Price transmission in three Italian Food Chains: a structural break approach

Carraro, A., Stefani G.
Dep. of Plant Soil and Environmental Sciences
University of Florence (Italy).

Price transmission in three Italian food chains: a structural break approach

Abstract

Recently a wide instability of food prices has been observed in world and European agricultural and food markets. Both media and policy makers have dealt with the unsatisfactory patterns of marketing margins and price transmission along the food chain which may bring about distributive issues and affect inflationary trends. Although price transmission and margins dynamics have attracted so much interest at the policy level, few Italian studies deal with this topic.

Our aim is to provide a first analysis of the price transmission mechanism in three Italian agri-food chains (lamb, pork and pasta), within a structural change framework.

Results show that structural breaks in the price transmission mechanism are an issue in the food chain of pasta and pork with the regime change arising in occasion of the price bubble of 2007-2008.

Price transmission in three Italian food chains: a structural break approach

1. Introduction

Recently a wide instability of food prices has been observed in world and European agricultural and food markets. Both media and policy makers have dealt with this issue highlighting the unsatisfactory patterns of marketing margins and price transmission along the food chain. At European level, studies about price transmission were commissioned by Institutions such as the European Parliament (Agra CEAS, 2007) or the UK Department for Environment Food and Rural Affaires (London Economics, 2004) concerned about the possible impacts of the new Common Agricultural Policy on consumer prices. According to a recent Commission Communication (COM 2009(591)) "a better functioning food supply chain is crucial for consumers and for ensuring a sustainable distribution of value added along the chain, thus contributing towards raising its overall competitiveness".

A related issues concerns the degree of pass through from raw commodity prices to consumer food prices and its impact on inflationary (or deflationary) trends as food account for about 20% of Euro area consumption (Ferrucci et al. 2010; National Bank of Belgium, 2008)).

Italian Institutions have commissioned studies on this topic too. The Ministry for Economic Development recently published a research on the dynamic of prices along the wheat chain (IPI, 2008). The Italian Antitrust Authority (AGCM, 2007; Giangiulio and Mazzantini, 2010), concerned about the presence of market power in the food chain, launched an inquiry on the food retail sector to check for anti-competitive practices along the marketing chain.

A theoretical question related to the above issue is about the nature and causes of the observed patterns of price transmission along the food chain. Surprisingly, although price transmission and margins dynamics have attracted so much interest at the policy level, few Italian studies deal with this topic. Frey and Manera (2007) in a recent literature reviewed list only 4 papers on Italian markets out of 64 studies, and they were about gasoline.

Most of the literature on price transmission is about estimating elasticities and possibly detecting the presence of asymmetries in the transmission mechanism. This field of econometrics has witnessed during the 80's the so called unit root revolution when the concept of cointegration was introduced and applied to empirical studies on price transmission (Meyer and Von Cramon Taubadel, 2004).

Only recently, the potential confounding between non stationarity of time series and structural changes was highlighted (Boetel and Liu, 2008). However, few studies have applied this framework to food prices so far. Non stationarity or lack of cointegration for a number of agricultural price was questioned by Wang and Tomek (2007): once structural breaks were accounted for, most series previously considered integrated of order one turned to be stationary. Boetel and Liu (2008), which provide a brief review of earlier studies, found evidence of cointegrating relationships along the pork and beef

chains in the U.S only after accounting for structural breaks. Adachi and Liu (2009) identified several regimes in the Japanese pork retail-farm price relationship.

The recent food price instability urges applied economists to take into account structural breaks in analysing long term price relationships in the food sector. Our aim is to provide a first analysis of the price transmission mechanism in three Italian agri-food chains (lamb, pork and pasta), within a structural change framework.

The paper is organized as follows. Section 2 presents a brief discussion of the recent structural change methodologies employed in applied research. We considered in turn the Zivot and Andrews (1992) test for stationarity, tests for cointegration (Gregory and Hansen, 1996; Carrion y Silvestre, 2006) and procedures to test for presence of single or multiple structural breaks and to identify breaks dates (Bai and Perron 2003).

Section 3 contains an empirical application to three Italian food chains price data. We checked for stationarity of price series and, whenever applicable, we tested for cointegration between the couples of price series. Finally, we analyzed the long run price relationship within each regime delimited by the break dates estimated with the Bai and Perron procedure.

Section 4 concludes and provides a summary of major results and suggestions for future research.

2. Structural change vs non stationarity in linear regression

Unit root tests

As clearly stated by Perron (2005) testing for a unit root against trend stationarity is equivalent to addressing the following question: "do the data support the view that the trend is changing every period or never?". The problem with conventional unit root tests, such as the widely used ADF - Augmented Dickey Fuller (1979), is precisely that the unit root null is tested against the extreme alternative of a trend that always changes, discarding the case of a trend that changes only "sometimes". Once allowance is made for one or more changes in the trend function the question addressed by the test is "do the data favor a view that the trend is 'always' changing or is changing at most occasionally?" (Perron, 2005, p. 48-49).

As a standard procedure to test the non stationarity of a series the ADF test is based on the regression:

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

non stationarity is refused when the test suggests that α is different from 1. However, when a structural break is present in the data generating process the conventional ADF test is biased toward the acceptance of the null resulting in a dramatic loss of power.

Considering the case of a one time change Perron (1989, 1990) proposes a modified version of the ADF test applicable to four main cases of structural change:

    a) non trending series: change in level
    b) trending series: change in level
    c) trending series: change in slope
    d) trending series: change in level or in slope

Each of the four cases can be modelled as if the change occurs instantaneously (additive outlier) or gradually (innovation outlier). The test proposed by Perron was based on know break dates, that is the date when the shift occurs must be known by the researcher before analysing the data. As this procedure may lead to some sort of data mining (for example when the break date is inferred by looking at the graphs of the time series), Zivot and Andrews (1992) suggested to determine the break date endogenously via a search algorithm. They proposed to search for the break date that gives the minimal value of the t statistics of the adjusted ADF:

$$t_{\hat{a}}[\hat{\lambda}_{inf}] = \inf_{\lambda \in \Delta} t_{\hat{\alpha}}(\lambda) \qquad (2)$$

where λ is the fraction $(T\tau/T)$ of the structural break point with respect to the whole sample, and Δ is a closed subset of (0,1).

The adjusted ADF statistics of Zivot and Andrews (henceforth ZA) is based on variations of the following regression which refers to case (d) above with gradual shifts (innovation outlier):

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \qquad (3)$$

where $DU_t (\lambda) = 1$ and $DT^*_t = T - \lambda T$ for $t > \lambda T$ and zero otherwise. In the simple change in level case (b above), DT is always zero, while in the change in slope case (c above) it is DU that is equal to zero. One possible drawback of the ZA test is that it considers a null of unit root process with no break against the alternative of a stationary process with one break. When a break is actually present under the null, the test involves size distortion. However such distortions are relevant only with implausible large shifts and they are hardly relevant in practice (Perron, 2005). Methodological developments in the area of unit root tests with structural breaks include generalizations to multiple breaks, to unit root with breaking trend null hypothesis and extension to tests for the null of stationarity such as the Kwiatkowski et al. (1992) or KPSS test (see Perron (2005) for a review and Adachi and Liu (2009) for an application to food price series).

Cointegration tests

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

$$y_t = \mu + \alpha x_t + u_t \qquad (4)$$

where x_t is a vector of independent variables. The system is cointegrated if the errors u_t are *I*(0). In that case the relation (4) may be interpreted as a long run equilibrium toward which the process y_t tends.

Within the conventional Engle and Granger (1987) test for cointegration the static equation (4) is first estimated via OLS and then the stationarity of the residuals of this relationship is tested via an ADF test using the critical values proposed by MacKinnon (1991). If the residuals are found to be stationary, the two series are maintained to be cointegrated. Gregory and Hansen (1996) extend the residual test to take into account a possible break in the long-run relationship of unknown date. As in ZA, the test statistic is the minimal value of the t statistics across all possible break dates. The authors consider three modified version of equation (4) that includes dummies for the structural change :

C $\qquad y_t = \mu + \theta DU_t + \alpha x_t + u_t \qquad (5a)$

C/T $\quad y_t = \mu + \theta DU_t + \beta t + \alpha x_t + u_t$ $\qquad$ (5b)

C/S $\quad y_t = \mu + \theta DU_t + \alpha_1 x_t + \alpha_2 DU_t x_t + u_t$ $\qquad$ (5c)

Model C entails a level shift in the equilibrium relationship, model C/T adds a trend to the previous model whilst model C/S deals with regime shift by adding a change in the slope coefficients. The authors provide asymptotically critical values for both the ADF test and the Phillips et al. (1988) Z_a and Z_t statistics.

Cointegration tests in the presence of structural breaks have been also framed maintaining the reversed null of cointegration . Carrion y Silvestre and Sanso (2006) discuss six modified versions of equation 4 adding to the models in (5) three further specifications[1]:

B $\quad y_t = \mu + \beta_1 t + \beta_2 DT_t + \alpha x_t + u_t$ $\qquad$ (6a)

C $\quad y_t = \mu + \theta DU_t + \beta_1 t + \beta_2 DT_t + \alpha x_t + u_t$ $\qquad$ (6b)

E $\quad y_t = \mu + \theta DU_t + \beta_1 t + \beta_2 DT_t + \alpha_1 x_t + \alpha_2 DU_t x_t + u_t$ $\quad$ (6c)

As pointed out by Perron (2005), the LM-type test statistics proposed by the authors is a modification of the Gardner's (1969) Q statistics[2]. The residuals for the Q statistics are obtained from the OLS estimation of equations (5) or (6) scaled by an estimate of the long run variance. In the general case, when x_t is allowed to be endogenous, the dynamic least squares (DOLS) estimator (Stock and Watson, 1993) is used. The break date is estimated by minimizing the sum of squared residuals.

Recently, both the Gregory and Hansen and the Carrion and Sanso tests have been extended to the two breaks case (Abdulnasser, 2008 ; Carrion and Sanso, 2007).

Estimation of structural breaks and break dates

Bai and Perron (1998) provide a procedure to estimate structural changes in a linear model with stationary variables. The procedure may be illustrated looking at a simpler pure structural change model, that is a model were all coefficients are subject to *m+1* regime changes:

$y_t = x'_t \delta_i + u_t$ for t = $T_{i-1}+1$, …., T_i and i = 1, …., m+1 $\qquad$ (7)

where x_t is a vector of regressors (among which possibly a constant and/ or a trend), the corresponding vector of coefficients δ_i is indexed over the *m* regimes and the indices T are the break points.

For a given *m* partition of the sample (T_1,….T_m), δ_i can be estimated by OLS minimizing the sum of squared residuals:

$$S_T(T_1,\ldots,T_m) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} \left[y_t - x'_t \delta_i \right] \qquad (8)$$

Then an estimate of the *m* break points is given by the set of break points that minimize the estimated $\hat{S}_T$ over all possible *m*-partitions of the sample (provided that a minimum size condition for each segment is fulfilled).

---

[1] Model An, A and D by Carrion i Silvestre and Sanso correspond respectively to models C, C/T and C/S by Gregory and Hansen.

[2] A similar test is proposed by Arai and Kurozumi (2005)

Bai and Perron (2003) proposes a dynamic programming algorithm in order to reduce the dimension of the search problem to manageable size. They also demonstrate that break dates are asymptotically independent when all variables in the model are stationary[3] providing methods to calculate confidence intervals.

As far as the number of breaks is considered, a number of tests are discussed by Bai and Perron (1998). Sup F test is given by the maximum value of a Wald test for the null hypothesis of no structural change versus the alternative of k changes, calculated over all possible k partitions with a common minimal length for each segment.

Double maximum tests are used to make inference without specifying a given number of breaks. As the name suggests they are the (weighted or not) maximum of the previous sup F test across all possible number of breaks up to a pre-specified maximum. Their use is advised since the power of sup-F test may be low when the actual number of breaks is greater than the one specified (Perron, 2005).

A sequential testing procedure can be based on the test of the null of l breaks against the alternative of $l+1$ breaks. Each step requires to carry out $l+1$ sup F test for one break in each of the $l+1$ segments obtained with the usual minimization of the sum of square residuals in (8). The hypothesis of one additional break is retained if the overall minimum value of $\hat{S}_T$ across all $l+1$ break models is sufficiently smaller than the sum of squared residuals from the l break model.

Finally, the number of breaks can be estimated with information criteria such as:
Schwarz's Bayesian Information Criteria (BIC, Schwarz, 1978) and Akaike (1973) Information Criteria (AIC). According to Bai and Perron (2003) the AIC performs always badly. In absence of serial correlation and when the breaks are actually present the BIC performs reasonably well. However when serial correlation is an issue or a lagged dependent variable is included in equation (7) none of the criteria is adequate.

3. An application to three Italian food chains

We applied the methodologies illustrated in section 2 to three Italian food chains for which price series are available at producer, wholesale (or industrially transformed) and consumer stages: Pasta, Lamb and Pork. First we checked for stationarity of price series. Second, we tested for cointegration between couples of price series. Finally, break dates were estimated with the Bai and Perron procedure and the long run price relationships were analysed within each regime.

Data

Datasets on farm, wholesale and retail monthly prices for pasta, lamb and pork were sourced from Datima and from the household panel ISMEA-Nielsen provided by ISMEA. The dataset spans from January 1994 through December 2008 for the farm and wholesale prices and from February 2000 through June 2010 for the retail price series. Datima is a collection of statistical databases including foreign trade and agricultural markets data, whereas ISMEA-Nielsen is a household panel designed to analyse the growth of domestic food consumption. Price data refers to aggregated product categories.

---

[3]   Methods to construct confidence intervals when the variables are I(1) but integrated are provided in Kerjwal and Perron (2008). In this case distributions of break dates are not asymptotically independent
.

Monthly farm prices of lamb and hogs are in Euro per Kg and in Euro per Kg slaughter weight respectively, as well as wholesale and retail prices about lamb and pork cuts. Durum wheat, semolina and pasta prices are in Euro per kg.

The analysis has been carried out on the period from February 2000 to June 2010 for which data are available at all market stages. All series were transformed in natural logarithms and deseasonalised by regressing the transformed series on monthly dummies.

Testing for unit roots

Stationarity of the series were first checked with the conventional ADF test that does not allow for any break in the data generation process. Results were compared with those obtained with the Zivot and Andrews (1992) test, described in section 2[4].

In order to select the appropriate number of lags for our analysis we applied the Schwert (1989) criterion:

$$Lag_{max} = 12 * (T/100)^{1/4} \tag{9}$$

Schwert criterion provides a larger number of lags with respect to other criteria such as AIC or BIC. However, a larger number of lags makes the actual size of the test closer to its nominal value (Harris, 1999, p.36).

Results from the ADF tests for pasta, lamb and pork meat are illustrated in table 1, 2 and 3 respectively. In the no-break column of each table the ADF test with both trend and drift and with drift alone is reported, whereas the one break column displays the ZA test results. The ADF test suggests the presence of a unit root in all series since the null hypothesis cannot be refused at the 5% level of significance everywhere.

On the contrary, the ZA test provides a different picture. Once a break in the deterministic trend is allowed for, the null hypothesis of a unit root process is rejected in three out of nine series. We run the test in the three versions illustrated in section 2, that is including a deterministic trend and allowing a shift either in the intercept or in the slope of the trend or in both.

In the pasta chain series a structural break was found only at the retail stage. The estimated date is January 2006 with a model fitted with a drift and a change in the trend slope. Also farm prices of lamb appear to be stationary with a break in March 2006 (a trend is included in the model).

In the case of pork meat we reject the null hypothesis of presence of a unit root both in retail and wholesale prices (ZA test with drift and with both change in trend slope and drift, respectively). These series appear to be stationarity with a structural break, showing a structural change in September and June 2007 respectively. Producer prices are found to be $I(1)$ confirming the ADF test.

---

[4]   ADF and ZA test were carried out in R, an open sources statistical software, employing respectively the functions *ur.df* from package **uroot** and *ur.za* from package **urca**.

Table 1 - Unit root test results for *pasta* series

Durum Wheat Semolina Pasta		No Break		One Break		
		Augmented Dickey & Fuller (ADF)		Zivot & Andrews (ZA)		
		Trend & drift	*drift*	*w/change in drift*	*w/change in Trend*	*w/ change in Trend & Drift*
wheat prices	*Test value*	-2,69	-2,99	-4,08	-4,71	-4,70
	Break date	n.a.	n.a.	n.a.	n.a.	n.a.
semolina prices	*Test value*	-2,25	-0,42	-4,31	-2,81	-2,86
	Break date	n.a.	n.a.	n.a.	n.a.	n.a.
pasta prices	*Test value*	-2,69	-1,15	-5,11**	-2,93	-5,39**
	Break date	n.a.	n.a.	gen-06	n.a.	gen-06

** Null hypothesis of non stationarity rejected at 5% of significance. t indicates a trend included in the ADF model
d indicates a drift included in the ADF model

Table 2 - Unit root test results for *lamb* series

Lamb		No Break		One Break		
		Augmented Dickey & Fuller (ADF)		Zivot & Andrews (ZA)		
		Trend & drift	*drift*	*w/change in drift*	*w/change in Trend*	*w/ change in Trend & Drift*
Farm prices	*Test value*	-2,97	-2,89	-4,92	-3,69	-5,55**
	Break date	n.a.	n.a.	n.a.	n.a.	mar-06
Wholesale prices	*Test value*	-3,04	-0,81	-4,27	-3,48	-4,31
	Break date	n.a.	n.a.	n.a.	n.a.	n.a.
Retail prices	*Test value*	-3,18	-1,77	-4,54	-3,38	-4,88
	Break date	n.a.	n.a.	n.a.	n.a.	n.a.

** Null hypothesis rejected at 5% of significance. t indicates a trend included in the model
d indicates a drift included in the model

Table 3 - Unit root test results for *pork* series

Pork		No Break		One Break		
		Augmented Dickey & Fuller (ADF)		Zivot & Andrews (ZA)		
		Trend & drift	*drift*	*w/change in drift*	*w/change in Trend*	*w/ change in Trend & Drift*
Farm prices	*Test value*	-1,87	-1,77	-3,84	-2,09	-4,44
	Break date	n.a.	n.a.	n.a.	n.a.	n.a.
Wholesale prices	*Test value*	-2,23	-1,03	-4,29	-2,12	-5,36**
	Break date	n.a.	n.a.	n.a.	n.a.	giu-07
Retail prices	*Test value*	-1,94	-1,76	-5,35**	-3,65	-4,58
	Break date	n.a.	n.a.	set-07	n.a.	n.a.

** Null hypothesis rejected at 5% of significance. t indicates a trend included in the model
d indicates a drift included in the model

Cointegration analysis

Although some of the series we checked for unit roots were found to be stationary with a breaking trend, we conservatively run tests of cointegration for all possible couple of series at the three marketing stages: producer-consumer, producer-wholesale and wholesale-consumer[5]. We first run the conventional test of Engle and Granger (EG) and then the two tests that account for a break in the cointegration relationship described in section 2: Gregory and Hansen (GH) and Carrion y Silvestre and Sanso (CS)[6]. In order to maintain comparability across tests we always included in the equilibrium equation both constant and trend[7].

We set the maximum lag to 6 and we used the BIC to select appropriate lag lengths. Specifications adopted in this procedure allow for a level shift in a trending equation or level and regime shifts (see equation 5b and 5c above) in the case of GH test and for level and regime shift only in the CS case (equation 5c displayed above). ADF statistics for EG and GH as well as estimated break dates are shown in table 4.

We also report break dates and LM-type statistics[8] for the CS test. However, it must be recalled that the break date estimated through the CS procedures is obtained by minimizing the sum of squared residuals of the long run equation and is not consistent with the one of the GH test that is selected through minimization of the t statistic of the ADF.

The cointegrating relationships are confirmed either in EG, GH or CS tests for the wheat-semolina and lamb wholesale-retail couples of series. Interestingly, we cannot reject the null of no-cointegration in the pork farm-wholesale case according to EG, whilst this couple of series appears to be cointegrated once a break is allowed for.

With the GH test we found a stronger evidence of a cointegrating relationship (rejecting the null hypothesis of absence of cointegration) when a regime change rather than a trend is included in the model. Overall, the CS test confirms the GH but for the lamb farm-retail and the pork farm-retail and wholesale-retail couples.

As far as break dates are concerned, GH test statistics agree with the CS tests for all the series with the sole exception of the lamb wholesale-retail couple.

---

[5] In the case of the pasta marketing chain the wholesale stage refers to the product of the first transformation durum wheat semolina.

[6] All estimates were obtained using the software R . To this purpose, we ported in R from Gauss both the script made available by Hansen in his webpage (http://www.ssc.wisc.edu/~bhansen/) and the code posted by Carrion y Silvestre (http://riscd2.eco.ub.es/~carrion/Welcome.html). R codes are available from the authors upon request.

[7] This implies running a simple ADF test on residuals not including neither a trend nor a constant when looking for stationarity in residuals of the equilibrium equation with the EG test.

[8] We used DOLS residuals following Carrion y Silvestre and Sanso (2006) suggestion that this version of the test has better size and power properties irrespective of the endogeneity or exogeneity of regressors.

Table 4 – Cointegration Tests

Chains			Engle & Granger (BIC)		Gregory & Hansen (BIC)		Carrion-i-Silvestre & Sansò (min SSR)
			Drift	*Trend & Drift*	*C/T*	*C/S*	*D*
Pasta	Wheat/Pasta	Test value	-2.7	-2.82	-4.19	-4.02	0.30**
		Break date	n.a.	n.a.	Oct-08	Oct-08	Oct-07
	Wheat/Semolina	Test value	-3.35	-4.15**	-5.26**	-5.96**	0.04
		Break date	n.a.	n.a.	Aug-07	Jun-07	Jun-07
	Semolina/Pasta	Test value	-2.53	-2.65	-4.66	-4.54	0.27**
		Break date	n.a.	n.a.	Oct-08	Oct-08	Dec-07
Lamb	Farm/Retail	Test value	-4.80**	-4.93**	-6.91**	-5.91**	0.34**
		Break date	n.a.	n.a.	Aug-01	Apr-04	Dec-08
	Farm/Wholesale	Test value	-3.56	-3.86	-7.12**	-4.61	0.19**
		Break date	n.a.	n.a.	Mar-03	Nov-04	Sep-04
	Wholesale/Retail	Test value	-4.67	-6.43**	-6.79**	-7.29**	0.04
		Break date	n.a.	n.a.	Aug-01	Agu-01	Oct-00
Pork	Farm/Retail	Test value	-2.66	-3.03	-2.94	-4.24	0.05
		Break date	n.a.	n.a.	Jul-01	Jan-06	Jul-05
	Farm/Wholesale	Test value	-2.58	-3.32	-6.52**	-6.87**	0.1
		Break date	n.a.	n.a.	Feb-07	Feb-07	Jan-07
	Wholesale/Retail	Test value	-1.82	-1.99	-3.87	-3.94	0.04
		Break date	n.a.	n.a.	Nov-01	Jul-05	Jan-06

** Reject the null hypothesis at the 5% of significance; Null for Carrion y Silvestree and Sanso test is cointegration.

Estimating structural breaks and break dates

To estimate break dates for the cointegration relationships above, we employed the Bai and Perron (2003) dynamic programming algorithm as implemented by Zeileis et al (2005)[9]. The algorithm provides confidence intervals for break dates that are valid only for relationships among $I(0)$ variables. However, point estimates remain consistent even for $I(1)$ cointegrated variables (Kejriwal and Perron, 2008).

For each possible breaking date, we carried out a sequence of F statistics based on the null hypothesis of no shifts against the alternative of a single-shift. The corresponding sup-F_T statistics provides a test for structural change against a single break alternative of

---

[9]   All estimates were obtained using the R package **strucchange**.

unknown timing. We also carried out tests for the number of breakpoints based on information criteria, notably BIC and simple Residual Sum of Squares (RSS)[10].

Once the break dates have been estimated, a pure structural change model for the price transmission equilibrium (or long run) relationships is given by equation (7):

$$y_t = x'_t \delta_i + u_t \tag{7}$$

where the subscript *i* refers to the *m*+1 regime delimited by the break dates. In our case *x* is a bi-dimensional vector with a constant (x_{1t}) and the logarithm of the upstream price series (x_{2t}).

Table 5 – Bai & Perron test – break dates, regimes and confidence intervals for *pasta*

Pasta - Bai & Perron test - (*Farm - Wholesale*)				
n° of breaks	break-dates	Partitions	Lower bound (2.5%)	Upper bound (97.5%)
1	2007(6)	2000(2) - 2007(6)	2007(5)	2007(8)
		2007(7) - 2010(4)		

Table 6 – Bai & Perron test – break dates, regimes and confidence intervals for *lamb*

Lamb - Bai & Perron test - (*Wholesale - Retail*)				
n° of breaks	break-dates	Partitions	Lower bound (2.5%)	Upper bound (97.5%)
1	2001(7)	2000(2) - 2001(7)	2000(8)	2001(12)
		2001(8) - 2010(4)		

Table 7 – Bai & Perron test – break dates, regimes and confidence intervals for *pork*

Pork - Bai & Perron test - (*Farm - Wholesale*)				
n° of breaks	break-dates	Partitions	Lower bound (2.5%)	Upper bound (97.5%)
1	2007(1)	2000(2) - 2007(1)	2006(12)	2007(2)
		2007(2) - 2010(4)		

---

[10] We set the trimming rate for both tests to 0.15, whereas the maximal number of breaks is built by default from the trimming parameter (generally five breaks maximum).

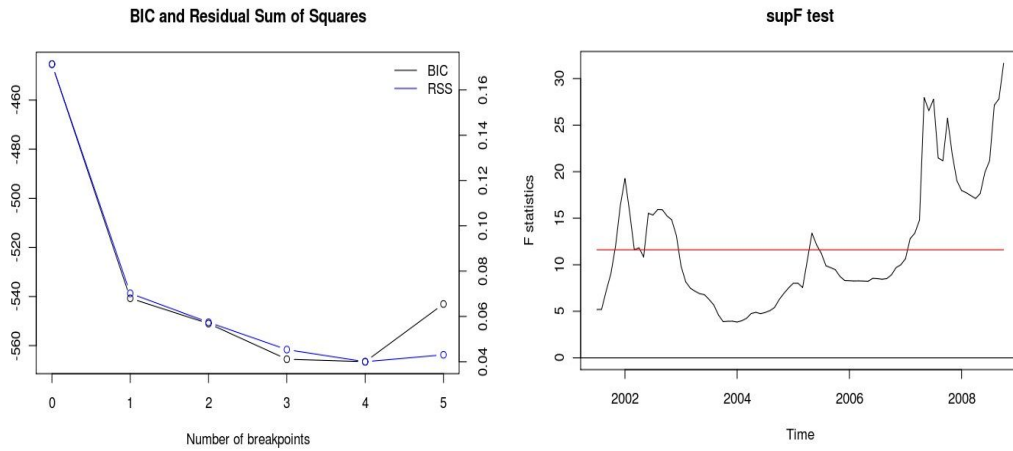**Figure 1 – BIC, RSS models for *m* break points and SupF test plots for *wheat-semolina***

Figure 2 – BIC, RSS models for *m* break points and SupF test plots for lamb wholesale retail seriews
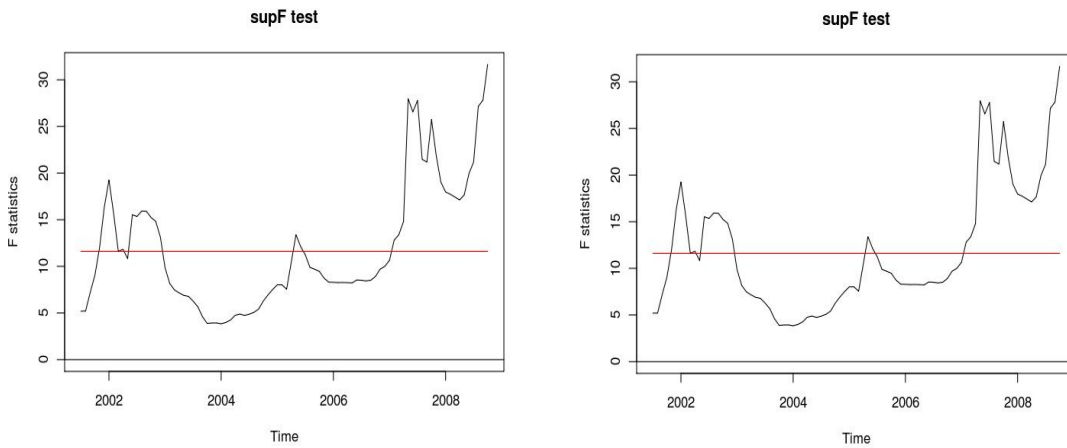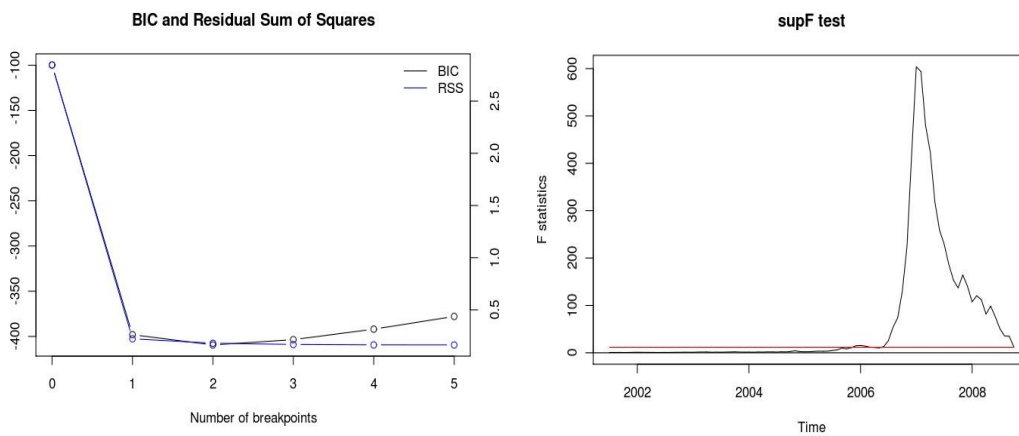

Figure 3 – BIC, RSS models for *m* break points and SupF test plots for hog farm wholesale series

Results are reported with reference to the 3 cointegrated couple of series. Although BIC or RSS criteria would have suggested for some series (notably lamb and pasta) a larger number of breaks we retained a single break given the relatively small number of observations. Tables 5, 6, and 7 show the estimates of optimal break dates and related regime partitions. Coefficient estimates for each regime are provided in table 9 where the i the coefficients $\delta_{1,i}, \delta_{2,i}$ refer respectively to the constant and the upstream price in the long run equation. Plots of the BIC and RSS values as well as SupF curves are provided in figures 1 to 3 while figures 4 to 6 show the original series with regime delimiters.

Table 9 - Bai & Perron test – Coefficients estimates

Pasta - Bai & Perron test - coefficients estimates			
Farm/Wholesale			
Coefficient	Estimate	std error	Partitions
$\delta_{1,1}$	-0.05	0.04	2000(2) - 2007(6)
$\delta_{1,2}$	0.25	0.02	2007(7) - 2010(4)
$\delta_{2,1}$	0.70	0.02	2000(2) - 2007(6)
$\delta_{2,2}$	0.83	0.01	2007(7) - 2010(4)
Lamb - Bai & Perron test - coefficients estimates			
Wholesale/Retail			
Coefficient	Estimate	std error	Partitions
$\delta_{1,1}$	1,03	0.20	2000(2) - 2001(7)
$\delta_{1,2}$	1,55	0,06	2001(8) - 2010(4)
$\delta_{2,1}$	0,59	0,12	2000(2) - 2001(7)
$\delta_{2,2}$	0.33	0.03	2001(8) - 2010(4)
Pork - Bai & Perron test - coefficients estimates			
Farm/Wholesale			
Coefficient	Estimate	std error	Partitions
$\delta_{1,1}$	1,03	0.01	2000(2) - 2007(1)
$\delta_{1,2}$	1,36	0.02	2007(2) - 2010(4)
$\delta_{2,1}$	0,81	0,04	2000(2) - 2007(1)
$\delta_{2,2}$	0.70	0,11	2007(2) - 2010(4)

Notably, we do not find a clear evidence of the correct number of breaks from the SupF test plot which shows a major break in 2007 and two minor structural changes in 2002 and, possibly in 2005 .

The single break estimate appears to be related to the beginning of the commodity bubble (june '07 break). Equilibrium coefficients estimates suggest that elasticity of transmission as well intercept estimates increase in the commodity bubble period. This results in a larger reactivity of semolina prices to changes in prices of durum wheat. The overall effect on margins is ambiguous as log of prices are negative in this case.

For the lamb wholesale-retail couple a single breakpoint is also suggested by the BIC criterion and by the sup-F test graphic that shows a single major spike at the beginning of

the series. Elasticity of transmission decrease after the break. However, the intercept increases. In the second regime margins tend to decrease as the growth of wholesale prices is only partially transmitted to the retail stage.

Finally, for the hog farm-wholesale transmission equation both the BIC and the graphical pattern of the sup-F test (the latter with a clear single spike) suggest a single breakpoint. The change occurs at the beginning of the commodity bubble (January 2007) that led to dramatic increase of prices in the food sector (fig.6).

The second regime is characterized by a larger intercept and a smaller elasticity of transmission. As the farm price shows no definite trend this results in wider margins. Along the period considered the Italian pork marketing chain has witnessed a move toward concentration of the slaughtering industry, this process may have impacted also the price transmission mechanism in the context of the 2007-09 price crisis as we observe a shrinking of the share of wholesale value accruing to farming.

Figure 4 - **Nominal price series in natural logarithms –** *wheat-semolina*
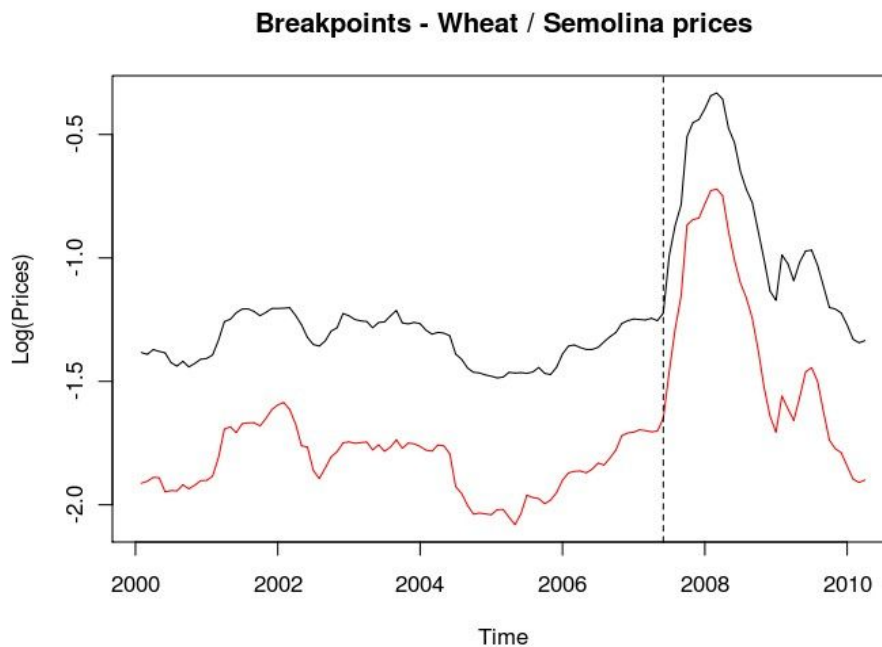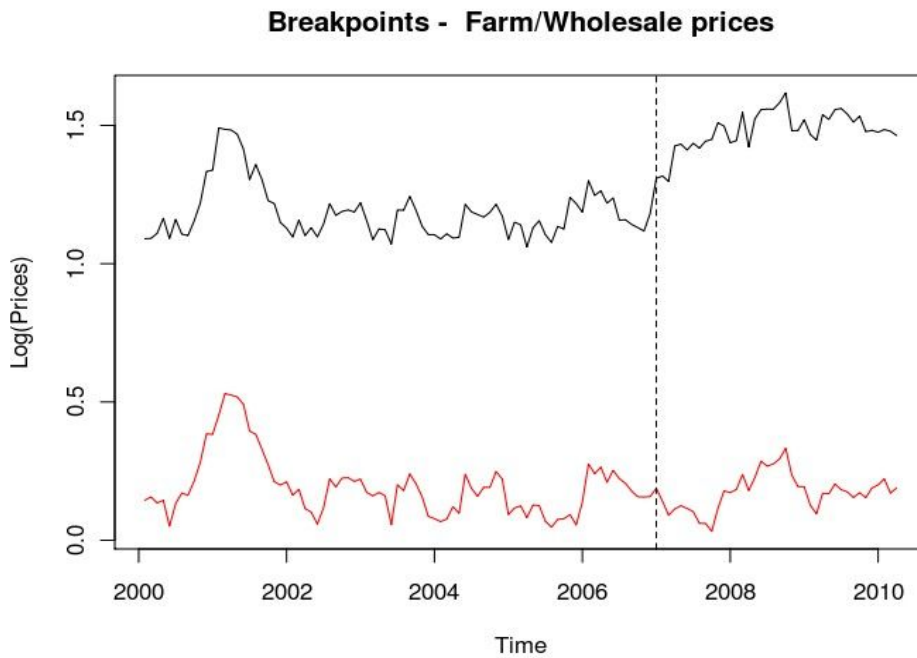

Note: farm price (in black) / wholesale price (in red)

Figure 5 - Nominal price series in natural logarithms – Wholesale / Retail - *lamb*
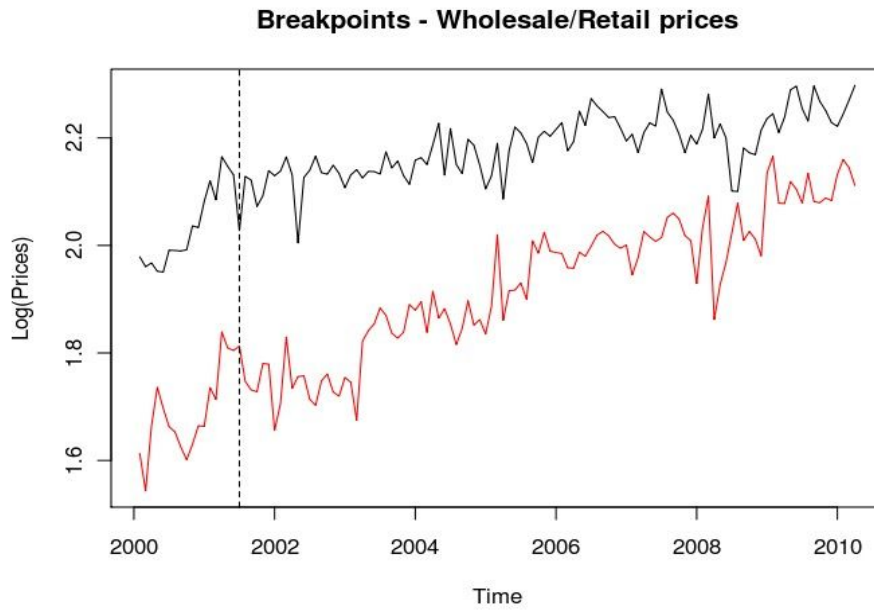
Note: Wholesale price (in black) / Retail price (in red)

Figure 6 - Nominal price series in natural logarithms – Farm / Wholesale - *hog*

Note: farm price (in black) / wholesale price (in red)

3. Conclusions

This article investigated long run transmission elasticities in three Italian food chains (pasta, lamb and pork) accounting for structural breaks. We analysed three price transmission relationships for each chain: farm-wholesale, wholesale-retail and farm retail. Within the classical cointegration framework we found evidence of equilibrium relationships only for the durum wheat-semolina and lamb wholesale-retail series. However, once structural breaks are accounted for a long run relationship emerges also for the hog farm-wholesale series. For each cointegrated couple of prices we examined the presence of structural breaks.

Main changes in the long run relationship were found for the pasta and the pork chain. Both equations show a break at the beginning of the price bubble that altered the transmission mechanism across different stages of the food chain notably in the direction of a larger elasticity of transmission in former case and a smaller elasticity in the latter .

Although, long run transmission elasticities are valuable findings *per se*, the analysis could be easily extended to the study of possible asymmetries in the transmission mechanism and to modelling of short run relationship via error correction models. Further research is also needed to study food prices using a wider range of unit root and cointegration tests within the multiple breaks framework that has been recently developed and applied in the fields of financial and macroeconomic series.

4. References

Abdulnasser Hatemi-J (2008), Tests for cointegration with two unknown regime shifts with an application to financial market integration, *Empir Econ* 35:497–505.

Adachi K., Liu D.J. (2009), Estimating Long-Run price relationship with structural change of unknown timing: an application to the Japanese pork market, *American Journal of Agricultural Economics*, 91: 1440-1447

AGCM (2007). *Indagine conoscitiva sulla distribuzione agroalimentare*, Indagini Conoscitive, 28, Autorità garante per la Concorrenza e il Mercato, Roma.

Agra CEAS (2007).*The Gap between Producer Prices and the Price paid by the Consumer*, European Parliament, Brussels.

Akaike H (1973) Information theory and an extension of the maximum likelihood principle. In: Petrov BN., Csake F. (eds) *Second international symposium on information theory.* Akademiai Kiado, Budapest, Hungary: 267–281

Andrews D.W.K. (1993). Tests for parameter instability and structural change with unknown change point, *Econometrica*, 61: 821–856.

Bai J., Perron P. (1998). Estimating and Testing Linear Models with Multiple Structural Changes, *Econometrica*, 66: 47-78.

Bai J., Perron P. (2003). Computation and Analysis of Multiple Structural Change Models, *Journal of Applied Econometrics*, 18: 1-22.

Boetel, Brenda L. & Liu, Donald J. (2008). *Incorporating Structural Changes in Agricultural and Food Price Analysis: An Application to the U.S. Beef and Pork Sectors*, Working Papers 44076, University of Minnesota, The Food Industry Center.

Carrion-i-Sylvestre, J.L., Sansó-i-Rosselló, A.S., (2006). Testing the null hypothesis of cointegration with structural breaks, *Oxford Bulletin of Economics and Statistics*, 68: 623-646

Carrion-i-Sylvestre, J.L., Sansó-i-Rosselló, A.S., (2007). The KPSS test with two structural breaks, *Spanish Economic Review*, 9(2):105-127

Commission of the European Communities (2009) - A *better functioning food supply chain in Europe* – Brussels 28.10.09 - COM 2009(591). Available at http://ec.europa.eu/economy_finance/publications/publication16061_en.pdf

Dickey, D.A., Fuller, W.A., (1979). Distribution of the estimators for autoregressive time series with a unit Root, *Journal of the American Statistical Association*, 74: 427-431.

Engle R. F., e Granger C.W.J. (1987). Co-integration and error correction: representation, estimation and testing, *Econometrica*, 55, 251–276.

Ferrucci G, Jiménez-Rodríguez R., Onorante L. (2010) - *Food price pass-through in the euro area, the role of asymmetries and non-linearities*. European Central Bank Working Paper Series n° 1168 - April 2010. Available at http://www.ecb.europa.eu

Frey G., Manera M. (2007). Econometric Models Of Asymmetric Price Transmission, *Journal of Economic Surveys* 21(2): 349-415.

Gardner, L.A., 1969. On detecting changes in the mean of normal variates, *The Annals of Mathematical Statistics* 40: 116-126.

Giangiulio D., Mazzantini G. (2010). *Il trasferimento dei prezzi lungo le filiere agroalimentari: una possibile metodologia di analisi per i profili di interesse Antitrust*, XLVII Convegno SIDEA, Campobasso

Gregory, A.W., and B.E. Hansen. (1996). Tests for cointegration in models with regime and trend Shifts, *Oxford Bulletin of Economics and Statistics* 58:555–60.

Harris, R. (1999), *Using Cointegration Analysis in Econometric Modelling*, London, Prentice Hall.

IPI (2008). *Le dinamiche dei prezzi nella filiera del frumento*, Istituto per la Promozione Industriale.

Kejriwal, M., and Perron, P. (2008). The limit distribution of the estimates in cointegrated regression models with multiple structural changes, *Journal of Econometrics* 146: 59–73.

Kwiatkowski, D., Phillips, P. C. B., Schmidt, P., and Shin, Y. (1992). Testing the null hypothesis of stationarity against the alternative of a unit root, *Journal of Econometrics* 54, 159-178.

Lee, J., Strazicich, M.C. (2003). Minimum LM unit root test with two structural breaks, *Review of Economics and Statistics* 63:1082-1089.

London Economics (2003), *Investigation of the determinants of farm-retail price spreads*. Report prepared for the Department for Environment, Food and Rural Affairs. Available at http://statistics.defra.gov.uk/esg/reports/pricespreads/default.asp.

MacKinnon J. (1991), Critical Values for Cointegration Tests. In Engle, R. and Granger, C. (eds), *Long-Run Economic Relationships*, Oxford University Press, Oxford.

Meyer, J. and S. von Cramon-Taubdel. (2004). Asymmetric Price Transmission: A Survey. *Journal of Agricultural Economics* 55:581-611.

National Bank of Belgium (2008). Processed food: inflation and price levels, *NBB Economic Review, Special Edition,* April: 53-72.

Peltzman, S. (2000). Prices Rise Faster than they fall, *Journal of Political Economy* 108(3): 466-502.

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis, *Econometrica* 57: 1361-1401

Perron, P. (1990). Testing for a Unit Root in a Time Series with a Changing Mean, *Journal of Business and Economic Statistics* 8: 153-162.

Perron, P. (2005). *Dealing with Structural Breaks*, Mimeo forthcoming in the Vol. 1 Handbook of Econometrics: Econometric Theory.

Phillips, P.C.B., Perron, P., (1988). Testing for a unit root in time series regression, *Biometrika* 75: 335-346.

Schmidt, P. and Phillips, P.C.B. (1992). LM Test for a unit root in the presence of deterministic Trends, *Oxford Bulletin of Economics and Statistics* 54(3): 257–287.

Schwarz G (1978). Estimating the dimension of a model, *Annals of Statistics* 6(2): 461–464 .

Stock, J.H., Watson. M.W., (1993). A simple estimator of cointegrating vectors in higher order integrated systems, *Econometrica*, 64: 783-820.

Wang, D., and W.G. Tomek. (2007). Commodity Prices and Unit Root Tests, *American Journal of Agricultural Economics*, 89: 873-889.

Zeileis A., Kleiber C., Krämer W., Hornik K. (2003). Testing and Dating of Structural Changes in Practice, *Computational Statistics and Data Analysis* 44:109-123.

Zeileis A, Leisch F, Kleiber C, Hornik K (2005). Monitoring Structural Change in Dynamic Econometric Models, *Journal of Applied Econometrics*, 20(1): 99-121.

Zivot, E., Andrews, D.W.K., 1992. Further evidence on the great crash, the oil price shock and the unit root hypothesis, *Journal of Business and Economic Statistics*, 10: 251-270.