports results of cointegration of the Gregory Hansen test between oil and US agri-commodity prices. As far as brent and cus price relation is concerned, ADF* fails to reject the null hypothesis of no cointegration with model C, C/T and C/S whereas Z_t^* and Z_α^* type test results indicate the rejection of the null for all the three models; the significant breaking periods are in May and September 2004 (week 22 and 37) and May 2010 (week 19). In the case of soybeans and brent, all these three tests do not reject the null hypothesis of cointegration presenting a structural break in 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, the tests do not support evidence on the existence of a cointegration relationship. In general, a possible explanation is the wheat prices were heavily influenced by weather events, that were reflected in the expectation about the stock levels overcoming the effect of the input prices more related to oil prices; in any case the results of Table 10 are not contrasting these last findings as they were related to a shorter period.

The results of the Gregory Hansen tests do not support the evidence of cointegration among the brent and the Italian commodity prices confirming the result obtained with the Johansen test. These results are coherent with our hypothesis of market integration that give priority to the US market signals.

Table 12. Cointegration test with structural break between It and US agri 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)

[12] 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 (Akaike, 1974, 1987).

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

The results of the statistics reported in table 12 confirm the evidence of cointegration between the Italian and US agri-commodity markets; more robust for wheat and soybean commodities. These results are in line with those obtained by running the cointegration test without structural breaks; however for corn, the evidence of cointegration is supported by Z_t^* and Z_α^* tests.

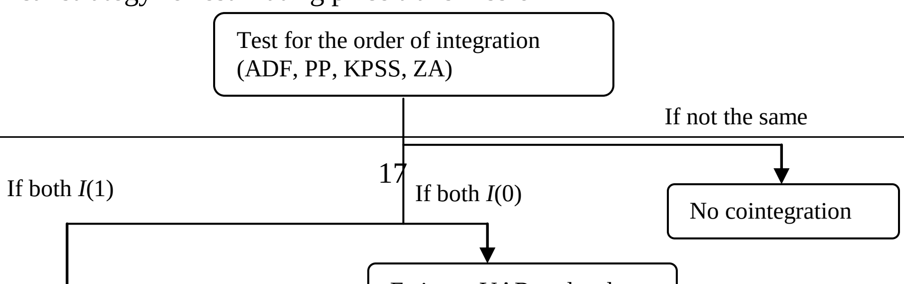
5.2. Price transmission

After the price cointegration, the analysis of price transmission between Italian prices (assumed to be the domestic market) and US ag commodity prices is performed and motivated by the hypothesis that the price co-movement in different markets is a condition of efficient, competitive market conditions, while the opposite reveals market inefficiency caused by scarce or asymmetric information, transport costs and others. Following the approaches suggested by Rapsomanikis *et al.* (2006) and Minot (2011), an econometric test of the impact of US food prices on Italian food prices has been carried out . The Minot procedure requires rather stringent assumptions, i.e. homogeneous cereal products, competition among numerous small traders, perfect information, no trade taxes or other policy barriers to trade, no transportation or transaction costs. In the above mentioned analysis, it is considered the vector error correction model (VECM) which assumes the domestic food price affected by the world price and examines three relationships: corn, soybeans and wheat. The VECM is appropriate if the following two conditions are satisfied: i) all variables are nonstationary and integrated *I*(1) following a random walk, but the first difference (X_t - X_{t-1}) is stationary *I*(0); ?? ii) the variables are cointegrated, meaning that there is a linear combination among the variables that satisfy the stationary condition . The cointegration equation is:

$$P_1 = \alpha + \beta P_2 + \varepsilon \qquad (6)$$

where the error term ε is stationary and the equation is comparable to eq (2) of the previous section.
For each pair of domestic and international prices, the analysis consists of more steps (partially investigated previously) outlined in Figure 2.The first one is to determine whether the individual variables are nonstationary or *I*(1); if both prices are not *I*(1), they cannot be cointegrated; if they are both stationary or *I*(0) they can be tested with the vector autoregressive (VAR) model that is a general framework used to describe the dynamic interrelationship among stationary variables. If the series are both *I*(1), the null hypothesis that they are not cointegrated is tested (in our case using the Johansen procedure). Finally, if the Johansen test indicates that there is a long-run relationship between the two variables, the vector error correction model (VECM) can be estimated. If the results of the test suggests no cointegration between the two variables, then they can be studied with a VAR on differences.

Figure 2. Basic empirical strategy for estimating price transmission



Source: Own elaboration based on Rapsomanikis et al. (2006)

The VECM tests for the impact of one variable over each other variable. In this study, the two-variable VECM tests the effect of world prices on domestic prices and the opposite. Since Italy is a "small country" in the staple foodcrop markets, there is little value in testing the effects of domestic prices on world prices; in addition, tests indicate that one lagged term is generally sufficient. Actually this research is focused in only one portion of the VECM. This portion can be simplified as follows:

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

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 and the term in parenthesis (p^d_{t-1}- βp^w_{t-1}) represents the long term transmission of international prices on Italian prices. The following two terms measure the short term impact of the lagged increments (Δ) of the natural logarithm of international and domestic prices. As β is unknown at first the equation (6) and after the equation (7) are estimated.

The coefficients in the error correction model can be interpreted as follows: the cointegration factor (β) describes how one price reacts to changes in the other in the long run[13]. The expected value for imported commodities is $0 < \beta < 1$, but for exports, it may be greater than 1. Thus, if $\beta = 0.5$, this implies that 50% of the proportional change in the international price will be transmitted to the domestic price in the long run (Minot, 2011).

The error correction coefficient (θ) reflects the speed of adjustment. It is expected to fall in the range of $-1 < \theta < 0$. 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 (that is, the closer to -1), the more quickly the domestic price (p^d) will return to the value consistent with its long-run relationship to the world price (p^w). The coefficient on change in the world price (δ) is the short-run elasticity of the Italian price

[13] Since prices are expressed in logarithms, β can be interpreted as the long-run elasticity of the domestic price with respect to the international price; it is the long-run elasticity of price transmission.

relative to the US price. In this case, it represents the percentage adjustment of domestic price one period after a one percent shock in international price. The expected value is $0 < \delta < \beta$.

The coefficient on the lagged change in the domestic price (ρ) is the autoregressive term, reflecting the effect of each change in the Italian price on the change in the same price during the next period. The expected value is $-1 < \rho < 1$.

Table 13 provides a summary of the results for the transmission of US prices to Italian ones: the unit root tests [results in section 4.1.] indicate all that domestic prices are not stationary and this result allows to to proceed with the Johansen cointegration test to see if there is a long-run relationship between the Italian and the international price for the same commodity [results in section 5.1.]. The cointegration test indicates that all the local prices have a long-run relationship with international corresponding prices.

The long-run elasticity of price transmission is statistically significant and high for all the commodities especially for soybeans (0.96) and wheat (0.74) meaning that a high percentage of the proportional change in the US price is transmitted 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 the all pairs of commodities. The auto-regressive term is statistically significant for all the variables and is higher for corn.

Table 13. Transmission of US food prices (world) to Italian food prices (domestic)

	Unit root in Italian prices?	Long run relationship?	Error correction model			
Commodity	ADF PP KPSS ZA	Johansen test	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

Summarizing the result obtained from the transmission model, the long run relationship prevails on the short run transmission for all the commodities in terms of value and significance. An important role takes also the autoregressive term meaning that approximately 42% of the change of the Italian corn price will be transmitted to the domestic price of the commodity in the subsequent period (20% for soybeans and almost 30% for wheat). Italian soybeans and soft wheat seem more connected to the international market than corn in the long run due to the high quantity of import[14].

The following steps will be performed for testing the existence and direction of causality among commodities.

5.3. Granger causality

5.3.1. Linear approach

After having demonstrated the evidence of cointegration, the next step consists to infer into causality using the vector autoregressive VAR to know how prices are spatially transmitted . For this reason the linear

[14]During the 2009-10 commercial campaign, Italy imported 60% of soft wheat, 87% of soybeans and 20% of maize (Associazione Nazionale Cerealisti, 2011)

causality tests is performed over the entire sample period, as well as on sample subperiods, to analyze whether the dynamic relationships between oil and ag-commodities prices change over time.

Table 14 reports the results of the Granger causality-test using the Durbin Watson diagnostic test (1950) to detect the presence of autocorrelation in the residuals from the regression analysis. The residuals of most of the estimated model confirm that they are free from autocorrelation (except the models obtained from the regression between oil and Italian agri commodities whose DW statistic results are substantially less than 2, with clear evidence of positive serial correlation).

The Ramsey Regression Equation Specification Error Test (RESET test; Ramsey, 1969), a general specification test for the linear regression model, clearly shows that the functional forms for the models are appropriately specified (with some exception in the comparison between oil and Italian commodities).

Table 14. Linear Granger causality test

Upper section	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-test	Probability	F-test	Probability	F-test	Probability	F-test	Probability
oil →cus	0.71	0.55	1.76	0.18	2.65*	0.07	0.25	0.61
cus →oil	0.92	0.43	1.05	0.35	0.35	0.71	1.01	0.32
Durbin-Watson	2.22		2.15		2.11		2.28	
RESET test	0.08	0.78	1.47	0.23	1.13	0.29	0.12	0.73
oil →sus	0.18	0.67	1.24	0.27	0.69	0.50	3.42**	0.03
sus →oil	0.38	0.53	0.83	0.36	0.17	0.84	1.46	0.23
Durbin-Watson	2.27		2.54		1.88		2.25	
RESET test	0.24	0.63	0.00	0.97	0.99	0.32	0.00	0.99
oil →wus	1.45	0.23	0.77	0.38	0.08	0.78	5.03***	0.01
wus →oil	0.77	0.38	1.29	0.26	0.61	0.44	0.51	0.60
Durbin-Watson	2.14		2.04		1.97		2.12	
RESET test	0.90	0.34	0.80	0.37	0.09	0.76	0.50	0.48

Mid section	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-test	Probability	F-test	Probability	F-test	Probability	F-test	Probability
oil →cit	2.08	0.15	0.17	0.68	0.00	0.97	0.24	0.79
cit →oil	0.34	0.56	1.71	0.19	1.19	0.28	0.07	0.93
Durbin-Watson	1.14		1.43		1.32		1.07	
RESET test	4.50	0.03	0.66	0.42	0.22	0.64	2.00	0.16
oil →sit	2.13	0.14	0.14	0.71	0.01	0.93	0.64	0.53
sit →oil	1.15	0.28	0.66	0.42	0.08	0.77	1.02	0.36
Durbin-Watson	1.61		1.43		1.67		1.75	
RESET test	11.15	0.00	0.30	0.59	0.48	0.49	3.11	0.08
oil →wit	0.45	0.64	0.32	0.57	2.36	0.13	0.83	0.43
wit →oil	0.02	0.98	0.03	0.86	1.61	0.21	1.19	0.31
Durbin-Watson	1.33		1.54		1.19		1.57	
RESET test	2.57	0.11	1.25	0.27	0.21	0.66	3.05	0.08

Lower section	1999w1-2012w21		1999w1-2004w29		2004w30-2008w40		2008w41-2012w21	
	F-test	Probability	F-test	Probability	F-test	Probability	F-test	Probability
cus → cit	2.36**	0.04	0.88	0.45	3.72**	0.03	1.77*	0.07
cit → cus	0.56	0.72	1.10	0.35	0.64	0.53	0.55	0.85
Durbin-Watson	1.14		1.40		2.08		2.34	
RESET test	0.01	0.92	0.89	0.45	2.51	0.12	1.01	0.32
sus →sit	4.49***	0.00	4.82**	0.03	14.89***	0.00	16.54***	0.00
sit →sus	0.19	0.31	3.34*	0.07	0.83	0.44	1.99	0.14
Durbin-Watson	1.76		1.46		1.90		2.01	
RESET test	0.30	0.58	0.52	0.47	0.14	0.71	0.71	0.67
wus → wit	14.10***	0.00	0.44	0.51	4.61***	0.00	7.44***	0.01
wit →wus	8.64***	0.00	0.01	0.93	1.86	0.12	1.68	0.20
Durbin-Watson	1.42		1.52		1.35		1.62	
RESET test	1.92	0.17	0.05	0.82	2.41	0.12	0.01	0.92

→ means non Granger causality hypothesis; ***/**/* denote statistical significance at 1%, 5% and 10% level, respectively. The optimal lag length was selected by Schwarz information criterion.

In the upper section of the table 14, values of F-statistic are reported with different probability levels for the null hypothesis of no causality between brent and the US agricultural prices (and *vice versa*). Results suggest the absence of causality from the oil prices to the agricultural prices, and the neutrality hypothesis in the entire period of observation; for the sub-samples analysis, these results are similar to those obtained by Nazlioglu (2011). The middle section reports the results for the Granger Causality analysis between oil and the Italian commodities. In general it is observed the absence of linear Granger causality between oil and agri-commodity prices consistent with Zhang and Reed (2008) results for local agri-commodities.

The lower section of the table reports the interaction between US and Italy markets: there is evidence of linear Granger causality going from US to Italy markets, with exception for wheat showing feedback. In all three cases the US ag-commodities do Granger-cause the Italian prices for the entire sample period and sub-sample ones.

5.3.2. Non linear approach

The non linear causality test is carried out in two steps: firstly it is applied to stationary series, and then, to the estimated residual series to remove any linear dependence, using the VAR model applied to the pairwise variables of interest. “By removing linear predictive power with a linear VAR model, any remaining incremental predictive power of one residual series for another, can be assumed as non linear predictive power” (Hiemstra and Jones, 1994).

The tests are performed for different lag values depending on the length of the sample, and the data are normalized to unit variance before running the test; the bandwidth value that plays an important role on the detection of non linear causality, is set to 1, as it is one time the standard deviation (Dicks and Panchenko, 2006). Because nonparametric tests rely on asymptotic theory, causality tests on sample subperiods are not performed in this case. Table 15 reports the t values for Diks and Panchenko’s test statistic applied to the variables and to residuals in both directions and for different lag lengths (1–2 lags) to give evidence of non linear Granger causality.

Table 15. Nonlinear Granger causality (Diks –Panchenko test)

		Raw data		Residuals	
	lags	**1**	**2**	**1**	**2**
oil-us_commodities	oil →cus	2.019**	2.788**	1.434*	2.282**
	cus →oil	0.210	0.812	0.450	0.144
	oil →sus	1.686**	1.826**	1.570*	1.796**
	sus →oil	0.395	0.172	0.311	0.250
	oil →wus	1.290*	1.282*	0.968	1.329*
	wus →oil	1.969**	1.829**	2.140**	2.135**
oil-it_commodities	oil →cit	1.491 *	0.844	1.133	1.238
	cit →oil	0.765	0.435	0.674	0.732
	oil →sit	0.546	0.633	1.702 **	1.166
	sit →oil	0.446	1.670 **	0.360	1.284 *
	oil →wit	2.627 ***	2.473 ***	1.133	1.238
	wit →oil	1.061	1.724 **	0.674	0.732
us-it commodities	cus →cit	3.574 ***	3.596 ***	3.585 ***	3.578 ***
	cit →cus	2.682 ***	2.716 ***	3.076 ***	3.079 ***
	sus →sit	2.704 ***	2.697 ***	3.081 ***	3.065 ***
	sit →sus	2.011 **	1.983 **	1.746 **	1.732 **
	wus →wit	2.547 ***	2.537***	2.252 **	2.221 **
	wit →wus	1.613 *	1.596 *	2.236 **	2.202 **

→ means non Granger causality hypothesys. ***/**/* denote statistical significance at 1%, 5% and 10% levels of significance, respectively.

By observing the the first part of the table, there is evidence of unidirectional causality going from **oil to corn** and **soybean** for raw data, at one and two lags, confirmed after filtering the series with VAR model and testing the residuals a condition that persisted in the long period. The economic meaning is that the oil price volatility is transmitted to corn and soybean commodity prices confirming the results obtained by Nazlioglu (2011). Until early 2007, corn prices were not affected by the crude oil prices; since then, corn prices have been growingly responsive to changes in crude oil prices justified by the growing amount of US corn used for fuel ethanol, now being around 40% of total corn production. With the rapid growth of the ethanol industry in the last few years, corn has become very much an energy crop as well as the world's most important source of feed grains for production of livestock, poultry, and dairy products.

The causality between **oil and wheat**, tested with raw data shows a weak statistical evidence of causality going from oil to wheat and stronger evidence for the reverse direction. Further investigation with the VAR residuals confirm these results at lag one and two. These findings emphasize the market differences between wheat and the other two commodities; the minor dependence of this cereal from the oil market is justified by the prevailing use in the food industry. Apart from oil influence, there are other exogenous factors determining the price volatility of wheat; weather uncertainty as long periods of droughtness, becoming more frequent in last years could have had a greater impacts on supplies and prices, as well as extremely wet periods, and wider range of temperature fluctuations.

The second section of table 16 reports the results for the nonlinear causality analysis between *brent* and the *Italian commodities:* the general evidence is that there is no dependence between brent and agri-commodity prices. In the case of cit and wit, after removing the linear dependence the non parametric test results support the neutrality hypothesis; with respect to the non linear causal linkages between brent and sit prices, results suggest non clear relationship.

Finally, the last section of the table indicates presence of non linear causality between the US and the Italian commodity prices. Raw data provide a feed-back evidence of relationship between the prices and the similar results are obtained after filtering the series. Such misleading results could be explained with the fact that there is a one way strictly *linear causality* among the examined variables (see Table 15), whereas a unidirectional non linear causality from the American to the Italian commodity prices does not seems so evident.

6. CONCLUSION

The increasing world-wide market volatility of agricultural commodities in these last years with prices hiking in 07-08, suggested to infer into the causes generating wider fluctuation in crude oil prices, as the declining supply, higher extraction costs and greater financial speculation, and transmitted to US and Italian agricultural commodity markets. The relationship between food and oil prices has attracted much of the experts' attention because of the increase in ag-commodity prices observed in the first part of 2008 and in the second part of 2010. Oil is affecting directly (fertilizer, fuel) or indirectly (machinery) most of the inputs used in agricultural production. Another possible cause of higher price fluctuation could have been induced by biofuel policies that stimulated the production of agri-commodities used for the biofuel industry with a sharp increase of the demand against a quite rigid supply. These policies may or may not hinder the market integration depending if they are directed to protect domestic markets, however this do not seems the case as a lot of feedstock is currently traded around the world and only in few cases Argentina and Russia have applied for security purposes protectionist policies. The exchange of ag-commodities is quite fluid between US and Italy supporting the hypothesis of market integration between the two countries.

Time series analysis has been used to test the hypothesis of market integration, price transmission and to verify the inherent dynamic market relationships due to factors as discontinuity or non linearity arising from market inefficiencies in arbitrage. To find the cause-effect relations, the research has been directed to test the existence of cointegration, price transmission and Granger causality among the prices. The results of the linear Granger causality analysis suggest to accept the presence of neutrality hypothesis in the US markets which means that the prices of oil and the US agricultural commodities do not cause each other in a strictly linear sense. Similar results are evident for oil and the Italian market; anyhow, Italian prices are co-integrated with US prices of similar agri-commodities and it has been found evidence of linear unidirectional Granger causality running from US to Italian markets or from global to local agri-commodity markets. These results confirm the Law of One Prices then US and Italian ag-commodity markets are working efficiently.

Non linear components of price trends are observed for oil, particularly evident in the period of hike in price. The Diks Panchenko test confirm the existence of non linear relationship between oil and the agri-commodities. For US the crude oil is confirmed to be a representative *leading indicator* also used for the formation of agri-commodity prices.

However, the oil price is a destabilizing factor for corn and soybeans markets because these markets are now closely linked to the biofuel prices. The case of wheat prices is completely different: the world has consumed more wheat than has been produced in the last six or seven years. The resulting drawdown in wheat stocks is the main responsible for the increase in wheat prices. This perception of food insecurity, due to the diminishing supply of flours, has brought wheat prices to surge upwards dramatically for the financial speculation prevailing on market fundamentals.

With the current large size of the ethanol industry, corn prices have become closely related to crude petroleum prices because corn is now the major energy crop in the northern hemisphere as well as the world's most important source of feed grains for production of livestock, poultry, and dairy products. This market evolution has strengthen the relationship between corn prices, crude oil and ethanol.

Volatility becomes an important issue for policy analysis when it induces risk averse behavior that leads to inefficient investment decisions creating problems that are beyond the capacity of producers, consumers or nations to cope with. To be effective, the market policies need to have unbiased information about the agri-food supply chain, producers, consumers and traders to reduce the risk of market volatility. It is necessary to focus on the policy options designed to prevent or reduce price volatility and mitigate its consequences: some would help to avert a threat, others are in the nature of contingency plans to improve readiness, while still others address long-term issues of resilience.

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