UNIVERSITÀ POLITECNICA DELLE MARCHE

DIPARTIMENTO DI SCIENZE ECONOMICHE E SOCIALI

AGRICULTURAL PRICE TRANSMISSION ACROSS SPACE AND COMMODITIES DURING PRICE BUBBLES

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QUADERNO DI RICERCA n. 367*

Novembre 2011

*** La numerazione progressiva continua dalla serie denominata:**
Quaderni di Ricerca - Dipartimento di Economia

AGRICULTURAL PRICE TRANSMISSION ACROSS SPACE AND COMMODITIES DURING PRICE BUBBLES

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Abstract

The objective of this paper is to investigate agricultural price transmission during price bubbles and to assess whether the implemented trade policy measures did eventually play a role. We study horizontal cereal price transmission both across different market places and across different commodities. The analysis is performed using Italian and international weekly spot (cash) prices in the years 2006-2010, a period of generalized exceptional exuberance and consequent rapid drop of agricultural prices. Firstly, the properties of the price series are explored to assess which data generation process may lie behind the observed patterns. Secondly, the interdependence across prices is estimated adopting appropriate cointegration techniques. Results suggest that most prices behave as I(1) series, though some also show either fractional integration in the first differences or explosive roots. A long-run (cointegration) relationship occurs among prices of the same commodity across different markets but not among prices of different commodities. In both long-run and short-run relationships the "bubble" seems to have played a role as well as the consequent policy intervention on import duties.

Key-words: Price Transmission, Price Bubbles, Time Series Properties, Cointegration
JEL Classification: Q110, C320

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Contents

Agricultural Price Transmission Across Space and Commodities During Price Bubbles

1. Introduction: objectives and data description

The analysis of agricultural price transmission during intense market turmoil represents a major research challenge. The investigation may be empirically problematic due to the particular stochastic properties assumed by the price series during these periods of turbulence; however, it is clear that studying the co-movement of prices under these conditions is particularly useful to understand to which extent prices are linked together. The main motivation of the present study is to investigate the properties of agricultural price series over the recent price rally (European Commission, 2008; Irwin and Good, 2009)[1] to better analyse how price shocks were transmitted horizontally, that is across market places and commodities. We use cereal weekly spot prices from May 2006 to December 2010. We opt for this time coverage for three major reasons. Firstly, this period fully contains the bubble: the bubble firstly inflated, then completely deflated and finally started raising again in the second half of 2010 (Figure 1). Secondly, we can assume an almost-constant policy regime in the European Union (EU). In 2006, the 2003 reform of the Common Agricultural Policy (CAP) was entirely in place, included its limited implications in terms of price policy and market intervention; we can then assume that the domestic policy regime remained constant over the years 2006-2010. Concerning international trade policy, the only relevant policy regime change has been the temporary suspension of EU import duties on cereals from January 2008 to October 2008, as a reaction to the exceptionally tight situation on world markets. This temporary measure might have altered the price transmission mechanisms. Therefore, while investigating price transmission during the bubble, it is also possible to assess the role played by this single and temporary policy measure. Thirdly, concentrating on this period facilitates international price comparisons as the cumulative inflation rate has been quite limited and relatively similar in Italy and in North-America (the two areas under study here). Therefore, comparisons of agricultural prices across different countries do not require the deflation of nominal prices into a common real base.

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[1] Henceforth, the "agricultural price bubble". Heuristically, and generically, the bubble clearly appears in Figure 1. From a more rigorous point of view, however, we will formally define and test the presence of a price bubble in section 2.3.

We analyse weekly spot price series observed in different Italian locations (from North to South; source: ISMEA) and international (North-American) markets (source: International Grain Council, IGC) (Esposti and Listorti, 2011). The reason why we focus on durum wheat and corn is that they represent somehow opposite cases. Amongst the main cereals, durum wheat experienced the largest price rise (and, then, decline) during the bubble, while corn prices showed the smallest variation (see Figure 1 and Esposti and Listorti, 2011). Comparing these two extreme cases is particularly insightful to understand whether some common features of the price movements can be found even across commodities showing quite diverse market fundamentals. Significant differences between durum wheat and corn may be found on both the demand and the supply side, at least in the Italian case. On the demand side, the prevailing domestic (Italian) uses are very different for the two cereals: almost exclusively human consumption for durum wheat, prevalent feed use for corn. Therefore, we can deduce a pretty limited interdependence among them, since they behave neither as strong complements nor as strong substitutes. On the supply side, Italy is the largest EU durum wheat producing country (durum wheat is one of the characteristic products of the Italian agriculture) while, on the contrary, corn is a less typical production. Therefore, we can presume a different linkage between the national and the international prices in the two cases.

The dataset has a Tx(NxK) dimension where: T=244 weeks (from the first week of May 2006 to the last week of December 2010); K=2 commodities (durum wheat and corn); N=5 market places, that is, North-Italy (Milan), Centre-North Italy (Bologna), Centre-South Italy (Rome), South-Italy (Foggia), US and Canada (or Rotterdam; see below). The codification and description of the NxK=10 price series under investigation are reported in Table A.1 (Annex 1).

The North-American prices are FOB prices. For agricultural commodities, freight rates are normally quite relevant. Moreover, due to volatility in energy, namely oil, prices, they also considerably oscillated during the commodity price bubble. For these reasons, freight rates (source: IGC) were added to US and Canadian prices in order to obtain the respective CIF prices. The freight rates used in this study are those from US Gulf (or Canada, where applicable) to Amsterdam/Rotterdam/Antwerp/Hamburg destinations. In practice, in our analysis the North-American prices actually serve as international prices taken at Rotterdam, thus as EU-reference prices. Henceforth, we will also refer to international prices as Rotterdam prices (Table A.1). These US and Canadian CIF prices have been finally converted from US dollars to Euros by using the weekly official $/€ exchange rate provided by Eurostat-ECB.[2]

2. Some basic evidence on agricultural price behaviour

Before analysing price interactions and the reciprocal transmission of price shocks, the time series properties of the data have been assessed. This allows identifying the common features of the price series, in order to achieve a proper specification of price transmission/interdependence relations.

Let us consider agricultural prices observed over three different dimensions: space, commodity and time. Therefore, the generic price is $p_{i,k,t}$ where: i=1,…,j,…,N is the (local) market place (spatial dimension); k=1,…,h,…,K is the commodity; t=1,…,s,…,T is the period of observation (time dimension). By more conventionally distinguishing between a cross-sectional dimension, given by the combination of the dimensions *ik*, and a time dimension *t*, we can identify any generic price observation as $p_{ik,t}$ (scalar) and any generic price series (vector) as $\mathbf{p}_{ik}$. The logarithms of prices are here considered. This monotonic transformation facilitates the economic interpretation of results, in particular considering that regression coefficients may be interpreted as elasticities. Therefore, henceforth $\mathbf{p}_{ik}$ identifies the time series of the price logarithm of the *k-th* commodity in the *i-th* market place.

The time-series properties of the prices are analysed in the following sections by testing, in sequence, stationarity, persistence (long memory or fractional integration) and explosiveness to assess whether these features may be invoked as possible causes of the observed exuberance.[3]

2.1. Stationarity

The presence of unit and/or explosive roots is critical to understand the behaviour of price series especially in periods of such dramatic exuberance and drop. Stationary (i.e. I(0)) series can be hardly reconciled with the presence of the bubble. Even stationarity around a drift (a constant term) and/or a deterministic trend is not evidently helpful in this respect. As clarified below, testing for the presence of a temporary explosive pattern cannot be simply achieved

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[2] Using prices already converted in the same currency is a widely used procedure; it follows that adjustment to the exchange rate is assumed instantaneous.

[3] Other time-series properties that could generate non-linear dynamics in price patterns, in particular non-normality and seasonality, have been tested and generally excluded (see Esposti and Listorti, 2011, for more details).

through conventional unit-root tests. Nonetheless, albeit not sufficient, these tests still allow to assess a necessary outcome of nonlinearities within price series: the series are expected to be I(1) and not I(0) (Diba and Grossman, 1988; Evans, 1991; Phillips and Magdalinos, 2009; Philips et al., 2009; Phillips and Yu, 2009).

Table A.3 (Annex 2) reports unit root tests on $\mathbf{p}_{ik}$ and on their respective first differences, $\mathbf{\Delta}_1\mathbf{p}_{ik}$. Four different unit-root tests are run (Enders, 1995; Stock and Watson, 2011). The first is the conventional ADF (Adjusted Dickey-Fuller) test. This test, however, encounters known limitations particularly for series like those under consideration here and, thus, may provide misleading evidence on their underlying stochastic properties. The second test is the PP (Phillips-Perron) test, that is expected to be more robust than the ADF test under heteroskedasticity which may, in fact, occur during periods of exuberance. Furthermore, two other tests are performed, both aiming at taking into account that the conventional ADF test tends to accept the null hypothesis of unit root also for series that are, in fact, only "near unit root" processes. This latter case is likely to occur in the price series under consideration here, especially in their first differences. Firstly, an ADF-GLS test is performed. This modified ADF test has significantly greater power than the conventional test especially under near unit root processes and small sample size. Secondly, a KPSS test (Kwiatkowski, Phillips, Schmidt and Shin) is performed. In this test, the null hypothesis is that the series is stationary, while the alternative hypothesis is that it is I(1). Therefore, the KPSS test is expected to reveal those series that the conventional ADF test tend to accept as I(1) while, in fact, they are only near unit-root processes. The combination of these four unit-root tests should provide an exhaustive picture on the real underlying stochastic processes of the price series thus allowing for a robust and conclusive answer about their order of integration.

In general terms, looking at Table A.3 it seems that, once the proper specification has been selected (in terms of number of lags and of the presence of drift and deterministic trend), all $\mathbf{p}_{ik}$ series show a unit root. If the conventional ADF test is considered, the evidence about the $\mathbf{\Delta}_1\mathbf{p}_{ik}$ series is more mixed, though I(1) series still prevail on I(2). As mentioned, however, conventional ADF tests may fail in detecting stationary series when their roots are near to unity. The more robust PP, ADF-GLS and KPSS tests, in fact, show that all first differenced series are stationary. The conclusion would be that all price series can be considered I(1) but not I(2).

I(1) series (random walks possibly with a drift and/or a deterministic trend), however, are apparently at odds with the evidence of a price bubble, as they can hardly generate the temporary exuberance observed in all markets. Apparently, these prices are "something more" than I(1) series. Two plausible explanations are the following. The first explanation is that $\boldsymbol{\Delta}_1\mathbf{p}_{ik}$, though I(0), may show long memory, therefore long persistence of shocks. A second explanation is that temporary explosive roots (the "bubble") actually occurred in $\mathbf{p}_{ik}$. As conventional unit-root tests can not really assess (or exclude) these two hypotheses, *ad hoc* tests are needed.

2.2. Long memory (fractional integration)

As emphasized by Wei and Leuthold (1998), agricultural prices (mostly, in fact, future prices) are often characterised by long memory, which may also generate non-linear patterns quite close to chaotic processes. In such cases, price series are neither I(0) nor I(1) but rather I(d) processes, with $0<d<1$ (hence the term "fractional integration"). Fractional integration implies that price series, though not behaving as random walks (where shocks never vanish over time), still keep the memory of a given shock for a long period. Roughly speaking, price shocks decay very slowly over time. This kind of stochastic process is here of specific interest since when fractional integration occurs in the first differences, then I(1) price series can generate the observed nonlinearities. In such cases, though a unit root cannot be detected in the first-differenced series, a long-memory process would still imply strong persistence of shocks.

Unfortunately, conventional unit-root tests may fail in detecting such property: they assess whether time series are I(0) or I(1) while, in fact, they are I(d). The presence of long memory within the price series can instead be tested following the approach originally proposed by Geweke and Porter-Hudak (1983) and then modified by Phillips (1999a,b). This test is based on a particular representation of the stochastic process generating the price series $\mathbf{p}_{ik}$, or its first differences $\Delta_1\mathbf{p}_{ik}$, called ARFIMA(p,d,q) (Autoregressive Fractionally Integrated Moving Average) model, where p and q express, as usual, the orders of auto-regressive and the moving-average parts, respectively, and d the order of (fractional) integration. The procedure proposed by Phillips (1999a,b), and adopted here, tests the value of parameter d then distinguishing stationary, unit-root and fractionally integrated processes. The procedure produces two test statistics, one for the null d=0 and one for d=1. If d=0 is accepted the series is stationary; if d=1 is accepted the series has an unit root. If both are rejected (namely, $0<d<1$), then fractional integration (long memory) is accepted.

Results of these tests are reported in Table A.4 (Annex 2). They confirm that, when applied to the original series, a unit root is evidently observed. When applied to first differences, the presence of a unit root is always rejected but in three national markets (two for durum wheat, one for corn) the presence of fractional integration cannot be excluded.

2.3. Testing explosiveness: recursive unit root tests

In time-series econometrics, nonlinear patterns can naturally emerge in non-stationary variables whose first differences are fractionally integrated processes or, *a fortiori*, second-order integrated, I(2), series (Engsted, 2006). Nonetheless, a price series showing a period of explosive behaviour (a temporary, or periodically collapsing, bubble) is not necessarily an I(2) series. On the contrary, a I(2) process would imply a permanent exuberance of prices while, in fact, the observed bubble inflates and deflates within a relatively limited period of time. In fact, bubbles induce a temporary explosive root in price series in addition to a unit root. If this additional root is not appropriately considered, conventional testing may fail in detecting the real underlying stochastic process (Evans, 1991).

Recent works by Phillips and Magdalinos (2009), Philips et al. (2009) and Phillips and Yu (2009) have provided an appropriate framework for assessing the presence of an explosive root within processes that would otherwise be ruled as I(1). They propose a test for the presence of bubbles in which forward recursive ADF tests are run on the price series. These sequential tests allow assessing period-by-period the possible nonstationarity of the price series against an explosive alternative. The forward recursive test is based on a conventional ADF regression where in the first recursion only $T_o = [r_oT]$ observations are used, where r_o is a fraction of the total sample T.[4] In subsequent regressions this initial data set is supplemented by successive observations, each time using a sample of size $T_r = [rT]$ for $r_o \leq r \leq 1$. For any recursive sub-sample T_r, the respective ADF test is computed. Of these forward recursive ADF tests (ADF_r), the test of explosiveness considers the maximum observed value: $SADF = \sup_{r\in[r_0,1]} ADF_r$. Under the null hypothesis of unit root (H_0: ρ_{ik} =1) and against the right-tailed alternative hypothesis of an explosive root (H_1: ρ_{ik} >1), we accept that the series contains an explosive root if the estimated test *SADF* is higher than the respective critical

[4] The brackets in $[r_o\, T]$ indicate that the integer part of the argument is taken.

values (reported, for different sample sizes, in Phillips et al., 2009, and Phillips and Yu, 2009). Here, r_o has been alternatively fixed at 0.10 (24 observations) and at 0.20 (48 observations) (see Phillips et al, 2009, and Phillips and Yu, 2009), and then incremented by each single following observation.

These *SADF* tests have been performed on $\mathbf{p}_{ik}$ including the constant term and 12 lags[5]. Table A.4 (last two columns) reports the results of these tests for both r=0.10 and r=0.20. The latter case, however, provides a more robust evidence since it uses larger sub-samples, thus reducing the risk of poor test performance in the first runs of the recursive process.[6] For this reason, a restrictive 1% critical value is used to test the presence of explosive roots. On this basis, we can conclude that a temporary explosive behaviour is definitely found only in durum wheat international (Rotterdam) price. The presence of an explosive root is doubtful in few cases, while it can be excluded in all other price series.

This kind of test may be particularly helpful also to understand the timing of price exuberance to eventually date the beginning and the end of the price bubble. To locate the origin and the conclusion of the exuberance, one can display the series of the above mentioned forward recursive ADF_r test and check if and when ADF_r exceeds the right-tailed critical values of the asymptotic distribution of the standard Dickey–Fuller t-statistic (Phillips et al. 2009). By adopting a restrictive 1% critical value, the price bubble turns out to be limited, especially in national markets, to a very few cases. Furthermore, it lasts for a short period of time, anticipating and only partially overlapping the period of suspension of the EU import duties on cereals (Figure 2). According to these results, and following these restrictive criteria (as explosiveness may be easily confused with different processes generating similar patterns), we will henceforth consider that only the international durum wheat price, *cwad_can*, clearly contains an explosive root. Taking this price as reference, we assume that the bubble begins the first time we observe ADF_r exceeding the 1% critical value (the first week of July 2007) and ends the last time we observe this limit being exceeded (the last week of March 2008).

[5] The longest significant lag found in performing conventional ADF tests is 12. Therefore, 12 are the standard lags considered in the testing procedures for both explosiveness and fractional integration (section 2.2 and Table A.4 in Annex 2).

[6] Due to this poor performance in the first runs, Figure 2 reports test results from January 2007.

3. Modelling price interdependence as price transmission mechanisms

In the previous section, evidence emerged in favour of the presence of "something more" than a simple I(1) process in the price series under study. This additional stochastic property may be the presence of fractional integration in $\Delta_1 \mathbf{p}_{ik}$ or of explosive roots in $\mathbf{p}_{ik}$. Nonetheless, this evidence is not concordant across all cereal prices while, in fact, visual inspection of price patterns (Figure 1) suggests that they all tended to move together over time, though with marked differences. Therefore, to directly look for interdependence across the price series might be much more insightful than examining their individual properties in search of some common feature.

3.1. A general model of price transmission

Let us consider the generic agricultural price $p_{i,k,t}$. The behaviour of $p_{i,k,t}$ over its three dimensions might be evidently represented within appropriate structural models as the combination of market fundamentals such as supply, demand and stock formation. However, such models are inherently very complex and hardly tractable in the empirical analysis, whereas the investigation of price evolution and linkage is more frequently afforded within reduced-form models. For example, Fackler and Goodwin (2001) provide a common template based on linear excess demand functions and embracing all dynamic regression models from which an estimable reduced-form model (in their case, a VAR in prices) can be derived. When all the three dimensions are explicitly considered, reduced-form models are actually and by far more feasible and of immediate use to generate price predictions, i.e. the estimation of $E\left(p_{i,k,t} \middle| p_{j,k,t-s}, p_{i,h,t-s}, p_{i,k,t-s}\right)$, given the available observations.

By distinguishing a cross-sectional dimension, *ik*, and a time dimension, *t*, a generic reduced-form model of price formation and transmission over these two dimensions is the following:[7]

(1a) $$p_{ik,t} = \alpha_{ik} + \sum_{s=1}^{s=S<T} \rho_s p_{ik,t-s} + \sum_{s=0}^{s=S<T} \sum_{jh \neq ik} \omega_{ik,jh}^{s} p_{jh,t-s} + \varepsilon_{ik,t}$$

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[7] If we want to maintain the original three-dimension specification, equation (1a) can be more extensively written as: $p_{i,k,t} = \alpha_{i,k} + \sum_{s=1}^{s=S<T} \rho_s p_{i,k,t-s} + \sum_{s=0}^{s=S<T} \sum_{j \neq i} \omega_{ij}^{s} p_{j,k,t-s} + \sum_{s=0}^{s=S<T} \sum_{i=1}^{N} \sum_{h \neq k} \varpi_{kh}^{s} p_{i,h,t-s} + \varepsilon_{i,k,t}$ with $\varepsilon_{i,k,t} \sim N\left(0, \sigma_{i,k,t}^{2}\right)$.

where S is the maximum time lag and $\varepsilon_{ik,t} \sim N\left(0, \sigma^2_{ik,t}\right)$. In a more compact matrix form equation (1a) can be written as:

$$\text{(1b)} \quad \mathbf{P} = \boldsymbol{\alpha} + \sum_{s=0}^{s=S<T} \mathbf{P}_s \mathbf{W}_s + \boldsymbol{\varepsilon}_t$$

where $\mathbf{P}$, $\mathbf{P}_s$ and $\boldsymbol{\varepsilon}_t$ are (Tx(NxK)) matrices, s expresses the time lag, $\boldsymbol{\alpha}$ is a (Tx(NxK)) matrix of time invariant parameters, that is, $\alpha_{ik,t} = \alpha_{ik,t-s} = \alpha_{ik}$, $\forall i, j, s$ (any column of $\boldsymbol{\alpha}$ contains T elements with constant value α_{ik}) and $\boldsymbol{\varepsilon}_t \sim N\left(\mathbf{0}, \boldsymbol{\Omega}_t\right)$. $\mathbf{W}_s$ is a ((NxK)x(NxK)) matrix of unknown parameters incorporating the correlation across prices within both the time and cross-sectional (space-commodity) dimensions. The diagonal elements, $\omega^s_{ik,ik}$, actually indicate the auto-correlation over time, with the exclusion of the matrix $\mathbf{W}_0$, where diagonal elements are evidently $\omega^0_{ik,ik} = 0, \forall ik$. The off-diagonal elements, $\omega^s_{ik,jh}$, represent the cross-sectional dependence of prices; in other words, they express the interdependence among the different prices and, therefore, the degree and the direction of transmission of the price shocks.

In particular:

- if $h=k$ but $i\neq j$, we are considering the price transmission for the same commodity across space, that is, different market places. In this case, under perfect spatial arbitrage, the validity of the Law of One Price (LOP) implies that $\omega_{ik,jk} = 1$;
- if $i=j$ but $h\neq k$, we are considering the price transmission between two different commodities in the same market. In this case, elements $\omega_{ik,ih}$ indicate the degree of substitutability between the different goods (Dawson et al., 2006). $\omega_{ik,ih}$ will be close to 1 (-1) under perfect substitutability (complementarity) between h and k, while it will be close to 0 under low substitutability (complementarity).

As $\mathbf{p}_{ik}$ indicates the logarithms of prices, the elements of $\mathbf{W}_s$ actually express the price transmission elasticities. Within this logarithmic form, the implicit assumption is that all factors possibly contributing to price differentials but not explicitly taken into account in the model (for example, transportation and transaction costs) are a constant proportion of prices.

These constant multiplicative terms (that can be naturally intended as percentages) apply to price $p_{ik,t-s}$ to obtain $p_{jh,t}$ and are captured by the elements of $\boldsymbol{\alpha}$.[8]

If the matrix of unknown parameters, $\mathbf{W}_s$, contains all the information about price linkages over the three dimensions, we can expect that the transmission equations (1a-c) get rid of possible autocorrelation and heteroskedasticity across both the time and cross-sectional dimensions; that is, we can assume that spherical error terms are restored: $\boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \sigma \mathbf{I})$ and $E(\boldsymbol{\varepsilon}_t, \boldsymbol{\varepsilon}_{t-s}) = 0$. The proper specification of (1a-b) aims indeed at restoring such conditions.

3.2. Model specification

The first specification issue concerning model (1a-b) has to do with its size. With 10 price series, the size of $\mathbf{W}_s$ becomes large especially whenever several lags have to be admitted due to the use of weekly data. To reduce the number of parameters to be estimated, and also to facilitate the economic interpretation of the results, the analysis of the price interdependence must be "confined" and "segmented". In this respect, assumptions can be made about the relevant interactions to be considered. In particular, we firstly assume that price transmission only occurs within the same commodity (p_{ik} and p_{jk}) and within the same market place (p_{ik} and p_{ih}). The consequent assumption is that p_{ik} and p_{jh} have no direct linkage. This implies fixing at 0 some of the elements of $\mathbf{W}_s$. Secondly, the relation across prices can be studied by sub-groups of commodities segmenting the analysis within the fixed-commodity/cross-space(market) dimension (p_{ik}, p_{jk}) from the analysis within the fixed-space(market)/cross-commodity dimension (p_{ik}, p_{ih}). Table A.2 in the Annex 1 shows in detail the sub-groups of prices within which interdependence has been considered.

When specifying and estimating model (1a-b), the fact that all $\mathbf{p}_{ik}$ series can be considered I(1) processes, with some also showing long memory in the first differences and one (the international durum wheat price) showing a temporary bubble, must be appropriately taken into account. Since the seminal work of Ardeni (1989), cointegration techniques have been extensively used for the study of agricultural price transmission mechanisms. Cointegration models presuppose that I(1) variables are linked by a long run (LR) relation, whose residuals are stationary. When fixed-commodity and cross-market price relations are considered, under

[8] See Fackler and Goodwin (2001) for a comprehensive explanation.

perfect spatial arbitrage, this relation is the LOP, which is expected to hold in the LR while in the short run (SR) prices are allowed to deviate from it.

In such circumstances, the price transmission equation takes the form of a standard Vector Error Correction Models (VECM):

$$(2)\ \boldsymbol{\Delta}_1\mathbf{p}_t = \boldsymbol{\alpha\beta}'\mathbf{p}_{t-1} + \sum_{i=1}^{S-1}\Gamma_i\boldsymbol{\Delta}_1\mathbf{p}_{t-i} + \boldsymbol{\varepsilon}_t$$

where $\mathbf{p}_t$ now is the (Vx1) vector containing the logarithms of the V prices at time t over the selected sub-group and dimension (space ij, commodity kh);[9] $\boldsymbol{\beta}$ is the cointegration matrix containing the long-run coefficients (the degree of price transmission); $\boldsymbol{\alpha}$ is the loading matrix containing the adjustments parameters (a measure of the speed of price transmission); $\boldsymbol{\Gamma}_i$ are matrixes containing coefficients that account for short-run relations; $\boldsymbol{\varepsilon}_t$ are white-noise error terms. The rank of $\boldsymbol{\alpha\beta'}$ gives information about the presence of cointegration amongst the variables. As the model is expressed in logarithms, the fundamental assumption underlying (2) is that price spreads (and also, all components which account for price spreads) are a stationary proportion of prices.

As already mentioned, however, we cannot exclude the presence of explosive behavior in some of the price series. Nonetheless, Engsted (2006) and Nielsen (2010) show that the Johansen (1995) approach to test and estimate cointegration relationships still holds its validity. Indeed, the cointegrated VAR model developed by Johansen turns out to be an "ideal framework" for analyzing the linkage between variables that have a common stochastic trend (they are cointegrated), but in which one of the series also has an explosive root. The Johansen method makes it possible to estimate the cointegrating relationship even though the relationship contains this explosive component. Under these circumstances, it is possible to rewrite equation (2) in a form that admits two structural relations. The first contains the usual cointegrating parameters (their linear combination is not I(1)); the second contains the co-explosive ones (their linear combination is not explosive) (Engested, 2006, 157):

$$(3)\ \boldsymbol{\Delta}_1\boldsymbol{\Delta}_\rho\mathbf{p}_t = \boldsymbol{\alpha}_1\boldsymbol{\beta}_1'\boldsymbol{\Delta}_\rho\mathbf{p}_{t-1} + \boldsymbol{\alpha}_\rho\boldsymbol{\beta}_\rho'\boldsymbol{\Delta}_1\mathbf{p}_{t-1} + \sum_{i=1}^{S-2}\Gamma_i\boldsymbol{\Delta}_1\boldsymbol{\Delta}_\rho\mathbf{p}_{t-i} + \boldsymbol{\varepsilon}_t$$

[9] As evident in Table A.2, V always ranges between 2 and 4.

where $\mathbf{\Delta}_{\rho} = (1 - \rho L)$ and ρ is the explosive (ρ>1) root. The conventional cointegration vector is $\boldsymbol{\beta}_1$ as it can be demonstrated that $\boldsymbol{\beta}_1 = \boldsymbol{\beta}$ (Nielsen, 2010), while $\boldsymbol{\beta}_{\rho}$ contains the co-explosive parameters. All other parameter matrices can be interpreted accordingly.

As the standard Johansen estimation procedure maintains its validity, and since we are interested in the long-run relationship among prices, we can simply proceed in estimating (2). However, in our analysis we also admit that the period of exuberance (or the shock that generated it) influenced the cointegration relationship, $\boldsymbol{\beta}_1 = \boldsymbol{\beta}$, itself. This can be done by allowing for the presence of structural breaks within the cointegration relationship. In this respect, Johansen et al. (2000) generalized the standard Johansen cointegration test by admitting up to two predetermined breaks in the cointegration space. They propose a model where breaks in the deterministic terms occur in known points in time. The time series is divided in *q* sub-periods, separated by the occurrence of the structural breaks, where *j* denotes any generic sub-period. The general VECM becomes:

$$(4)\quad \mathbf{\Delta}_1 \mathbf{p}_{\mathrm{t}} = \boldsymbol{\alpha} \begin{bmatrix} \boldsymbol{\beta} \\ \boldsymbol{\mu} \end{bmatrix}' \begin{bmatrix} \mathbf{p}_{\mathrm{t-1}} \\ \mathrm{t}\mathbf{E}_{\mathrm{t-1}} \end{bmatrix} + \boldsymbol{\gamma}\mathbf{E}_{\mathrm{t}} + \sum_{i=1}^{S-1} \mathbf{\Gamma}_i \mathbf{\Delta}_1 \mathbf{p}_{\mathrm{t-i}} + \sum_{i=1}^{S} \sum_{j=2}^{q} \mathbf{k}_{\mathrm{i,j}} \mathbf{D}_{\mathrm{j,t+k-i}} + \sum_{m=1}^{M} \mathbf{\Theta}_{\mathrm{m}} \mathbf{w}_{\mathrm{m,t}} + \boldsymbol{\varepsilon}_t$$

where *S* is the lag length of the underlying VAR. $\mathbf{E}_t = [E_{1t} \quad E_{2t} \quad \ldots \quad E_{qt}]'$ is a vector of *q* dummy variables that take the value 1, i.e. $E_{jt} = 1$, if the observation belongs to the j^{th} period (j = 1, …, q), and 0 otherwise. $\mathbf{D}_{\mathrm{j,t+k-i}}$ is a so-called impulse dummy that equals 1 if the observation *t* is the i^{th} of the j^{th} period and 0 otherwise, is included to allow the conditional likelihood function to be derived given the initial values in each sub-period. w_t are the so-called intervention dummies (up to *M*) included to obtain well-behaving residuals.[10] The short run parameters are included in matrices $\boldsymbol{\gamma}$ (Vx*q*), $\mathbf{\Gamma}$ (VxV), $\mathbf{k}$ (Vx1) for each *j* and *i*, and $\mathbf{\Theta}$ (VxV). $\boldsymbol{\varepsilon}_t$ are assumed to be i.i.d. zero-mean disturbances with symmetric and positive definite variance, $\mathbf{\Omega}$. $\boldsymbol{\mu} = [\mu_{1t} \quad \mu_{2t} \quad \ldots \quad \mu_{qt}]'$ is the vector containing the long run drift parameters and $\boldsymbol{\beta}$ contains the usual long run coefficients in the cointegrating vector. The cointegration hypothesis is thus assessed by testing the rank of $\boldsymbol{\pi} = \boldsymbol{\alpha} \begin{bmatrix} \boldsymbol{\beta} \\ \boldsymbol{\mu} \end{bmatrix}'$.

[10] See Johansen et al. (2000) for more details.

3.3. Structural breaks: the bubble and the policy regime switching

Within this cointegration framework, we also want to investigate how the long-run price transmission relationships have been affected by the price "bubble" and the EU suspension of the import duties. Both shocks enter the model as temporary structural breaks; that is, following (4), a "bubble" and a "policy" dummy have been included in the cointegration space as exogenous variables. In this form, both regime changes are assumed to affect the constant term of the LR relation among prices.

Based on the test results on the presence and timing of the explosive behaviour (section 2.3), the first dummy has been given the value 1 for all weekly observations situated between the first week of July 2007 and the last week of March 2008, and zero otherwise. The second dummy mimics the suspension of the EU import duties on cereals. Indeed, it is well known that the trade policy regime may have a major role in price transmission mechanisms (Listorti 2007). It must be recalled that the EU protection mechanism for cereals, even after the Uruguay Round Agreement on Agriculture (URAA) converted all border measures into import duties, for a long period resulted in a wide gap between entry (border) and intervention (domestic) prices and, consequently, high duties. During the 2007-2008 price bubble, the European Union suspended import duties for cereals though, in fact, they were already set at very low levels due to the high world prices. The suspension began in January 2008, and was then prolonged until June 2009; finally, the reintroduction of duties was anticipated at the end of October 2008.[11] Therefore, the policy dummy takes the value 1 for all weekly observations between January 2008 and October 2008, and 0 otherwise (Figure 2). Within the adopted model, this dummy is expected to take into account how this policy intervention affected the transmission between international (Rotterdam) and national (Italian) prices.

4. Econometric procedure

The model specifications discussed in the previous section require an appropriate estimation procedure. This is repeated for both fixed-market/cross-commodity and fixed-commodity/cross-market sub-groups (Table A.2). First of all, cointegration among prices within the sub-group is assessed using the conventional Johansen (trace) test. If cointegration is found, then the respective VECM is estimated following specifications (2) and (4). If cointegration is not found, and no explosive root is present within the price group, a first-

[11] See Reg. CE 1/2008, Reg. CE 608/ 2008 and Reg. CE 1039/2008.

difference VAR is estimated.[12] Finally, if an explosive root is present without cointegration, no model specification is suitable as first order differentiation itself can not ensure the removal of the explosive pattern. Nonetheless, also in this case we estimate the respective first-difference VAR, as these results may provide further information on the presence of co-explosiveness. Table A.2 shows the model specifications adopted for all the five price groups under consideration: two fixed-commodity/cross-market cases (durum wheat and corn); three fixed-market/cross-commodity cases (Central-Northern Italy, Central-Southern Italy and International markets).

In all cointegration tests and VECM estimates, the "restricted constant case" (i.e., allowing for a constant in the cointegration space) of the Johansen procedure has been considered. This because the series don't show any linear trend in levels (Figure 1), but both theory and visual inspection of the data imply the presence of a constant term in the LR relationship, accounting for all elements contributing to price differentials not explicitly modelled in the price transmission equations.

Each model is estimated with and without the bubble and policy structural breaks. In VECM models, following Johansen et al. (2000), these dummies are assumed to have an impact on the constant term only inside the cointegrating space. As a consequence, in equation (4), t =1, γ=0, and a constant term is included in the cointegration space, with the corresponding elimination of one of the q dummy variables; the coefficients of the structural break dummies have then to be interpreted as relative to the constant term valid over the whole period. For these VECM estimates, the underlying assumption is that the rank of the cointegration matrix remains the same with or without the two structural breaks. In fact, the Johansen, et al. (2000) procedure doesn't allow testing for the cointegration rank with the number of breaks here considered. Also for this reason, conventional ADF tests are run on the residuals of the cointegration relation to check if the rank selected without the breaks can be confirmed *ex post* after their introduction.

If prices turn out not to be cointegrated, the bubble and policy dummies are simply introduced as exogenous dummy variables in a standard first-difference VAR model, thus allowing for a shift in the constant term of the VAR equations. In such circumstances, however, the response to a price shock is a SR adjustment can not be interpreted in respect to

[12] In both VECM and VAR models the lag length is selected according to the conventional information criteria (HQI, AIC and SBIC). This lag length is then confirmed by testing whether the adopted specification removes autocorrelation of the residuals (LM test of autocorrelation).

a LR relationship, since no evidence supports the existence of a LR pattern. Therefore, the dummies themselves can not be interpreted as structural breaks occurring within a LR relationship.

In both the VECM and first-difference VAR specifications, post-estimation allows assessing the presence of a residual explosive component in the estimated relationship. The stability condition (i.e., modulus of the largest eigenvalue < 1) and, in the VECM models, the stationarity (tested with ADF tests) of the estimated residuals of the LR relationship indicate that, although an explosive root may be found in some individual price series, explosiveness can be ruled out in the estimated equations, possibly due to the structural breaks. Consequently, a specification including co-explosiveness, like (3), is not needed.

Weak exogeneity tests (i.e., conventional t-tests on the coefficients of α) for the estimated VECM, and Granger causality tests for the VAR are performed to assess how price horizontally transmits from p_{ik} to p_{jk} or from p_{ik} to p_{ih}. For all price groups, this allows to identify the existing causal relationships, or, in other words, the "central" (i.e., leader in price formation) and local or "satellite" (follower) markets (Verga and Zuppiroli, 2003). The size, direction and timing of these significant causal relationships are finally analysed by computing the respective Impulse Response Functions (IRF).

5. Results

5.1. Cross-market transmission

This section discusses the estimation results for the two fixed-commodity/cross-market cases, that is, durum wheat and corn. In the case of durum wheat, the rank of the cointegration matrix is equal to one (Table 1). Though the international (Rotterdam) price presents strong evidence of explosive behaviour, the residuals from the cointegration relation remain stationary and the stability condition is respected. Therefore, we can conclude that the presence of co-explosiveness in the VECM model can be excluded. This may be attributed to the presence of the structural breaks that may take into account the period of more intense price turmoil though, in fact, results suggest that an explosive root can be excluded even when structural breaks are not included in the model specification.

When no break is included, in the cointegration vector both coefficients of the national prices (two "satellite" markets) are significant while the coefficient of the Rotterdam price is not and is much lower than the others. When structural breaks are included, however, only the coefficient associated to the Southern-Italian price remains statistically insignificant. With or without the breaks, both the Rome and the Foggia price adjustment coefficients are significant, whereas the Bologna and Rotterdam prices are weakly exogenous. The conclusion can be that, if the Rotterdam price can be interpreted as the driving price, the same holds for the Bologna price at least in the national durum wheat market. The bubble and the policy dummies do not substantially affect these findings. Their coefficients are positive in sign (0.033 for the bubble dummy and 0.018 for the policy dummy), but not statistically significant. Since the constant term in the cointegration vector is positive, this would indicate that in both time frames covered by the breaks the distance between the prices increases.

In the case of corn (Table 2), two cointegration vectors emerge. In estimating the VECM, however, we impose only one cointegration vector. The residuals form the cointegrating relation are stationary and the stability condition is met. Even in such case, therefore, we may exclude co-explosiveness as could be expected since no corn price clearly shows an explosive root (Table A.4). In the VECM without the bubble and policy dummies the coefficient of the Bologna price in the cointegration vector is significant and, in absolute value, is the largest and close to one. The coefficient of the Rome price is positive and significant, whereas the coefficient of the Rotterdam price is negative and not significant. The adjustment coefficients of the Milan and Rome prices are significant while the Bologna and the Rotterdam prices behave as weakly exogenous, thus confirming that both can be considered as driving markets.

When the structural breaks are included, within the cointegration vector, the coefficient of the Rome price changes its sign. In fact, all coefficients are barely statistical significant suggesting that, when the structural breaks are taken into account, the LR relationship among prices becomes weaker and price linkages mostly concern SR adjustments. The adjustment coefficients confirm that the international price behaves as weakly exogenous while all national prices now show statistically significant coefficients with the expected sign. The coefficient of the bubble dummy is positive (0.023), whereas the policy dummy has a negative coefficient (-0.027). So they almost reciprocally offset and, considering the negative sign of the constant term, these estimates would indicate that during the bubble the distance between the prices widened, while it diminished during the suspension of the EU import duties. However, even in this case, both are not statistically significant at the 5% confidence level.

According to these results we can conclude that, in the durum wheat market, Rotterdam and Bologna behave as the leader markets transmitting price shocks to the local markets of Rome and Foggia. In the case of corn, Rotterdam behaves as the driving price and transmits shocks to all national prices. The IRFs computed for the endogenous prices provide visual evidence on how this price transmission works. In durum wheat (Figure 3) we can notice that all response are positive and do not die out (as implied by cointegration vector), the only exception being the negative response of the Foggia price to a shock in the Rome price, the two evidently behaving as local or “satellite” markets. In general, the responses to shocks on the international prices are lower than those to shocks on the domestic ones. In this respect, Bologna seems to behave as the most influencing market. This is confirmed by the IRFs for corn (Figure 4). They show that the response of Milan to a shock in Bologna is higher than the response to a shock in Rotterdam, which is the only weakly exogenous price, whereas the response of Milan to a shock in Rome is negative as may expected for two “satellite” markets. Even in this case, Bologna confirms its dominant leading role at the national level. The response to shocks in international prices in normally lower than the one to the national ones.

5.2. Cross-commodity transmission

The results concerning the fixed-market/cross-commodity price relations are reported in Tables 3-5. The linkage between durum wheat and corn prices is investigated in three market places: Central Northern Italy (Bologna), Central-Southern Italy (Rome), and International (Rotterdam). In all cases, the cointegration rank turns out to be 0. Consequently, a first-difference VAR specification is estimated. This means that no long-run relationship can be detected between corn and durum wheat prices in any of the market places considered. This doesn’t imply that no linkage exist across prices but rather that, if present, it is limited to short-run responses to other price’s shocks. Furthermore, only in the Bologna market we notice a clear statistically significant linkage across the two prices; the durum wheat price is endogenous as it is Granger-caused by the corn price. In the other two cases (Rome and Rotterdam) the two prices are independent or only weakly dependent. In the Rome market durum wheat price is Granger-caused by corn price at 10% significance level. The opposite occurs in the Rotterdam market, where durum wheat prices Granger-cause corn prices at the 10% significance level. This latter effect, however, vanishes whenever structural breaks are included. We conclude that the only clear linkage across commodity prices is the dependence of durum wheat on corn prices in the Italian markets.

The parameters associated to structural breaks (dummies) confirm what observed in the cross-market analysis. The price bubble tends to significantly increase the price variations in response to exogenous shocks, but this effect is entirely compensated by the opposite effect of the policy. The impact of these dummies is however quite limited in magnitude.

The IRFs reported in Figure 5 provide a clearer picture of the relevant linkages occurring in the same market across different commodities. For the two significant relationships, both concerning the impact of a corn price shock on durum wheat price in national markets, the pattern of the response over time generally indicates a rapid decay of the response in few weeks. In a first-difference VAR model, this suggests that a shock in the logarithm of one price (that is, in its price growth rate) only temporarily affects the growth rate of the other price, which eventually comes back to the initial growth rate. It is worth noticing that the lack of a cointegration prevents from interpreting this response in terms of a reversion to a LR relationship. In both markets, durum wheat prices positively responds to a shock on corn prices, which suggests a sort of substitutability between corn and durum wheat. However, this relationship only holds in the SR: the response vanishes after few weeks and, as mentioned, no LR linkage can be detected.

An additional comment has to be made with respect to the explosive behaviour of some prices. Since the Rotterdam durum wheat price is the only case for which explosiveness is definitely observed, we may conclude that, in this international market, the first-difference VAR estimate is expected to show explosiveness, thus invalidating the respective parameter estimates. Nonetheless, Table 5 shows that no explosiveness is observed: in all market places residuals are indeed stationary and the stability condition is met.

We can summarize these results by arguing that cross-commodity price transmission is much weaker than in the cross-market case. Between the two international prices, no LR relationship emerges and no SR response to shocks is observed. In national markets, however, it would seem that a shock on the corn price has an impact on the durum wheat price. Given the results reported and discussed in previous section, this is particularly interesting in the case of Bologna, as this is the national driving market. The casual chain eventually emerging seems to be the following: shocks on the corn price are transmitted from the international to the national driving market and then to durum wheat and to the national "satellite" markets.

6. Some final remarks

This paper aims at analysing the horizontal agricultural price transmission (across space and commodities) during the years 2006-2010, a period of extreme market turbulence. Since during market turmoil weekly price series may follow peculiar and hardly tractable stochastic processes, this research objective is particularly challenging. Our evidence, although preliminary, confirms that all the price series considered (durum wheat and corn prices in both Italian and international markets) behave as I(1) processes. However, some of them actually seem to be "something more" than simple I(1) series. Fractional integration in the first differences and, even more significantly, explosive roots can not be excluded in some of the price series examined. Nonetheless, it must be acknowledged that, while I(1) processes alone can not explain the observed exuberance, no other common pattern emerges across all prices. In other words, the apparently analogous behaviour of all prices during the price bubble can not be explained by some common stochastic process, but must be related to reciprocal price interaction.

The stochastic properties of the series are of particular relevance while finding the appropriate specification of the price transmission equations (VECM or first-difference VAR models). In general terms, when the two structural breaks (the "bubble" and the suspension of the EU import duties) are included in the cointegration space, a LR relationship among prices only emerges for the same commodity across different markets but not across different commodities. Even in the former case, however, this LR linkage tends to be statistically significant only for durum wheat. This would indicate that, at least when such turbulent periods are considered, the relationship between prices mostly concerns SR responses to shocks while in the LR prices appear to be relatively more independent. Even in this prevailing SR horizon, however, a casual chain seems to emerge. Exogenous shocks prevalently come from the corn international price and are then transmitted to the pivotal national market (Bologna). Then, this latter transmits shocks to the national "satellite" markets and to the durum wheat price. If we exclude the dependency emerging among these "satellite" markets, price responses (across places or commodities) always move in the same direction of the shocks and this may explain why a single shock may be transmitted and even amplified downstream in the other markets.

A final comment concerns the role of the two structural breaks considered, the "bubble" and the policy intervention (the suspension of the EU import duties). Econometric findings suggests a rather unexpected evidence: although the inclusion of these structural breaks often affects the results, the two dummies are barely significant in the cross-market transmission while are usually significant in the cross-commodity transmission. The temporary trade-policy measure may have indeed played a role in price transmission, but it seems to be limited to the cross-commodity case, while its alleged aim is actually relative to the cross-market dimension. Furthermore, in all statistically significant cases the “bubble” and the policy always operate in opposite directions. The former amplifies price variation in response to external shocks while the latter actually reduces this response. This not only indicates that these two structural breaks tend to reciprocally offset but also that, in fact, the policy intervention eventually played a role in reducing the magnitude of the price response to shocks.

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Table 1 – Cross-market price linkage: 3-lag VECM estimates (standard errors in parenthesis) – durum wheat[a]

	Without breaks	With breaks
Trace (Johansen) test		
Rank = 0	60.572	
Rank = 1	28.101†	
Rank = 2	10.759	
Rank = 3	3.168	
Rank = 4	-	
Cointegrating vector (β)	**Without breaks**	**With breaks**
fd_fi_bo	1.000	1.000
fd_fi_ro	-0.333* (0.129)	-0.315 (0.160)
fd_fi_fo	-0.741* (0.123)	-0.824* (0.157)
cwad_can	0.030 (0.028)	0.044 (0.035)
"bubble" dummy		0.033 (0.020)
"policy" dummy		0.018 (0.019)
Constant	0.236* (0.051)	0.501* (0.146)
Adjustment vector (α)		
fd_fi_bo	0.042 (0.078)	0.103 (0.071)
fd_fi_ro	0.138* (0.067)	0.177* (0.060)
fd_fi_fo	0.255* (0.067)	0.252* (0.060)
cwad_can	0.114 (0.128)	0.100 (0.117)
ADF test on residuals of long-run relation (asymptotic p-values in parenthesis) [b]	2.709* (0.01)	-2.560* (0.01)
Stability condition (highest eigenvalue, modulus) [c]	0.886	0.792

[a] Estimated coefficients of the first-difference terms are available upon request. The 3-lag specification of the VECM has been selected according to the conventional information criteria. Whereas HQIC, AIC and SBIC indicated 2 lags, however, 3 lags were necessary in order to remove autocorrelation in the residuals (LM test). The values reported for the Johansen test refer to 3 lags, as well.
[b] The ADF test specification includes 12 lags.
[c] The VECM specification imposes unit root modulus, not reported here.
† Accepted rank: lowest rank whose test result is lower than 5% critical values.
*Statistically significant at 5% confidence level.

Table 2 – Cross-market price linkage: 4-lag VECM estimates (standard errors in parenthesis) – corn [a]

Trace (Johansen) test		
Rank = 0	62.523	
Rank = 1	36.453	
Rank = 2	15.714†	
Rank = 3	2.573	
Rank = 4	-	
Cointegrating vector (β)	**Without breaks**	**With breaks**
mais_mi	1.000	1.0000
mais_bo	-1.184* (0.100)	-0.492 (0.113)
mais_ro	0.225* (0.110)	-0.632 (0.126)
mais_us	-0.039 (0.026)	0.060 (0.036)
"bubble" dummy		0.023 (0.014)
"policy" dummy		-0.027 (0.014)
Constant	-0.007 (0.102)	0.338 (0.157)
Adjustment vector (α)		
mais_mi	-0.137* (0.067)	0.172* (0.061)
mais_bo	0.041 (0.085)	0.298* (0.076)
mais_ro	-0.227* (0.99)	0.403* (0.091)
mais_us	0.269 (0.162)	-0.026 (0.151)
ADF test on residuals of long-run relation (asymptotic p-values in parenthesis) [b]	-4.044* (0.00)	-3.394* (0.00)
Stability condition (highest eigenvalue, modulus) [c]	0.850	0.792

[a] Estimated coefficients of the first-difference terms are available upon request. The 4-lag specification of the VECM has been selected following the conventional information criteria: 2 lags were indicated by HQIC and SBIC, 4 lags by the AIC. When one cointegration vector was imposed, 4 lags were preferred as they allow removing autocorrelation in the residuals (LM test). The values reported for the Johansen test refer to 4 lags, as well.
[b] The ADF test specification includes 12 lags.
[c] The VECM specification imposes unit root modulus, not reported here.
† Accepted rank: lowest rank whose test result is lower than 5% critical values.
*Statistically significant at 5% confidence level.

Table 3 – Cross-commodity price linkage: 5-lag first-difference VAR estimates (standard errors in parenthesis) – Central-Northern Italy (Bologna) [a]

Trace (Johansen) test		
Rank = 0	11.60†	
Rank = 1	3.42	
Rank = 2	-	
Short-run Granger Causality tests (χ^2)	**Without breaks**	**With breaks**
Durum Wheat on Corn	4.28	5.42
Corn on Durum Wheat	19.49*	16.89*
VAR coefficients		
Durum Wheat: bubble dummy	-	0.011* (0.005)
policy dummy	-	-0.010* (0.004)
Corn: bubble dummy	-	0.005 (0.004)
policy dummy	-	-0.011* (0.004)
ADF test on residuals (asymptotic p-values in parenthesis) [b]		
Durum Wheat	-3.74 (0.00)*	-4.27 (0.00)*
Corn	-4.83 (0.00)*	-5.29 (0.00)*
Stability condition (highest eigenvalue, modulus)	0.790	0.749

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag (5 weeks) has been selected according to the conventional information criteria. A constant term is included in VAR equations.
[b] The ADF test specification includes 12 lags.
†Accepted rank: lowest rank whose test result is lower than 5% critical values.
*Statistically significant at 5% confidence level.

Table 4 – Cross-commodity price linkage: 3-lag first-difference VAR estimates (standard errors in parenthesis) – Central-Southern Italy (Rome) [a]

Trace (Johansen) test		
Rank = 0	16.06†	
Rank = 1	5.24	
Rank = 2	-	
Short-run Granger Causality tests (χ^2)	**Without breaks**	**With breaks**
Durum Wheat on Corn	3.63	2.39
Corn on Durum Wheat	6.23	7.83*
VAR coefficients		
Durum Wheat: bubble dummy	-	0.014*(0.004)
policy dummy	-	-0.011*(0.004)
Corn: bubble dummy	-	0.003(0.006)
policy dummy	-	-0.011*(0.005)
ADF test on residuals (asymptotic p-values in parenthesis) [b]		
Durum Wheat	-3.24* (0.00)	-4.12* (0.00)
Corn	-4.56* (0.00)	-4.91* (0.00)
Stability condition (highest eigenvalue, modulus)	0.714	0.553

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag (3 weeks) has been selected according to the conventional information criteria. A constant terms in included in VAR equations.
[b] The ADF test specification includes 12 lags.
†Accepted rank: lowest rank whose test result is lower than 5% critical values.
*Statistically significant at 5% confidence level.

Table 5 – Cross-commodity price linkage: first-difference VAR estimates – International markets (Rotterdam) [a]

Trace (Johansen) test		
Rank = 0	11.35†	
Rank = 1	3.02	
Rank = 2	-	
Short-run Granger Causality tests (χ^2)	**Without breaks**	**With breaks**
Durum Wheat on Corn	2.87	3.35
Corn on Durum Wheat	6.49	3.06
VAR coefficients		
Durum Wheat: bubble dummy	-	0.016* (0.007)
policy dummy	-	-0.017* (0.007)
Corn: bubble dummy	-	0.007 (0.008)
policy dummy	-	-0.010 (0.007)
ADF test on residuals (asymptotic p-values in parenthesis) [b]		
Durum Wheat	-3.65* (0.00)	-4.61* (0.00)
Corn	-4.16* (0.00)	-4.14* (0.00)
Stability condition (highest eigenvalue, modulus)	0.681	0.583

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag (3 weeks) has been selected according to the conventional information criteria. A constant terms in included in VAR equations.
[b] The ADF test specification includes 12 lags.
†Accepted rank: lowest rank whose test result is lower than 5% critical values.
*Statistically significant at 5% confidence level.

Figure 1 – The price bubble: durum wheat and corn price series over the period between May 2006 and December 2010 (see Annex 1 for price codes).

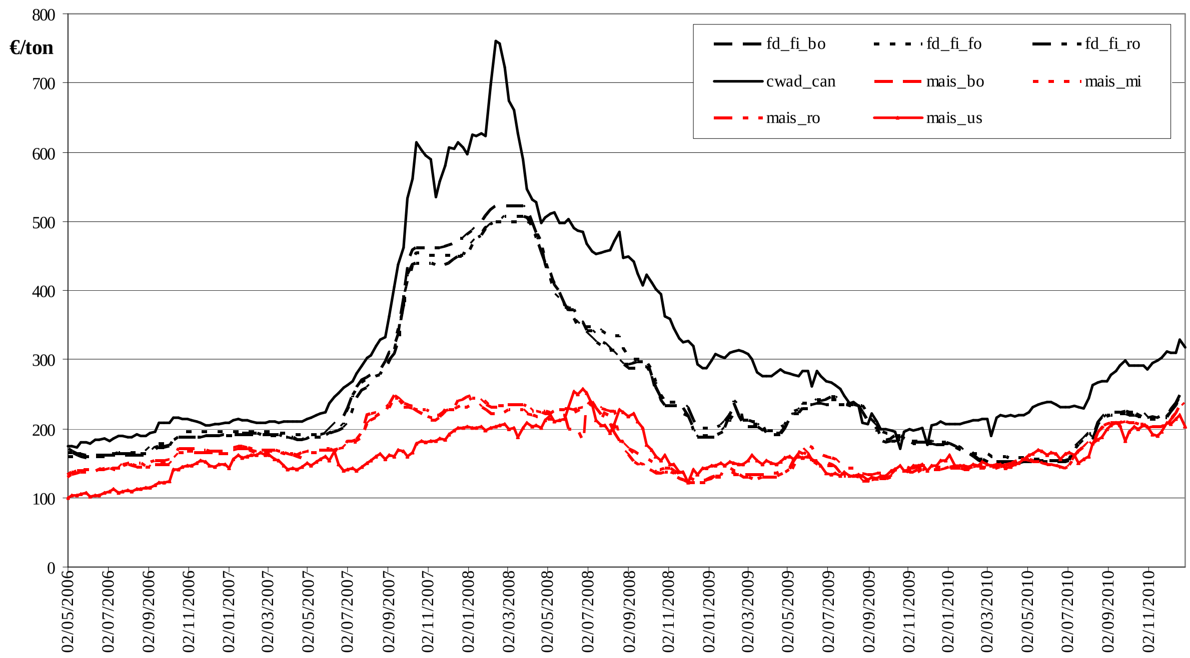


Figure 2 – Dating the bubble: time series of the forward recursive ADF t-statistic for the durum wheat and corn price series (logarithms) (see Annex 1 for price codes).

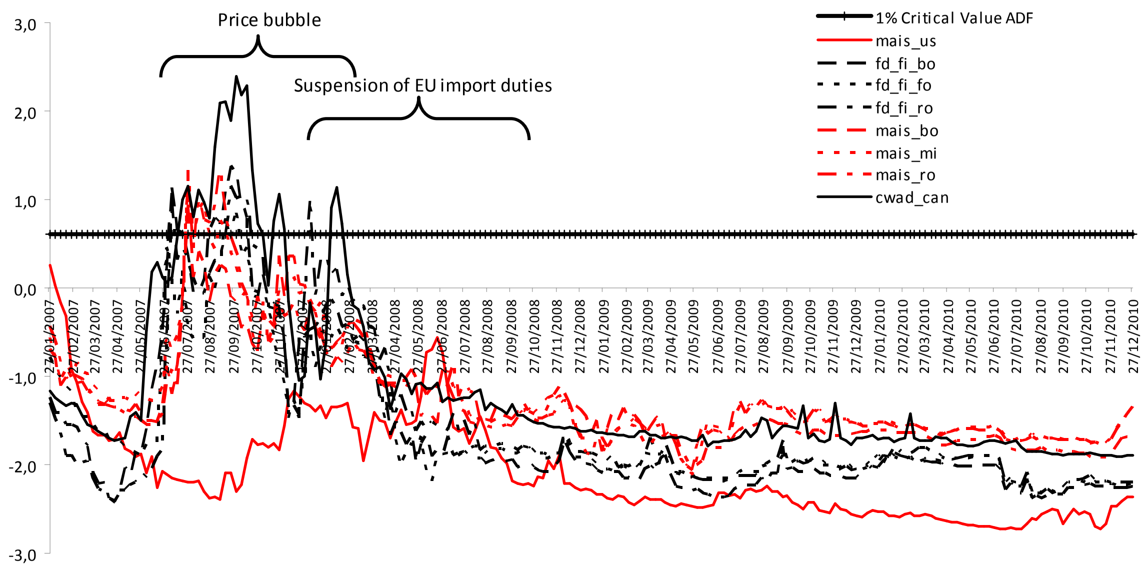


Figure 3 – Impulse Response Functions of the endogenous prices (Rome and Foggia) in durum wheat markets (ordering: Bologna, Rome, Foggia, Rotterdam; model with bubble and policy dummies).

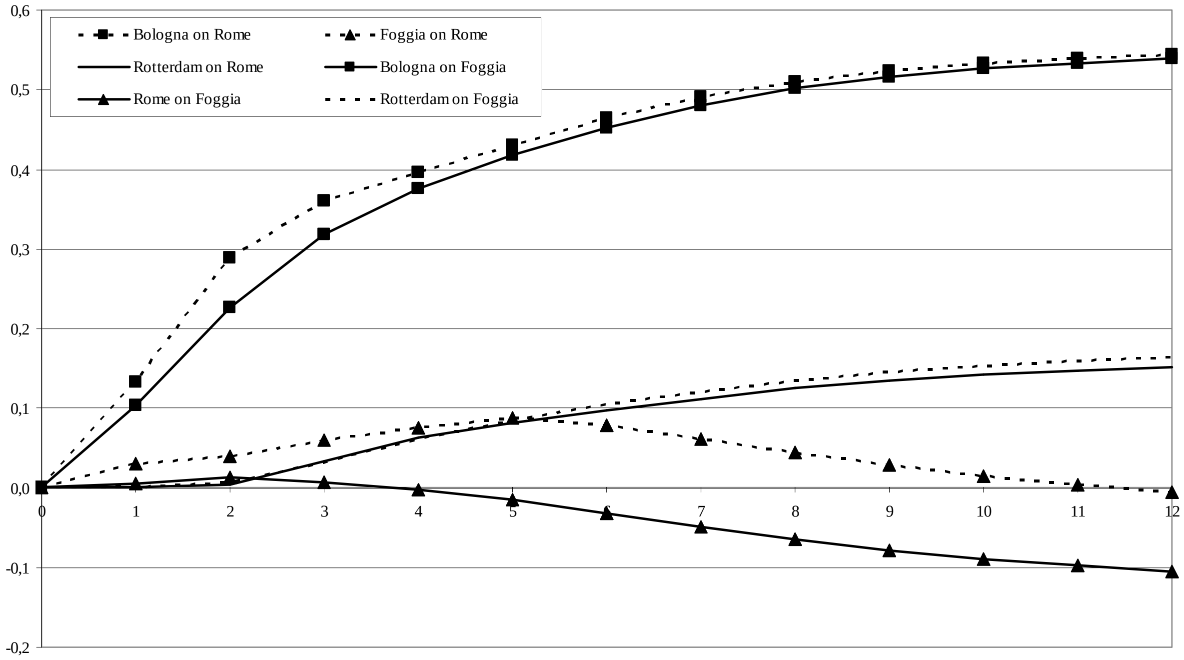


Figure 4 – Impulse Response Functions of the endogenous prices (all national prices) in corn markets (ordering: Milan, Bologna, Rome, Rotterdam; model with bubble and policy dummies).

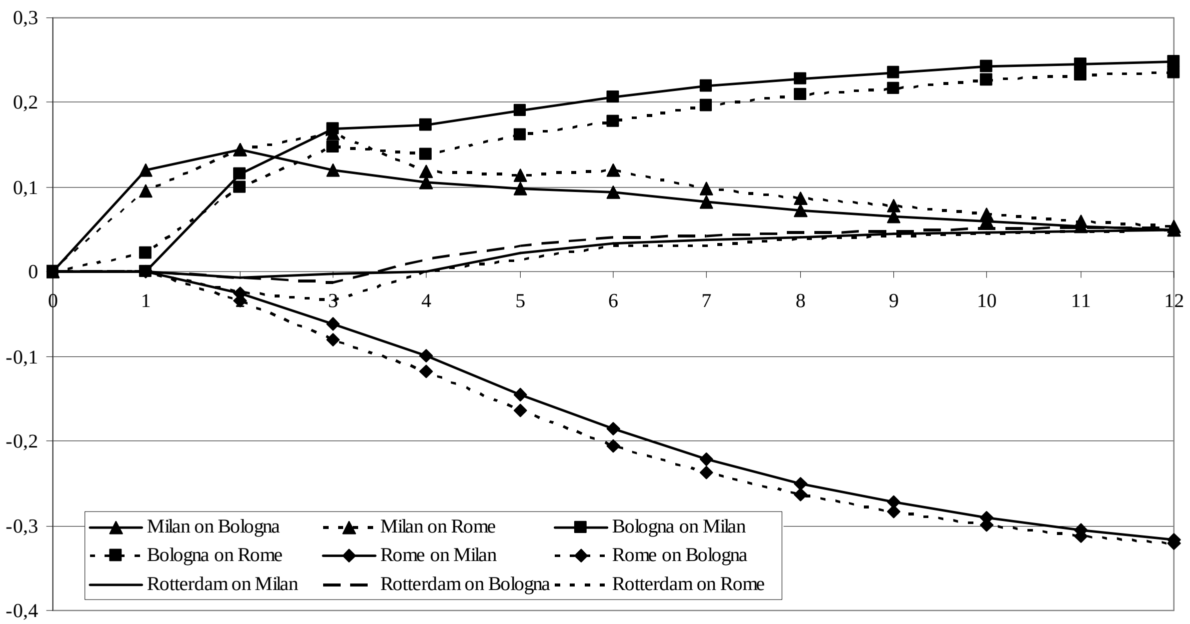
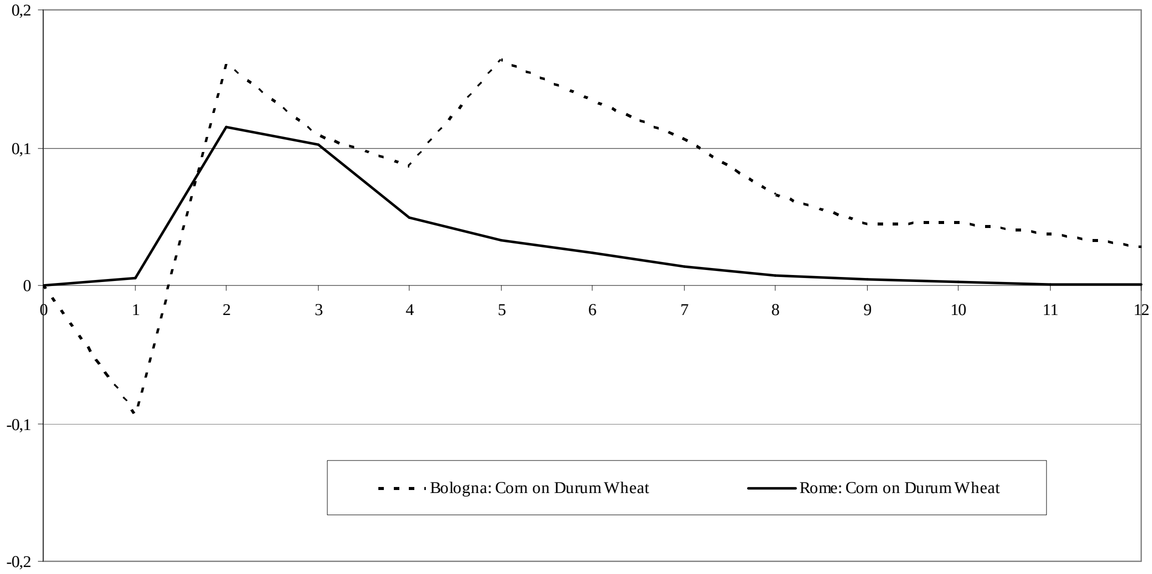


Figure 5 – Impulse Response Functions of the endogenous price (durum wheat) in Bologna and Rome markets (ordering: durum wheat, corn; model with bubble and policy dummies)



ANNEX 1 – Price commodities and groups under analysis

Table A.1 – Codification and description of the prices adopted in the analysis

Price Code	Product Description	Market Place
fd_fi_bo	Durum Wheat, Fino	Bologna (Central-Northern Italy)
fd_fi_fo	Durum Wheat, Fino	Foggia (Southern Italy)
fd_fi_ro	Durum Wheat, Fino	Rome (Central-Southern Italy)
mais_bo	Maize, Ibrido Nazionale	Bologna (Central-Northern Italy)
mais_mi	Maize, Ibrido Nazionale	Milan (Northern Italy)
mais_ro	Maize, Ibrido Nazionale	Rome (Central-Southern Italy)
cwad_can[a]	Wheat, Canada Western Amber Durum (CWAD)	Canada, St Lawrence/Rotterdam
mais_us[a]	Maize, #3 Yellow Corn (3YC)	US, Gulf/Rotterdam

[a] CIF price

Table A.2 – Price groups for the analysis of price linkages (VECM or first-difference VAR models)

Group of Interdependent Commodities/Markets	Property of the Series	Estimated Model
Fixed Commodity-Cross Market		
DURUM WHEAT		VECM
fd_fi_bo	I(1)	
fd_fi_ro	I(1)	
fd_fi_fo	I(1)	
cwad_can[a]	I(1) + explosive root	
CORN		VECM
mais_mi	I(1)	
mais_bo	I(1)	
mais_ro	I(1)	
mais_us[a]	I(1)	
Fixed Market-Cross Commodity		
CENTRAL-NORTHERN ITALY (Bologna)		First-difference VAR
fd_fi_bo	I(1)	
mais_bo	I(1)	
CENTRAL-SOUTHERN ITALY (Rome)		First-difference VAR
fd_fi_ro	I(1)	
mais_ro	I(1)	
INTERNATIONAL (Rotterdam)		First-difference VAR
cwad_can[a]	I(1) + explosive root	
mais_us[a]	I(1)	

[a] CIF price

ANNEX 2 – Unit and explosive roots testing

Table A.3 – Unit root tests on $\mathbf{p}_{ik}$ *and* $\boldsymbol{\Delta}_1\mathbf{p}_{ik}$ *: Adjusted Dickey-Fuller (ADF)*[a]*, Phillips-Perron (PP)*[b]*, Adjusted Dickey-Fuller GLS (ADF GLS)*[c] *and Kwiatkowski, Phillips, Schmidt e Shin (KPSS)*[d] *test; p-values in parenthesis; the values for which the null is rejected are in bold (10% critical values)*

Price	$\mathbf{p}_{ik}$				$\boldsymbol{\Delta}_1\mathbf{p}_{ik}$			
	ADF	**PP**	**ADF GLS**	**KPSS**	**ADF**	**PP**	**ADF GLS**	**KPSS**
fd_fi_bo	-1.075	-1.112	-1.196	**0.368**	**-5.428**	**-6.501**	**-3.686**	0.171
	(0.728)	(0.710)	(0.213)	**(0.091)**	**(0.000)**	**(0.000)**	**(0.000)**	(>0.100)
fd_fi_fo	-1.887	-1.226	-1.098	**0.376**	-2.128	**-6.616**	**-2.446**	0,190
	(0.339)	(0.662)	(0.247)	**(0.088)**	(0.234)	**(0.000)**	**(0.014)**	(>0.100)
fd_fi_ro	-1.909	-1.124	-1.330	**0.376**	-2.005	**-6.218**	**-2.586**	0.180
	(0.328)	(0.705)	(0.170)	**(0.087)**	(0.285)	**(0.000)**	**(0.009)**	(>0.100)
mais_bo	-1.534	-1.656	-0.298	**0.377**	**-3.494**	**-7.619**	**-2.841**	0.168
	(0.516)	(0.454)	(0.579)	**(0.087)**	**(0.008)**	**(0.000)**	**(0.004)**	(>0.100)
mais_mi	-1.534	-1.615	0.388	**0.369**	**-3.545**	**-8.694**	**-1.966**	0.175
	(0.516)	(0.475)	(0.544)	**(0.091)**	**(0.007)**	**(0.000)**	**(0.047)**	(>0.100)
mais_ro	-1.616	-1.815	-0.461	**0,374**	**-2.704**	**-11.388**	**-2.213**	0.155
	(0.474)	(0.373)	(0.516)	**(0.089)**	**(0.073)**	**(0.000)**	**(0.026)**	(>0.100)
cwad_can	-1.067	-1.374	-0.637	**0.365**	**-2.831**	**-14.008**	**-2.839**	0.242
	(0.731)	(0.595)	(0.442)	**(0.093)**	**(0.054)**	**(0.000)**	**(0.004)**	(>0.100)
mais_us	-2.234	-2.399	0.125	**0.377**	-2.487	**-15.302**	**-1.703**	0.160
	(0.194)	(0.142)	(0.722)	**(0.087)**	(0.118)	**(0.000)**	**(0.084)**	(>0.100)

[a] H_0: unit root. The test specification includes a constant term, and all significant lags “testing down” up to a maximum of 12.
[b] H_0: unit root. The test specification includes 12 lags and a constant term.
[c] H_0: unit root. The test specification includes a constant term and all significant lags “testing down” up to a maximum of 12.
[d] H_0: no unit root. The test specification includes 12 lags and a constant term; p-values are interpolated

Table A.4 – Test of fractional integration on $\mathbf{p}_{ik}$ *and* $\boldsymbol{\Delta}_1\mathbf{p}_{ik}$ *according to Phillips (1999a,b)*[a] *and SADF tests (forward recursive regressions) of explosive roots on* $\mathbf{p}_{ik}$ *according to Phillips et al.(2009)*[b]*; p-values in parenthesis; the cases for which the null is rejected are in bold (5% critical values, respectively). All test specifications with a constant term and 12 lags*

Price	Test of fractional integration				SADF test on $\mathbf{p}_{ik}$	
	$\mathbf{p}_{ik}$		$\boldsymbol{\Delta}_1\mathbf{p}_{ik}$			
	t (H_0: d=0)	z (H_0: d=1)	t (H_0: d=0)	z (H_0: d=1)	(r=0.1)	(r=0.2)
fd_fi_bo	**11.359**	1.569	**2.367**	**-4.532**	1.376	1.376
	(0.000)	(0.117)	**(0.029)**	**(0.000)**	(>0.050)	(>0.050)
fd_fi_fo	**10.410**	1.565	1.567	**-4.774**	**1.789**	0.888
	(0.000)	(0.118)	(0.134)	**(0.000)**	**(<0.050)**	(>0.100)
fd_fi_ro	**15.347**	1.687	**2.728**	**-4.132**	1.370	1.154
	(0.000)	(0.092)	**(0.014)**	**(0.000)**	(>0.050)	(>0.100)
mais_bo	**4.306**	1.913	1.799	**-4.343**	**7.018**	0.997
	(0.000)	(0.056)	(0.089)	**(0.000)**	**(<0.010)**	(>0.100)
mais_mi	**4.314**	1.741	**2.564**	**-3.751**	**11.865**	1.365
	(0.000)	(0.082)	**(0.020)**	**(0.000)**	**(<0.010)**	(>0.050)
mais_ro	**5.087**	0.862	1.742	**-4.303**	1.339	1.339
	(0.000)	(0.389)	(0.100)	**(0.000)**	(>0.050)	(>0.050)
cwad_can	**9.765**	1.641	2.081	**-3.912**	**2.396**	**2.396**
	(0.000)	(0.101)	(0.052)	**(0.000)**	**(<0.010)**	**(<0.010)**
mais_us	**5.484**	0.320	0.671	**-5.689**	**2.653**	-0.559
	(0.000)	(0.749)	(0.510)	**(0.000)**	**(<0.010)**	(>0.100)

[a] If d=0 the series is I(0); if d=1 the series is I(1); if $0<d<1$, the series is I(d) (long memory process). As a deterministic trend has been excluded, the original test has not been detrended. The test runs under alternative possible values of the arbitrary power parameter (see Phillips, 1999a,b) here assumed equal to 0.55. Test robustness is performed using a set of values of the power parameter ranging from 0.4 to 0.75. Test results for these alternative values of the power parameter are available on request. They do not significantly differ from those presented here.
[b] H_0: no explosive root. Critical values for both sample sizes of 100 and 500 give the same test outcome.