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PRICE TRANSMISSION MECHANISMS: A POLICY INVESTIGATION OF INTERNATIONAL WHEAT MARKETS

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PRICE TRANSMISSION MECHANISMS: A POLICY INVESTIGATION OF INTERNATIONAL WHEAT MARKETS

di Giulia Listorti[1]

Abstract

This work focuses on soft wheat price transmission mechanisms between the United States and the European Union. In particular, by performing a cointegration analysis, it aims at analyzing if and to which extent the prices in the two countries were related in the years from 1978 to 2003, provided that the market of this commodity was deeply influenced by the Common Agricultural Policy. The issue of how domestic and international policy regime changes affected price transmission elasticities is also explored.

Key words: international price transmission; cointegration; Law of One Price; Common Agricultural Policy.

EconLit Classification: Q110, Q170, Q180.

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1. Introduction

The study of price transmission mechanisms amongst different markets and regions is an important tool to understand the relations existing between them. Domestic and border policies are likely to play an important role in this respect. While price transmission mechanisms for agricultural markets have, so far, received considerable attention in literature, the use of price data only has too often implied the use of increasingly sophisticated techniques and the lack of attention for the role played by policy factors. In econometric models, on the other side, policy variables are explicitly added as regressors in the relevant equations, but the price transmission mechanisms often rely on simplistic hypothesis, like the exogeneity of the world price for the European Union (EU) in the AGMEMOD model[2].

This works aims at suggesting some new ways of analysing price transmission mechanisms in international markets, and at providing useful and simple tools that account for the major political changes occurred in the past years.

This paper focuses on soft wheat. Soft wheat is a heavily traded commodity. The main exporters are Argentina, Australia, Canada, United States (US) and the EU; in most years, they account for about 90% of world wheat exports (FAO 2007). Algeria, Brasil, Egypt, Japan and the EU have been some of the biggest importers in the past years, even if the import demand is much less concentrated than the export supply (USDA 2005).

US and EU policy actions are then very likely to have an influence on each other's agricultural policies as well as on world market prices (Barassi and Ghoshray 2007). As Ghoshray et al. (2000) point out, policy regimes play a significant role in soft wheat production and export shares; the CAP is a case in point, since during the 1980s the EU emerged as the second larger exporter of wheat, having previously been a net importer.

1.1 The evolution of the political context

In the past 30 years, the EU Common Agricultural Policy has considerably evolved. Four major periods can be identified (see also Thompson 1999):

<u>1971:01-1988:6:</u> These years are characterized by the full functioning of the Common Market Organizations (CMOs): a number of institutional prices regulate internal markets. Intervention mechanisms ensure that domestic prices never fall below the intervention price.

[2]AGMEMOD is an econometric, dynamic, multi-commodity partial equilibrium model of the European Agriculture (Chantreuil et al., 2005; AGMEMOD Partnership 2007a; 2007b; www.agmemod.org).

Variable levies and export subsidies insulate the domestic market from the world market (see Figure 1). This system led, in the 1980s, to surpluses growth and budgetary costs escalation. The EU emerged as a net exporter having previously been a net importer.

1988:07-1993:06: First CMO Reforms concerning arable crops are put in place. New measures aimed at reducing the production surpluses and budgetary costs are introduced, such as co-responsibility levies (deductions from farmers to pay for the cost of surplus production), stabilizers (if production exceeded a maximum guaranteed quantity, co-responsibility would increase and intervention price would be reduced the following year), voluntary set-aside for cereals.

1993:07-2000:06: The small effects of the already implemented reforms lead to the introduction of new policy measures with the objective of ensuring a progressive return to market mechanisms. The Mac Sharry Reform implements substantial cuts in intervention prices to re-align them with the world prices; compensations to farmers through direct subsidies per hectare are put in place. Set aside requirements are established for producers of more than 92 tonnes[3]. The old variable levy and export subsidy structure, on the other side, keeps on insulating the EU market from the world ones.

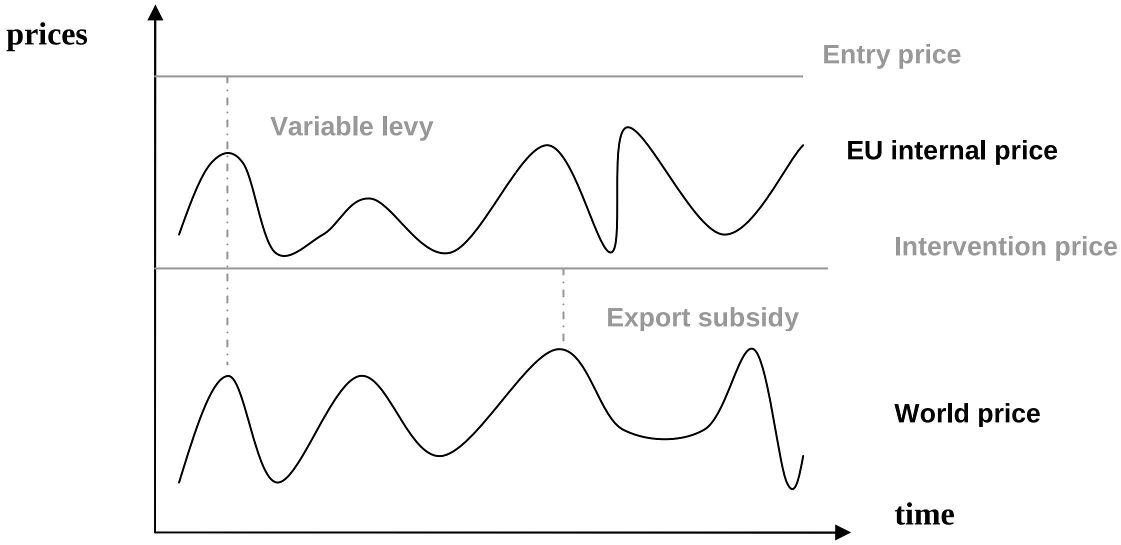
It has been argued that the changes of the 1992 CAP Reform essentially apply to the grain sector and, to a lesser extent, to the beef sector (Mahe 1996). The reason lies in the pressure put by the US and other competitors in the Uruguay Round Agreement on Agriculture (URAA) on the EU in a sector characterized by major policy interdependencies.

2000:06- 2007:07: With the Agenda 2000 Reform, both a 15% reduction in two years of the intervention price for cereals and the introduction of decoupled payments are decided. The set-aside regime remains in force, too. The Fischler Reform, in 2003, strengthens the decoupling of payments. Intervention prices are not further reduced but the monthly seasonal adjustments applied to them are halved[4].

[3] Those producers owning an agricultural area capable of producing more than 92t of product, according to local average yields.

[4] The two Reforms are put together despite the much bigger impact of the Fischler Reform due to the lack of data in the most recent years.

Figure 1: Cereals' Common Market Organization's functioning up to 1992.



As far as international trade agricultural politics is concerned, the most relevant event in the period examined has been the institution of the World Trade Organization in 1995, and the following implementation of the URAA, the first multilateral agreement explicitly referred to agriculture. The URAA regulates both domestic support (but the EU always kept a safe margin with respect to the maximum allowed) and export subsidies. For wheat, these limits were never binding for the EU (Anania 2007). As far as market access is concerned, the URAA abolished the possibility of keeping threshold prices and variable levies; all border measures were converted into import duties to be lowered in the following six years (-36%, with a minimum reduction of 15% for each "8 digit" product; in the case of agriculture, customs duties are for the most part specific duties (expressed in EUR/t)). Due to the high levels of protections in the reference period (the years 1986-1988), with regard to the main cereals, the tariff equivalents notified at the WTO were so high that a system for capping duties was deemed necessary for the EU, in order to put a ceiling on the entry price. This was then capped at 155% of the intervention price, if the sum of the duties would make it go above this threshold. The entry price turned out to be almost always capped, thus eliminating any real difference between the old variable levy system and the new one (Table 1).

Table 1: Elements for calculating the duty on cereals (Gallezot 2007).

Import duty =155% intervention price – CIF reference price

CIF reference price = average US quotation for the reference variety during the receding two weeks
+ Gulf premium or Great Lakes premium*
+ Gulf Rotterdam or Great Lakes-Rotterdam freight premium*

Intervention price = price in force including monthly increases at the time the duty is applied.

*The first freight premium corresponds to the cost of freight across the United States to the Gulf (« Gulf premium ») if the commodity is quoted in Chicago or Kansas City, or to the Great Lakes (« Great Lakes » premium) if the commodity is quoted in Minneapolis. The second freight premium corresponds to freight from the United States (Gulf or Great Lakes) to the port of Rotterdam in the Netherlands.

Summing up, the most relevant EU trade policies for cereals are export refunds[5] and variable levies, with the objective of ensuring the maintenance of high prices on domestic markets. In the 1970s and the 1980s the EU heavily subsidized wheat exports by adding export subsidies to cover the difference between its internal support price and the world price (Barassi and Ghoshray 2007). In 1985, the US retaliated with the Export Enhancement Program (EEP), a targeted export subsidy program for wheat[6]. The EEP didn't alter substantially the EU-US price relationship, since, after their introduction, the EU appeared to set its export subsidies in relation to US prices (Mohanty et al. 1999; Barassi and Ghoshray 2007). In practice, the EEP has not been used after August 1995. Tight world supplies and high world prices implied that the EU didn't use export refunds than to a very limited extent, and the US didn't re-activate the EEP. In the very last years, commodity world markets have been characterized by a boom in prices caused by both transitory (such as adverse meteorological conditions) and structural factors (namely, the increase in food demand from emerging economies and the biofuels demand).

In this paper, an examination of the mechanisms determining price transmission for soft wheat in international markets is provided. In paragraph 2, the analytical framework for the study of price transmission is briefly revised, suggesting how policy variables may be introduced while using a cointegration analysis; in paragraph 3, the data are presented.

[5] For short periods of time, usually no more than a few months, when world prices were substantially higher than internal prices, export taxes have also been used.

[6] Brooks et al. (1990) show that the impact of the EEP on other exporters' wheat trade and importers' demand has been small relative to the magnitude of total EEP sales: over the period 1986-1989, the displacement of sales ranged from 87% to 92%, while additional exports were only 8% to 13% of the total.

Paragraph 4 and 5 report the results of the analysis. In paragraph 6, some final considerations are made.

2. International price linkages and the Law of One Price: the analytical framework

2.1 The general framework

Studying price transmission mechanisms implies referring to some basic economic concepts for which, unfortunately, no common definition exists in literature (Fackler and Goodwin 2001).

In a nutshell, the *spatial arbitrage condition* implies that the difference between prices in different locations will never exceed transport costs[7], or otherwise the profiting opportunities would be immediately exploited by arbitrageurs[8]. Actual prices may diverge from this relation in the short run, but the actions of the arbitrageurs will make it valid in the long run, moving the price spread toward the transport cost. This is at the basis of the Law of One Price (LOP): markets linked by trade and perfect arbitrage (and competition) will have a unique price, when expressed in the same currency, net of transportation costs. This has also bell defined the "strong version" of the LOP, since price spreads are assumed to be exactly equal to (and not minor than) transport costs.

This concept has a very long tradition in economics which dates back to Marshall but, nevertheless, most of the empirical tests are against it (Fackler and Goodwin 2001).

The concept of *spatial market integration* generally indicates the degree of co-movement shown by prices in different location, which could be explained also by factors that have nothing to do with commercial integration (such as seasonality, influence of a common factor etc.). The definition proposed by Fackler and Goodwin (2001) is a measure of the degree to which demand and supply shocks arising in one region are transmitted to another one. This is a more restrictive concept than the LOP: even if the LOP is satisfied, if transport costs are large and volatile, prices don't move together. Furthermore, the LOP can hold even if the price transmission ratio between two regions is less than one[9].

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[7] In this paper the term "transport costs" is meant to include all relevant costs of arranging transports between spatially separate locations.

[8] This condition is referred to by someone as "spatial market efficiency", since markets should produce prices that accurately reflect all the available information about demand and supply conditions as well as transport costs.

[9] Shocks have generally a bigger effect in the originating region than in the other one (Fackler and Goodwin 2001).

Transport costs play a crucial role while investigating markets efficiency; this is particularly true in agriculture, since they are relevant if compared to the traded commodities' value. If inefficiency in price transmission mechanisms is found, it might depend on the fact that markets are inefficiently integrated (the information available is not properly used), but also on the assumptions made about transport costs. It could be argued that any value of price transmission elasticities is consistent with integration depending on the value of transaction costs.

A number of different econometric techniques have been used in the past decades to investigate price transmission mechanisms: simple regression and correlation analysis; dynamic regression models based on a point location model; switching regime models; rational expectations models[10]. A considerable amount of literature has developed increasingly sophisticated econometric devices to deal with the fact that, usually, prices are the only data available to examine spatial relationships. Unfortunately, as Barrett (1996) notes, "agricultural economists' toolkits have changed nearly as rapidly and dramatically as have developing economy markets, but these methodological refinements have not been accompanied by conceptual advance".

Furthermore, despite the fact that so many attempts have been made to investigate the LOP, and that the LOP is the building block of international trade theory, what actually emerges is that this Law has been violated by empirical tests probably more than any other economic laws (Miljkovic 1999).

A number of factors prevent prices from convergence. Conforti (2004) identifies six groups of factors that influence the LOP:

- transport and transaction costs: unless certain assumptions are made, their treatment is not easy;
- market power;
- increased returns to scale in production;
- product homogeneity and differentiation;
- the extent to which changes in the exchange rates are "passed through" on output prices;
- border and domestic policies.

Trade policies play an important role. Variable levies, non tariff barriers, tariff rate quotas, prohibitive tariffs, technical barriers play a strong role, whereas ad valorem and fixed tariffs

[10] A detailed description of the advantages and weaknesses associated with each of these methods is provided by Fackler and Goodwin (2001) and goes beyond the objectives of this paper.

should behave as proportional and fixed transaction costs (Conforti 2004). Even the very existence of specific trade agreements that create different trading blocks with different degrees of integration can prevent domestic prices from convergence in different countries, and make the simplicity of the LOP rather questionable.

2.2 The cointegration analysis

As soon as they were developed, cointegration analysis techniques were assumed to be a sort of "natural tool" to investigate price transmission mechanisms. Cointegration models presuppose that observable variables exhibiting non-stationary behaviour will nonetheless be linked by a long-run relationship. This long run relation is, in this case, due to the LOP.

By taking into account cointegration, we will have the following Vectorial Error Correction Model (VECM):

$$\mathbf{\Delta p_t} = \boldsymbol{\alpha\beta}'\mathbf{p_{t-1}} + \sum_{i=1}^{m} \mathbf{\Gamma_i \Delta p_{t-1}} + \boldsymbol{\varepsilon}_\mathbf{t},$$

where $\mathbf{p}_t$ is a vector containing the prices, $\boldsymbol{\beta}$ is the cointegration matrix which contains the long-run coefficients (the degree of price transmission, Conforti 2004), $\boldsymbol{\alpha}$ is the loading matrix which contains the adjustments parameters (a measure of the speed of price transmission, Conforti 2004), $\mathbf{\Gamma}$ is a matrix containing coefficients that account for short-run relations, and $\boldsymbol{\varepsilon}_t$ are white noise errors. When prices are expressed in logs, the coefficients included in $\boldsymbol{\beta}$ can be read as price transmission elasticities[11]. If y_t contains only two log-prices, say $p_{1,t}$, the domestic price, and $p_{2,t}$, the world price, we will have the following long-run relation between them: $\boldsymbol{\beta}'\mathbf{p_{t-1}} = [\beta_0 \quad \beta_1 \quad \beta_2] \begin{bmatrix} 1 \\ p_{1,t-1} \\ p_{2,t-1} \end{bmatrix} = \beta_0 + \beta_1 p_{1,t-1} + \beta_2 p_{2,t-1} = u_t$, where $u_t \approx I(0)$. Normalizing with respect to β_1 and rearranging the terms, we have $p_{1,t-1} = -\frac{\beta_0}{\beta_1} - \frac{\beta_2}{\beta_1} p_{2,t-1} + u_t$, were $\frac{\beta_2}{\beta_1}$ is the long-run price transmission elasticity: it indicates the percentage change in the domestic price in response to a one-percent change in the world price. In international markets, price transmission elasticities show the extent to which changes in world prices are transmitted back to country prices; Thompson and Bohl (1999) argue that they can indeed be interpreted

[11] When logs are used, the LOP is implicitly assumed to hold in a multiplicative form in levels, i.e. $P_{1,t} = T^{\beta_0} P_{2,t}^{\beta_1}$, from which $\log P_{1,t} = \beta_0 \log T + \beta_1 \log P_{2,t}$ and, holding the first term constant, $p_{1,t} = t + \beta_1 p_{2,t}$, where $p = \log P$. The underlying assumption is then that transport costs are stationary around a proportional constant of prices.

as a measure of the degree of market insulation, or the extent to which border policies are transmitted to domestic market. Price transmission is affected by trade liberalization and by trade policies: generally speaking, trade liberalization will contribute to greater price transmission elasticities. Nevertheless, we should keep in mind that the transmission parameters summarizes the overall effect of al the factors that affect price signals, including, for example, the existence of market power, the degree of product homogeneity, etc. (Conforti 2004).

The existence of a stable relation between two prices, i.e. bivariate cointegration, has been assumed as a necessary condition for integrated markets (Ardeni 1989). Others claim that the strongest condition that prices differences are stationary is needed, and thus impose instead of estimating the cointegration relationship (Baffes 1991). In a system with *n* prices, the number of the cointegrating vectors has been taken as an index of the degree of integration of the markets.

Nevertheless, the use of cointegration techniques in investigating price transmission mechanisms presents a number of shortcomings. First of all, even in the simplest model (where the arbitrage condition ensures that the price spread between two markets is stationary, i. e. P_1-P_2 = T, where p are the prices in markets 1 and 2 and T the transport costs) it is clear that cointegration is not necessary at all for markets to be efficiently integrated, since T could be a non-stationary process. If transport costs are non stationary, cointegration is then an unnecessary condition for market integration (Barrett 1996)[12].

Secondly, if the transport costs band is large and P_1-$P_2 \cong I(0)$ but this difference remains within the band, prices could result cointegrated even if markets are actually not integrated, since the difference between prices would not justify any shipment of the commodities. Barret (1996) points out that market segmentation can arise because of inter-market margins larger than transfer costs ("absence of rational arbitrage") or of margins less than transfer costs ("rational absence of arbitrage"): in both cases there is no efficient exchange between markets, but cointegration tests identify the former one only.

Thirdly, Barrett (1996) notices that cointegration could be consistent with a negative relationship between prices, whereas market integration suggests a positive correlation

[12] By using a simple point location model, Mc New and Fackler (1997) show that neither efficiency nor market integration necessarily lead to linearly related prices. They demonstrate that in the case where prices in two regions are not cointegrated (and the underlying forces affecting supply and demand are not as well) arbitrage alone doesn't guarantee that prices exhibit cointegration, even if the LOP holds, especially as transport rates increase in size and volatility.

between them, and that many reported coefficients have magnitudes implausibly far from unity.

Fourthly, trade flow discontinuity could also represent a problem: at the break points, the relation between prices is zero, whereas in others it could be roughly one. In the case where demand and supply forces are themselves cointegrated, prices in two regions could be cointegrated even in the absence of trade flows.

At this point, it is evident that a number of problems basically arise from the use of price data only.

In this regard, Barrett and Li (2002) notice that in traditional analysis it is not possible to distinguish between market integration (i.e., the tradability of products between spatially distinct markets), and competitive market equilibrium (in which extraordinary profits are exhausted by competitive pressures). They say that “the fundamental weakness of much of the existing literature is that it attempts inference off just a subset of relevant variables, typically just prices, and then focuses on just the special case of perfect integration, when two markets are both integrated and in competitive equilibrium. Yet actual market relationships are messy”. The model they develop, based on both prices and transport costs and trade flows data, is an answer to this drawback as they identify four basic regimes according to the presence or the absence of integration and equilibrium.

Despite all this, to date cointegration has been one of the most widely used technique for studying price transmission mechanisms: in fact, the tests for cointegration became very popular methods in LOP studies since Ardeni (1989) argued that conventional LOP tests had disregarded the time series properties of the data. The cointegration approach has been used quite a number of papers, as reported in Annex 1 (Goodwin and Fackler 2001)[13].

Shortly enlisting those who use with soft wheat monthly prices to investigate international markets, Thompson and Bohl (1999) by using a threshold cointegrating technique, find that German soft wheat producer prices and US Gulf Dark Northern Spring (DNS) wheat prices are indeed cointegrated.

Ghoshray et al. (2000) analyze 13 series of FOB wheat prices, different for origin and quality. They find that the wheats for human consumption embody a common price trend, while feed wheats share another one. The EU always acts as a price follower but (differently from Barassi and Ghoshray 2007), not with US soft wheat.

[13] Miljkovic (1999) presents a critical review of them.

Barassi and Ghoshray (2007) test cointegration with structural change (time of the break unknown) to analyze the nature of the long-run relationship between US and EU wheat export prices over the years 1981-2000. They carry on their analysis on EU standard wheat, US Soft Red Wheat (SRW) and US Hard Red Winter Wheat (HRW). They find that the breakpoint occurred after the Mac Sharry CAP Reform was implemented. Before, there was no long-run relationship between EU and US wheat prices. After, they find two long-run relationships: one between US HRW and US SRW wheat prices, and the other between US SRW and EU wheat prices. In both cases, the price transmission elasticity is equal to unity. The fact that EU prices are linked to SRW prices may be due to their similar end uses, and to competition of both countries in the North African markets, which, before the EU emerged as a net exporter of wheat, used to be regular importers of US SRW. While there is no price leader between the two US wheats, they find evidence that the US SRW wheat has been acting as a price leader for the EU, as it would be expected given that the EU has retained the use of export subsidies.

Finally, Verga and Zuppiroli (2003) use weekly data for the years 1990-2002 for four European markets, US prices and EU institutional prices and find that the European soft wheat markets are strongly cointegrated amongst themselves but not with the US one.

2.3 The introduction of policy variables

In this paper, an attempt of taking explicitly into account policy variables in made. The presence of a set of commercial policies in the EU (namely, import variable levies, and export subsidies) aimed at insulating the internal market (to make the intervention mechanism effective) has, in practice, always been in place. Considering this, we want to check whether the presence of a co-movement between the EU internal price and the world price can nonetheless be tested.

On the one side, while testing the presence of a co-movement between the EU prices on world markets (on top of which an export subsidy has been paid) and the other countries' ones, one should expect to find cointegration, right because export subsidies are established in order to cover the difference between the EU intervention price and the world ones. On the other side, testing the presence of cointegration between EU domestic prices and the world ones should bring no sensible results, since the two are separated by commercial policy measures.

A possible way of overcoming this problem is focusing on the major objective the EU market policy has had in the past 20 years, which we argue is keeping a minimum price (=the intervention one) on domestic markets. We might think that the intervention price acts as a threshold below which the world price has no effect on the EU domestic ones. In this way, by creating this "lower threshold", we are implicitly identifying the presence of different regimes. If the world price is below the intervention one, then the intervention mechanism is expected to be "active", and the EU internal prices to follow the intervention ones. If the world price is above the intervention price, we might think that the EU domestic price might actually follow the behaviour of the world one, while the presence of the export taxes should eventually prevent it from rising too much.

While one might think of the use of threshold models to accurately account for discontinuous adjustments to the long run equilibrium (like the Enders and Siklos methodology used in Thompson and Bohl, 1999, and the models presented in Balke and Fomby, 1997), the model presented here could nonetheless be interpreted as a particular threshold cointegration model, in which the switch in the policy regime is implicitly considered *via* the simple introduction of a composite variable (constituted by the maximum between the intervention price and the US price), as explained later. The intervention price acts as a threshold for the US price.

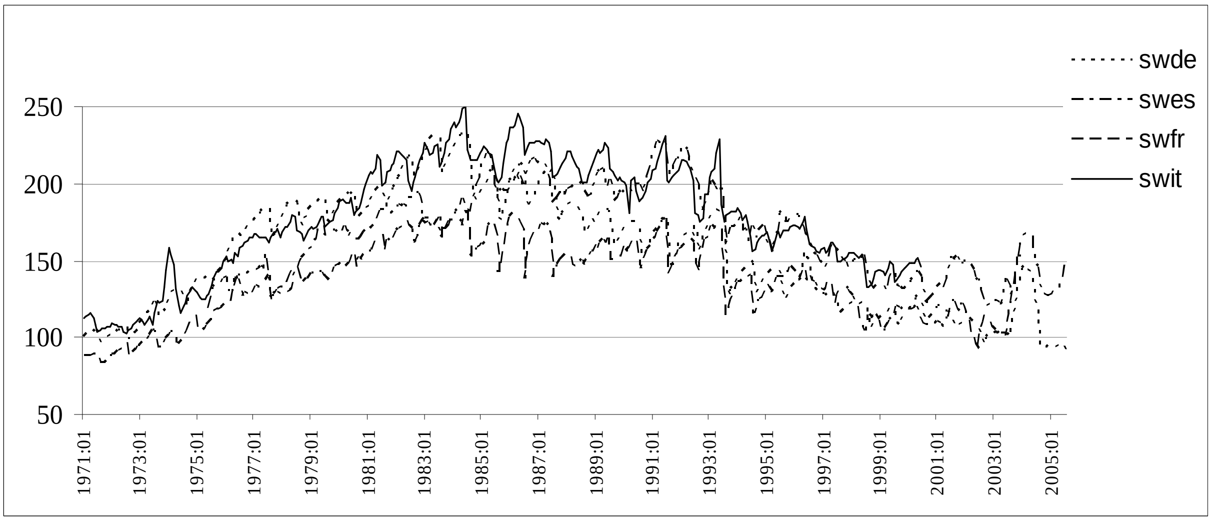
Not only changes in policy regime are then implicitly taken into consideration *via* the composite variable but, as a further preliminary investigation, the model presented is also tested in two sub-samples identified according to the presence of a different political framework (see Thompson and Bohl, 1999; Barassi and Ghoshray, 2007). While this allows considering structural breaks as well, the specific use of cointegration models accounting for structural breaks is not considered in this paper and needs to be addressed by future research.

Though the analysis is still performed within the framework of the LOP validity, we could say that, in this context, finding cointegration in itself would be a relevant result, whereas commenting the magnitude of the coefficients could bring some interpretative results. In fact, amongst some obvious simplifications of the model there are the assumptions that both the intervention price and the US price are measured at the same point in space, and that perfect competition exists on international markets.

3. The data

The dataset used in this paper consists of wheat monthly prices for the US and France for the period 1978:12 to 2003:12 (i.e., from December 1978 to December 2003; 301 observations). The French price is assumed to be the representative EU price, following the template of the AGMEMOD model. From Figure 2, indeed, we can see how the European prices from the four major agricultural producer countries move together over the past 30 years. As far as Italy and France (amongst the most important EU countries in this respect) are concerned, Verga and Zuppiroli (2003) find out that the strong relationships existing between them are not affected by the Mac Sharry Reform, which would imply that, besides sharing the same political context, the two countries are tied by strong commercial linkages (see also Zanias 2003).

Figure 2: Soft wheat prices in Germany, Spain, France and Italy (EUROSTAT, EUR/t).



In this work, the US price is considered to be the world price, as it happens in the AGMEMOD model. The US are a major player on international wheat markets: since 1975-76, the US share of global exports fluctuated between 25 and 45 percent of world markets (USDA 2005). Nevertheless, to whom the "price leadership" on world wheat markets belongs is an open question, to which empirical works have given different answers. Amongst the five world major players (US, Canada, Argentina, Australia and the EU), Mohanty et al. (1999)

show that there is no distinct leader[14]. However, taking the US price as representative of the world markets seems a reasonable assumption.

Data concerning US Gulf FOB HRW[15] prices and freight rates were obtained from the International Grains Council; the freight rates used in this study are those from US Gulf to ARAH (ARAH stands for Amsterdam/Rotterdam/Antwerp/Hamburg destinations), which were added to the US price in order to obtain CIF prices (*hrw*).

Soft wheat French prices (*swfr*) are available on-line from the EUROSTAT database. EUROSTAT monthly bilateral exchange rates have been used to convert in EUR prices and freight rates expressed in US $[16]. Using prices already converted in the same currency and not introducing the exchange rate itself as a regressor is a widely used procedure; it follows that adjustments to the exchange rate are considered instantaneous[17]. Prices were not deflated and are thus nominal. The EU intervention price time series (*pint*) has been reconstructed by adding to the prices established annually by the European Commission monthly seasonal adjustments, the latter obtained from European Commission regulations[18;19].

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[14] They find out that, while the US price is affected by the Canadian and the Australian only, the EU responds to US and Canadian changes but doesn't have any influence on their prices. The EU responds to the US, but the reverse is not true. Previous studies (all reported in Mohanty et al. (1999) some of which could be misspecified for not taking into account cointegration) get different results: some show that there is no significant leadership role between the US and Canada, while others find a strong leadership role for Canada for durum and Hard Red Spring markets. Goodwin and Schroeder (1991) find that both the US and Canada have a significant effect on the price of competing exporters but also that, while the US has an effect on Canada, the reverse is not true (the wheats considered are US Gulf Hard Wheat Ordinary Winter and Canadian No. 1 Western Red Spring Wheat); they consider the US as a price leader in the international wheat market. Dawson and Sanjuan, 2007, analyze the relations between the prices of Canadian Western Red Spring, US Dark Northern Spring, and find that Canada is the price leader. In Ghoshray and Lyoid (2003) Canada is found to be the price leader in the North American market for hard wheat exports.

[15] That wheat is a non homogeneous product is a well known issue (for a table of the various classes of wheat, see Ghoshray, 2000). Quality differences, and namely protein content, could influence international price linkages by making varieties of wheat imperfect substitutes of one another (Mohanty et al., 1999). Larue, 1991, shows that wheat is differentiated by end use and by country of origin, and that wheat protein content has a significant influence on prices. Ghoshray and Lyoid, 2003, investigate the relations between wheat prices of different exporters according to wheat type and port location. Wheat type depends on hardness, a milling characteristic (usually determined by protein content: we have hard and soft wheats), and on dough strength, which is a baking quality (we distinguish between strong and weak wheats). While the EU soft wheat is a medium protein wheat, weak, the HRW has a higher protein content (normally 12.5 %) and is classified as strong. The HRW has commonly been used to represent the world price, while on the other side it has been argued that the SRW would be indeed a closer substitute for European soft wheat (Barassi and Ghoshray, 2007; Verga and Zuppiroli, 2003). In this work, the HRW is chosen as the representative price of the US market since the Hard Red Winter wheat quotations are indeed used to determine import duties by the European Commission for medium quality soft wheat (Gallezot, 2007). Moreover, wheat markets should be interrelated to the extent that individual wheat types are close substitutes in consumption and thus respond to global supply and demand conditions.

[16] Though the use of US dollars is more common in studying international agricultural markets, euros were preferred because of the analysis of EU policies which will be performed later.

[17] Goodwin et al., 1990, argue that this allows focusing directly on the LOP validity.

[18] The monthly adjustments for 1985 have been calculated as arithmetical average.

Some descriptive statistics of the data are provided in Table 2.

Table 2: Descriptive statistics of the price time series (EUR/t).

	Time	Average	Standard deviation	Skewness	Kurtosis
Soft wheat French price	**1978:12-2003:12**	**144.16**	**22.673**	**-0.292**	**-1.023**
	1978:12- 1988:06	161.18	12.886	-0.228	-1.244
	1988:07-1993:06	158.41	8.64	0.14	-0.79
	1993:07- 2000:06	126.30	11.57	-0.17	-1.10
	2000:07- 2003:12	112.94	12.61	1.48	2.80
Intervention price	**1978:12-2003:12**	**151.49**	**32.813**	**-0.126**	**-1.532**
	1978:12- 1988:06	179.14	15.923	-0.345	-0.908
	1988:07-1993:06	174.23	6.58	0.28	-0.54
	1993:07- 2000:06	119.99	5.46	-0.96	0.47
	2000:07- 2003:12	106.31	5.07	0.74	-0.63
HRW US price	**1978:12-2003:12**	**148.23**	**29.748**	**0.481**	**-0.129**
	1978:12- 1988:06	157.51	36.161	0.129	-1.031
	1988:07-1993:06	133.30	24.320	-0.015	1
	1993:07- 2000:06	140.73	21.249	1.089	1.507
	2000:07- 2003:12	159.16	16.395	1.321	2.573

We can notice that both *swfr* and *pint* average values show a decrease in time. On the other side, *hrw* decreases in the second sub-period, and then increases in the other two to higher levels than those in the beginning of the sample.

hrw's variability decreased over time; *swfr*'s variability, which is always smaller than the *hrw's* one, decreases, too, until 1993, to increase afterwards. Bale and Lutz (1979) demonstrate how policy measures and international trade distortions do affect the variability of prices: so, one might then argue that *swfr's* variability was kept low thanks to protectionist EU agricultural policies, and then rose after 1993, when substantial political reforms were introduced (Thompson 1999). Verga and Zuppiroli (2003) on the other side, assert that the increased volatility for the EU prices realized after 1993 doesn't depend on a major interaction with international prices, but only on the fact that, as the EU intervention price was reduced, prices could fluctuate more.

The distributions tend to be skewed and platykurtic, which are common features of non-stationary price series (Ghoshray 2000).

[19] EUROSTAT price data are normally producer prices, net of transport costs to the storage centre. This is very likely to contribute to the fact that the French price is so often below the intervention one. In the analysis, we consider this together with the other transport costs.

4. Time series analysis

In this section, the objective of the analysis is to find out whether a long-run relationship exists between the French soft wheat price (*swfr*) and the US CIF price (*hrw*). Prices are expressed in logs. The analysis is performed both on the whole sample (301 obs., 1978:12 to 2003:12), and on two sub-samples, obtained as described in Table 3. Considering the limited power of unit roots and cointegration tests, it was not deemed appropriate to split the whole sample into the four sub-samples described in the introduction.

Table 3: The two sub-samples used in the cointegration analysis.

Time period	Description
1978:12- 1993:06 *(175 obs.)*	Precedent to substantial CAP Reforms (Regular CMO functioning, Reform of 1988).
1993:07- 2003:12 *(126 obs.)*	Following substantial CAP Reforms (Mac Sharry Reform, Agenda 2000, Fischler Reform).

Besides the prices mentioned above, a composite variable, the "EU external reference price", has been calculated as follows. Since the target of the EU price policy is to keep *at least* the intervention price level in the internal market, each month the "EU external reference price" is calculated as the maximum between the intervention price and the world price. What is argued is that the EU price follows the US price, when it is above the intervention price, and *vice versa* (Figure 3). The intervention price acts as an implicit downward threshold for the US price, that is not-active when below such a threshold. As expected, the US price tends to be higher than the intervention price from the 1990s on (Figure 4).

Figure 3: Soft wheat French price, US HRW price, intervention price, whole sample (EUR/t).

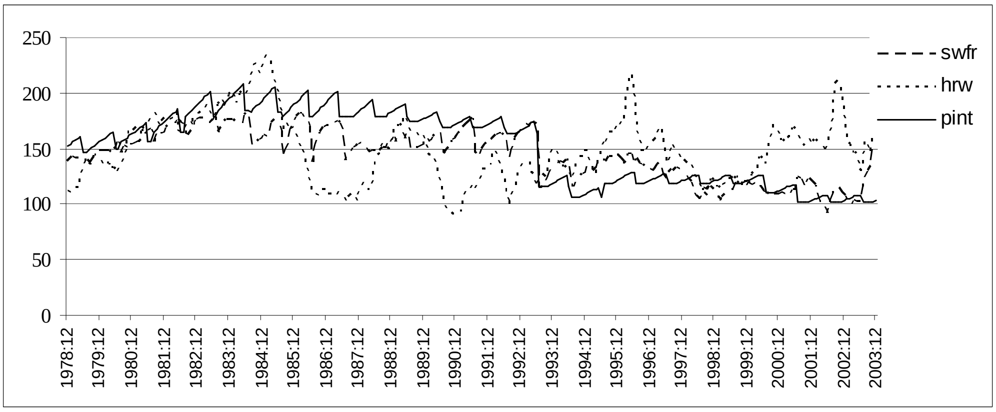
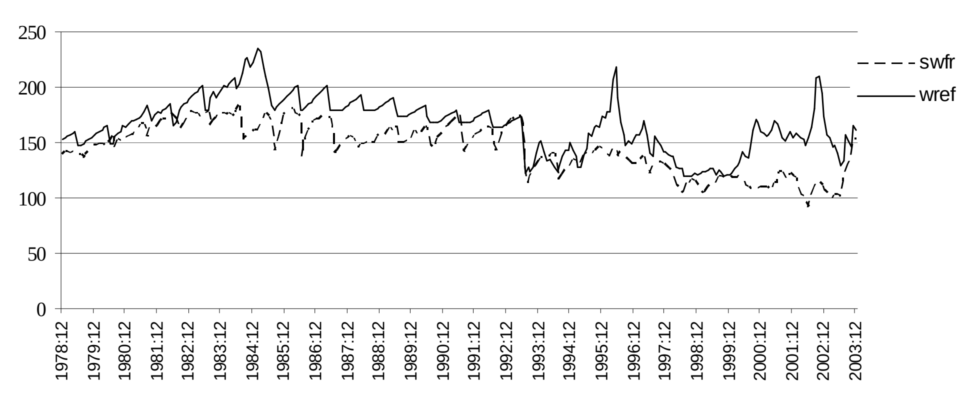


Figure 4: Soft wheat French price and world reference price, whole sample (EUR/t).



Unit root tests have been repeated both in the whole period and in the two above mentioned sub-periods. Augmented Dickey Fuller (ADF) tests have been run by adding lags until the last significant one up to a maximum of 18 lags, due to the monthly nature of the data. Phillips-Perron (PP) tests were run with a number of lags determined by minimizing the Schwarz's Bayesian information criterion (SBIC) in a number of autoregressive specifications up to a maximum of 18 lags. The choice of the lag length, which was generally unaffected by the deterministic trend, was done also for the log-differentiated time series. This number of lags was used for the KPPS (Kwiatkowski, Phillips, Schmidt e Shin) tests, as well[20]. While on logs the null hypothesis of unit roots could not be rejected, for log-differences it was possible to do so. This means that the series can be considered I(1) processes, as it is the case in all the literature revised, despite the fact that they are nominal prices (Fanelli and Bacchiocchi 2005). This is generally true for all the series considered, both in the whole sample and in the two sub-samples.

It could be argued that the presence of seasonality in the monthly price series might undermine the ability of ADF tests to verify unit roots. For this reason, monthly dummies were introduced into the ADF tests estimations. Results overall confirmed that the processes are I(1), and are reported in Annex 2.

Provided that the prices are first-difference stationary, a cointegration analysis has been performed. The series have been analyzed in pairs. The optimum lag-length for the VAR has been chosen according to the minimization of the Schwartz Bayesian Criterion up to a maximum of 24 lags, for the no-constant, constant and trend options. Generally speaking, the optimum number of lags selected was not or minimally affected by the deterministic trend

[20] Both STATA® and GRETL softwares were used for the econometric estimations.

chosen. The Johansen and Juselius procedure has been used to estimate the rank of the cointegrating matrix.

First of all, in the whole sample, the rank of the cointegrating matrix was estimated between *swfr* and *hrw,* to see whether the LOP holds between the EU and the US prices.

The optimal lag selection was 2 lags. As expected from visual inspection of the two time series and from policy considerations, the rank of the cointegration matrix turned out to be zero. This means that, in the period examined, the French soft wheat price and the HRW Gulf US price are not linked by any long-run relationship. We conclude that the each of the two prices follows its own pattern and that the influence of the policy instruments is so strong that they are not related to each other. This should result not come as a surprise, since the very existence of the European border policies (namely, variable levies and export subsidies) is very likely to bring about such a result.

Other studies (Barassi and Ghoshray 2007; Ghoshray et al. 2000) show instead the presence of cointegration between the EU price and the US price. These different results are simply explained by the fact that they use EU Rotterdam FOB prices. These prices are indeed very close to the world ones, since it is on top of them that export subsidies are paid to European exporters. So, there should be no surprise also in their findings that the EU is a price taker in world markets, right because it applies export refunds on the basis of the world market prices.

Thompson and Bohl (1999) use producer’s level German prices and US FOB DNS prices and find that the two time series are cointegrated. They use a threshold cointegrating technique, which could partly explain this difference in results. On the other side, the selection of the DNS price (DNS has even a higher protein content than HRW) was not appropriate for the purposes of this paper.

At this point, the cointegration analysis was repeated for *swfr* and the series *wref*, which contains 162 times the intervention price over a total of 301 months (54% of the observations are constituted by the intervention price).

The rank of the cointegration matrix, tested with 2 lags in the restricted constant option[21], turned out to be one. The VECM was first estimated with two lags; considering the monthly nature of the data, the insertion of twelfth lagged differences and seasonal dummies (selected with specification tests) was necessary in order to remove the presence of autocorrelation in the residuals[22]. The resulting model had the constant very big in magnitude and negative in sign. Since the introduction of seasonal dummies actually lowered the VAR optimum lag selection to 1 (according to the SBIC criterion), we tried by removing the first differenced lag (which was scarcely significant, as well).
Estimates are reported in Table 4.

Table 4: VECM coefficient, 1978:12-2003:12, *swfr* and *wref*.

	Δswfr (t)	**Δwref (t)**	**Long-run relation**
α	-0.043**	0.031	
Δswfr (t-12)	0.482***	0.107*	$swfr_t = \underset{(0)}{0.149} + \underset{(1.196)}{0.942}\, wref_t$
Δwref (t-12)	-0.065	-0.026	
D6	-0.008	-0.053***	Standard errors in parentesis.
D7	-0.019**	-0.018**	
D8	-0.003	0.029***	

* 10% significance; ** 5% significance; *** 1% significance.

The resulting long run relationship is $swfr_t = 0.149 + 0.941 wref_t$. The price transmission elasticity's coefficient between the time series has the right sign and is indeed very close to one[23]. The constant term accounts for transaction costs; remembering that the equation is written in logarithms, this implies that transaction costs are equal to a constant proportion of

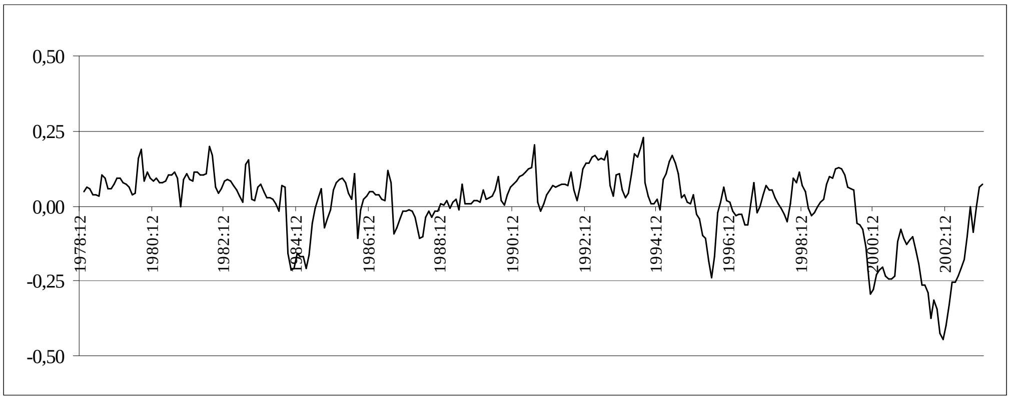
[21] Visual inspection of the data shows that there is no trend in the series (Figure 3), and both theory and visual inspection of the data imply the presence of a constant in the long-run relationship, accounting for transport costs. This means that, even if there are no linear time trends in the level of the data, the cointegrating relation has a constant mean. The general form of the Vectorial Error Correction Model (VECM) which is estimated is $\mathbf{\Delta y_t = \alpha\beta' y_{t-1}} + \sum_{i=1}^{p} \mathbf{\Gamma_i \Delta y_{t-1} + \nu + \delta t + \varepsilon_t}$ where $\mathbf{\nu = \alpha\mu + \gamma}$ and $\mathbf{\delta t = \alpha\rho t + \tau t}$ account for the presence of deterministic trends. Restrictions are imposed on the coefficients so that the model we actually estimate is $\mathbf{\Delta y_t = \alpha(\beta' y_{t-1} + \mu)} + \sum_{i=1}^{p} \mathbf{\Gamma_i \Delta y_{t-1} + \varepsilon_t}$.
[22] Autocorrelation was tested by a LM test with the null hypothesis of no-autocorrelation up to the 18th lag.
[23] According to its t-statistic the coefficient is not statistically significant but, most importantly, restricting the cointegrating vector to [1 0 -1] is not rejected with a p-value of 0.129. This means that imposing perfect price transmission elasticity doesn't alter the cointegration rank.

prices[24]. Johansen's test of weak exogeneity (Goshray 2000) is a test of the statistical significance of the error correction coefficients. An insignificant loading coefficient indicates that the price is weakly exogenous. Here the *wref* series seems to perform as the weakly exogenous one. This could have to do with the inclusion of the intervention price (a policy measure which is exogenously determined) in this composite variable.

The values of the elasticities found by Thompson and Bohl, 1999, whose data are of a kind comparable to those used in this study, are 0.54 for the long-run response of the German price to US Dark Northern Spring price. The value of Germany's adjustment coefficient is -0.03.

The values of the cointegrating relationship are shown in Figure 5. The residuals of the cointegration relation show indeed a stationary behaviour (the ADF test in the no-constant case rejected the null hypothesis of a unit root with a p-value of 0.021). Three major deviations from the equilibrium occurred when the French price was substantially lower than the external reference price. In the first two, occurring in 1985 and 1996, the explanation of the disequilibrium lies in the taxes on exports imposed by the European Commission to prevent domestic prices from rising. In 2002, this was not the case; it has been argued that, despite the high international prices, EU prices were depressed by large inflows of Russian and Ukrainian grains, which could take advantage of the low tariffs in place because of the high US prices (which soon led to the creation of a new European tariff rate quota for low-quality grains).

Figure 5: Cointegration relationship, long run errors, whole sample.



[24] The underlying equation in levels is $SWFR_t = e^{0.149} WREF_t^{0.941}$ where $swfr = \ln SWFR$ and $wref = \ln WREF$; this means that, in levels, the presence of transaction costs implies multiplying the external reference price for the constant 1.161.

The analysis was then repeated for the first sub-sample. In the period 1978:12-1993:06, the external reference time series turns out to be constituted almost exclusively by the intervention price with a few exceptions (the intervention price is above the US HRW price 145 times over 175 months, i.e. 83% of the observations).

The rank of the cointegrating matrix was estimated for *swfr* and *hrw*. The optimum lag selection was 2 lags without dummies and 1 with. Cointegration was then tested using Johansen's maximum likelihood procedure with two lags in the restricted constant option. A VECM is estimated in the restricted constant option with the addiction of a twelfth differentiated lag in order to deal with the problem of autocorrelation. The transmission elasticity coefficient is very low (see estimates reported in Table 5) and the constant very big.

Table 5: VECM coefficient, 1978:12-1993:06, *swfr* and *hrw*.

	Δswfr (t)	**Δhrw (t)**	**Long-run relation**
Z (t-1)	-0.208***	-0.025	
Δswfr (t-1)	0.011	0.314***	$swfr_t = \underset{(0)}{4.800} + \underset{(0.265)}{0.057}\, hrw_t$
Δhrw (t-1)	0.070	0.224***	
Δswfr (t-12)	0.514	-0.048	Standard errors in parentesis.
Δhrw (t-12)	-0.208	-0.124--	

* 10% significance; ** 5% significance; *** 1% significance.

The existence of cointegration between *swfr* and *wref* was then checked. The optimal VAR lag-selection was 2 lags without dummies and 1 with. Cointegration was tested using Johansen's maximum likelihood procedure with two lags in the restricted constant options; the cointegration rank turned out to be one. A VECM is estimated with two lags in the restricted constant option; like in the whole sample, a twelfth lagged differentiated term and seasonal dummies were included in order to deal with autocorrelation[25]. Estimates are reported in Table 6.

[25] The first lagged difference was not significant and was excluded from the regression, together with the seasonal dummies that turned out not to be significant.

Table 6: VECM coefficient, 1978:12-1993:06, *swfr* and *wref*.

	Δswfr (t)	Δwref (t)	Long-run relation
α	-0.224***	0.064**	
Δswfr (t-12)	0.326	0.032	$swfr_t = \underset{(0)}{3.484} + \underset{(0.632)}{0.310}\, wref_t$
Δwref (t-12)	0.114	-0.082	
D1	0.007	0.012**	Standard errors in parentesis.
D2	0.0006	0.012**	
D6	0.009	-0.072***	
D7	-0.038***	-0.008	
D8	-0.039***	0.022***	
D9	0	0.013*	
D11	0.008	0.013**	

* 10% significance; ** 5% significance; *** 1% significance.

The long run relations is $swfr_t = 3.484 + 0.310\, wref_t$. In this case, none of the variables performs as weakly exogenous. The value of the price transmission coefficient is smaller than when all observations are included in the sample.

Finally, the analysis was repeated for the second sub-sample. In the period 1993:07-2000:06 the external reference time series turns out to be constituted almost exclusively by the US HRW price (which is lower than the intervention price only 17 months over 126, which means 13% of them).

swfr and hrw turned out not to be cointegrated (see Table 7). Summing up, we can say that the two series seem not to share any long-run common trend (the cointegrating relation found in the first sub-sample is indeed very weak).

The existence of cointegration between *swfr* and *wref* was then checked. The optimal lag-selection was 1 lag both with and without seasonal dummies. Cointegration was then tested using Johansen’s maximum likelihood procedure with one lag in the restricted constant case. The cointegration rank was zero (see Table 7). This is an expected result, since the *wref* time series is mainly constituted by *hrw*, which is not cointegrated with the French price.

So, *wref* is cointegrated with the French price in the whole sample and in the first sub-sample, but not in the second one.

Table 7. Results of the Johansen's tests in the period 1993:06-2003:12.

	Cointegration tests 2 lags, restricted constant option	
	swfr and *hrw*	*swfr* and *wref*
$H_0(H_1)$	trace statistics (pvalue)	
r = 0 (r = 2)	14.173 (0.2841)	15.102 (0.2254)
r = 1 (r = 2)	3.9864 (0.4262)	4.0327 (0.4193)

Summing up, in the whole sample, while the French price turned out not to be cointegrated with the US HRW price, it showed indeed to share a long-run relation with the "EU external reference" price, composed by the maximum value over the intervention price and the US HRW price. The intervention price is implicitly assumed to be a threshold for the US price to influence the EU one. Despite the fact that a commercial policy insulating the domestic market from the world one has been practically always in place, what we observe is a co-movement of the European price and the US one under certain conditions, namely, the latter to be higher than the intervention price.

The cointegration relation between the French price and *wref* was present in the period 1978:12-1993:06, too, but not in the 1993:07-2003:12 one. We could think that the transition between the two sub-samples in itself does have a role in explaining price transmission elasticities. But, since it is actually in this second sub-sample that the US price is above the intervention price, an obvious consideration is that the linkage between *swfr* and *wref* could be nothing but a linkage between the French price and the intervention price[26]. The answer to this question is not straightforward, first of all because the relationship present in the whole sample seems to be very different from the one present in the first sub-period. Moreover, in the second sub-sample finding no cointegration relationship could depend on the fact that the analysis doesn't explicitly take into account what happened in 2002, when the Ukrainian price, much lower than the US one, is very likely to have been the "real" world price for EU operators. This problem might gain weight as the number of the observations is reduced in passing from the whole sample to the second sub-sample.

[26] Very preliminary analysis showed no cointegrating relations existing between *lpint* and *swfr* on the whole sample, some evidence of cointegration in the first sub-sample and no relation in the second one.
This results agree with those of Verga and Zuppiroli (2003): by using weekly data, they find that US Soft Red Winter Wheat Rotterdam CIF price is never cointegrated with domestic European prices, and that the intervention price is cointegrated in the period 1990-2002 but not in the sub-period 1995-02. This could be explained by the instability of the relation between intervention prices and EU internal prices. In 1995-2002, when both the intervention price and the US price are put together in the cointegrating relation, they find evidence of cointegration. They suggest that EU quotations could be linked to an "average" the two prices.

5. Linear regression analysis

In the AGMEMOD model, for each commodity the price transmission mechanism is represented by a price formation and a set of price transmission equations, all estimated with OLS techniques (time-series properties of the data are not considered). The first equation regresses what is called the “EU key price”, i.e. the price of the representative country for that specific product (that is, for soft wheat, France) on the world price, together with other policy-relevant variables. The world price is then assumed to be exogenous. This “key price” equation for a given commodity market is also a function of the EU self-sufficiency rate, reflecting the endogenous development of the EU internal balance for the commodity concerned. For example, the EU soft wheat key price (i. e., the French price) is modelled as a function of the US price, the EU soft wheat intervention price, some relevant trade policy variables (tariff rate quotas, limits on export subsidies) and the self-sufficiency rate for wheat in France. Then, through the price transmission equations, each country’s price is then regressed on this key price.

The underlying assumption is that the EU is a “small economy” which doesn’t affect world prices. If the world price has to become endogenous in the model, thus relaxing the small country hypothesis and moving toward the more realistic assumption that the EU is a large country affecting the world prices’ level, the time series approach proposed in the previous section is an interesting alternative to overcome this drawback of the model.

As far as OLS techniques are concerned, also recently Thompson et al. (2000), using annual data, test whether the EU internal price really doesn’t depend on the world price (as it should be the case because of its domestic and border policies), and make a regression of the German price on the US price.

Here also an OLS approach is used, but, since *swfr* proved not to be cointegrated with *hrw*, any OLS regression between them is likely to be spurious. On the other side, since we know that *wref* and *swfr* are cointegrated, we can further investigate the long-run relationship between the two variables with standard OLS estimates, checking how it depends on policy regime changes.

From econometric theory we know that, in case the series exhibit a non- stationary behaviour, static models estimated by OLS will be biased or inconsistent. But, if the two processes are cointegrated, then OLS estimates are instead super-consistent, which means that they converge at faster rate than normal[27]. So, provided the simple static regression is not spurious, dynamic misspecification is not necessarily a problem.

The fundamental difference from the cointegration approach is that in this case one of the two variables is implicitly assumed to be exogenous. This has been claimed to be one of the major drawbacks of linear regressions to study price transmission mechanisms, and is one criticism to the so called “Ravallion’s model” (Baulch 1997). Keeping this in mind, here we want to see how the dependency of the EU price on the EU external reference price (the maximum between the US price and the intervention price) is affected by policy regime changes.

The model which will be estimated is, in its most general form, $swfr_t = \alpha + \beta_0 wref_t + \beta_1 R_1 + \beta_2 R_2 + \beta_3 R_3 + \beta_{40} R_4 + \beta_4 R_4 * wref_t + \varepsilon_t$, where all prices continue to be expressed in logs[28]. The EU policy dummies are defined in Table 8:

Table 8: The policy dummies used.

Time period	Variable name	Description
1978:12- 1988:06	dropped	Regular CMO functioning
1988:07-1993:06	**R1**	First CMO Reforms
1993:07- 2000:06	**R2**	Mac Sharry Reform
2000:06- 2003:12	**R3**	Agenda 2000 Reform and Fischler Reform
1995:01-2003:12	**R4**	URAA implementation

Thanks to the length of the sample four sub-periods can be identified, which are characterized, as explained in the introduction, by progressively more market-oriented policies. These policy reforms are assumed to have had an effect on the level of prices: through the reduction of the intervention prices, EU domestic prices are expected to have fallen. On the other side, the last dummy accounts for changes in the elasticity of price transmission, which should be the consequence of the implementation of the URAA agreement.

Thompson et al. (2000) just insert one dummy which accounts for the Mac Sharry Reform and another one for the URAA implementation. They find that the first one has a negative coefficient, as implied by the reduction of intervention prices. Concerning the URAA dummy, they demonstrate that with effective tariffication in place the elasticity of transmission between word prices and domestic prices is equal to 1, while it is equal to zero if the “155% of the intervention price” rule is in place, or if the country fixing the domestic price faces a

[27] Nonetheless, super-consistency is a large sample result, and coefficients might be biased in finite samples.

[28] This means that the underlying model written in levels is $P_{fr,t} = AP_{wref,t}^{\beta_0 + \beta_4 R_4} e^{(\beta_1 R_1 + \beta_2 R_2 + \beta_3 R_3 + \beta_{04} R_4)}$, where $\alpha = \ln A$, $swfr = \ln P_{fr,t}$, and $P_{fr,t}$ and $P_{wref,t}$ indicate the French price and the EU external reference price expressed in levels.

net export situation. This means that the coefficient of the URAA dummy should help to understand how effective tariffication was.

In this study, care should be taken in the interpretation of the coefficients because *wref* already takes into account intervention prices, that are assumed to be the "threshold" above which US prices have an influence on the EU ones.

Results of the estimates are reported in Table 9.

Table 9: Estimate's result, OLS estimator (301 obs.).

Regressors	α	β_0	β_1	β_2	β_3	β_{04}	β_4	R^2 adj.	BIC
Model 1: wref	1.007***	0.774***						0.500	-432.481
*Model 2: wref, R1,R2, R3, R4,R4*$p_{wref,t}$*	2.781***	0.441***	0.003	-0.102***	-0.264***	-0.178	-0.029	0.821	-718.98
Model 3: wref, R1,R2, R3	2.740***	0.449***	-0.004	-0.128***	-0.295***			0.820	-726.262
*Model 4: wref,R4,R4*$p_{wref,t}$*	1.918***	0.607***				1.623***	-0.357***	0.739	-619.699

**indicates 5% significance level of the parameters, ***1% significance level.

Differently from Thompson et al., 2000, a time trend was not included, right because it was deemed not to be possible to distinguish between increasing "budgetary pressures" which a time trend should account for and the progressive reduction of the intervention price.

Model 1 represents the cointegration relationship identified in the previous section. This means that a 1% change in the external reference EU price causes a 0.774% change in the French domestic price.

Some general considerations about the meaning of the dummies can be drawn. First of all, we can notice that the number of months in which the intervention price was actually above the US price decreased over time, as the CAP changes would lead us to think (Table 10; see also Figure 6).

Table 10: Number of months in which the intervention price was above the US price.

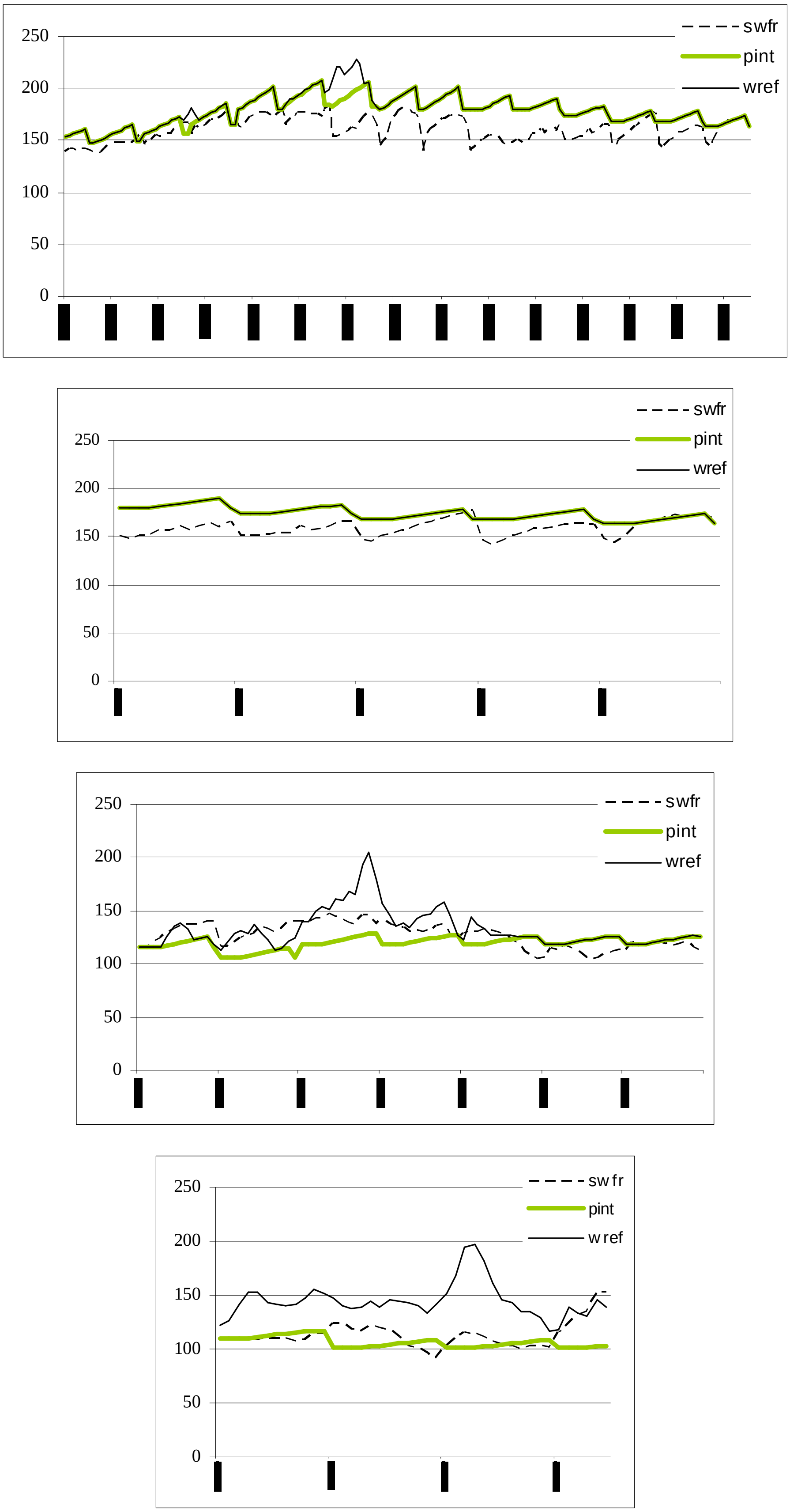
Time period		Number of months in which pint>hrw	%
1978:12-1988:06		85/115	74%
1988:07-1993:06	R1	60/60	100%
1993:07- 2000:06	R2	17/84	20%
2000:06- 2003:12	R3	0/42	0%

The first dummy, R1, which refers to the period 1988:07-1993:06, seems to very close to zero and not significant. This is somehow an expected result, since in these months the intervention price was always above the US price, and is thus already included in *wref*.

R2, covering the years 1993:07-2000:6, and R3, for the period 2000:07-2003:12, are both negative. This implies a reduction in the level of prices, following the implementation of the EU CAP Reforms. In this two last sub-periods we notice that the external reference price series contains an increasing number of observations of the US HRW price, because it is higher than the intervention price. The EU internal price is then deemed to follow the US price rather than the intervention price. Despite the fact that HRW prices are not only higher than intervention prices, but also increasing in their absolute value, what we see is that EU prices decrease relative to their initial values.

R4, which covers the URAA implementation, has a negative sign. This means that, from 1995 on, the price transmission elasticity decreased. This results is somehow difficult to be interpreted: we might argue that the URAA implementation was ineffective. But this result might also depend on the fact that *wref* is constituted by the US price instead of the intervention price an increasing number of times after 1995 (which causes some interpretative problems, as it happened in the previous section).

Figure 6: Soft wheat French price, intervention price and EU external reference price in the four periods (EUR/t).



6. Some final considerations

This paper aims at suggesting some further ways of exploring the nature of international price transmission relationships for agricultural commodities. Differently from most of the previous works, domestic EU prices for soft wheat have been used together with US HRW prices. This allowed evaluating the impact of both internal and commercial policies. French prices are assumed to be representative of EU prices. Soft wheat is a commodity which is heavily traded internationally and whose market has been heavily regulated by the CAP. The analysis has been performed in the general framework of the LOP validity.

The cointegration analysis carried out for the period 1978:12-2003:12 showed that, as expected, there is no long-run relation between the US CIF price and the EU domestic price. EU policies prevented prices from sharing the same pattern.

When the validity of the LOP between the EU domestic price and the "EU external reference price" (constituted by the maximum between the US price and the intervention price), was tested, evidence of the presence of a long-run relation is found, and it is consistent with the LOP. This means that domestic and commercial EU policies for soft wheat actually played a strong role in insulating the internal markets from the world ones, especially until 1993. After this date, the reduction of the intervention price and the evolution of international markets caused US prices to be much more often above the intervention price. The analysis shows that this allowed the US prices to have a stronger effect on the EU internal ones, despite the fact that the same system of border policies kept being in place. We might then argue that the reduction in intervention prices, more than the (non)-reduction of policy barriers, was what increased the degree of interaction between the EU internal prices and the US ones (together with increasing the variability of EU internal prices).

The impact of the major EU domestic policy reforms and international trade setting changes was also evaluated. Despite the increasing relevance of US prices, and their increase in absolute value, EU domestic prices show a constant decline over the years. The URAA implementation seems to have had no or negative effect on price transmission elasticities, as it would be expected provided the fact that major internal reforms had already been implemented years before it was in place, and that no real reduction of border policies was implemented after it.

For sure, the model here presented is over-simplified, in that it doesn't explicitly take into account a number of policy instruments that were in place in the past 25 years and are very likely to have had an impact. Nevertheless, policy developments are taken into consideration implicitly *via* the construction of an *ad hoc* composite variable, and the price transmission equations developed can be inserted and tested in more broad econometric models (Listorti and Esposti, 2008). The use of more sophisticated econometric techniques, like threshold models or cointegration models accounting for structural breaks, needs further research. The analysis needs for sure to be updated once more recent data will be available in order to disentangle the dynamics of the last years, right because of the major influence that market mechanisms have shown to have had.

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ANNEX 1 : Some previous studies on price transmission in agricultural markets.

METHODOLOGY	DATA	MAJOR FINDINGS
CORRELATION		
Stigler and Sherwin, 1985		
They investigate the relations existing between differentiated series of prices. They use correlation analysis to assess the belonging of prices to the same market.	Monthly wholesale prices (logs, indexes, levels) for a number of items: silver futures, flour, wheat, gasoline, interest rates, wages… Various years.	The study, which covers different kinds of markets (future markets, commodity markets, capital and labour markets), discusses fundamental issues related to market integration.
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: RAVALLION'S CRITERIA		
Aldermann,1992		
Applies both Ravallion's radial dynamic model of market integration and cointegration techniques, adapting them to study markets of commodities that are close substitutes. Cointegration relations are analyzed within the same market between commodities that can be considered close substitutes. His objective is to verify whether the connections existing among markets of different agricultural commodities allow governments in developing countries to use policy measures addressed to only one of them.	Monthly wholesale maize, sorghum and millet prices (levels) from three markets in Ghana. 1977-1990.	The basic assumption is that, if markets are efficiently linked, the lagged price of one commodity should not contain information not contained in the past price of another one (provided it is a close substitute): the information conveyed in p_{jt} will not improve the prediction in p_{it+1} over what contained in p_{it}. A single market is efficient if the prices of two commodities are not cointegrated, since cointegration would imply GC in at least one direction, which can be interpreted as inefficiency. He finds functional efficiency in Ghana's coarse grain markets.
Ravallion, 1986		
He proposes a procedure for testing market integration. The price series for each local market have their own autoregressive structure and a dynamic relationship with market prices in a central region. His approach permits to distinguish between short and long run market integration.	Monthly rice price data in five districts of Bangladesh (levels). Dummies to account for seasonality, the famine in 1974, and a time trend are included. July 1972- June 1975.	Departures from the conditions of both short run and long run integration are found in Bangladesh. They would not be revealed by the use of static correlation techniques.
Richardson, 1978		
He tests the Law of One Price using twice differenced prices, to avoid a number of problems related to serial correlation and omitted explanatory variables.	Monthly observations on Canadian and US price indexes (twice differenced logs). The exchange rate is included as a regressor. 1965-1974.	The Law of One Price fails. Canadian prices respond in the same way to exchange rates and US prices.
Thompson et al., 2000		
They want to test if the elasticity of transmission between EU and world prices was zero under the old CAP, and how it was affected by policy regime changes. They also build an econometric model. They analyze the effect of the change in the variability of prices on consumers' welfare. Germany is assumed to be a small country for which the world price is exogenous.	Annual data are derived by averaging monthly data (logs are used). The world wheat price is in US $ from the USDA, converted into DM using IMF exchange rates. The German producer price is from CRONOS dataset of EUROSTAT. Both price series are deflated by	They find that the change in the internal policy regime caused a reduction in the levels of EU prices, but also that the URAA didn't have any significant effect on price transmission elasticities (which means that tariffication was not effective). The existence of low but not zero price transmission elasticities before

	the CPI of their respective country. 1976-1998.	the URAA was in place suggest that the domestic support prices were not completely independent from the world ones. The impact of policy changes on producers' welfare is mostly due to cuts in protection (transfer) than to income instability (risk). Through compensative hectare payments, EU farmers are overcompensated.
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL (GC TESTS, IRFS, FEV)		
Goodwin and Schroeder, 1991a		
They estimate a VAR in levels. Their analysis focuses on Forecast Variance Error (FEV) decomposition and Impulse-Response Functions (IRF).	Monthly FOB prices (in levels) for six exporting and importing wheat markets: US, Canada, Argentine, Australia, Japan and the EU. Missing observations are obtained regressing the series on a price series of a closely related market. Special Drawing Rights are used as a measure for exchange rates effects. A unique freight rate is calculated as the arithmetic average of various ones. July 1975 - December 1986	According to the FEV, US and Canada seem to be dominant markets. A large proportion of FEVs are explained by transportation costs and the US and Canadian prices. Adjustments to innovations in freight rates are quite slow to occur. The IRFs show rapid price adjustments to exchange rate shocks, to US price shocks, to Canadian shocks (but not for the US and Argentina). Shocks in freight rates take two or more months to produce effects, but the response is large and persistent.
Gordon et al., 1983		
They use a modification of Granger Causality Test, the Holmes and Hutton causality test based on the rank ordering of each variable. They test both bivariate models and trivariate ones.	Weekly prices (levels) of lamb: French lamb, Anglo-Irish lamb price on the French market, UK lamb price in the UK. 1983-1986.	British and French markets are integrated in the sense that price shocks in one market are eventually and fully transmitted on price changes in the other market. However, long orders of lag specifications are necessary, indicating a slow response to price incentives.
Gupta and Mueller, 1982		
They use GC tests (Haugh test to assess the dependence-independence between series; Sims test to ascertain the direction of causality) to see whether markets are perfectly price-efficient. Differently from the use of correlation coefficients, which only report the association between prices, this methodology allows to test if markets are independent, interdependent, or if lead-lag relations exist.	Weekly prices (differentiated logs) for slaughter hogs in three German. Week 1:1977 - week 50:1980	They find that markets are price-efficient, since the tests show that they are interdependent.
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Ardeni, 1989		
For the first time, cointegration is proposed to analyze the LOP.	Quarterly export prices (in logs) for wheat, wool, beef, sugar, tea, tin and zinc for Australia,	The LOP holds only in a small number of cases (but is nonetheless valid for US-

	Canada, the UK, the US. Data are adjusted for the exchange rates. January 1957-January 1986 (from 79 to 117 obs., depending on the commodity).	Australian wheat, US-Canadian wheat); deviations from the pattern are permanent.
Baffes, 1991		
Using the same dataset and the same techniques as Ardeni, he shows instead that the LOP holds. The difference in the findings is explained by the fact that he imposes the cointegration parameters and uses the variables in levels.	It is the same as Ardeni, but variables are used in levels.	The LOP cannot be rejected as a maintained hypothesis. When it fails, transport costs might be the explanation (non-stationarity of freight-rates might explain non-stationarity of prices differentials).
Barassi and Ghoshray, 2007		
They test cointegration with structural change between the EU and the US. They use a novel cointegration method which allows the time of the break to be unknown.	Monthly FOB prices for EU wheat, US Hard Red Winter Wheat and US Soft Red Wheat (used in logs). July1981-July 2000.	A structural break occurred after the 1992 CAP Reforms were implemented. Before that, the only cointegrating relation existing is between the two US wheats. After that, they find a cointegrating vector between the US HRW and the US SRW, and another one between the US SRW and the EU. The EU acts as a price follower, as it would be expected since it makes use of export subsidies.
Dawson and Sanjuan, 2007		
They apply Johansen's procedure allowing for structural breaks. The time of occurrence of the structural break is known.	Monthly FOB prices (in logs): Canadian Western Red Spring, US Dark Northern Spring. January 1974-December2001.	They find that cointegration exists and that there are two structural breaks corresponding to the beginning and the end of the US Export Enhancement Programme.
Ghoshray and Lyoid, 2003		
By using the method of irreducible cointegrated vectors, they investigate the relations existing amongst wheat prices of different countries according to wheat type and port location.	Monthly FOB export prices (in logs) of 13 different price series (from Argentina, Australia, Canada, the US, the EU). July 1980-December 1998.	They provide statistical support for examining wheats that have different end uses (namely hard wheats), separately.
Ghoshray et al., 2000		
They analyze the cointegrating relationships existing in 12 pairs of prices constituted by the EU price plus another one.	Monthly FOB export prices (in logs) of 13 different price series (from Argentina, Australia, Canada, US, EU). July 1980-December 1998.	They find that cointegration exists in all the pairs but in those including Canadian feed wheat. They argue that this means that all wheats used for human consumption show a common trend.
Goodwin, 1992		
He applies Johansen's multivariate cointegrating testing procedure, arguing that bivariate Engle Granger cointegration tests are limited.	Monthly FOB export prices (in levels) for the US, Canada, Australia, CIF prices for Japan and the EU. Monthly average freight rates from the US Gulf to Rotterdam and Japan. January 1978-December 1989.	He argues that, if transport costs are explicitly taken into account, the Law of One Price is valid. The inclusion of freight rates in the model (by subtracting them to the CIF prices of the EU and Japan) allows to find one cointegration vector.

Goodwin and Schroeder, 1991b		
By using cointegration tests, they evaluate spatial linkages in regional cattle markets. They conduct seven different Engle-Granger cointegration tests for two specifications of ten market comparisons over four periods. By using bootstrap regression techniques, they assess the influence on the test statistics of concentration ratios, average annual slaughter volumes and the distance between markets.	Weekly price series for slaughter steers for eleven regional US markets (in levels). January 1980-September 1987.	Their results can be summarized as follows: increased market concentration leads to higher cointegration; relative slaughter volume has a negative impact on cointegration (smaller markets exhibit a smaller degree of spatial dependence than bigger ones); the degree of price cointegration is negatively affected by bigger distances between markets. Increasing cointegration during the 1980s is explained by structural changes in the livestock industry.
Margarido et al., 2004		
They investigate the elasticities of transmission in the soybeans market through the use of cointegration techniques (one VECM model for all the prices). They then calculate IRFs and FEVs.	Monthly soybeans prices (in logs): CIF Rotterdam Port, FOB Argentina, FOB US. October 1995-October 2003.	The LOP is valid in the long run. Brazil and Argentina can be seen as price takers. Seasonal differences may explain the pattern of the response of Brazilian prices to shocks in the international market.
Mohanty and Langley, 2003		
They examine the integration between US and Canadian grain prices using cointegration techniques and ECMs in four different sub-periods: pre-post NAFTA, pre-post the Western Grain Transportation Act of Canada.	Monthly prices for wheat and barley (in logs). For Canada and the US. June 1986 - July 1999.	They find that the series are always cointegrated, and that the coefficients of transmissions are higher after the implementation of the two agreements.
Mohanty et al., 1999		
They analyze world wheat markets using cointegration techniques and an ECM model which includes only the variables significant at 10% level.	Monthy FOB prices (in logs) for the US, Canada, Australia, Argentina and the EU. January 1981 to June 1993.	They find that there is no distinct leader in the international wheat market. This differs with the results of a number of different studies, that could nonetheless have been mis-specified (for example, because they didn't take cointegration into account).
Sanjuan and Gil, 2001.		
Cointegration tests are applied to study long-run relationships; GC tests are then used to obtain the general pattern of influences. FEV analysis is used to analyze the strength of price interdependence.	Weekly prices (in levels) for pork and lamb carcasses for seven EU countries (DE, DK, ES, FR, IL, NL,UK). 1988-1995 (418 obs.).	Pork markets shows a high degree of integration; a more limited degree of integration is observed in the sheep market.
Thompson and Bohl, 1999		
They perform a cointegration analysis to check how policy regime changes affected international price transmission elasticities. They identify three different sub-periods corresponding to the main European policy changes. They use a threshold cointegrating technique (Enders and Siklos) that allows for the cointegrating relationship to be locally inactive and then become active once the system gets too far from the	Domestic producer selling prices for Germany are obtained from CRONOS database of EUROSTAT (DM). World prices (Dark Northern Spring Wheat, CIF Rotterdam, US$) are obtained from the USDA. Monthly exchange rates from the IMF were used for the conversion. Nominal prices are used since real price didn't show	The price series are integrated. Long run transmission elasticities range from 0.18 during the 70s and 80s and 0.30 during the post URAA. They argue that reforms made to the CAP had an effect in increasing internal price variability. World price volatility decreased over the period considered.

equilibrium relationship.	any statistical difference. All prices are expressed in logs. June 1976 to December 1998.	
Verga e Zuppiroli, 2003		
They apply cointegration analysis to four European soft wheat markets to test the relations between them, with the institutional EU prices and with the international prices. With the "directed graphs" technique, they also analyze contemporaneous relations between the prices. They tests are run on the whole subsample and in the sub-period 1995-2002.	Weekly prices (in levels) for two main Italian markets and two main French makets; CIF Rotterdam price of US Soft Red Winter wheat; intervention price; EU import price. January 1990-December 2002.	The European market is strongly cointegrated, but there is no cointegrating relationship with the US price. The series are cointegrated with the institutional prices (but they claim this relation is unstable and thus distorted) but not in the subsample chosen. Higher variability of EU prices is not due to bigger linkages with the international markets but to the lowering of the intervention price.
Viju et al., 2006		
They asses the accession of Austria, Finland and Sweden to the EU from the perspective of market integration. They use cointegration techniques in two different samples, pre and post accession. For each product, for each period, they run cointegration analysis in pairs with the German price.	Monthly data (in logs) for rye, oats, barley, soft wheat and potatoes for Austria, Finland, Sweden, Germany (=EU reference price). Data are obtained from CRONOS, EUROSTAT, and all converted to ECU/EUR. January 1975-December 2004.	For soft wheat, there is evidence of market integration with the exception of Finland. For the couple Germany-Austria there was evidence of market integration before Austria joined the EU. The observed convergence in prices for soft wheat might be influenced by operation at a minimum price, even thought this situation would not likely be observed without markets being linked through arbitrage.
Zanias, 1993		
For four products (for which different levels of political support exist), they run cointegration tests amongst all possible couples of countries for which data are available. Both the unrestricted and the restricted version of the LOP are tested.	Monthly prices from CRONOS, EUROSTAT, for soft wheat (1980:1-1990:12), cow's milk (1983:1-1990:12), pig carcasses grade I(1986:1-1990:12), potatoes (1983:4-1990:12), for a number of countries, depending on data availability (BE,DE, FR,IT, NL, UK). All prices are expressed in ECU.	The LOP holds in about half of the cases considered. Monetary Compensatory Amounts play an important role. For soft wheat, the existence of a minimum intervention price be very important for in market integration, but efficient arbitrage is also necessary.
SWITCHING REGIME-THRESHOLD MODELS		
Balcombe, 2007		
They develop a generalized Threshold Error Correction Model to test for the presence and behaviour of price transmission.	Monthly Brazilian, US and Argentine prices for wheat (1988-2001), maize (1986-2001), soybeans (1988-2001). All prices are used in logs and converted in US $.	Evidence of thresholds is found in three out of the five commodity pairs investigated.
Baulch, 1997		
He develops the Parity Bound Model (PBM), an alternative methodology in which information on transfer costs is used in addition to food prices. According to the size of price spreads	Monthly wholesale price (in levels) for Philippine rice coming from eight markets. January 1980-June 1993.	Monte Carlo experiments show that the PBM is statistically reliable. An application to the Philippine rice markets demonstrates that the PBM

and of transfer costs, three regimes are derived (within, at and outside the arbitrage band).		detects efficient arbitrage when other tests do not.
Goodwin, Piggot 2001		
They use a threshold model, in which regime switching is triggered when deviations in prices exceed the "neutral band" represented by transport costs. The evaluations are made pairwise.	Daily (in logs) corn and soybean prices in 4 North Carolina terminal markets. 2 January 1992-4 March 1999.	Thresholds turn out to be bigger when there is a bigger distance between markets. It is found that models that explicitly recognize the presence of thresholds effects imply faster adjustments to deviations from equilibrium than when threshold behaviour is ignored.
Negassa and Myers, 2007		
The standard PBM model is extended to allow for dynamic shifts in regime probabilities (the probability of being in a particular trading regime is not time invariant) in response to changes in marketing policy. This allows seeing whether changes in policies have increased or not spatial efficiency. It is applied to seven market, analyzed in pairs.	Weekly wholesale prices (in levels) converted into monthly prices for maize and wheat in eight Ethiopian markets. August 1996-August 2002.	The results highlight the importance of allowing for adjustment to policy changes.
Sexton et al., 1991		
Their estimation is based on a switching regression model with three regimes (the direction of trade flows is fixed, but arbitrage conditions might be violated): efficient arbitrage, relative shortage, relative glut. Only price data are needed.	Weekly prices (in levels) of US celery in Florida, California in 6 terminal markets. January 1985-December 1988.	The methodology allows for investigating market integration, arbitrage efficiency, magnitude of marketing margins, product substitutability, and competitiveness on markets.
RATIONAL EXPECTATIONS MODELS		
Goodwin et al., 1990		
Typical analyses of the LOP overlooks temporal elements of trade, and assume that parity should hold contemporaneously. They use two different approaches to test the LOP: GMM to estimate rationally formed expected future prices (expectations augmented version of the LOP; actual freight rates are used as a proxy for transport costs); nonparametric analysis of price parity using actual freight rates.	Monthly prices (levels) for 34 commodities from various countries, all converted in dollars. The period considered is from July 1973 to December 1985 (72 to 128 obs., depending on the price).	They conclude that using a simple-augmented expectations-model produces greater support for the LOP than using contemporaneous prices. Results provide strong support for a rational expectations version of the LOP. For the wheat market, adherence to the standard version of the LOP is limited (only two out of six markets support parity); for the expectations augmented version, the LOP is instead rejected only in one case.

ANNEX 2 : Unit root tests

Variable	n° of lags *logs (diff. logs)*	ADF statistics -18 (1978:12-2003:12) Logs	Differentiated logs
swfr			
n	*12(11)*	-0.042 (0.669)	-3.761 (0.0001)
c	*12(11)*	-1.823 (0.369)	-3.733 (0.004)
t	*12(11)*	-3.256 (0.074)	-3.606 (0.029)
c+s	*12(11)*	-1.737 (0.412)	-3.761 (0.003)
t+s	*12(11)*	-3.183 (0.088)	-3.637 (0.027)
hrw			
n	*1(0)*	0.172 (0.736)	-12.544 (0)
c	*1(0)*	-3.255 (0.017)	-12.528 (0)
t	*1(0)*	-3.325 (0.062)	-12.511 (0)
c+s	*1(0)*	-3.107 (0.026)	-12.477 (0)
t+s	*1(0)*	-3.183 (0.088)	-12.461 (0)
wref			
n	*12(12)*	-0.101 (0.649)	-5.149 (0)
c	*12(11)*	-2.163 (0.220)	-4.864 (0)
t	*12(11)*	-3.133 (0.099)	-4.878 (0)
c+s	*4(5)*	-2.301 (0.172)	-8.719 (0)
t+s	*1(5)*	-3.663 (0.025)	-8.720 (0)

n= no constant; *c*= constant; *t*=trend ; *s*= seasonal dummies included; p-values are based on MacKinnon (1996) and reported in parenthesis.

Variable	n° of lags *logs (diff. logs)*	ADF statistics -18 (1978:12-1993:06) Logs	Differentiated logs
swfr			
n	*12(11)*	0.668	-3.679
		(0.860)	(0.0001)
c	*17(11)*	-2.720	-3.737
		(0.070)	(0.003)
t	*17(11)*	-2.882	-3.740
		(0.168)	(0.019)
c+s	*12(11)*	-2.502	-3.025
		(0.115)	(0.032)
t+s	*12(11)*	-2.566	-2.988
		(0.296)	(0.135)
hrw			
n	*1(0)*	-0.072	-9.223
		(0.659)	(0)
c	*11(0)*	-2.328	-9.196
		(0.163)	(0)
t	*11(0)*	-2.984	-9.25
		(0.136)	(0)
c+s	*4(0)*	-2.319	-8.957
		(0.166)	(0)
t+s	*4(0)*	-3.061	-9.027
		(0.116)	(0)
wref			
n	*12(11)*	0.141	-3.461
		(0.726)	(0.0005)
c	*12(11)*	-1.886	-3.450
		(0.339)	(0.009)
t	*12(11)*	-1.860	-4.872
		(0.675)	(0.002)
c+s	*2(11)*	-1.902	-4.243
		(0.332)	(0.0005)
t+s	*12(11)*	-1.762	-4.721
		(0.723)	(0.0006)

n= no constant; *c*= constant; *t*=trend ; *s*= seasonal dummies included; *p-values are based on MacKinnon (1996) and reported in parenthesis.

Variable	n° of lags *logs (diff. logs)*	ADF statistics -18 (1993:07-2003:12) Logs	Differentiated logs
swfr			
n	*12(11)*	0.546	-1.926
		(0.8343)	(0.0516)
c	*12(11)*	-1.807	-1.808
		(0.377)	(0.3772)
t	*12(11)*	-2.245	-1.798
		(0.464)	(0.706)
c+s	*12(0)*	-1.574	-8.327
		(0.496)	(0)
t+s	*12(0)*	-2.113	-8.308
		(0.537)	(0)
hrw			
n	*2(1)*	0.272	-7.950
		(0.765)	(0)
c	*1(1)*	-3.081	-7.928
		(0.028)	(0)
t	*1(1)*	-3.102	-7.9023
		(0.106)	(0)
c+s	*1(1)*	-2.843	-7.269
		(0.052)	(0)
t+s	*1(1)*	-2.853	-7.246
		(0.178)	(0)
wref			
n	*2(1)*	0.282	-7.943
		(0.7678)	(0)
c	*1(1)*	-3.215	-7.922
		(0.019)	(0)
t	*1(1)*	-3.240	-7.898
		(0.076)	(0)
c+s	*1(1)*	-2.985	-7.186
		(0)	(0)
t+s	*1(1)*	-2.994	-7.166
		(0.134)	(0)

n= no constant; *c*= constant; *t*=trend ; *s*= seasonal dummies included; *p-values are based on MacKinnon (1996) and reported in parenthesis.

Variable	n° of lags *logs (diff. logs)*	5%critical values	PP statistics (1978:12-2003:12) Logs	Differentiated logs
swfr	*13(12)*			
n		-8.000	0.013	-180.846
		-1.950	0.107	-16.087
c		-14.000	-9.253	-180.937
		-2.878	-2.160	-16.044
t		-21.339	-24.587	-181.209
		-3.428	-3.503	-15.997
hrw	*2(1)*			
n		-8.000	0.049	-209.427
		-1.950	0.229	-12.582
c		-14.000	-14.941	-209.561
		-2.878	-2.814	-12.567
t		-21.339	-15.387	-209.646
		-3.428	-2.871	-12.550
wref	*2(1)*			
n		-8.000	-0.004	-227.208
		-1.950	-0.026	-13.740
c		-14.000	-16.898	-227.211
		-2.878	-2.935	-13.716
t		-21.339	-24.101	-227.232
		-3.428	-3.576	-13.694

Variable	n° of lags *logs (diff. logs)*	5%critical values	PP statistics (1978:12-1993:06) Logs	Differentiated logs
swfr	*13(12)*			
n		-7.949	0.036	-101.352
		-1.950	0.620	-15.626
c		-13.848	-25.548	-101.061
		-2.885	-3.883	-15.688
t		-20.996	-25.669	-100.736
		-3.440	-3.863	-15.706
hrw	*2(1)*			
n		-7.949	-0.006	-113.330
		-1.950	-0.041	-9.180
c		-13.848	-5.894	-113.333
		-2.885	-1.738	-9.152
t		-20.996	-8.467	-114.695
		-3.440	-2.338	-9.218
wref	*1(0)*			
n		-7.949	0.009	-152.693
		-1.950	0.117	-11.513
c		-13.848	-11.621	-152.729
		-2.885	-2.548	-11.479
t		-20.996	-11.161	-153.563
		-3.440	-2.432	-11.516

Variable	n° of lags *logs (diff. logs)*	5%critical values	PP statistics (1993:07-2003:12) Logs	Differentiated logs
swfr	*2(1)*			
n		-7.917	0.055	-93.872
		-1.950	0.503	-8.619
c		-13.750	-9.682	-94.210
		-2.888	-1.988	-8.607
t		-20.800	-13.062	-94.528
		-3.447	-2.082	-8.597
hrw	*2(1)*			
n		-7.917	0.043	-90.899
		-1.950	0.301	-8.267
c		-13.750	-13.225	-91.014
		-2.888	-2.641	-8.239
t		-20.800	-13.508	-91.024
		-3.447	-2.646	-8.207
wref	*2(1)*			
n		-7.917	0.044	-91.046
		-1.950	0.313	-8.729
c		-13.750	-14.284	-91.167
		-2.888	-2.759	-8.252
t		-20.800	-14.636	-91.185
		-3.447	-2.766	-8.220

Variable	n°of lags *logs (diff. logs)*	5% critical value	KPSS statistics (1978:12-2003:12)	
			Logs	**Differentiated logs**
swfr	*13(12)*			
c		*0.463*	1.679	0.106
hrw	*2(1)*			
c		*0.463*	0.743	0.061
wref	*2(1)*			
c		*0.463*	3.949	0.037

Variable	n°of lags *logs (diff. logs)*	5% critical value	KPSS statistics (1978:12-1993:06)	
			Logs	**Differentiated logs**
swfr	*13(12)*			
c		*0.463*	0.184	0.133
hrw	*2(1)*			
c		*0.463*	1.540	0.221
wref	*1(0)*			
c		*0.463*	1.650	0.154

Variable	n°of lags *logs (diff. logs)*	5% critical value	KPSS statistics (1993:07-2003:12)	
			Logs	**Differentiated logs**
swfr	*2(1)*			
c		*0.463*	1.939	0.151
hrw	*2(1)*			
c		*0.463*	0.428	0.048
wref	*2(1)*			
c		*0.463*	0.425	0.047