Agricultural Price Transmission Across Space and Commodities During Price Bubbles

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Abstract

This paper analyses the horizontal transmission of cereal price shocks both across different market places and across different commodities. The analysis is carried out using Italian and international weekly spot (cash) price data and concentrating the attention on years 2006-2010, a period of generalized exceptional exuberance and consequent rapid drop of agricultural prices. The work aims at investigating how price transmission may be affected during price bubbles. The properties of price time series are firstly explored to assess which data generation process may have eventually produced the observed patterns. Secondly, the interdependence across prices is specified and estimated adopting appropriate cointegration techniques.

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

1. Introduction: objectives and data description

The main motivation of this study is to understand the properties of agricultural price series over the 2007-2008 price rally (Ec, 2008; Irwin and Good, 2009)[1] to better analyse how the prices transmitted horizontally, i.e., across market places and commodities, during the bubble. This analysis of the horizontal price transmission during price exuberance concerns agricultural weekly spot prices and is limited to the period going from May 206 to December 2010. We opt for this restriction of the time coverage for three major reasons. First of all, this period fully contains the bubble: the bubble firstly inflated, then completely deflated and finally started raising again in the second half of year 2010 (Figure 1) (see Esposti and Listorti, 2010, for more details on price series). Secondly, we can assume an almost-constant policy regime in the EU over this period. In 2006, the 2003 CAP Reform was entirely in force, included its limited implications in terms of price policy and market intervention; the regime then remained constant over the whole 2006-2010 period. The only relevant policy regime change over the period under consideration has been the temporary suspension of EU import duties on cereals from January 2008 to October 2008. This change, in fact, was justified by price movements themselves (the bubble). Therefore, in investigating price transmission during the bubble it is also possible to assess the role played by this single policy measure whose application is confined into a limited number of months. Finally, concentrating on this period facilitates international price comparison 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, comparison of agricultural prices across different countries does not require the deflation of nominal prices into a common real base.

The dataset here adopted is made of cereal weekly spot price series observed in different Italian locations (from North to South) (source: ISMEA) and in international (North-American) markets (source: International Grain Council, IGC). Commodities under consideration are durum wheat and corn. The attention is here limited to durum wheat and corn because they somehow represent opposite situations. During the price bubble, among main cereals, durum wheat experienced the largest rise (and, then, decline) while corn price showed a relatively limited variation. Moreover, prevalent domestic (Italian) use are very different for the two cereals: almost exclusively human consumption for durum wheat, prevalent feed use for corn. Finally, Italy is the largest EU durum wheat producing country and this cereal is one of the characteristic product of the Italian agriculture while, on the contrary, corn is a less typical production. Therefore, it is here of major interest to

[1] Henceforth, the "agricultural price bubble". Heuristically, and generically, the bubble clearly appear in Figure 1. From a more rigorous point of view, however, we will formally define and test the presence of a bubble in price series in section 2.2.

understand if these two "extreme" cases among cereals may show a different behaviour over the price rally and which is their reciprocal linkage.
Therefore the dataset dimensions are: T = 244 weeks (from firs week of May 2006 to 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). In practice, the dataset has a Tx(NxK) size. Table A1 (Annex) details the NxK = 10 price series under investigation and their respective acronyms.
To facilitate comparisons and analysis of transmission, international and national prices have to be expressed in a common currency. Therefore, US and Canadian prices have been converted from US dollars to Euros by taking the weekly official $/€ exchange rate provided by Eurostat-ECB.[2] These North-American prices actually are FOB prices. For agricultural commodities, the ratio of freight rates on FOB prices is quite high and it also considerably oscillated during the commodity price bubble (especially due to volatility in energy, namely oil, prices). According to this, 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 ARAH (i.e, Amsterdam/Rotterdam/Antwerp/Hamburg destinations). In practice, in our analysis the North-American prices actually serve as international price taken at Rotterdam, thus as EU-reference prices. Henceforth, we will refer to international prices as Rotterdam prices (see Table A.1).

2. Some basic evidence on agricultural price behaviour

Before entering the analysis of how market prices interact, thus reciprocally transmitting price shocks, it seems rational to firstly assess the time series properties of the prices under study. This analysis not only allows identifying those prices showing common features but it is also needed to achieve a proper specification of price transmission/interdependence relations.
Let consider agricultural prices observed over these three different dimensions: space, commodity and time. The generic price is, therefore, $p_{i,k,t}$ where i=1,…,j,…,N is the (local) market (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 $\mathbf{p}_{ik}$. Logarithmic transformations are here considered rather than levels as this allows to more directly refer to price linkages as elasticities. Therefore, henceforth $\mathbf{p}_{ik}$ will actually identify the logarithm of the price of k-th commodity in the i-th market place.
This section analyses the time-series properties of prices by testing, in sequence, stationarity and explosiveness to assess whether these features may be invoked as possible causes of the observed exuberance. At the same time, they also imply different stochastic processes and price interdependence.[3]

2.1. Stationarity

What really matters in understanding the behaviour of price series especially in a period of such dramatic exuberance and drop, is the presence of unit and/or explosive roots. Stationary (i.e. I(0)) series can be hardly reconciled with the formation of the bubble. Even stationarity around a drift (a constant term) and/or a deterministic trend is not evidently helpful in this respect. As will be clarified below, testing for the presence of a temporary explosive pattern can not be simply achieved through conventional unit-root tests. Nonetheless, albeit not sufficient, these latter tests still allow assessing a necessary outcome of nonlinearities within price series: series turn out to be

[2] It implies that adjustments to the exchange rate are considered instantaneous.
[3] Other time-series properties that could generate non-linear dynamics in price patterns have been tested and generally excluded (non-normality, AR CH effects, seasonality, fractional integration) (see Esposti and Listorti, 2001, for more details).

I(1) and not I(0) (Diba and Grossman, 1988; Evans, 1991; Phillips and Magdalinos, 2009; Philips et al., 2009; Phillips and Yu, 2009; Gutierrez, 2010).
Table 1 thus reports unit root tests on $\mathbf{p}_{ik}$ and respective first differences, $\mathbf{\Delta}_1\mathbf{p}_{ik}$. Two different tests are run, the ADF and the PP tests (Enders, 1995), as the latter is expected to be more robust under heteroskedasticity which may, in fact, occur. In general terms, it looks like that, once the proper specification has been selected (in terms of admitted lags and of presence of drift and deterministic trend), all $\mathbf{p}_{ik}$ series show a unit root. On the contrary, $\mathbf{\Delta}_1\mathbf{p}_{ik}$ series are stationary if PP tests are considered. 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 temporary explosive behaviour. More generally, it remains to explain how such I(1) $\mathbf{p}_{ik}$ series may have eventually generated the observed price pattern in all markets. A plausible explanation is that temporary explosive roots (the "bubble") actually occurred. As conventional unit-root tests can not really assess (or exclude) the presence of temporary explosive roots, more appropriate *ad hoc* tests should be adopted.

2.2. Testing explosiveness: recursive unit root tests

A more formal and rigorous definition of a temporary (or periodically collapsing) bubble in price or financial series consists in the presence of temporary explosive roots. Such temporary explosive patterns present a problem to standard time series analysis. The problem, once more, is that price series containing a temporary exuberance do behave neither as I(1) processes nor as I(2) processes and, therefore, 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 (see also Gutierrez, 2010) have provided an appropriate framework for assessing the presence of an explosive root (a bubble) within processes that would be otherwise ruled as I(1). They propose a test for the presence of bubbles where 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 like (1) where in the first recursion only $T_o = [\ r_o\ T]$ observations are used, and r_o is a fraction of the total sample T.[4] In subsequent regressions the originating data set is supplemented by successive observations any time giving a sample of size $T_r = [nr]$ 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 consider 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 presence of explosive root (H_1: ρ_{ik} >1), if the estimated test *SADF* is higher than the respective critical values, we accept that the series contains an explosive root. Critical values are reported in Phillips et al. (2009). Here, r_o has been alternatively fixed at 0.10 (24 observations) at 0.20 (48 observations) (see Phillips et al, 2009, and Phillips and Yu, 2009) and then incremented by each single following observation. The *SADF* tests have been repeated on price series, including the constant term and with 12 lags and then compared with the critical values provided, for different sample sizes, in Phillips et al. (2009) and Phillips and Yu (2009). Table 1 reports results of these *SADF* tests. Results are reported for both r=0.10 and r=0.20. The latter case, however, provides more robust evidence because it uses larger sub-samples thus reducing the risk of poor test performance in the first runs of the recursive process. Moreover, a restrictive 1% critical value is used to test the presence of explosive roots. On this basis, we can conclude that a temporary

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

explosive behaviour is definitely found only in durum wheat international (Rotterdam) price. The presence of a an explosive root is doubtful in few cases, while it can be excluded in most of the other price series.

Although this kind of test may be not conclusive with respect to the presence of explosive roots, the underlying approach is particularly helpful in understanding the timing of price exuberance and, therefore, 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 abovementioned forward recursive ADF_r test and check if, when and how long ADF_r exceeds the right-tailed critical values of the asymptotic distribution of the standard Dickey–Fuller t-statistic (Phillips et al. 2009). Still adopting a restrictive 1% critical value, Figure 2 reports this evidence and clearly shows that the price bubble is limited, especially in national markets, to very few cases and lasts for a very short period According to these results, and following restrictive criteria as explosiveness may be easily confused with different processes generating similar patterns, we will henceforth consider that only price *cwad_can* (see Table A1) clearly contains an explosive root.

Table 1 – Adjusted Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests on price series (H_0: unit root; p-values in parenthesis)[a] *– and Phillips et al.(2009) SADF tests of explosive roots on logarithms of prices (H_0: no explosive root)*[b] *– in bold values for which the null is rejected (5% confidence level) and values greater than 1% critical values for a sample size of 500, respectively*

Price	$\mathbf{p}_{ik}$		$\Delta_1\mathbf{p}_{ik}$		**SADF test (Forward Recursive Regressions)**	
	ADF	**PP**	**ADF**	**PP**	(r=0.1)	(r=0.2)
fd_fi_bo	-1.075 (0.728)	-1.112 (0.710)	**-5.428** **(0.000)**	**-6.501** **(0.000)**	1.376	1.376
fd_fi_fo	-1.887 (0.339)	-1.226 (0.662)	-2.128 (0.234)	**-6.616** **(0.000)**	1.789	0.888
fd_fi_ro	-1.909 (0.328)	-1.124 (0.705)	-2.005 (0.285)	**-6.218** **(0.000)**	1.370	1.154
mais_bo	-1.534 (0.516)	-1.656 (0.454)	-2.595 (0.094)	**-7.619** **(0.000)**	**7.018**	0.997
mais_mi	-1.534 (0.516)	-1.615 (0.475)	**-3.545** **(0.007)**	**-8.694** **(0.000)**	**11.865**	1.365
mais_ro	-1.616 (0.474)	-1.815 (0.373)	-2.704 (0.073)	**-11.388** **(0.000)**	1.339	1.339
cwad_can	-1.067 (0.731)	-1.374 (0.595)	**-4.794** **(0.000)**	**-14.008** **(0.000)**	**2.396**	**2.396**
mais_us	-2.234 (0.194)	-2.399 (0.142)	-2.487 (0.118)	**-15.302** **(0.000)**	**2.653**	-0.559
		Critical values (Phillips et al., 2009):			Sample size = 500	Sample size = 100
		10% confidence level			1.180	1.191
		5% confidence level			1.460	1.507
		1% confidence level			2.004	2.190

[a] The ADF test specification includes a constant term, all significant lags "testing down" up to a maximum of 12, and seasonal dummies. The ADF tests have been repeated also without the inclusion of a constant term, and, both with and without the constant term, without seasonal dummies. Phillips-Perron tests have been repeated with 4 lags (number determined with the Newey-West procedure) and 12 lags, in both cases with and without the inclusion of a constant term. The table shows the results of the specification with 12 lags and a constant term. Results do not substantially differ between the various test specifications.

[b] Tests are performed including a constant term, assuming 12 lags. The correspondent rolling regressions (Phillips et al, 2009) are also available upon request.

3. Modelling price interdependence as price transmission mechanisms

In the previous section, some evidence emerged in favour of the presence of "something more" than a simple I(1) process in price series, but this evidence is not concordant across all prices. Still, however, visual inspection of price patterns (Figure 1) suggest that cereal prices clearly moved together, despite differences, over time. Therefore, it may be insightful to directly look for linkage (interdependence) across prices rather than examining their individual properties then looking for some common properties.

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. But such models are inherently very complex and hardly tractable in the empirical analysis, whereas the analysis of price evolution and linkage is more frequently afforded within reduced-form model. For example, Fackler and Goodwin (2001) provide a common template embracing all dynamic regression models, based on linear excess demand functions, from which an estimable reduced-form model (in their case, a VAR in prices) can be eventually derived within this framework. Reduced-form models are actually and by far more feasible and of immediate use to generate price predictions on the basis of the available observations and when the three dimensions are explicitly considered, i.e., to estimate $E\left(p_{i,k,t} \middle| p_{j,k,t-s}, p_{i,h,t-s}, p_{i,k,t-s}\right)$.

A generic reduced-form model of price formation and transmission over the three dimensions can be represented as follows:

(1a) $$p_{i,k,t} = \alpha_{i,k} + \sum_{s=1}^{s=S<T} \rho_s p_{i,k,t-s} + \sum_{s=0}^{s=S<T} \sum_{j \neq i} \omega_{ij}^s p_{j,k,t-s} + \sum_{s=0}^{s=S<T} \sum_{i=1}^{N} \sum_{h \neq k} \varpi_{kh}^s p_{i,h,t-s} + \varepsilon_{i,k,t}$$

where $\varepsilon_{i,k,t} \sim N\left(0, \sigma_{i,k,t}^2\right)$. Equation (1a) can be rewritten, as mentioned, in a more conventional form by distinguishing a cross-sectional dimension (given by the combination of the dimensions *ik*) and a time dimension, *t*. Therefore, we are interested in a reduced-form model predicting the expected value of the generic price $E\left(p_{ik,t} \middle| p_{jh,t-s}, p_{ik,t-s}\right)$ and (1a) becomes:

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

where $\varepsilon_{ik,t} \sim N\left(0, \sigma_{ik,t}^2\right)$. Equation (1b) can be written in a more compact matrix form:

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

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 space-commodity dimensions. The diagonal elements, $\omega_{ik,ik}^s$, actually represent the auto-correlation over time, with the exclusion of the matrix $\mathbf{W}_0$, where diagonal elements are evidently $\omega_{ik,ik}^0 = 0, \forall ik$. The off-diagonal elements, $\omega_{ik,jh}^s$, represent the cross-sectional dependence of prices; in other words, they express the degree of common movement shown by different prices and, therefore, the degree and the direction at which price shocks are transmitted.[5]

In particular:

- if *h=k* but *i≠j*, we are considering the spatial price transmission for the same commodity across different market places. In this case, under perfect spatial arbitrage, the Law of One Price (LOP) holds and we may expect $\omega_{ik,jk}$ =1 ;
- if *i=j* but *h≠k*, we are considering the price transmission between different commodities in the same market. In this case, elements $\omega_{ik,ih}$ indicate the degree of substitutability between the

[5] See Fackler and Goodwin (2001), for a comprehensive description. Indeed, the study of price transmission mechanisms implies referring to a number of economic concepts for which no common definitions exist in literature.

different goods (Dawson et al., 2006). $\omega_{ik,ih}$ will be close to 1 under perfect substitutability between *h* and *k*, while it will be close to 0 under low substitutability.

If model (1a-c) is specified in $\mathbf{lP}_t$ (the matrix of price logarithms), elements of $\mathbf{W}_s$ indicate the price transmission elasticities that is, how much of a percentage variation in $p_{ik,t-s}$ is transmitted to $p_{jh,t}$. Within 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. Elements of $\boldsymbol{\alpha}$ indicate these constant multiplicative terms (that can be naturally intended as percentages) applied to price $p_{ik,t-s}$ to obtain $p_{jh,t}$.

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

3.2. Model specification

The first specification issue concerning model (1a-c) has to do with its size. With 10 price series, the size of $\mathbf{W}_s$ becomes large especially whenever several lags (due to weekly data) had to be admitted. To reduce the size of the problem (and the number of parameters to be estimated), assumption can be made about the relevant interactions to be considered. Therefore, also to facilitate the economic interpretation of the results, the analysis of the price interdependence have to be "confined" and "segmented". In particular, we may firstly assume that price transmission only occurs within the same commodity (that is, p_{ik} and p_{jk} are interdependent) and within the same market place (that is, p_{ik} and p_{ih} are interdependent). The consequent assumption is that p_{ik} and p_{jh} have no direct linkage. This implies fixing at 0 some of the elements of the $\mathbf{W}_0$ matrix. Secondly, the relation across prices can be separately studied by group 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 details the groups of prices whose interdependence has been considered.

The conclusion achieved in the testing section, that all series $\mathbf{p}_{ik}$ can be considered I(1) processes with some of them also showing a temporary bubble, must be appropriately taken into account when developing, specifying and estimating model (1a-c). Actually, agricultural commodity price series are often found to be I(1); for this reason, since the seminal work of Ardeni (1989), cointegration techniques have been extensively used for the study of price transmission mechanisms. Cointegration models presuppose that variables exhibiting nonstationary behaviour will nonetheless be linked by a long run relation, whose residuals are stationary. When fixed-commodity and cross-market price relations are considered, under perfect spatial arbitrage, this relation is the LOP, which is expected to hold in the long-run (LR) despite in short-run (SR) prices are allowed to deviate from it.

The standard Vector Error Correction Models (VECM) equation can be written as

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

where $\mathbf{p}_t$ now is the (Vx1) vector containing the logarithms of the V prices at time t over the selected dimension (space ij, commodity kh).[6] $\boldsymbol{\beta}$ is the cointegration matrix which contains the long-run coefficients (the degree of price transmission). $\boldsymbol{\alpha}$ is the loading matrix which contains the adjustments parameters (a measure of the speed of price transmission,). $\boldsymbol{\Gamma}_i$ are matrixes containing coefficients that account for short-run relations, and $\boldsymbol{\varepsilon}_t$ are white-noise error terms. The rank of $\boldsymbol{\alpha\beta'}$ gives information about the presence of cointegration amongst the variables. The fundamental assumptions underlying (2) is that, the model expressed in logarithms, price spreads (and also, all components which account for price spreads) are a stationary proportion of prices[7].

As already mentioned, we cannot exclude the presence of explosive behavior in some of the price series. However, Engsted (2006) and Nielsen (2010) have shown that the use of the Johansen (1995) approach to test and estimate cointegration relationships is still admitted. 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 that is not I(1)); the second the co-explosive ones (their linear combination that is not explosive) (Engested, 2006, p. 2006):

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

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

The standard Johansen estimation procedure holds its validity, and since we are prevalently interested in the usual cointegrating relationship between prices, we can simply proceed in estimating (2).

Furthermore, in considering the bubble in our analysis of price linkages, we can assume that the period of exuberance (or the shock that generated it) influenced the cointegration relationship, $\boldsymbol{\beta}_1 = \boldsymbol{\beta}$, itself. In particular, we may be interested in taking into account 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, and proposed a model where breaks in the deterministic terms are allowed at known points in time. The sample is divided in q periods, separated by the occurrence of the structural breaks, where j denotes each period. The general VECM becomes:

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

where k is the lag length of the underlying VAR. $\mathbf{E}_t$ is a vector of q dummy variables that take the value 1, i.e. $E_{jt} = 1$, if the observation belongs to the jth period (j = 1, …, q), and 0 otherwise; that is, $\mathbf{E}_t = \begin{bmatrix}E_{1t} & E_{2t} & \dots & E_{qt}\end{bmatrix}'$. D_t is an impulse dummy that equals unity if the observation t is the ith of the jth period, and is included to allow the conditional likelihood function to be derived given the initial values in each period. w_t are the intervention dummies (up to d) included to obtain well-behaving residuals. The short run parameters are included in matrices $\boldsymbol{\gamma}$ (2 x q), $\boldsymbol{\Gamma}$ (2 x 2), $\mathbf{k}$ (2 x 1) for each j and i, and $\boldsymbol{\Theta}$ (2 x 2). $\boldsymbol{\varepsilon}_t$ are assumed to be i.i.d. with zero mean and symmetric and positive

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

[7] For a general review of the implications of the use of cointegration techniques in price transmission analyses see Listorti, (2009).

definite variance, Ω. $\boldsymbol{\mu} = \begin{bmatrix} \mu_{1t} & \mu_{2t} & \dots & \mu_{qt} \end{bmatrix}'$ is the vector containing the long run drift parameters and $\boldsymbol{\beta}$ contains the usual the long run coefficients in the cointegrating vector.

The cointegration hypothesis is formulated by testing the rank of $\boldsymbol{\pi} = \boldsymbol{\alpha} \begin{bmatrix} \boldsymbol{\beta} \\ \boldsymbol{\mu} \end{bmatrix}'$; its asymptotic distribution can not be generalized as it depends on the number of non-stationary relations, on the location of breakpoints and on the trend specification. The Johansen et al. (2000) procedure directly stems from the Johansen framework.

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

Two structural breaks are taken into account in the present study as both might have influenced the LR price transmission relations. For this reason, a "bubble" and a "policy" dummy have been included in the cointegration space as exogenous variables following (4). Based on the results obtained from tests on presence and timing of the explosive behaviour (Figure 1), the former 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.

It is well known that agricultural markets are often subject to considerable policy intervention (Listorti 2007, 2009). In particular, the trade policy regime may have had a role because border policy for cereals changed during the years of observation. It must be recalled that the EU protection mechanism for cereals, even after the 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, as a reaction to the exceptionally tight situation on the world cereals markets, 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.[8] This temporary measure might have altered the price transmission mechanisms between international (Rotterdam) and national (Italian) prices. To take into account this aspect within the adopted model, the policy dummy takes the value 1 for all weekly observations between January 2008 and October 2008, and 0 otherwise.

3.4. Econometric procedure

Following from the considerations presented in the previous paragraph, an appropriate econometric procedure has been put in place. This methodology has been repeated for fixed-market/cross-commodity or fixed-commodity/cross-market prices (in this case, both with and without the international prices). First of all, cointegration among prices within the group is assessed using the conventional Johansen (trace) test. If cointegration is found, then the respective VECM of the prices is estimated.[9] If cointegration is not found, and no explosive root is present within the price group, a first-difference VAR is estimated. Finally, if an explosive root is present, without cointegration, no model specification should be adopted as first order differentiation itself can not ensure the removal of the explosive pattern. Nonetheless, we still report estimates of the respective first-difference VAR in cases where an explosive root is detected (see Table A.2 and Table 6) as these results may provide further information on the presence of co-explosiveness.

In all the cointegration tests and VECM estimates, the "restricted constant case" of the Johansen procedure has been considered. The series don't show any linear trend in levels, but both theory and visual inspection of the data imply the presence of a constant in the long-run relationship,

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

[9] The lag length is selected according to the conventional information criteria (HQIN, SBIC, AIC). In all models, autocorrelation has been tested with a LM test up to the fourth lag.

accounting for all elements contributing to price differentials not explicitly modelled in the price transmission equation. This means that, even if there are no linear deterministic trends in the level of the data, the cointegrating relation has a constant mean.

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. Therefore, in equation (4), t has been fixed =1 and γ=0, a constant term included with the corresponding elimination of one of the q dummy variables. The coefficients of the structural break dummies have to be interpreted as relative to the constant valid over the overall 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. As a matter of fact, the Johansen, et al. (2000) procedure doesn't allow testing for the cointegration rank with the number of breaks here considered. For this reason, unit root 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.[10]

The introduction of the bubble and policy dummies is straightforward when prices are not cointegrated. Within the standard first-difference VAR to be estimated in this case, structural breaks simply enter as exogenous dummy variables in the VAR, thus allowing for a shift in the constant term of the VAR equations. In such case, however, the response to a price shock can not be interpreted as a reversion to some LR relationship because no evidence support the existence of such LR pattern.

Weak exogeneity tests (i.e., convention t-tests on coefficients of vector $\boldsymbol{\alpha}$) for the estimated VECM and Granger causality tests for the VAR, are eventually performed to assess how price horizontally transmits from p_{ik} to p_{jk} or p_{ik} to p_{ih}. Through these tests, for any group of prices we can identify which causal relationships emerge, that is, which are the "central" (leader in price formation) and "satellite" (follower) markets (Verga and Zuppiroli, 2003).

4. Results

4.1. Cross-market transmission

In this section, the models estimated in the fixed-commodity/cross-market case are analysed. Both for durum wheat and for corn, the models have been estimated considering both national and international (Rotterdam) prices. In the case of durum wheat, the rank of the cointegration matrix is equal to one (Table 2). Though the 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 can be excluded from the VECM model and no co-explosiveness specification (as in (6)) has to be estimated. This may be attributed to the presence of the structural breaks that may take into account the period of more intense price turmoil. Nonetheless, results suggest that an explosive root can be excluded even when structural breaks are not included in the model specification.

Both coefficients of the national prices within the cointegration space are significant (-0.333 for the Rome price and -0.741 for the Foggia price), but the coefficient of the Rotterdam price is much lower than the others and has not the expected sign (0.030). Both the Rome and the Foggia price adjustment coefficients are significant (they are, respectively, 0.138 and 0.255), 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 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

[10]According the discussion above, however, due the presence of an explosive ρ>1 we can not exclude that the linear combination $\boldsymbol{\beta}'\mathbf{p}_{t-1}$ contains an explosive component.

dummy), but not significant. Since the constant in the cointegration vector is positive (0.501), this would indicate that in both time frames the distance between the prices has increased.
In the case of corn (Table 3), two cointegration vectors emerge. In estimating the VECM, however, we impose only one cointegration vector. The residuals form the cointegrating relations are stationary and 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 over the whole period (Table 1). In the VECM without the bubble and policy dummies the coefficient of the Bologna price is very close to one, and significant (-1.184). The coefficient of the Rome price is positive and significant (0.225), whereas the coefficient of the Rotterdam price is negative and not significant (-0.039). The Bologna and the Rotterdam prices behave as weakly exogenous, thus confirming that both can be considered as driving markets, while the adjustment coefficients of the Milan and Rome price are significant (-0.137 and -0.227), although the coefficient of the Milan price has not the correct sign. When the structural breaks are included, within the cointegration vector, the adjustment coefficient of the Bologna price is -0.492, the one of the Rome price is -0.632, and the one of the US price is 0.060. None of them is significant. The coefficient of the bubble dummy is positive (0.023), whereas the policy dummy has a negative coefficient (-0.027) but both are not statistically significant. Considering the sign of the constant term, this would mean that during the bubble the distance between the prices widened, while it diminished during the suspension of the EU import duties. Only the Rotterdam price has not a significant adjustment coefficient.

Table 2 – Cross-market price linkage: VECM estimates (standard errors in parenthesis) – durum wheat, national and international prices [a, c]

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	2.709*	-2.560*
Stability condition (highest eigenvalue, modulus) [c]	0.886	0.792

[a] Estimated coefficients of the first-difference terms are available upon request. The optimal lag of the VECM has been selected according to the conventional information criteria. Whereas HQIC, AIC and SBIC indicated 2 lags, 3 lags were necessary in order to remove autocorrelation in the residuals of the VECM (LM test up to the 4th lag). 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 3 – Cross-market price linkage: VECM estimates (standard errors in parenthesis) – corn, national and international prices [a, c]

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	-4.044*	-3.394*
Stability condition (highest eigenvalue, modulus) [c]	0.850	0.792

[a] Estimated coefficients of the first-difference terms are available upon request. The optimal lag of the VECM has been selected according to 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 to remove autocorrelation in the residuals (LM test up to the 4th lag). 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

4.2. Cross-commodity transmission

The results of models concerning the fixed-market/cross-commodity (durum wheat vs corn) relationships are reported in Tables 4-6. Three market places have been analysed: Central Northern Italy (Bologna), Central-Southern Italy (Rome), and International (Rotterdam). In all cases we observe that cointegration rank is 0 and a first-difference VAR specification is thus estimated. Therefore, no long-run relationship can be detected between the corn and durum wheat prices in any of the market places considered. This does not mean that no linkage occurs across prices but, if present, this linkages is limited to short-run responses to other price's shocks. Only in the Bologna market we notice a 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 Rotterdam market where is durum wheat price to Granger-cause corn price at 10% significance level. This latter effect, however, vanishes whenever structural breaks are included. The conclusion could be that the only clear linkage across commodity prices is the dependence of durum wheat on corn prices in the Italian markets.

Parameters associated to structural breaks' dummies tend to confirme what observed in the cross-market analysis. The price bubble tends to significantly increase the price variations in response to exogenous shocks but then this effect is entirely compensated by the opposite effect of the policy. In magnitude, however, the impact of these dummies is quite limited. An additional comment has to be made with respect to the presence of explosive behaviour of some prices. As we consider the Rotterdam durum wheat price the only case for which explosiveness is definitely observed, we may conclude that the first-difference VAR estimates in the case of the international market are expected

to show explosiveness and this would hamper the validity of the respective parameter estimates. Nonetheless, Table 6 shows that no explosiveness is observed in this case, as residuals are stationary and stability condition is met as occurs also in the other two market places where explosiveness has been, in fact, excluded. Even in this case, therefore, there is no reason to look for an alternative specification taking into account co-explosiveness.

Table 4 – Cross-commodity price linkage: first-difference VAR estimates – 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[b]		
Durum Wheat	-3.74*	-4.27*
Corn	-4.83*	-5.29*
Stability condition (highest eigenvalue, modulus)	0.790	0.749

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag ($s=5$) 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 – Central-Southern Italy (Rome) [a,c]

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	6.83	6.23*
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[b]		
Durum Wheat	-3.24*	-4.12*
Corn	-4.56*	-4.91*
Stability condition (highest eigenvalue, modulus)	0.714	0.553

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag ($s=5$) 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

As this paper aims at analysing the horizontal transmission (across space and commodities) of agricultural prices during a period of extreme market turbulence, namely, years 2007-2008, we can now combine these results together and come back to the original research questions. First of all, econometric evidence would suggest that the "bubble", if and when present, does not seem to affect very much the analysis of price interdependence. Not only because appropriate econometric instruments are now available to take explosiveness into account, but also because no evidence of co-explosiveness emerges from VECM-VAR model estimations. Even when present in single prices, explosiveness is partially captured by structural breaks and by price linkages themselves.

The temporary trade-policy measure (suspension of the EU import duties on cereals) may itself have had a role as it behaves as a shifter that reduce the cross-market and cross commodity price gaps and variations, thus compensating the opposite effect induced by the price bubble. Estimates suggest strong cross-market linkages, with the international market behaving as the driver together with the Bologna market at least in the durum wheat case. On the contrary, weak cross-commodity linkages are observed, if not in the national markets with durum wheat behaving as endogeneous. As the international durum wheat price is the one showing the highest price peak during the "bubble" these results would suggest that cross-market and cross-commodity linkages do not favour the transmission of this turbulence to the national prices and the corn price; the opposite would rather hold true.

Table 6 – 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	3.63	3.52
Corn on Durum Wheat	6.83	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 [b]		
Durum Wheat	-3.65*	-4.61*
Corn	-4.16*	-4.14*
Stability condition (highest eigenvalue, modulus)	0.681	0.583

[a] The other estimated coefficients of the VAR are available upon request. The optimal lag (s=5) 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

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Figure 1 – The price bubble: cereal price series over the four years between May 2006 and December 2010 (see Annex for price codes)

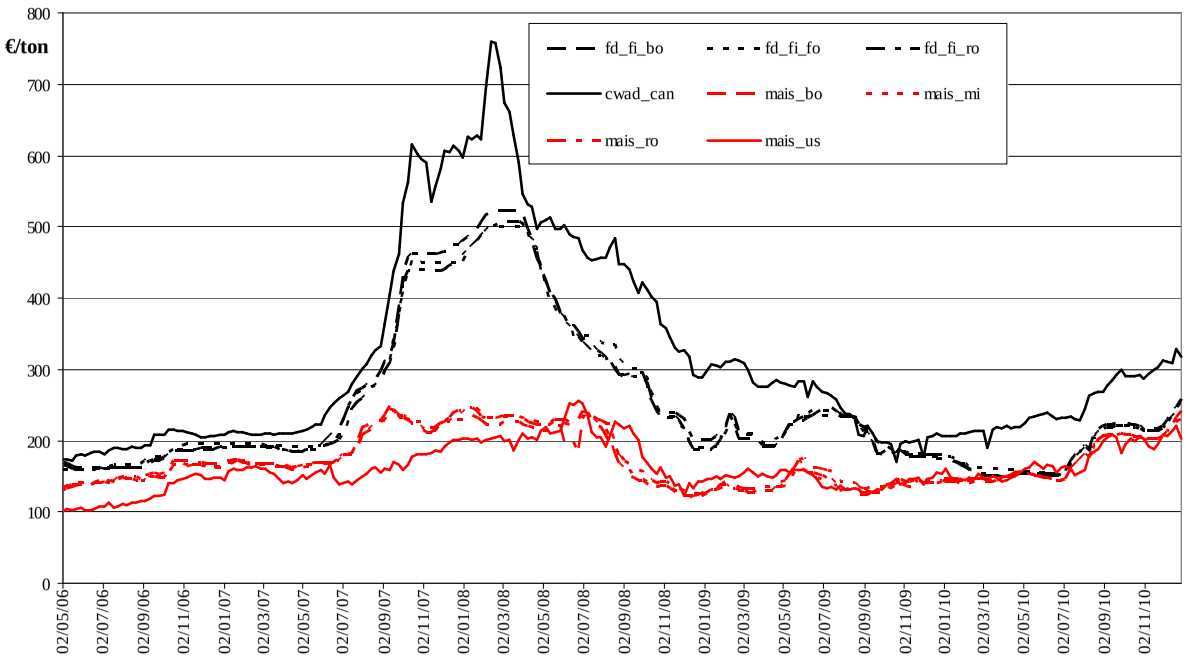
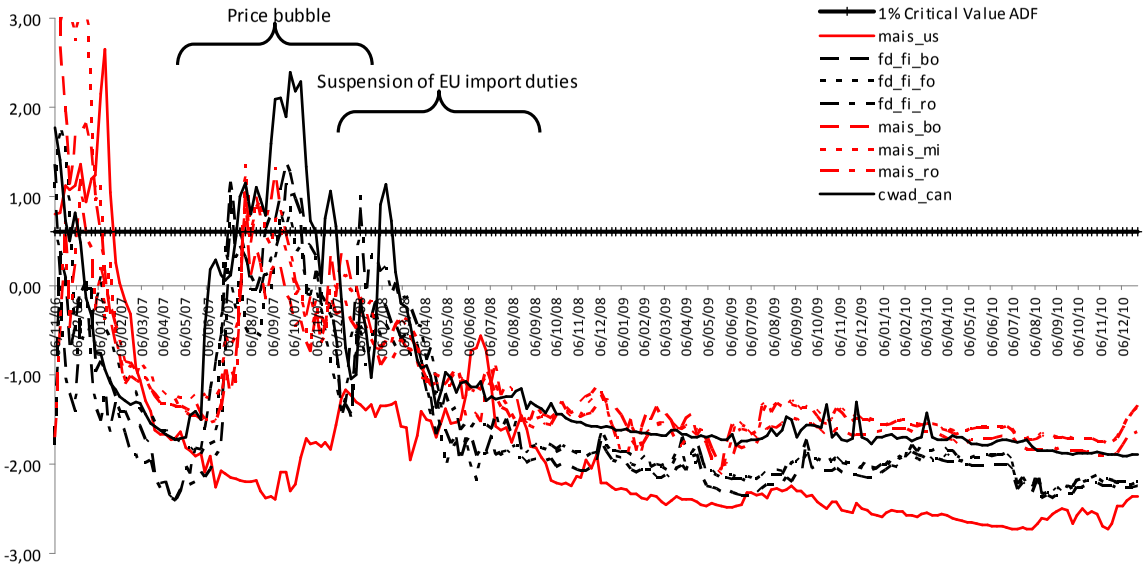


Figure 2 – Dating the bubble: time series of forward recursive ADF t-statistic for the cereal price series (logarithms) in Italian and American market (see Annex for price codes)



ANNEX

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
fd_fi_fo	Durum Wheat, Fino	Foggia
fd_fi_ro	Durum Wheat, Fino	Rome
mais_bo	Maize, Ibrido Nazionale	Bologna
mais_mi	Maize, Ibrido Nazionale	Milan
mais_ro	Maize, Ibrido Nazionale	Rome
cwad_can[a]	Wheat, Canada Western Amber Durum (CWAD)	Canada, St Lawrence
mais_us[a]	Maize, #3 Yellow Corn (3YC)	US, Gulf

[a] CIF price - Rotterdam

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 Space		
DURUM WHEAT		VECM
fd_fi_bo	I(1)	
fd_fi_ro	I(1)	
fd_fi_fo	I(1)	
cwrs_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 Space-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 - Rotterdam