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TESTING INTERNATIONAL PRICE TRANSMISSION UNDER POLICY INTERVENTION. AN APPLICATION TO THE SOFT WHEAT MARKET

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PhD Studies

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PhD Studies Series: Volume 6, 2009

This work wouldn’t have been possible without the support of many people I have met in these years. Apart from their suggestions and advice, which I’m sure were essential for its continuous improvement, I would say that probably the biggest gift of this thesis was the possibility of meeting them, and of being honoured of their friendship.

I would then thank Julian Binfield, William Meyers, Wyatt Thompson and Patrick Westhoff of the Food and Agricultural Policy Research Institute (FAPRI), University of Missouri-Columbia, for rendering my stay in the US a fruitful and unique experience.

My acknowledgements also go to Frederic Chantreuil and to the whole AGMEMOD partnership for their interesting comments throughout the whole writing of this work; to Soeren Kissmeyer-Nielsen and Yves Plees for having assisted me during my internship at the EU Commission; to Rupert Somerscales of the International Grains Council for kindly (and timely) providing the data.

I am also grateful to Riccardo (Jack) Lucchetti for having introduced an Agricultural Sciences graduate student to the study of Econometrics, and on the other side to Angelo Frascarelli, for having suggested me to apply for a PhD programme in Economics.

Last, but not least, I would like to thank Roberto Esposti and Franco Sotte for their continuous assistance and advice. Working with them has been an extraordinary opportunity, human and scientific, and in both respects they really taught me more than I could either have imagined or write here.

And, finally, the biggest acknowledgments go to my boyfriend, to my family and to my friends, for their love and support.

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1 INTRODUCTION

The notion of "price transmission" refers to the co-movement shown by prices of the same good in different locations. In the early literature already, spatial price analysis was devoted to defining markets (Fackler and Goodwin 2001, p.973): an economic market is the spatial area "*within which the price of a good tends toward uniformity, allowance being made for transportation costs*" (Stigler 1966 cited Fackler and Goodwin 2001, p.974).

When spatially separated markets are considered, price transmission analyses play a crucial role in trying to assess how efficiently integrated they are, i.e. to which extent rational arbitrage operates. Most works essentially aim at verifying whether the Law of One Price (LOP) holds between different markets; this Law simply states that homogeneous goods in spatially separated locations will have a unique price, when expressed in the same currency, net of transport costs.

International price transmission is an issue which has received considerable attention, and different econometric techniques have been used in this respect. In particular, cointegration models have shown to have an "intuitive appeal", as they consent to disentangle short and long run dynamics. Prices are assumed to be linked in the long run despite being allowed to diverge from each other in the short term.

Nonetheless, despite the use of increasingly sophisticated techniques, evidence supporting the LOP is weak, and empirical findings are extremely mixed.

This is not surprising, given that there is wide acknowledgment that the hypothesis of the LOP are quite unlikely to hold in practice: the LOP is simply expected to regulate spatial price relations in a "*frictionless undistorted world"* (Conforti 2004, p.1), while, in fact, a number of factors are expected to prevent prices from convergence.

This work, in particular, will focus on how domestic and border policies affect international price transmission.

In this respect, (often) highly policy-regulated agricultural commodities markets are extremely interesting for the analysis. Indeed, traditionally, agricultural commodities' markets, especially in developed countries, have been

characterized by a high degree of policy intervention, whose motivations deeply hinge on some inherent characters of agricultural production.

First of all, agriculture produces food, which is of strategic importance. Moreover, the agricultural production is characterized by a high degree of uncertainty, being dependant on natural production factors.

Secondly, agricultural enterprises are small if compared to the size of the market, and their structural adjustment process is normally slow and constrained by the rigid supply of some of the production factors.

Thirdly, agricultural markets are often characterized by a high degree of competition, and both market demand and supply are typically inelastic, rendering both producers' yields and consumers' expenditures extremely volatile (Hallett 1968).

In addition to these, and principally in developed countries, other reasons have been added to justify the presence of policy subsidies, concerning the specific functions agriculture performs in the rural environment (the so called "multifunctional" character of agriculture[1]).

However, in the past months, it has been the dramatic increase that commodities prices have experienced on international markets that drew the attention of both the media and of economic analysts on the agricultural sector.

The rise in food commodity prices is due to a number of factors. The increasing demand of the emerging economies, especially for livestock products, the rising biofuels demand, and, on the supply side, the steady increase in energy prices can be classified as structural factors, whereas, as short term factors, we can mention the adverse meteorological conditions occurring during last year.

The sudden increase in agricultural prices, together with the high levels of volatility they experienced, brought, if possible, an even more renewed interest in the policy debate about the reform agricultural policies should undergo to tackle an evolving global context, in which supply and demand forces are characterized by major interdependences.

In this context, the analysis of how prices are transmitted, and of how intervention policies are likely to affect such mechanisms, is crucial. Understanding how price shocks are transmitted between different countries, and how such transmission has been affected, in the past, by policy regimes, is a prerequisite for policy analysis and future intervention. For the European Union (EU), this basically means analyzing to which extent, provided the presence of its regulatory framework, the world price has affected and has interacted with the domestic one.

[1] "Beyond its primary function of supplying food and fibre, agricultural activity can also shape the landscape, provide environmental benefits such as land conservation, the sustainable management of renewable natural resources and the preservation of biodiversity, and contribute to the socio-economic viability of many rural areas" (OECD 2001).

These recent phenomena occurring on world markets, together with the controversial empirical evidence mentioned before, make the analysis of the effects of policy intervention on price transmission a challenging issue.

However, in economic analyses, while policy variables are explicitly considered in econometric models, this is often not the case in empirical works.

In fact, in literature, even if price transmission mechanisms for agricultural commodities have received considerable attention, analyses often use commodity prices only. This has implicated the use of increasingly sophisticated econometric techniques, and frequently not appropriate concern for policy factors. In econometric models, on the other side, explicit use of policy variables is made, but some of them disregard the time series properties of the data and rely on simplistic assumptions (which is basically due to their dimension and inherent complexity), like the exogeneity of the world prices for the EU in the AGMEMOD model (AGMEMOD Partnership 2007a; 2007b).

The objective of this work is to bridge this gap, that is, to include policy considerations in a consistent theoretical framework and to provide empirical applications.

This work studies price transmission mechanisms between EU an international soft wheat markets in the years 1978-2003. While the domestic French price is assumed to represent the European one, the Unites States' (US) price is taken as representative of the world one. This commodity is the most traded grain, internationally. Domestic and border policies of the major producer and consumer countries are expected to have a deep influence on world markets (Ghoshray *et al.* 2000, p.3).

Empirical evidence concerning the existence and the evolution of price transmission for soft wheat between the US and the EU is extremely mixed. Either no evidence of co-movement is found, or it turns out to be low, and in this case being either affected or not by international trade policy reform.

The object of this analysis is then twofold: on the one side, developing a consistent theoretical framework which allows to understand under which conditions co-movement between commodity prices in international markets is expected, after considering the presence of domestic and border policies: namely, the European Common Agricultural Policy (CAP) and the implementation of the Uruguay Round Agreement on Agriculture (URAA). The first is expected to have lowered the distance between EU and world prices, whereas the latter to have increased price transmission elasticities by ensuring a higher degree of liberalization.

On the other, from this framework different econometric models, all making use of cointegration techniques, are subsequently derived, in order to test whether such a co-movement exists, and to which extent it has been affected by policy reforms.

This work is organized as follows. Chapter 2 provides a general introduction to the study of price transmission. The basic concepts and definitions are presented, followed by some general considerations concerning all factors that affect the possibility of the LOP to hold in practice.

In chapter 3, the econometric techniques used for the study of price transmission are revised. Cointegration models are explained in detail, as these will be used in the empirical sections of this work. Some observations about the properties of the models relying on price data only conclude the chapter.

Chapter 4 aims at providing a general overview of international soft wheat markets. The institutional framework is presented, and the relevant international and EU policies for soft wheat are described.

This allows to develop a theoretical framework which aims at understanding under which policy regime conditions co-movement between EU and world prices is actually expected, provided the presence of the CAP measures regulating the EU market, and of international trade regulations. Once the various combinations of market prices and policy regimes are analyzed, it is argued that it is the relative position of the intervention price, which is the basic EU instrument to regulate markets, and of the world price that allows to distinguish between the different policy regimes in place.

The description of the dataset concludes chapter 4, together with considerations on issues related to the price leadership on world wheat markets, product homogeneity, transport costs assumptions.

Subsequently, the models presented in chapters 5 and 6 allow to test through empirical applications the theoretical framework developed.

All models presented in chapter 5 share the common characteristic of the use of a policy variable, the intervention price, together with monthly EU and US prices, to account for policy regime changes. The models presented represent an original attempt of combining policy and price data, and all yield consistent results: the EU and the US price interaction can be adequately assessed only when policy regimes are explicitly considered.

However, some interpretative problems remain. For this reason, in chapter 6, which concludes this work, a final model is presented to explore the relation between just the EU and the US price. This time, policy regime changes are modelled by using cointegration techniques allowing for structural breaks, which have only recently been applied to agricultural commodities markets (the underlying assumption being that the two prices have always been linked despite domestic and border policies, which are on the other side assumed to have affected this relation).

Once again, empirical evidence confirms the presence of co-movement between the two prices only when policy regimes are appropriately considered.

Whereas the Common Agricultural Policy reforms have reduced the distance between domestic and international prices, liberalization reforms appear to have positively influenced price transmission.

2 PRICE TRANSMISSION AND THE LAW OF ONE PRICE

2.1 Introduction

This chapter aims at providing a general overview of the theoretical framework for the study of price transmission for a certain good between spatial separate locations[2]. This topic has received considerable attention, and this is true especially for agricultural commodities. These are characterized by being produced over an extensive spatial area, and costly to transport relative to their total value.

Spatial price analysis allows to gain insights into the working of markets and to evaluate their performances (Fackler and Goodwin 2001, p.973). Indeed, market integration affects economic growth, induces structural change and alters the location of economic activities (Vollrath and Hallahan 2006, p.56). Potential losses and benefits crucially depend upon markets receiving price signals; this is why the topic has attracted so much empirical work (Conforti 2004, p.1). In addition to this, this issue gains probably nowadays particular relevance considering the dramatic increase, due to both contingent (adverse meteorological conditions, speculations) and structural factors (the increase of the demand originating from emerging economies, the biofuels demand), that food commodity prices have experienced since 2007. Understanding how prices are transmitted between different countries is a challenging issue, and a prerequisite for policy intervention.

This chapter has the objective of providing the basic theoretical concepts relevant to the study of price transmission, in order to fully understand how empirical models have been developed in the existing literature, and to evaluate the innovations that will be introduced and tested in the remainder of this work.

[2] This work then focuses on "horizontal" price transmission, whereas "vertical" price transmission refers to the transmission of prices between producers, intermediaries and consumers (see for example Meyer and von Cramon-Taubadel 2004).

The fundamental definitions are given (paragraph 2.2), underlying the relations existing between them, and the appropriateness of their use in the analysis. In paragraph 2.3, some major problems due to the reliability of the hypothesis at the basis of the LOP to hold in practice are described, which is strongly interlinked with a first assessment of the problems inherent to empirical models relying only on price data (which will be extensively dealt with in chapter 3). Paragraph 2.4 concludes.

2.2 Some basic concepts

The study of price transmission mechanisms implies referring to a number of economic concepts for which, unfortunately, no common definitions exist in literature (Fackler and Goodwin 2001, p.976). The most important ones will here be briefly revised.

The basic notion is the *spatial arbitrage condition*, which can be formalized as in equation (2.1),

$$P_j - P_i \leq R_{ij} \tag{2.1}$$

where P indicates the prices in the two spatially separated locations i and j, and R_{ij} is the cost of moving the good considered from i to j. The spatial arbitrage condition implies that the difference between prices in different locations will never exceed transport costs[3], or otherwise the profiting opportunities would be immediately exploited by arbitrageurs: they would buy the goods in the market in which the price is lower, and sell them in the one where it is higher. Other things equal, the price would then go up in the first market due to the increased demand, and go down in the second one because of the increase of supply, towards equilibrium. In the short run, actual prices may diverge from the spatial arbitrage condition, but the actions of the arbitrageurs are expected to make it valid in the long run, moving the price spread toward the transport cost[4]. More realistically, explicit arbitrage for nowadays markets is the outcome of the actions of a number of dispersed producers who evaluate the market conditions existing in several terminal markets and sell in the market with the highest price. Net returns and prices are equalized across markets; prices are expected to differ by no more than the difference in the costs of selling in one market versus another (Goodwin and Piggott 2001, p.302).

[3] In this paragraph, for the sake of simplicity, the term transport cost is meant to include all costs of arranging transactions between spatially different locations. See paragraph 2.3 for a more detailed explanation.

[4] In this respect, Stigler defines the market as the spatial area "*within which the price of a good tends toward uniformity, allowance being made for transportation costs*" (1966 cited Fackler and Goodwin 2001, p.974).

The *Law of One Price* (LOP) directly follows from the spatial arbitrage condition: in markets linked by trade and arbitrage, prices expressed in the same currency will be equalized, net of transport costs. The LOP is based on international commodity arbitrage, implying that "*in the assumed absence of transport costs and trade restrictions, perfect commodity arbitrage insures that each good is uniformly priced (in common currency units) throughout the world* " (Isard 1977, p.942)[5]. Some Authors distinguish between a *weak* and a *strong* version of the LOP: the former is the same as the arbitrage condition, while the latter is the relation (2.1) taken with an equality sign. Goodwin *et al.* (1999, p.192) provide a comprehensive definition of the LOP. The economic agent is assumed to have the opportunity to sell in two different markets[6], and to aim at maximizing his profit, *V*, which is given by:

$$V(q_{1t}, q_{2t}) = \sum_t \left\{ \delta^t \left[P_{1t+k} q_{1t} + P_{2t+j} q_{2t} - C(Q_t) - \tau_1 q_{1t} - \tau_2 q_{2t} \right] \right\} \tag{2.2}$$

where q_{it} is the quantity sold in market i ($i = 1,2$) at time t, $P_{i,t+k}$ is the price received upon delivery in market i (at k and j periods after t), $C(\cdot)$ is a general arbitrage cost function, Q_t is the total traded quantity ($Q_t = q_{1t} + q_{2t}$), δ is the real discount factor, and τ_{it} is the per unit transaction cost. For all $s > t$, the first order conditions are given by:

$$E_{t+s} = \left\{ \delta^s \left[\delta^k P_{1t+k+s} - C'(Q_{t+s}) - \tau_{1t+s} \right] \right\} = 0 \tag{2.3}$$

$$E_{t+s} = \left\{ \delta^s \left[\delta^j P_{2t+j+s} - C'(Q_{t+s}) - \tau_{2t+s} \right] \right\} = 0 \tag{2.4}$$

that can be arranged to yield

$$E_t = \left\{ \delta^k P_{1t+k} - \tau_{1t} - \delta^j P_{2t+j} + \tau_{2t} \right\} = 0 \tag{2.5}$$

Supposing there are no delivery lags (i.e. $k = j = 0$), this relation implies:

$$P_{1t} = P_{2t} + (\tau_{1t} - \tau_{2t}) \tag{2.6}$$

[5] Asche *et al*. (1999), note that the LOP is closely related to the Composite Commodity theorem of Hicks and Leontief. For both conditions to hold, in fact, prices must move proportionally over time.

[6] Storage is not explicitly considered, but may be added to the model without loss of generality; additional arbitrage conditions would exist among expected prices, storage costs, and transportation costs.

i.e., the LOP. [7; 8]

The LOP is at the basis of the Purchasing Power Parity (PPP), which can be considered as its aggregate version stated in terms of price indices, and whose assumptions are thus far more restrictive.

The LOP, whose the very "law" designation reflects the faith placed in its adherence (Fackler and Goodwin 2001, p.977) has a very long tradition in economics which dates back to Marshall (1890 cited Fackler and Goodwin 2001, p.977) but, nevertheless, most of the empirical tests are against it. Miljkovic (1999, p.126), asserts that "*although called a law, it has probably been violated probably more than any other economic laws (on the basis of the results of numerous empirical studies)*". A detailed description of why this might be the case will be given in paragraphs 2.3 and 3.4, as it might depend both on the strong assumptions underpinning it and on the inherent features of the empirical models used.

Another basic concepts is the one of *competitive market equilibrium*, that is a condition in which extraordinary profits are exhausted by competitive pressures, regardless of whether this results in physical trade flows between markets (Barrett and Li 2002, p.394)[9]. Analogously, the *spatial market efficiency* can be taken as a synonymous of the spatial arbitrage condition. Regional or interregional markets characterised by arbitrage opportunities can be considered as inefficient, since markets should produce prices that accurately reflect all available information about demand and supply conditions as well as transaction costs. Market efficiency implies that, in a competitive market with perfect information, arbitrage will ensure that price differentials in related spatial and temporal markets will reflect all marketing costs. In spatially related markets, price differentials will reflect transportation costs; in temporally related markets, the cost of storage (Hui-Shung and Griffith 1998, p.369). However, in developing countries, the concept of spatial market efficiency may also encompass an assessment of the size of the transaction costs (Fackler and Goodwin 2001, p. 980).

For the concept of *spatial market integration*, instead, Barrett and Li (2002, p.292) explicitly make clear that it just reflects the tradability of products between spatially distinct markets, irrespective of the presence or absence of spatial market equilibrium and efficiency. In other words, market integration is a quantity based indicator of tradability, while, as explained, efficiency is a price-based indicator,

[7] If transaction costs are a constant proportion of prices ($\tau_{1t} = \lambda_{1t}P_{1t}$), then the first order conditions imply $P_{1t} =(1- \lambda_2)/(1-\lambda_1) P_{2t} = \beta_0 P_{2t}$, which, as explained in the following paragraph, is a multiplicative version of the LOP which can be put in a linear form by taking the logarithms of the prices.

[8] If delivery lags are different, for example $k = 0$ and $j > 0$, it becomes $P_{1t} = \beta_0 E_t\{P_{2t+j}\}$; see paragraph 3.2.4.

[9] Markets may maintain spatial linkages even when direct trade doesn't occur between them. This happens if selling agents from both markets compete in a third one and perfect competition holds in all three markets. Even in the absence of trade, spatial market linkages might be disciplined under the threat of competition; prices differences are kept between the cost band (Goodwin *et al*. 1999, p.160).

based on economics concepts of equilibrium (Thompson and Bohl 2002, p.1042). Nevertheless, Fackler and Goodwin (2001, p.978) report that for the concept of spatial market integration no satisfactory definition exists: this has lead to some inconsistencies and confusions. Indeed, the term has been used loosely in the existing literature: to indicate, in general, the degree of co-movement shown by prices in different locations, and, also, the degree at which price shocks are transmitted among spatially separated markets (Goodwin and Piggott 2001, p.302). In this respect, Fackler and Goodwin (2001, p.978) propose the following indicator:

$$RA_{AB} = \frac{\delta P_B}{\delta \varepsilon_A} \Big/ \frac{\delta P_A}{\delta \varepsilon_A} \tag{2.7}$$

where ε_A is a shock occurring on market *A*, and RA_{AB} is the price transmission ratio between region *A* and region *B*. If the two markets are perfectly integrated, it should show a value of 1. This ratio, which could also not be symmetric between the two regions, allows to study how shocks are transmitted from one market to the other. The underlying idea is that if price shocks are transmitted, then trade exists, and the markets are integrated. Anyway, a number of objections can be raised, underlying the confusion arising when market integration is measured without using trade flows data. Indeed, also if two regions are not part of a common trading network, there could be price co-movement because of factors that have nothing to do with commercial integration: prices can share the same evolving pattern because of seasonality, common macroeconomic shocks, collusive behaviour and so on[10]. Indeed, it should be noticed that the indicator proposed by Fackler and Goodwin still relies on the use of price data only, which is the case for the overwhelming majority of the models which will be revised in the next chapter. Instead, market integration can be adequately assessed only if trade data are available (Barrett and Li 2002).

A final observation concerns the fact that, since the LOP can hold even if the price transmission ratio between two regions is less than one[11], the following hierarchy holds for the concepts defined so far (Fackler and Goodwin 2001, p.979; where the notion of "market integration" refers to equation 2.7):

$$\textit{Perfect market integration} \Rightarrow \textit{LOP strong} \Rightarrow \textit{LOP weak} \tag{2.8}$$

[10] This might happen also for unrelated commodities. Pindyck and Rotemberg (1990 cited Tomek and Myers 1993, p.89) note that the co-movement of prices has three main explanations: supply and demand shocks in one market may spill over into other markets; macroeconomic shocks can affect all prices together; speculators overreact to new information, and this causes spillovers between commodity markets.

[11] Shocks have generally a bigger effect in the originating region than in the other one (Fackler and Goodwin 2001, p. 988).

2.3 A theoretical problem: why should the LOP hold?

The previous paragraph provides a general outline of the fundamental theoretical concepts relative to price transmission analysis. As it will be clearer in chapter 3, where the empirical models which have been used for the study of price transmission are presented, what emerges is a controversial picture. Indeed, not only are even the basic definitions not fully agreed upon researchers, but empirical analysis has brought a very mixed evidence (see annex A for a review).

In this paragraph, some concluding remarks concerning the reliability of the LOP hypothesis to hold in practice will be given, which might be the cause of the absence of empirical evidence supporting it. This issue is nonetheless strictly entangled, as we will see in chapter 3, to the underlying hypothesis and properties of the models used, and this is particularly true if they rely only on price data (paragraph 3.4).

It is well recognized that, though denoting this concept a "law" reflects the considerable faith in its adherence (Fackler and Goodwin 2001, p.977), the simplicity of the LOP can be easily questioned, as its assumptions prove to be quite restrictive and unlikely to hold in practice. Williamson (1986 cited Fackler and Goodwin 2001, p.975) and Miljkovic (1999, p.126) note that the LOP has been violated by empirical tests probably more than any other economic laws.

The LOP is expected to regulate spatial price relations in a frictionless undistorted world; the premises of full price transmission and market integration correspond to those of the standard competition model (Conforti 2004, p.1). Following Miljkovic (1999) and Conforti (2004), some major groups of factors that prevent prices from convergence can be identified[12].

Transaction costs play a crucial role while investigating market efficiency, and this is particularly important in agriculture, since they are relevant if compared to the value of the commodities considered (Fackler and Goodwin 2001, p.973; Barrett 2001, p.27). Unless certain assumptions are made, their treatment is not easy; and, even if the LOP is satisfied, if transaction costs are large and volatile, prices don't move together. Transaction costs have many components. In addition to the *very* transport costs (normally, freight rates, i.e. per unit transportation costs), we also have variable transport costs (insurance, financing, hedging, contracting, technical barriers to trade), exogenous costs components (underwriting fees, testing charges), unit average duties, and other immeasurable transaction costs (such as search costs and exchange rate variability; Barrett 2001, p.23); in other words, all other information, negotiation, monitoring and enforcement costs (Williamson cited Conforti 2004, p.2). Linnemann (1996 cited

[12] In general, the conditions necessary for adherence to factor price equalization are even stronger than those required for output markets; evidence supporting factor market integration is weak (Fackler and Goodwin 2001, p. 976).

Miljkovic 1999, p.135), concludes that the effect of distance on trade consists of three components: transportation costs, time element (perishability, adaptability to market conditions, irregularities in supply in addition to interest costs), and "psychic" distance, i.e. familiarity with laws, institutions and habits. Both "distance" and "border" matter: some studies have shown that price differentials across countries for very similar goods are usually more volatile than price differentials within a country for very dissimilar goods (Engle and Rogers 1996, McCallum 1995, Mussa 1986, all cited Miljkovic 1999, p.128-129). This could be explained by the nominal stickiness of prices when expressed in national currencies, and by the relatively high volatility of nominal exchange rates (cross-border prices would fluctuate along with the exchange rate, but within country prices would be fairly stable), but also by the homogeneity of productivity shocks in the non-tradable factors of productions and differences in the demand elasticities between different countries (Milikovic 1999, p.129).

Border and domestic policies play a role. Policies aimed at insulating internal commodity prices will not allow changes in the world ones to be transmitted to the internal prices (Bredahl *et al.* 1979). Variable levies, export subsidies, non tariff barriers, tariff rate quotas, prohibitive tariffs, technical barriers to trade are expected to prevent prices from convergence, whereas ad valorem and fixed tariffs should behave as proportional and fixed transaction costs[13]. Even the very existence of specific trade agreements that create different trading blocks with different degrees of market integration can prevent domestic prices from convergence in different countries. According to Miljkovic (1999, p.136), art. XXIV of the General Agreement of Tariffs and Trade (GATT), which allows the existence of non multilateral Free Trade Agreements, sets the institutional support for violating the very existence of the LOP. The importance of domestic and border policies in affecting price transmission will be further discussed in chapters 3 and 4.

Market power can exist along production chains: depending on the degree of concentration of each industry, some agents may behave as price makers while some others as price takers. The presence of i*ncreased returns to scale in production* might be at the origin of market power. Miljkovic (1999, p.134) underlines that export demand elasticities vary by country, and this provides the impetus for the differential pricing of exports as opposed to price taking (we have pricing to market behaviour, i.e., the ability of exporting firms to discriminate prices across destinations).

Prices are not expected to converge also if *product homogeneity and differentiation* influence the substitutability in consumption.

[13] Mundlak and Larson (1992, p. 405) propose the explicit inclusion of stabilization policies in a linear model of the LOP. They are expected to lower the price transmission coefficient.

Exchange rates "pass-through" on output prices, which has been studied in relation to pricing to market behaviour, and *exchange rate risks* do have an effect on export prices (Miljkovic 1999, p.134-135).

Finally, *imperfect flows of information* can rise the cost of arbitraging.

Moreover, price transmission might also be characterized by asymmetry, i.e. it differs according to whether prices are increasing or decreasing (in magnitude, speed or both). Price transmission asymmetry implies a different distribution of welfare than what we would have under symmetry and might be a manifestation of market failure (for a survey see Meyer and von Craumon-Taubadel 2004). Asymmetric adjustment costs, asymmetric information and, of course, market power might be at the origin of horizontal price transmission asymmetry (see paragraph 3.3.2.2[14]).

At this point, it should be clear that distance and transportation costs, after adjusting for exchange rates, don't account for international price variability. Miljkovic (1999, p.130) notes that this is implicitly recognized in Stigler and Sherwin (1985) already: when claiming that a product's prices "tend" to move together, they are aware of the existence of multiple prices for a good in the same market. "*Because it is recognized that international markets are (generally) far from perfect, it is not clear why imposing the LOP in international agricultural trade modelling is so critical*" (Miljkovic 1999, p.130). Moreover, a number of the studies aiming at testing the LOP have national or at most interregional dimension (Baulch 1997; Sexton *et al.* 1991).

In most empirical models (see chapter 3), all sources of deviations from the LOP not explicitly considered amongst the regressors turn out to be included in the error term. This, in turn, implies strong assumptions on their behaviour. For example, we can anticipate that, in cointegration models, which will be extensively dealt with in the remainder of this work, they are assumed to be stationary (if the model is expressed in logarithms, stationary around a constant proportion of prices).

An important consequence is that, being the hypothesis at the basis of the LOP very strict, a transmission parameter will summarize the overall effect of a set of factors affecting price signals (transaction costs, the existence of market power and so on). The fact that some of this elements change proportionally with the prices while some others directly impact on price spreads, and that their effects are likely to interact with each other, further complicate the analysis. Nevertheless, still the value of the parameters and their significance level provides information about the extent to which markets share the same price shocks (Conforti 2004, p.5). At the same time, knowledge of the institutions and

[14] In agricultural economics, asymmetric price transmission is an issue extensively dealt with in vertical price transmission analyses (see for example Meyer and von Craumon-Taubadel 2004).

conditions pertinent to the markets in question will be necessary to properly interpret the results of empirical evaluations (Fackler and Goodwin 2001, p.1017).

2.4 Concluding remarks

Perhaps, the most surprising observation that emerges from a general outline of the concepts necessary to study price transmission mechanisms is that, unfortunately, sometimes no common definitions exist. In particular, only the most recent literature has shed light on the distinction between market integration and efficiency (Barrett and Li 2002, p.292), whereas various inconsistencies in some of the previous works remain.

However, when the theoretical definitions presented are translated into empirical models, most analyses basically aim at verifying the validity of the LOP between spatially separated markets by using price data only (as it will be extensively explained in chapter 3).

However, despite an abundant literature (see annex A for a review), the empirical evidence is nonetheless mixed. While, on the one side, when empirical tests rely only on price data, as it is often the case, it is just not possible to test separately the efficiency hypothesis and the strong assumptions underpinning the model specifications (Barrett 2001, p.29; see chapter 3), on the other, as a matter of fact the hypothesis needed for the LOP to hold are quite restrictive and unlikely to hold in practice, since many factors can prevent prices from convergence. These factors have been here shortly enlisted and described.

Clearly, domestic and border regulation policies play a fundamental role in influencing price transmission and market integration (Conforti 2004, p.2). In this respect, (often) highly policy-regulated agricultural commodities markets are a case in point: the political intervention to which agriculture is subject creates a gap between world and domestic prices, which generates cross-country variations in agricultural prices (Mundlak and Larson 1992, p.399).

The aim of this work is then to provide a consistent theoretical framework to consider policy regimes while testing for price transmission.

3 EMPIRICAL TESTS FOR SPATIAL PRICE ANALYSIS

3.1 Introduction

While chapter 2 provides a general outline of the fundamental theoretical framework underlying price transmission analysis, this chapter will focus instead on the econometric techniques which have been used so far in empirical applications (a review of the economic models of price determination, also surveyed in Fackler and Goodwin, 2001, is on the other side beyond the objectives of this work)[15].

A fundamental important premise is that, as we will see, prices are often the only data available to examine spatial relationships. Accordingly, increasingly sophisticated econometric devices have been developed. Barrett (1996, p.825) notes that "*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*". In much of the existing literature, indeed, the analysis typically relies on price data only, and focus on the special case of "*perfect integration, when two markets are both integrated and in competitive equilibrium. Yet actual market relationships are messy*" (Barrett and Li 2002, p.294). Clearly, in order to be informative, market (price) analysis will also have to be accompanied by a thorough study of the conditions in which trade takes place.

Amongst all empirical models used, time series analysis and in particular cointegration techniques have been considered as very appealing, as they allow to disentangle short and long run market dynamics and to remove the exogeneity hypothesis of one of the prices.

[15] Also, a parallel literature has emerged in "pricing to market" (PTM) models. The canonical PTM model compares fob prices from a single source country to multiple destination markets (Barrett 2001, p.25).

However, despite all the different models developed, empirical evidence is extremely mixed.

This chapter is structured as follows. In paragraph 3.2, the most important econometrical approaches used hitherto to investigate price transmission mechanisms are revised: simple regression and cointegration models, dynamic regression models, switching regime models and rational expectation models. Paragraph 3.3 is entirely devoted to cointegration techniques, as these will be used in the empirical section of this work: issues related to unit root testing when applied to price series, the various cointegration techniques developed so far and the empirical evidence are presented. Finally, in paragraph 3.4, the common drawbacks of all empirical models relying only on price data are explicated. This is in a certain way complementary to paragraph 2.3, since, concerning the recurrent violations of the LOP found by empirical works, problems related to its strict assumptions are strictly intertwined with the characteristics of the model specifications. Paragraph 3.5 concludes.

3.2 Empirical tests: a general framework

In this section, following Fackler and Goodwin (2001), the existing literature concerning price transmission mechanisms will be reviewed according to the various econometric tools used:

- simple regression and correlation analysis;
- dynamic regression models based on a point location model;
- switching regime models;
- rational expectations models.

As soon as the corresponding econometric techniques were developed, they were used for the study of price transmission mechanisms. For each of them, a brief description of the approach used is given, underlying its main advantages and drawbacks.

3.2.1 Simple regression and correlation analysis

The one relying on simple regression and correlation analysis is the oldest approach used for the study of price transmission. Correlation and regression-based spatial market integration[16] analysis are nearly identical in terms of the

[16] As explained in paragraph 2.2, especially the early literature did not effectively distinguish amongst efficiency and integration (Barrett and Li 2001, p. 292). For this reason, in the remainder of the discussion, the word “integration” is used following Fackler and Goodwin (2001), and refers to the transmission of price shocks between markets rather than to the actual presence of trade.

mechanisms used to develop empirical tests, though the interpretation of the results is different[17].

The basic model is the following:

$$P_{1t} = \beta_0 + \beta_1 P_{2t} + \beta_2 T_t + \varepsilon_t \tag{3.1}$$

where P_t are the prices in locations 1 and 2 at time t, T represents transaction costs and ε is the error term. Arbitrage conditions are assumed to hold instantaneously (as no lags are included in the model). Markets are taken to be perfectly integrated if

$$\beta_1 = \beta_2 = 1 \tag{3.2}$$

$$\text{and } \beta_0 = 0 \tag{3.3}$$

These models can be evaluated both in levels or in logarithmic form. In the first case, the coefficient of the price term represents the marginal effect of the change of one price to the other; in the second one, this coefficient represents the price transmission elasticity. In both cases, a value of 1 is assumed for perfectly integrated markets[18]. In most of the revised literature price transmission equations are normally written in logarithmic form, i.e.

$$p_{1t} = \beta_0 + \beta_1 p_{2t} + \beta_2 t_t + \varepsilon_t \tag{3.4}$$

the underlying equation in levels being

$$P_{1t} = e^{\beta_0} P_{2t}^{\beta_1} T_t^{\beta_2} e^{\varepsilon_t} \tag{3.5}$$

where $p = log\ P$ and $t = log\ T$. β_1 is the elasticity of price transmission.

[17] Given two prices, P_1 and P_2, the correlation coefficient is given by $\rho_{12} = \sigma_{12}/(\sigma_1\sigma_2)$, where σ_{12} is the covariance between the two prices and σ_i is the standard deviation of price i. In a regression written in the general form $P_1 = \alpha + \beta P_2$, the least square estimate of β is σ_{12}/σ_{22}, where σ_{22} is the variance of P_2. Since $\beta = \rho_{12}(\sigma_1/\sigma_2)$, β and ρ_{12} are proportional and of the same sign.

[18] Sharma (2002) notes that, despite the complete pass-through of prices in absolute terms (what he calls "absolute price transmission"), the price transmission in proportional terms (or "proportional" price transmission, i.e. the percentage change in domestic *-1-* price divided by percentage change in world *-2-* price) is smaller than one in the import case and higher than one in the export case. This effect is larger the bigger the constant term; only if the constant is zero, the absolute and the proportional price transmission will be equal. In this work, following Fackler and Goodwin (2001), when talking about price elasticities we will refer to absolute price transmission.

Usually, it is assumed that $\beta_2 = 0$ and the assumption that $\beta_0 = 0$ is, in turn, relaxed, so that this term becomes a crude mean to represent transaction costs (or, generally, all other factors contributing to price differentials which are not included amongst the regressors; see paragraph 2.3, and, for an example, Thompson *et al.* 2002a, p.1043). In this way, the strong assumption which is implicitly made in all linear regression models is that all factors possibly contributing to price differentials but not taken into account in the model are fixed (if a linear specification is used; or a constant proportion of prices, when variables are expressed in logarithms[19]).

Moreover, in linear regression models, it is clear that either one of the two prices must be assumed to be exogenous (Richardson 1978), or otherwise simultaneity problems can be addressed with Instrumental Variables techniques (Goodwin *et al.* 1990).

Numerous different versions of the regression model can be developed. For example, Stigler and Sherwin (1985) use simple price correlation analysis amongst prices to establish their belonging to the same market. In Richardson (1978), log differenced prices are employed in a linear regression model. If the role of the exchange rate, *e* (defined as the ratio of the domestic -country *1*- versus foreign -country *2*- currency) is made explicit in the logarithmic form of the LOP, we have:

$$p_{1t} = \beta_0 + \beta_1 p_{2t} + \beta_2 t_t + \beta_3 e_t + \varepsilon_t \tag{3.6}$$

where *e* is expected to have a positive sign[20]. The underlying assumption is that the level of disaggregation allows the exchange rate to be taken as exogenous. Anyway, most commonly, commodity prices are directly expressed in the same currency[21].

A first shortcoming of this approach, that, as we will see, is common to all models relying on price information only, is that unknown common components could explain price co-movement regardless of the extent to which markets are linked. Moreover, in theory, any value of the correlation coefficient is consistent with integrated markets, depending on the size of transaction costs. A final, fundamental flaw is that what is here tested is the strong form of the LOP, and not the validity of the arbitrage condition (which, differently from the first, would also

[19] This hypothesis would seem to be quite reliable for the data used in this analysis; see figure 4.11.

[20] Equation (3.6) can be interpreted as an export pricing equation for the home country: a lower *e* (appreciation of the home country currency) reduces foreign sales unless the exporter lowers its price (p_1 decreases). Alternatively, viewing it as an import pricing equation for the home country, a higher *e* (depreciation of the home country) makes imports more expensive, allowing domestic producers to raise their prices (p_1 increases) (Vollrath and Hallahan 2006, p.61).

[21] When not, results are generally not altered (Thompson 1999; Bukenya and Labys 2005).

hold with an inequality sign). The importance of this flaw can be assessed only verifying the frequency at which price differences are within the band of the transport costs or exceed it.

3.2.2 Dynamic regression models based on a point location model

Dynamic models gained attention because price linkages might be of a non-contemporaneous nature. All dynamic regression models basically refer to the dynamic time-series properties of the data, using some version of a Vector Autoregression (VAR) model:

$$\mathbf{A_0 P_t} = \sum_{k=1}^{n} \mathbf{A_k P_{t-k}} + \mathbf{D X_t} + \mathbf{e}_t \tag{3.7}$$

where $\boldsymbol{P_t}$ is a (n X 1) vector of prices, $\boldsymbol{X_t}$ is a (n X 1) vector of exogenous factors, the $\boldsymbol{A_k}$ are the (n X n) matrixes of coefficients of the *k-th* included lagged prices, and $\boldsymbol{e_t}$ is a (n X 1) vector of unobservable serially independent market shocks.

A common template comprising all dynamic regression models is provided by Fackler and Goodwin (2001, p.996). Their economic model, a point location model[22], is based on the following linear excess demand functions (see figure 3.1)

$$q_{it} = b_i (a_{it} - P_{it}) \tag{3.8}$$

where q are the net exports in country i, a is a shock that causes parallel shifts in the excess demand (for example, a change in the autarchy price), P is the i–th price and b a coefficient.

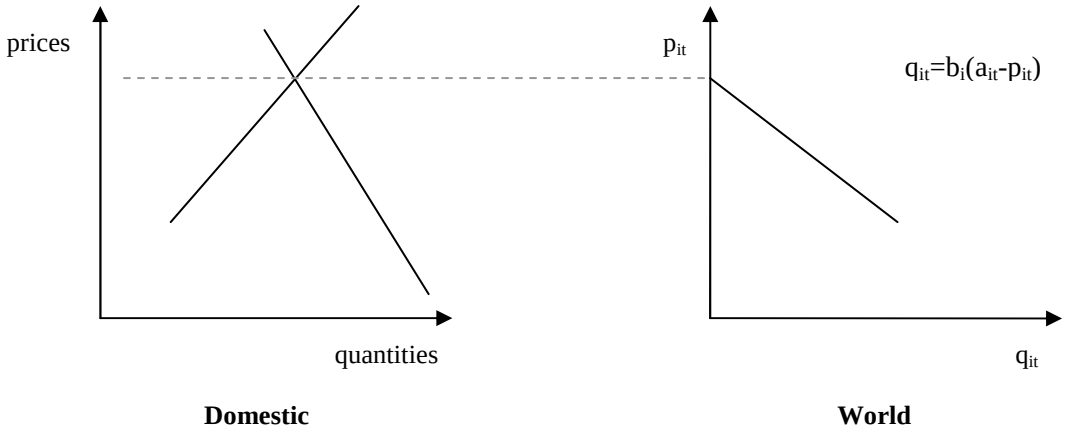
Indicating with r_t the transport cost from location *1* to location *2*, the equilibrium conditions for the two-location model in which *1* always exports to *2* can be written as:

$$\begin{bmatrix} b_1 & b_2 \\ -1 & 1 \end{bmatrix} \begin{bmatrix} P_{1t} \\ P_{2t} \end{bmatrix} = \begin{bmatrix} b_1 a_{1t} + b_2 a_{2t} \\ r_t \end{bmatrix} \tag{3.9}$$

where both net exports and prices, net of transport costs, are equalized.

[22] In point location models the network structure links serve only for commodity transportation flows. On the contrary, in agents-on-links models, markets or firms are located at network nodes and consumers or commodity producers are located continuously along the network links. These models are used to represent spatial oligopoly situations. The differences in behaviour attributed to this model and the point location model have then more to do with the competitive structure of the market than with the spatial one (Mc New and Fackler 1997).

Figure 3.1 Linear excess demand function for country i at time t



From (3.9), defining $x_{1t} = b_1 a_{1t} + b_2 a_{2t}$ and $x_{2t} = r_t$, supposing the three variables (a_{1t}, a_{2t}, r_t) can be written as a VAR, we have

$$\mathbf{x_t} = \sum_{\mathbf{k=1}}^{n} \mathbf{B_k x_{t-k}} + \mathbf{v}_t \tag{3.10}$$

eliminating the forcing variables results in the VAR in prices already described in equation (3.7):

$$\begin{bmatrix} b_1 & b_2 \\ -1 & 1 \end{bmatrix} \mathbf{P_t} = \sum_{k=1}^{m} \begin{bmatrix} B_{11k} & B_{12k} \\ B_{21k} & B_{22k} \end{bmatrix} \begin{bmatrix} b_1 & b_2 \\ -1 & 1 \end{bmatrix} \mathbf{P_{t-k}} + \mathbf{v}_t \tag{3.11}$$

Though, normally, it is the reduced form of the VAR presented in equation (3.7) which is analysed, at issue is what restrictions spatial equilibrium imposes on the coefficients of equation 3.11. Granger causality and lead-lag relationships, the so-called Ravallion market integration criteria, impulse response analysis and cointegration analysis all must be interpreted within the framework proposed.

3.2.2.1 Granger causality and lead-lag relationships

The Granger Causality (GC) concept refers to the notion of causality in terms of lead and lag relationships: significant coefficients of the lagged prices imply that shocks to prices in one market evoke significant responses in another after some lags.

Unidirectional causality between two prices can be seen as informational inefficiency, since the second price is not incorporating the information coming from the first one, and its values could be predicted on the basis of those of the second one (see, for example, Gupta and Mueller 1982, p.303). Nonetheless, as Fackler and Goodwin point out (2001, p.998), the dynamics in the price adjustments, owing for example to delivery lags, could make this assumption questionable.

GC analysis, together with IRFs (see paragraph 3.2.2.3) often accompanies dynamic model studies in empirical works.

3.2.2.2 Ravallion and Timmer market integration criteria

Ravallion's model, as presented in his original article (Ravallion 1986), builds up a radial structure with a central market and other satellites ones. Together with the Timmer model (which, as explained further in this paragraph, directly stems from Ravallion's one), it can be interpreted as a VAR model with tests of restrictions on the reduced-form parameters of the model (Fackler and Goodwin 2001, p.1000). Ravallion's model is described in equations (3.12) and (3.13). The price in market *1* is influenced by contemporary and lagged prices in all other markets and its own lags, while the price in any of the other *i* markets is influenced by the contemporary and lagged values of the price in market *1* and its own lagged values, only. Two equations describe the price transmission mechanisms, but, due to under-identification problems, only the second one is used in practice:

$$P_{1,t} = \sum_{j=1}^{n} a_{1,j} P_{1,t-j} + \sum_{k=2}^{m} \sum_{j=0}^{n} \beta_{i,j}^{k} P_{k,t-j} + X_{1,t} c_1 + e_{1,t} \tag{3.12}$$

$$P_{i,t} = \sum_{j=1}^{n} a_{i,j} P_{i,t-j} + \sum_{j=0}^{n} \beta_{i,j} P_{1,t-j} + X_{i,t} c_i + e_{i,t} \tag{3.13}$$

where $\boldsymbol{X_i}$ is a vector of other influences on local markets. Normally, it is equation (3.13) which is used; if $\beta_{i,j} = 0$, we have market segmentation, since the market *1* doesn't influence the market *i*. On the other side, if $\beta_{i,0} = 1$, immediate transmission is present. We have "strong" short run integration if $\beta_{i,0} = 1$ and $\alpha_{i,j} = \beta_{i,j} = 0$ for any $j > 0$ (the lagged prices have no influence), while we can talk about a "weaker" short run market integration if

$$\sum_{j=1}^{n} a_{ij} + \beta_{ij} = 0 \tag{3.14}$$

i.e., the effects of the lagged central price and of the lagged price itself cancel, on average, each other out. To have long run integration, the equivalence

$$\sum_{j=1}^{n} a_{i,j} + \sum_{j=0}^{n} \beta_{i,j} = 1 \tag{3.15}$$

must hold; under long run integration, the second equation can be re-written as:

$$\begin{aligned} \Delta P_{i,t} &= (a_{i,1} - 1)(P_{i,t-1} - P_{1,t-1}) + \sum_{j=2}^{n} a_{i,j}(P_{i,t-j} - P_{1,t-j}) + \beta_{i,0}\Delta P_{1,t} + \\ &+ \sum_{j=1}^{n-1} (\beta_{i,o} - 1 + \sum_{k=1}^{j} a_{i,k} + \beta_{i,k})\Delta P_{1,t-j} + X_{i,t} c_i + e_{i,t} \end{aligned} \tag{3.16}$$

where changes in local prices are attributed both to past spatial price differentials and to changes in central prices; the latter variables allow for the possibility that the markets are not observed in an integrated equilibrium at a given point in time, and so there are feedbacks from prior disequilibria. Fackler and Goodwin (2001, p.1001) derive an interpretation of Ravallion's model based on the structural VAR model they propose. The "strong" and the "weak" criterion are found not to refer to weaker equilibrium conditions but rather to weaker identification assumptions concerning the driving forces. They assert that the strong form of the short run integration implies transport rates to exhibit no persistence, which is a rather strong assumption, whereas in the weak form excess demand shocks are implicitly assumed not to have long run effects on the transport rate, which could be more reasonable.

The so-called Timmer's criterion is based on Ravallion's model, but assumes that central market prices are predetermined relative to hinterland prices and that a first-order model is sufficient to capture price dynamics. After some algebraic manipulations the following expression is derived (Alderman 1992, p.9):

$$P_{it} = (1 + b_1)P_{it-1} + b_2(P_{1t} - P_{1t-1}) + (b_3 - b_1)P_{1t-1} + cX_t + \mu_{it} \tag{3.17}$$

where

$$b_1 = \alpha_{i-1} \tag{3.18}$$

$$b_2 = \beta_{i0} \tag{3.19}$$

$$b_3 = \alpha_i + \beta_{i0} + \beta_{i1} - 1 \tag{3.20}$$

An Index of Market Connectiveness (IMC) is thus obtained, being

$$IMC_i = \frac{(1+b_1)}{(b_3 - b_1)} \tag{3.21}$$

Since long run equilibrium conditions bring $(P_{1t} - P_{1t-1})$ = 0, the numerator and the denominator are, respectively, the contributions of local and central market price history to current prices. If markets are integrated, the IMC index should be close to zero, since the lagged effects of regional market shocks are small relative to the central reference market ones.

Fackler and Goodwin (2001, p.1003) also in this case provide an interpretation of the model based on the structural VAR model they elaborate. They find out that a large value of the IMC may indicate that the locations are not integrated but may also indicate that they are integrated and that transport rates exhibit a high degree of persistence; a low IMC suggests instead that markets are not isolated but is unclear how connected they are. They conclude that both Ravallion's strong form criterion and Timmer's IMC index are helpful only if one has the confirmation that transport costs are white noise processes.

3.2.2.3 Impulse response analysis

In the moving average representation of a VAR system, Impulse Response Functions (IRFs) represent the effects of exogenous shocks to the variables, and allow to study their path of response. For a system of *n* prices, the set of impulse responses is given by

$$P_t = \sum_{k=0}^{\infty} M_k e_{t-k} \tag{3.22}$$

The most important aspect of the analysis of IRFs is the possibility of checking whether the price series converge quickly after an isolate, exogenous shock to one of them (Goodwin and Piggot 2001, p.315). If two markets are integrated, an exogenous shock to prices in one market should indeed evoke an equilibrating response in the other one (Goodwin *et al.* 1999, p. 168). IRFs provide a more general view of market integration than the standard "all or nothing" tests, since

the price response to disequilibria may take several periods to be complete (Goodwin *et al.* 1999, p.177), as delivery lags and other impediments to arbitrage and trade exist.

IRFs have been interpreted as both dynamic disequilibrium adjustments (which is justified only if shocks are serially uncorrelated) and equilibrium adjustments to ongoing changes in the economic fundamentals. If price data alone are available, due to identification problems it is difficult to decide which interpretation is correct (Fackler and Goodwin 2001, p.1004). In either case, shocks must be given an economic interpretation: a standard practice is to assume that the shocks are uncorrelated and that, in equation (3.7), $\mathbf{A_0}$ is triangular, implying that prices form a causally recursive system; this means that shocks on some prices have no effect on the others. Unfortunately, not only is this assumption untestable and rather strong, but it is like imposing a priori inefficiency on the market, while wanting to study it. Finally, for agricultural commodities, it must be noticed that seasonal differences (like those originating form different harvesting times in the two hemispheres of the world) may lie behind different pattern of responses (Margarido *et al.*, 2004).

3.2.2.4 Cointegration analysis

Starting from the seminal work of Ardeni (1989), cointegration techniques soon had an intuitive appeal for the study of price transmission mechanisms. Indeed, commodity price time series are often I(1) (integrated of order 1), which need to be differentiated before becoming stationary. Cointegration models presuppose that variables exhibiting nonstationary behaviour will nonetheless be linked by a long run relation, whose residuals are stationary. In this case, the latter is nothing but the LOP, which is expected to hold in the long-run despite short-run price variations.

As it is this methodology which will be adopted for the empirical estimates of this work, cointegration techniques will be extensively dealt with in paragraph 3.3.

3.2.3 Switching regime models

Switching regime models provide estimates of being in one of different trade regimes, determined on the basis of information on prices and transport costs. Indeed, conventional tests that rely on price data alone normally fail to recognize the role played by transport costs (see paragraph 3.4). Sexton *et al.* (1991), by using information on trade flows, study a market which is linked by unidirectional trade and develop a switching regime model in which arbitrage conditions may be violated; transport costs are estimated endogenously. The three regimes are

defined as the situation within, at or outside the arbitrage condition. Market efficiency is equivalent to the probability of the regime which satisfies the arbitrage condition to be one and the others to be zero. Baulch (1997) extends this model by adding explicit information on transaction costs and allowing trade flow reversals to take place. The use of this approach (called the Parity Bound Model) allows markets to be integrated in some periods but not in others (trade flows discontinuity), transfer costs to vary between periods, and prices to be contemporaneously determined.

Indicating with $|P_{1t} - P_{2t}|$ the price differential, with C_{12} the transfer cost, and with u_{It} and u_{Ot} the additional error terms inside and outside the parity bounds (assumed to be independently and half normally distributed), three regimes are derived according to the size of price spreads and of transfer costs:

- $|P_{1t} - P_{2t}| = C_{12}$. Regime 1 – efficient arbitrage: the price differential is exactly at the parity bound; transfer costs equal the market price differential and there are no impediments to trade between markets. Trade will cause prices to move on a one-to one basis and the arbitrage conditions are binding.
- $|P_{1t} - P_{2t}| = C_{12} - u_{It}$. Regime 2 – the price differential is inside the parity bound. In this case, the interpretations of Baulch and Sexton *et al.* are slightly different. For Baulch, when the price differential is inside the parity bound, it is because transfer costs exceed the inter-market spread: trade will not occur and the arbitrage conditions will not be binding (no profitable trade flows). For Sexton *et al.*, that assume unidirectional trade, if the price spread is within the arbitrage band it is because too much trade has occurred ("relative glut").
- $|P_{1t} - P_{2t}| = C_{12} + u_{Ot}$. Regime 3 – the price differential is outside the parity bound. In Baulch's model, if spreads exceed transfer costs, the spatial arbitrage conditions are violated whether or not trade occurs (non exploited profitable trade opportunity). In the interpretation of Sexton *et al.*, too little trade occurs ("relative shortage" case).

In the switching regime models presented, the probability distribution associated with the price spread is a mixture of three distributions, corresponding to the three regimes. The logarithm of the likelihood function of the sample is minimized and gives probability estimates for each of the three regimes[23]: the extent of market integration is expressed in terms of a continuous frequency of

[23] If transfer costs were known, testing for the efficiency of spatial arbitrage would reduce to arithmetically calculate whether trade occurred or not whenever the inter-market price differential exceeded or not transfer costs. But time series data on transfer costs are rarely available: either they must be estimated and inserted in the model, or be an endogenous parameter (Sexton *et al.* 1991).

each of the possible three regimes. The higher the incidence of regime 3, the lower the extent of market integration[24].

However, both regime models presented still rely on prices, only, and not on trade flows. Indeed, Sexton *et al.* (1991) only consider price data, and Baulch (1997) adds information on transactions costs.

The Barrett-Li (2001) method, instead, also makes use of trade flows data. Indeed, as mentioned in paragraph 2.2 already, Barrett (2001, p.20) and Barrett and Li (2002, p.292) claim that the measurement of integration should rely on flow-based indicators of tradability, whereas the estimation of efficiency should rely on price-based tests for market equilibrium, the most famous of which is the Enke-Samuelson-Takayama-Judge spatial equilibrium condition (ESTJ). According to the ESTJ equilibrium condition (Barrett 2001, p.21), the dispersion of prices in two locations for an otherwise identical good is bounded from above by the cost of arbitrage between the two markets and from below when trade volumes reach some ceiling value[25]. Trade is neither necessary nor sufficient for the attainment of the equilibrium in the ESTJ sense. This equilibrium condition uses four variables: prices, transaction costs, trade volumes, and trade volume quotas, "*yet economists too commonly (and falsely) claim an ability to evaluate market efficiency on the basis of only one variable: prices*" (Barrett 2001, p.21). So, Barrett and Li (2002, p.293), propose a new approach based on the maximum likelihood estimation of a mixture distribution model incorporating price, transfer cost, trade flow data, and a corroborating non parametric test. The combination of the various conditions of trade, *T*, and of the marginal returns, *R*, allows the definition of different regimes. In fact, price transmission might occur in absence of trade (segmented equilibrium), and trade might take place in absence of price transmission (imperfect market integration). They estimate different regimes that together define the four possible market conditions for any given pairs of markets:

- perfect integration: $R_{jit} = 0 \text{ and } T_{jit} \geq 0$
- segmented equilibrium: $R_{jit} < 0 \text{ and } T_{jit} = 0$
- imperfect integration: $R_{jit} \neq 0 \text{ and } T_{jit} > 0$
- segmented disequilibrium: $R_{jit} > 0 \text{ and } T_{jit} = 0$

Barrett and Li (2002, p.294) point out that most econometric applications are in fact aimed at testing the most restrictive condition (i.e., the presence of perfect

[24] When production and consumption are not specialized (i.e., both occur in both places), both regime 1 and 2 are consistent with market integration (Baulch 1997, p. 480).

[25] Indicating with *p* the prices in locations *1* and *2*, with τ the arbitrage costs, and with *q* trade values, we have that

$$\begin{array}{lll} p_1 \leq p_2 + \tau_{12} & \textit{if} & q_{12} = 0 \\ p_1 = p_2 + \tau_{12} & \textit{if} & q_{12} \in (0, \bar{q}_{12}) \\ p_1 \geq p_2 + \tau_{12} & \textit{if} & q_{12} = \bar{q}_{12} \end{array}$$

integration), in which both market integration and a competitive equilibrium are verified, and are not capable of fully capturing the "messy" character of market relationships.

Nevertheless, also "*these methodological enhancements are no panacea*" (Barrett 2001, p. 25). In fact, switching regime models have some common drawbacks. They are not dynamic, i.e. only offer static comparisons between the prices, and do not contain information about the speed of adjustment of prices; in addition to this, when trade data are not considered, violations of the spatial arbitrage conditions indicate lack of market integration irrespective of its causes. But a fundamental problem associated with switching regime models is that the believability of regime interpretation rests on "*very strong and rather unrealistic*" underlying distributional assumptions about which the economic theory has little to say (Barrett 2001 p.25; Fackler and Goodwin 2001, p.1012; Dercon 1999, p.2)[26].

3.2.4 Rational expectations models

The hypothesis behind the use of rational expectations models is that, since price linkages are not contemporaneous and may involve lags, agents have expectations that they must formulate about prices at the time of delivery. Goodwin *et al.* (1990) develop a rational expectation version of the LOP. Recalling equation 2.6 (see also note 7 of chapter 2), if we admit that delivery lags are different, such as in the case where $k = 0$ and $j > 0$, we will have

$$P_{1t} = \beta_0 E_t \{P_{2t+j}\} \tag{3.23}$$

That is simply the LOP holding when rational expectations are considered.

3.3 The use of cointegration techniques in spatial price transmission

Cointegration models presuppose that observable variables exhibiting nonstationary behaviour will nonetheless maintain long-run relationships. The residuals from such long run relationships are stationary. In our case, the long-run relationship is nothing but the LOP, which is assumed to be valid in the long run despite prices being allowed to diverge from it in the short run. In fact,

[26] For example, in Sexton's interpretation, the assumption of independent errors is difficult to understand, since it implies that there is no process of adjustment to the arbitrage errors. In the Baulch one, the same can be said for the assumption of a half-normal distribution inside the parity bounds, since it implies a higher density near the parity bound, even though the markets are not connected at that moment in time; an identical distribution of the error inside the parity bound would be more realistic (Dercon and Campenhout 1999, p.4).

commodity price time series are often found to be I(1) (integrated of order 1) processes, which need to be differentiated to become stationary.

It is easy to understand why cointegration models soon had an "intuitive appeal" for being used for price transmission analysis: they allow to disentangle short and long run dynamics. In addition to this, notably, and differently from linear regression models, none of the prices is required to be assumed as exogenous.

Starting from a from a standard VAR model with lag length *p*, or VAR(*p*), a Vector Error Correction Model (VECM) implying cointegration between the variables can easily be derived (see Enders 1995). Rewriting the general VAR(*k*) presented in equation (3.7) as

$$\mathbf{p_t} = \mathbf{A_0} + \mathbf{A_1 p_{t-1}} + \ldots + \mathbf{A_k p_{t-k}} + \boldsymbol{\varepsilon}_\mathbf{t} \qquad t = 1,\ldots,T \tag{3.24}$$

where *k* is the lag length, $\mathbf{p_t}$ (n X 1) is a vector of endogenous prices (where the lower cases indicate the use of logarithms); $\mathbf{A_t}$ (n X n) are matrices of unknown parameters; $\mathbf{p_{t-k}}$ (n X 1) is a vector of the *k*-th lagged value of p_t and $\boldsymbol{\varepsilon}_\mathbf{t}$ (n X 1) are white noise disturbance terms which may be contemporaneously correlated. A VECM with lag length *k*, VECM(*k*), can be obtained by simple re-parametrization:

$$\Delta\mathbf{p_t} = \boldsymbol{\Pi}_\mathbf{0} + \boldsymbol{\Pi}_\mathbf{1}\Delta\mathbf{p_{t-1}} + \ldots + \boldsymbol{\Pi}_\mathbf{k-1}\Delta\mathbf{p_{t-(k-1)}} + \boldsymbol{\Pi}\ \mathbf{p_{t-k}} + \boldsymbol{\varepsilon}_t \tag{3.25}$$

that is equal to

$$\Delta\mathbf{p_t} = \boldsymbol{\Pi}_\mathbf{0} + \sum_{\mathbf{i=1}}^{\mathbf{k-1}} \boldsymbol{\Pi}_\mathbf{i}\Delta\mathbf{p_{t-i}} + \boldsymbol{\Pi}\ \mathbf{p_{t-k}} + \boldsymbol{\varepsilon}_t \tag{3.26}$$

where

$$\boldsymbol{\Pi}_\mathbf{0} = \mathbf{A_0}\ ,\ \boldsymbol{\Pi}_\mathbf{i} = -\left(\mathbf{I} - \sum_{\mathbf{j=1}}^{\mathbf{i}} \mathbf{A_j}\right),\ j = 1,2,\ldots p-1 \tag{3.27}$$

$$\boldsymbol{\Pi} = -\left(\mathbf{I} - \sum_{\mathbf{i=1}}^{\mathbf{k}} \mathbf{A_i}\right) \tag{3.28}$$

and $\Delta\mathbf{p_{t-j}}$ is a vector of $\mathbf{p_{t-j}}$ in first differences, j =1,2,…*k-1* (see Enders 1995, pages 389-390). The rank of $\boldsymbol{\Pi}$ provides the basis for determining the presence of cointegration between the variables: if rank($\boldsymbol{\Pi}$) = 0, the variables are not

cointegrated and the model is equivalent to a VAR in first differences; if rank($\mathbf{\Pi}$) = n, the variables are stationary and the model is equivalent to a VAR in levels; if 0 < rank($\mathbf{\Pi}$) = r < n, the variables are cointegrated. $\mathbf{\Pi}$ can be decomposed as $\mathbf{\Pi} = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ is the (n X r) matrix of the speed of adjustment coefficients and $\boldsymbol{\beta}$ is the (n X r) matrix of the cointegration vectors. The basic assumption is that $\boldsymbol{\beta}'\mathbf{p_t} = z_t$, representing the long run disequilibrium error, is stationary. All variables of an error correction model (ECM) exhibit the property of stationarity under the conditions that the levels of the individual variables are integrated of order one and a cointegration relationship between the variables exists. Statistical significance of the error correction parameter implies in turn the existence of a cointegrating relationship (Thompson *et al.* 2002, p.1044).

There is also an alternative way to derive an ECM model which includes responses from a one-lagged long run disequilibrium term (Lucchetti 2008). Starting from the general VAR

$$\mathbf{p_t} = \mathbf{A_1 p_{t-1}} + \boldsymbol{\varepsilon_t} \tag{3.29}$$

subtracting $\mathbf{p_{t-1}}$ on both sides, we have that

$$\Delta\mathbf{p_t} = \mathbf{\Pi p_{t-1}} + \boldsymbol{\varepsilon_t} \tag{3.30}$$

where $\mathbf{\Pi} = \mathbf{A_1 - I} = \boldsymbol{\alpha\beta}'$, whose rank allows to determine the presence of cointegration amongst the variables (see above). An autoregressive term can be added to remove the short run persistence in the error tem; like in equation 3.26, in the following VECM

$$\Delta\mathbf{p_t} = \boldsymbol{\alpha\beta}'\mathbf{p_{t-1}} + \sum_{i=1}^{k-1}\Gamma_i \Delta\mathbf{p_{t-i}} + \boldsymbol{\varepsilon_t} \tag{3.31}$$

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, Prakash 1999 cited Conforti 2004, p.3), $\boldsymbol{\alpha}$ is the loading matrix which contains the adjustments parameters (a measure of the speed of price transmission, Prakash 1999 cited Conforti 2004, p.3), $\boldsymbol{\Gamma_i}$ are matrixes containing coefficients that account for short-run relations, and $\boldsymbol{\varepsilon_t}$ are white noise errors. Since the long run relation is the LOP, and being prices written in logarithmic form, the assumption is that price spreads (that is, all components which account for price spreads) are a stationary proportion of prices.

When prices are expressed in logs, the coefficients included in $\boldsymbol{\beta}$ can be read as price transmission elasticities. In fact, if p_t contains only two log-prices, say p_{1t},

the domestic price, and p_{2t}, the world price, we will have the following long-run relation between them:

$$\boldsymbol{\beta}'\mathbf{p}_{\mathbf{t-1}} = \begin{bmatrix}\beta_0 & \beta_1 & \beta_2\end{bmatrix}\begin{bmatrix}1\\ p_{1t-1}\\ p_{2t-1}\end{bmatrix} = \beta_0 + \beta_1 p_{1t-1} + \beta_2 p_{2t-1} = z_{t-1} \tag{3.32}$$

where $z_t \approx I(0)$. Normalizing with respect to β_1 and rearranging terms, we have

$$p_{1t-1} = -\frac{\beta_0}{\beta_1} - \frac{\beta_2}{\beta_1} p_{2t-1} + z_{t-1} \tag{3.33}$$

where β_2 / β_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, p.2) argue that they can indeed be interpreted 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: in general, trade liberalization will contribute to greater price transmission elasticities.

The point estimates of adjustment coefficients, which measure how much a price responds to the disequilibrium from the previous period (in other words, the larger the absolute value of the adjustment coefficient, the faster the convergence toward the LOP), have important economic content (Thompson *et al.* 2002, p.1044). The stationarity of price spreads requires $\alpha_1 < 0$ and $\alpha_2 > 0$. Moreover, the weak exogeneity of one price with respect to the long run parameters requires its adjustment coefficient to be zero[27]. Finally, the adjustment coefficient is a proxy for the short run transmission elasticity.

The existence of a stable relations between two prices, i.e. bivariate cointegration, has been deemed to be a necessary condition for integrated markets since the seminal work of Ardeni (1989). Others claim that the strongest condition that prices differences are stationary is needed, and thus impose instead of estimating the cointegration vector (Baffes 1991). Cointegration tests are run pair wise (Engle-Granger procedure) or with multivariate systems (Johansenn's

[27] In a VECM, weak exogeneity does not imply GC since GC involves statistical tests of significance also of the lagged terms in addition to the parameters of the error correction terms.

procedure)[28]. In a system with *n* prices, the number of the cointegration relations has been considered as an index of the degree of integration of the markets. Cointegration versions of Ravallion's model have been proposed, as well (Alderman 1992; Dercon and Campenhout 1999).

Despite their intuitive appeal, the use of cointegration techniques presents a number of shortcomings (see Barrett 1996; Miljkovic 1999, pp.132-133).

First of all, it is evident that cointegration is not a necessary condition for markets to be efficiently integrated, since transaction costs (as well as other elements not explicitly included in the LOP equation, and then contributing to price spreads; see paragraph 2.3) could be nonstationary processes[29; 30; 31]. So, price diverging far from each other could not imply any arbitrage possibility at all: if transaction costs are non stationary, cointegration is then an unnecessary condition for market efficiency (Barrett 1996; Mc New and Fackler 1997).

Secondly, if the transaction cost band is large but the price spread remains within the band, prices could result cointegrated even if markets are actually not linked, since the difference between the prices would not justify any shipment of the commodities. This means that, in order to use the cointegration approach, transaction costs should be small if compared to prices, which is often not the case for agricultural goods. Barrett (1996, p.827), 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

[28] Both the Engle-Granger and the Johansen procedure have been extensively used in the empirical literature (Miljkovic 1999, p. 132). Bivariate Engle Granger cointegration tests only allow pair wise comparisons, and require one of the two prices to be designated as exogenous. Also, Banejeree *et al*. (1986 cited Goodwin 1999, p. 118) raise concerns about the potential for small sample biases in parameter estimates obtained from such two-steps procedure, and the fact that Engle-Granger testing procedures do not have well-defined limiting distributions. Hall (1989 cited Goodwin 1999, p. 118) notices that Johansen's multivariate testing procedure, on the other side, suggests a maximum likelihood estimation procedure that provides estimates of all the cointegrating vectors existing among a group of variables. This approach uses test statistics that have an exact limiting distribution that is function of a single parameter.

[29] Baffes (1991), in those cases were cointegration between prices doesn't apply, finds out that the non stationarity of freight rates could be a possible explanation for the non stationarity of price differentials.

[30] Fackler and Goodwin (2001) note that in some alternative models (such as the switching regime model) price spreads depend on all the shocks; this means that, in order to ensure cointegration, transport rate stationarity is not sufficient, but also stationarity in excess demand shock differences is needed.

[31] 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, even when transport rates are assumed to be stationary, 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. This means that the arbitrage condition is not strong enough to ensure cointegration. As transport rates increase, and markets are thus less likely to be integrated, price spreads tend not to be stationary and price cointegration tends to be lower. They conclude that if we want to test the LOP, we need to know which the cointegration properties of the autarky prices are, that markets are well integrated and that transport rates are stationary before we can have faith in cointegration tests.

arbitrage): in both cases there is no efficient exchange between markets, but cointegration tests identify only the former kind.

Thirdly, Barrett (1996, p.827) notices that cointegration could be consistent with a negative relationship between prices, whereas market integration suggests a positive correlation between them. He furthermore notices that many reported coefficients have magnitudes implausibly far from unity. Moreover, when prices are expressed in levels, the relation between market prices is of exactly unit slope if costs are purely additive; if some component is proportional, it might differ from one (Miljkovic 1999, p.133; the same consideration holds when prices are expressed in logarithms and transaction costs are instead expected to be a proportion of them).

Fourthly, trade flow discontinuity could also represent a problem: at the break points, the relation between prices is indeed zero. And, if demand and supply forces are themselves cointegrated, prices in two regions could be cointegrated even in the absence of trade flows. Cointegration models then require continuous trade and no trade flow reversals. Moreover, since cointegration implies GC in at least one direction, fully efficient markets should not be cointegrated, or otherwise there would exist at least one GC relationship between the markets, implying that profits could be made from using information on past prices to predict present prices; cointegration is consistent with an interconnected market, but not with a fully efficient one (Engle and Granger 1987, Granger and Escribano 1987 cited Dercon and Campenhout 1999, p.7).

A final consideration is that a stable long run margin between markets could be consistent with monopolistic pricing and other causes of inefficiencies (Faminow and Benson cited Dercon and Campenhout 1999, p.2).

Despite these shortcomings, to date cointegration has been the most widely used technique to study market efficiency: 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, and results were flawed because of nonstationarity in the data, spurious regressions, erroneous use of first differences. The cointegration approach has been used so far on a huge number of empirical works (Goodwin and Fackler 2001); for a table summarizing the methods and the data used and the results obtained, see annex A. Miljkovic (1999), presents an interesting critical review.

In the next paragraphs, some issues related to the use of cointegration techniques for the study of price transmission will be further explored. Paragraph 3.3.1 will deal with unit root tests when used with price series. Then, in paragraph 3.3.2, the different kinds of cointegration models which have been proposed to the study of price transmission will be briefly presented. Finally, in paragraph 3.3.3, the empirical evidence will be commented.

3.3.1 Unit root tests applied to price series

Understanding the time series properties of the data is a prerequisite for any cointegration analysis which, as explained in the previous paragraph, normally requires the time series to be I(1).

Empirical tests, motivated by a vast literature on unit roots and error correction models, often find evidence of unit roots in commodity prices, though according to price theory price series need not have unit roots. Indeed, price theory states that commodity prices will be auto correlated convergent series, which arises from the biological nature of commodity production and the storage and arbitrage costs over time (Wang and Tomek 2007, p.873).

Moreover, the outcome of unit root tests is likely to be affected by a number of factors, first of all alternate test specifications (Wang and Tomek 2007, p.875). Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root tests have low power against the alternative that the series is stationary; for this reason, Kwiatkowski, Phillips, Schmidt e Shin (KPPS) tests, having the null hypothesis of stationarity, can be used as a consistency check, though, often, neither test will reject (Tomek and Myers 1993, p.189). Wang and Tomek (2007) find that unit root tests results are sensitive to the test equation specification, and that evidence favouring unit roots in commodity prices is not strong. Data transformations (using the series in nominal or in real terms, or in logs) should strictly depend upon the research questions which are asked. Data frequency also plays a role[32]. Finally, if the number of lags used for the unit root tests is too small, then the inference about the existence of a unit root is biased; if is too large, the finite sample properties of unit root tests are likely to deteriorate, and an inefficient estimate is obtained.

Structural breaks also represent an important object of analysis. It is well known that failure to allow for an existing structural break while testing for unit roots leads to a bias that reduces the ability to reject a (false) unit root null hypothesis. If a structural break has occurred, but is not modelled in the test equation, then standard ADF tests will be biased towards non-rejection of the null hypothesis. For this reason, Perron and Vogelsang (1992) have extended the ADF specification to allow for possible structural breaks in the time series. They propose two tests which allow for the existence of one endogenous structural break both under the null hypothesis of non-stationarity and under the alternative hypothesis of stationarity. The Additive Outlier (AO) test, by using a dummy

[32] Defining the total number of observations as T, which is given by the product between span of the sample S and the frequency f , the consistency and the power of the tests as f increases depend upon the assumptions on S. If S is fixed, as f increases the power of tests increases but ultimately levels off. If the assumption of a fixed S is relaxed, it appears that the frequency of the sample is more important than the number of observations in determining the power of the test, but this of course assumes a constant structure of the data generating process, which might not be the case as S increases (Wang and Tomek 2007, p.877).

variable to account for a shift in the mean, assumes that the structural change occurs at a specific point in time. The Innovational Outlier (IO) model assumes instead that the change affects the level of the series gradually. The time of the break is chosen by minimizing the *t*-statistic associated with the change in mean. Clemente *et al.* (1998) extend these tests by allowing for two endogenous changes in the mean. Zivot and Andrews (1992), on the other side, build a test in which the break date is selected where the t-statistic of the ADF test for unit root is at minimum (most negative); consequently, a break date will be chosen were the evidence is least favourable for the unit root null hypothesis. In this test, on the other side, the presence of a structural break is not considered under the null hypothesis; this might lead to the erroneous conclusion that the time series is trend stationary when in fact it is non-stationary with breaks. Other tests have also been developed for the case of more than one structural break in the series (for a review, see Glynn *et al.* 2007). A concern is that empirical approaches to find structural changes will result in the unit root hypothesis being falsely rejected, or, in other words, in finding a spurious structural change. Wang and Tomek (2007, p.877) suggest that in order to identify structural changes it is advisable to rely on logic, known causes and timing of the change; a simple inspection of graphs can then confirm it. As we will see in the next chapters, in the case study presented here there is clear evidence that at least some policy regime changes did indeed constitute structural breaks.

3.3.2 Different cointegration techniques

Over the past years, different cointegration models have been developed towards several directions for the study of price transmission. For example, threshold cointegration models have been used to represent the fact that prices show a tendency to return to their long run equilibrium only when price differentials exceeds a certain threshold, constituted by transaction costs (Goodwin and Piggott 2001, p.303). In asymmetric cointegration models[33] (Ghoshray 2002), instead, the speed of adjustment to the long run equilibrium is allowed to vary according to the sign and the magnitude of the variations of the error correction term. In the next paragraphs (3.3.2.1 and 3.3.2.2), threshold and asymmetric models will be briefly revised. Cointegration models allowing for structural breaks affecting the parameters of the cointegration vector have been developed, too; because they represent an interesting possibility of introducing policy regime changes in the model, they will be used for the empirical estimates presented in chapter 6, and are described in detail in paragraph 3.3.2.3.

[33] For a survey of asymmetric price transmission models see Meyer and von Craumon-Taubadel, 2004. Also in this respect, however, cointegration models have gained a prominent place.

3.3.2.1 Threshold models

In threshold cointegration models, the basic assumption is that prices show a tendency to return to their long run equilibrium only when price differentials exceed a certain threshold[34]. In other words, when deviations in prices exceed this "neutral band", which accounts for transaction costs, a sort of regime switching is triggered (Goodwin and Piggot 2001, p.303).

Some basic threshold cointegration models used in agricultural commodity markets studies are the Eq-TAR representation and the Band-TAR representation (see Balke and Fomby 1997; Brooks *et al.* 2005; Balcombe *et al.* 2007).

In the Eq-TAR representation (equation 3.34, where only the error correction term, z_{t-1}, and its adjustment coefficients, $\alpha_i^I(z_{t-1})$, are presented), when deviations from the long run equilibrium, z_t, exceed the threshold's width, λ, prices are attracted towards the centre of the threshold interval. The subscripts I and O denote the within and without π adjustment speed.

$$\begin{aligned} \alpha_i^I(z_{t-1}) &= \pi_{i,O} z_{t-1} \quad if \quad |z_{t-1}| > \lambda \\ &= \pi_{i,I} z_{t-1} \quad if \quad |z_{t-1}| < \lambda \end{aligned} \tag{3.34}$$

In general, the speed of adjustment is allowed to differ depending on prices to be inside or outside the interval; if $\pi_{i,I}$ collapses to zero, this means that, inside the band, the price spread will follow a random walk. In other words, this implies that if the price differential is smaller than the transport costs the arbitrage mechanisms will not be active and the two series will not be cointegrated; a price differential smaller than the transport costs will not justify any shipment of the commodities.

Differently from the Eq-TAR model, in the Band TAR model when prices are outside the threshold they are attracted to the edge of the threshold band, rather than to the middle of it. So, the Band -TAR representation is given by equation (3.35):

[34] Threshold models show the difference existing between cointegration and efficient arbitrage: if efficient arbitrage takes place, unit root behaviour in price margins should be observed (this happens up to and right at the parity bound); only when imperfect arbitrage takes place we should observe cointegration between the prices (since cointegration implies GC in at least one direction, profits could be made predicting one with the past values of the other; Dercon 1999, p.6; see also paragraph 3.3).

$$\begin{aligned}\alpha_i^{II}(z_{t-1}) &= \pi_{i,O}(z_{t-1}-\lambda) \quad if \quad z_{t-1}>\lambda \\ &= \pi_{i,O}(z_{t-1}+\lambda) \quad if \quad z_{t-1}<-\lambda \\ &= \pi_{i,I}(z_{t-1}) \quad if \quad |z_{t-1}|<\lambda\end{aligned} \tag{3.35}$$

The Eq-TAR and the Band TAR models can be easily put together (Balcombe 2007, p.312),

$$\begin{aligned}\alpha_i^{II}(z_{t-1}) &= \pi_{i,O}(z_{t-1}-\vartheta\lambda) \quad if \quad z_{t-1}>\lambda \\ &= \pi_{i,O}(z_{t-1}+\vartheta\lambda) \quad if \quad z_{t-1}<-\lambda \\ &= \pi_{i,I}(z_{t-1}) \quad if \quad |z_{t-1}|<\lambda\end{aligned} \tag{3.36}$$

where ϑ is a constant such that $\vartheta \in [0;1]$.

Dercon and Campenhout note that threshold error correction models "nest" a switching regime model and standard cointegration analysis (1999, p.19). In their work, they link the cointegration version of Ravallion's model with the switching regression approach of Baulch's Parity Bound Model. In an attempt to remove some of the unrealistic assumptions of both approaches, they note that the Band Threshold-Autoregression (Band-TAR) model allows for dynamic analysis as well as for trade discontinuities (and reversing trade flows), and explicitly models transaction costs.

Testing for threshold effects presents a number of challenges; normally, grid searches are used to determine the threshold level (both maximizing a likelihood function or minimizing a sum of squared errors are methodologies which can be used; Goodwin and Piggott 2001, p. 325).

Empirical evidence confirms that models that explicitly recognize the presence of thresholds effects imply faster adjustments to deviations from equilibrium than when threshold behaviour is ignored. Thresholds turn out to be bigger when there is a bigger distance between markets. Goodwin and Piggott (2001) also estimate a richer threshold model accounting for asymmetric effects[35]. A shortcoming of threshold models is that they are not capable of discriminating between transactions costs and other possible causes of threshold effects, such as structural breaks and policy shifts.

[35] Thresholds could be asymmetric. Positive and negative price shocks will in turn determine an asymmetric response, since shipment in one direction might be more costly than shipping in the other direction between two markets if, for example, one is normally importer and the other normally exporter (Goodwin and Piggot 2001, p.309).

3.3.2.2 Asymmetric adjustment models

The concept of "asymmetry" refers to the fact that adjustment coefficients are allowed to vary depending on the sign of price differentials. A very intuitive explanation of why this might be the case is that agents holding market power will pass-through only (or mostly) positive input changes (Ghoshray 2002).

In dynamic applications, the short-run adjustment term is then substituted by two separate coefficients indicating the response, respectively, to negative and positive deviations from the long run equilibrium.

Prakash *et al.* (2001 cited Conforti 2004, p.9), look at the significance of a dummy variable accounting for positive residuals in the static regression between the two price series involved. If this variable is significantly different from zero, and has a positive sign, transmission is asymmetric, since positive shocks are passed through faster than the negative ones (Conforti 2004, p.9).

In more sophisticated models, the speed of adjustment is expected to vary according to either the sign or the *variations* of price differentials. We start from the autoregressive model of the VECM error correction term, which we have indicated as z_t. In equation 3.37, rejecting the null hypothesis of no cointegration (H_0: $\gamma = 0$) implies that the residuals ω_t, are stationary (Ghoshray 2002, p. 304),

$$\Delta z_t = \gamma z_{t-1} + \omega_t \tag{3.37}$$

The Threshold Autoregressive Model (TAR model) can be represented as

$$\Delta z_t = I_t \gamma_1 z_{t-1} + (1 - I_t)\gamma_2 z_{t-1} + \omega_t \tag{3.38}$$

where I is the Heaviside indicator, which takes the following form:

$$I_t = \begin{matrix} 1 & if & z_{t-1} \geq 0 \\ 0 & if & z_{t-1} < 0 \end{matrix} \tag{3.39}$$

In this model, the adjustment coefficients are then allowed to vary according to the sign of the long-term residuals of the cointegration relationship. If the residuals are positive, the adjustment coefficient will be γ_1; if they are negative, it will be γ_2.

The sufficient conditions for the stationarity of z_t are γ_1 , $\gamma_2 < 0$.

If for example, $-1 < \gamma_1 < \gamma_2 < 0$, this means that the negative phase of z_t will tend to be more persistent than the positive phase.

In the Momentum Threshold Autoregressive Model (M-TAR model), the adjustment coefficients can vary according to the sign of the *variations* of price

differentials. The Momentum Heaviside indicator function appearing in the 3.38 is here defined as

$$I_t = \begin{matrix} 1 & if & \Delta z_{t-1} \geq 0 \\ 0 & if & \Delta z_{t-1} < 0 \end{matrix} \tag{3.40}$$

If, for example, $\gamma_1 < \gamma_2$, then the M-TAR model exhibits little adjustments for positive variations of the error term, i.e., increases tend to persist but decreases tend to revert quickly to the attractor.

3.3.2.3 Models with structural breaks

In analogy with unit root testing (see paragraph 3.3.1), disregarding the presence of structural breaks in the cointegration relation between the series might lead to erroneously accept the null hypothesis of nonstationarity in their linear combinations, concluding that they are non cointegrated. Indeed, what happens is that the cointegration residuals are incorrectly defined as non-stationary, while instead they are stationary around a broken level or trend.

In some approaches, here mentioned for the sake of completeness and not used in the remainder of this work, it is instead assumed that the structural break affects the cointegration rank, *r*. Barassi and Ghoshray (2007) do this by analyzing sequences of recursive and rolling likelihood based trace tests allowing for *k*-regime changes. They test the null hypothesis $H_o : (r = r_0)$ against the alternative $H_1 : (r > r_0)$ in some. $H_1 : (r > r_0)$ The time of the break is then unknown.

In other approaches, the structural breaks are instead assumed to affect the cointegration parameters. In fact, if structural breaks within the individual series occur, and they do not cancel each other out, cointegration must then be analyzed in a framework where breaks in the deterministic components are admitted (Dawson *et al.* 2006).

Johansen *et al.* (2000) generalize the standard Johansen cointegration test by admitting up to two predetermined breaks, and propose a model where breaks in the deterministic terms are allowed at known points in time. In the so-called "Johansen, Mosconi and Nielsen" procedure, the sample is divided in *q* periods, separated by the occurrence of the structural breaks, where *j* denotes each period. The general VECM model, considering only two prices, is:

$$\Delta \mathbf{p_t} = \boldsymbol{\alpha}\begin{bmatrix}\boldsymbol{\beta}\\ \boldsymbol{\mu}\end{bmatrix}'\begin{bmatrix}\mathbf{p_{t-1}}\\ \mathbf{tE_{t-1}}\end{bmatrix} + \boldsymbol{\gamma}\mathbf{E_t} + \sum_{\mathbf{i=1}}^{\mathbf{k-1}}\boldsymbol{\Gamma}_i\Delta\mathbf{p_{t-i}} + \sum_{\mathbf{i=1}}^{\mathbf{k}}\sum_{\mathbf{j=2}}^{\mathbf{q}}\mathbf{k_{i,j}D_{j,t+k-i}} + \\ + \sum_{\mathbf{m=1}}^{\mathbf{d}}\boldsymbol{\Theta}_\mathbf{m}\mathbf{w_{m,t}} + \boldsymbol{\varepsilon}_t \tag{3.41}$$

where k is the lag length of the underlying VAR. $\mathbf{E_t}$ is a vector of q dummy variables that take the value 1, i.e $E_{jt} = 1$., if the observation belongs to the jth period (j = 1, … q), and 0 otherwise; that is, $E_t = \lfloor E_{1t} \quad E_{2t} \quad ... \quad E_{qt} \rfloor'$[36]. $\mathbf{D_t}$ is an impulse dummy (with its lagged values) that equals unity if the observation t is the ith of the jth period, and is included to allow the conditional likelihood function to be derived given the initial values in each period (for example, if k = 3, impulse dummies will thereby have the value 1 at $t+2$, $t+1$, t, where t is the first observation of each period); $\mathbf{w_t}$ are intervention dummies (up to d) included in order to render the residuals well behaved. The short run parameters are $\boldsymbol{\gamma}$ (2 X q), $\boldsymbol{\Gamma}$ (2 X 2), $\mathbf{k}$ (2 X 1) for each j and i, and $\boldsymbol{\Theta}$ (2 X 2). $\boldsymbol{\varepsilon_t}$ are assumed to be i.i.d. with zero mean and symmetric and positive definite variance, Ω. $\boldsymbol{\mu} = \lfloor \mu_{1t} \quad \mu_{2t} \quad ... \quad \mu_{qt} \rfloor'$ is the vector containing the long run drift parameters and $\boldsymbol{\beta}$ are the long run coefficients in the cointegrating vector. The cointegration hypothesis is formulated by testing the rank of $\boldsymbol{\pi} = \boldsymbol{\alpha}\begin{bmatrix}\boldsymbol{\beta}\\ \boldsymbol{\mu}\end{bmatrix}'$; its asymptotic distribution depends on the number of non-stationary relations, the location of breakpoints and the trend specification.

It should be noticed that this framework includes two models: a first one is where there are no linear trends in the levels of the endogenous variables and the first differenced series have a zero mean; the broken level is restricted to the cointegration space (i.e., $\boldsymbol{\gamma}$ = 0, and the regime dummies are not multiplied by any trend in the cointegration vector). In the second case, a broken linear trend is accounted for in the cointegration vector but any long run linear growth is not accounted for by the model (i.e., $\boldsymbol{\gamma} \neq 0$ and $t \neq 0$ in the cointegration vector).

An application of the Johansen, Mosconi and Nielsen procedure to agricultural future and export prices is presented in Dawson *et al.* (2006) and Dawson and Sanjuan (2006), and will be used for the empirical analysis in chapter 6 since, as it will be clear, structural breaks can be a mean of representing the changes in policy regimes while testing for price transmission.

[36] For example, if q = 3, i.e. there are two structural breaks, we have $\mathbf{E'_t} = \lfloor E_{1,t} \quad E_{2,t} \quad E_{3,t} \rfloor = \lfloor 1 \quad 0 \quad 0 \rfloor$ if the observations of time t belong to the first period, $\mathbf{E'_t} = \lfloor 0 \quad 1 \quad 0 \rfloor$ if they belong to the second one, and $\mathbf{E'_t} = \lfloor 0 \quad 0 \quad 1 \rfloor$ otherwise.

3.3.3 Cointegration models and price transmission: empirical evidence

As explained before, cointegration techniques soon had an intuitive appeal for the study of price transmission, since they allow to disentangle short and long run market dynamics. The LOP is assumed to be valid in the long run despite prices diverging from it in the short run. For this reason, to date, cointegration has been one of the most widely used techniques for studying price transmission mechanisms.

Nevertheless, the empirical evidence is controversial, as the articles reviewed in annex A show. Here we will just mention a single example, that can be revealing. In his seminal work, Ardeni (1989) applied for the first time a cointegration methodology to test if the LOP holds, at least in the long run, for a number of commodities. He found that, quite uniformly, the LOP fails. But however, his results were overturned already by Baffes (1991) who, using the same dataset but a different model specification, found that the LOP was indeed holding in the markets considered.

There are countless applications of cointegration techniques to the study of price transmission. Miljkovic (1999, p.131) provides an interesting critical review of them (see annex A). Also amongst those who investigate soft wheat international markets, which are relevant to the analysis which will be carried out in this study, there is mixed empirical evidence.

Ghoshray *et al.* (2000) investigate 13 export wheat prices, different for origin and quality, and find that the wheats for human consumption embody a common price trend, while feed wheats share another one. Thompson and Bohl (1999) find that domestic German soft wheat prices and US Gulf Dark Northern Spring wheat prices are indeed cointegrated. Thompson *et al.* (2002a) use a Seemingly Unrelated Error Correction Model (SURECM) methodology and find evidence supporting the validity of the LOP amongst domestic EU prices (France, Germany and United Kingdom) and US export prices. They also find that price transmission is positively affected by market liberalization reforms. Verga and Zuppiroli (2003), on the other side, find that European (Italian and French) soft wheat markets are strongly cointegrated amongst themselves but not with the US one. Barassi and Ghoshray (2007) test cointegration with structural change to analyze the nature of the long-run relationship between US (Soft Red Winter wheat and Hard Red Winter wheat) and EU wheat export prices. After the breakpoint, occurred after the Mac Sharry reform of the European Common Agricultural Policy, the EU price is cointegrated with the US Soft Red Winter wheat price.

Although the differences in the data and methodologies used render this works not directly comparable to each other, what emerges is a complex picture in which results are controversial. While soft wheat export prices seem to share a common trend, this is not the case when European domestic prices are used: the transmission elasticity with world prices is either zero, or positive, or changing in

time. What we argue is that the presence of border and domestic regulation policies for the commodity considered (which makes export and domestic prices for the European Union intrinsically different, as we will see in chapter 4), when not adequately and explicitly modelled, increases the difficulties in comparing and assessing the results.

3.4 Problems in empirical models relying on price data

As explained in paragraph 2.3, any study of price transmission ultimately refers to verifying the validity of the LOP. Empirical evidence is controversial and, in most cases, the LOP fails. This can be due to the scarce probability for its hypothesis to hold in practice (paragraph 2.3), but this issue is also strictly entangled to the underlying hypothesis and properties of the models used, and this is particularly true if they rely only on price data.

Despite the use, over the past decades, of increasingly sophisticated techniques (amongst which, the very development of cointegration models represented a turning point), it is well recognized that none of the approaches described in this chapter is preferable in all cases, and that inferences should not be based upon a single test but, if possible, on a variety of inferential techniques. Models' misspecifications and, more generally, their inherent characteristics, have indeed been added as an explanation for not finding empirical evidence supporting the LOP.

What is common to all the empirical approaches which have been revised is that they basically rely on the use of price data, only, and of transport costs (normally freight rates) when available. Most of the literature normally uses traded prices (cost insurance freight, cif, or freight on board, fob) already converted in US dollars. Price series are usually nominal[37]. As far as exchange rates are concerned, they are likely to be endogenous only if highly aggregated data are used, so it is claimed that the appropriate level of disaggregation allows to treat them as exogenous variables (Richardson 1978; Goodwin *et al.* 1990; Mohanty and Langley 2003). Vollrath and Hallahan (2006) propose a "detailed" LOP model that explicitly accounts for changes in prices due to the exchange rate (see paragraph 3.2.1). Nevertheless, some attempts of separating prices and exchange rates have been criticised as too restrictive (Crouhy-Verac *et al.* 1982 cited Goodwin *et al.* 1990, p.686).

In all inferential techniques relying only on price data a critical problem is constituted by transaction costs, which are composed by a variety of elements which, as seen in the previous paragraph, hardly ever can be exactly addressed. When tests rely on prices only it is just not possible to separately test the

[37] Aldermann (1992), Goodwin *et al*. (1999) (both, for regional markets of the same country) and Thompson *et al*. (2000), use real prices; Thompson and Bohl (1999), and Thompson *et al*. (2002a), also use deflated prices but report no differences from the results obtained with the nominal ones.

efficiency hypothesis and the strong assumptions underpinning the model specifications (Barrett 2001, p.29). Inadequate treatment of these costs becomes crucial as they increase in size and volatility. Most commonly, quite strong assumptions are imposed: transaction costs are taken equal to zero, or constant (when, in linear models, prices are expressed in levels), or as a fixed proportion of prices (when prices are expressed in logs), and as serially uncorrelated, or, in cointegration models, stationary. Moreover, some variable costs depend on prices; so, there is always the problem of correlation between the unmeasured component of transaction costs, which is included in the error term, and prices (Barrett 2001, p.23). While it's worth noting that transport and transaction costs are a problem of all inferences, we will have to assess their magnitude relative to that of the prices to see which kind of distortion any assumption on them is likely to cause, remembering that they can be of a considerable size for agricultural commodities.

Moreover, in models that do not make use of trade data, the LOP should be interpreted as a test conditional on assumptions regarding trade linkages, since prices could move together for other reasons than market integration (Fackler and Goodwin 2001, p.992). Ability to test the null hypothesis of market efficiency is then severely constrained by data insufficiency (Barrett 2001, p.29).

Barrett (1996, p.825) proposes a hierarchy of all market analysis methods: those who use only price data (that cannot permit normative inference, Barrett 1996, p.828); those who combine transaction costs and price data; those who combine trade flows and price data. In fact, trade reveals the expected relative profitability of market arbitrage; by definition, the observation of trade implies integration (Barrett 1996, p.827). For linear models relying on price data, only, to be consistent with efficient arbitrage, trade is assumed to be continuous and unidirectional[38].

As reported in paragraph 2.2 (see also paragraph 3.2.3) Barrett (2001, p.19-20) and Barrett and Li (2002, p.292) claim that traditional analysis typically confuses two concepts: *market integration*, which reflects the tradability of products between spatially distinct markets, irrespective of spatial market equilibrium, and *competitive market equilibrium*, in which extraordinary profits are exhausted by competitive pressures, regardless of physical trade flows. The simple point location model presented in Fackler and Goodwin (2002), which is highly stylized, can generate any of the tests proposed by literature; but in interpreting the results of these so called "market integration" tests (the term "integration" here refers to the transmission of shocks between prices), the null hypothesis is that they are both efficient and perfectly integrated. Rejections of the tests are

[38] For many products trade is not only discontinuous but reverses direction, often seasonally, or is contemporaneously bidirectional due to intra-industry trade (in which case the products are not truly homogeneous and aggregation becomes an issue; Barrett 2001, p. 23). In markets in which transport links break down in some periods, a linear dynamic regression specification is probably not the most appropriate.

inherently incapable of determining if it is because of inefficiency in an integrated market or lack of perfect integration (Fackler and Goodwin 2001, p.1008; p.1016).

Conforti (2004, p.3) notices that the prevailing approach in many papers is non-structural, and just aims at testing the consistency of the empirical evidence with the competitive framework, without proposing an explicit behavioural model; many econometric applications have analyzed mostly the dynamics of price transmission, while elaborating less from a theoretical point of view. A structural approach is on the other side more frequent in the analysis of vertical price transmission.

A final consideration concerns issues related to how policy intervention is dealt with. Indeed, for agricultural commodities, border and domestic policies are in many cases expected to play a strong role. But, even if price transmission in agricultural commodities markets is an issue which has received considerable attention (Fackler and Goodwin 2001), despite the use of increasingly sophisticated econometric techniques appropriate concern for policy factors is still missing.

Some modelling approaches, on the other side, while making explicit use of policy variables in price transmission equations, still rely on some simplistic assumptions, like the "small country hypothesis" for the EU in the AGMEMOD model: the EU is assumed not to have an influence on world prices, which enter the model exogenously (AGMEMOD Partnership 2007a; 2007b).

Appropriate concern for policy factors seems to be of crucial importance for some agricultural commodities, as will be clear in chapter 4. The objective of this work is to develop a consistent theoretical framework explicitly considering policy variables and policy regimes (chapter 4), while empirical applications will be provided in chapters 5 and 6.

3.5 Concluding remarks

To test for price co-movement, a variety of empirical methods have been used as soon as new econometric techniques were developed: simple regression and correlation analysis, dynamic regression models, switching regime models and rational expectation models. In particular, amongst the dynamic models, cointegration techniques soon had an intuitive appeal for the study of price transmission, as they allow to disentangle short and long run dynamics. These models have been extensively used and developed towards several directions (such as the introduction of thresholds, the possibility of asymmetric adjustment, the introduction of structural breaks).

For agricultural commodities, price transmission mechanisms have been extensively studied: most works basically aim at verifying whether the LOP holds

between different markets, analyzing the existence of co-movement between prices in different locations.

However, despite an abundant literature, empirical evidence is nonetheless mixed. While, on the one side (see paragraph 2.3), the hypothesis needed for the Law of One Price to hold are quite restrictive and unlikely to hold in practice, since many factors can prevent prices from convergence, on the other, when empirical tests rely on prices only, as it is often the case, it is just not possible to test separately the efficiency hypothesis and the strong assumptions underpinning the model specifications (Barrett 2001, p.29).

Moreover, for many agricultural commodities, policy intervention strongly affects international price transmission.

This aspect has not been adequately assessed in most empirical applications; the introduction of policy variables and policy regimes has often been considered only marginally. Empirical simulation models, in turn, explicitly consider policy factors, but they often rely on simplistic assumptions, like the exogeneity of one of the prices.

To bridge this gap, appropriate techniques should be used to remove these simplistic hypothesis, while, at the same time, including policy regimes in the model.

In the remainder of this work, by using cointegration techniques (which allow to remove the exogeneity hypothesis and to separate short and long run dynamics), an explicit attempt will then be made at introducing policy regimes by developing a consistent theoretical framework. Accordingly, different empirical models will be derived.

4 INTERNATIONAL SOFT WHEAT MARKETS UNDER POLICY INTERVENTION

4.1 Introduction

In this chapter, a description of what constitutes the case study of this work, i.e. the international markets for soft wheat, is given. Internationally, wheat is the most traded commodity; the US and the EU are amongst the major exporting countries.

Agricultural markets are often subject to considerable policy intervention. For soft wheat, Ghoshray *et al.* (2000, p.3) point out that policy regimes play a significant role in production and export shares. The EU Common Agricultural Policy (CAP) is a case in point, since thanks to its measures during the 1980s the EU emerged as the second largest exporter of wheat, having previously been a net importer.

At the multilateral level, the Uruguay Round Agreement on Agriculture of the World Trade Organization represents the first attempt of implementing a multilateral trade agreement specifically related to agriculture, setting common provisions for all the member countries.

A detailed description of all these policy measures will make clear the need for taking in due account policy intervention when testing for price transmission. This has often not been the case, which might explain a very mixed empirical evidence.

For this reason, a consistent theoretical framework will be developed, which stems from the one outlined in chapter 2, but tries to extend it towards the inclusion of policy regimes.

It will be this framework which will be used to develop the empirical models presented in chapters 5 and 6.

Chapter 4 then represents a connection between the first two ones, where, respectively, the basic concepts necessary for the study of price transmission and the empirical models used in the relevant literature have been revised, and the

following two, devoted to the results of the empirical analysis after an alternative theoretical framework has been proposed.

The chapter is structured as follows. In paragraph 4.2, an introductory overview of the world wheat market is given, by describing international trade flows (paragraph 4.2.1), and by explaining the functioning of EU domestic and border policies for soft wheat, together with the implementation of the Uruguay Round Agreement on Agriculture of the World Trade Organization (paragraph 4.2.2). In paragraph 4.3, a theoretical framework accounting for policy regimes is developed, in order to test for international price transmission by considering the relevant policy measures. In paragraph 4.4, the data for the empirical analysis are presented together with some considerations on product homogeneity, price leadership, assumptions concerning transport costs. Paragraph 4.5 concludes.

4.2 Trade flows and policies in international wheat markets

This work aims at introducing policy variables in cointegration models for the study of international price transmission. The case study is constituted by soft wheat markets in the years 1978-2003. We expect both domestic EU policies and the ones related to international trade agreements to have significantly affected price transmission. To explain why is this the case, a general overview of the functioning of international soft wheat markets, as well as of the relevant policy measures in place, will be given.

In paragraph 4.2.1, the data concerning the world wheat trade over the past 50 years are analyzed; in paragraph 4.2.2, EU and international policy measures are described. They are relevant both to the elaboration of a consistent theoretical framework introducing policy regimes while testing for price transmission (which is the object of paragraph 4.3) and to the empirical analysis which will be performed in chapters 5 and 6.

4.2.1 International soft wheat markets: an overview

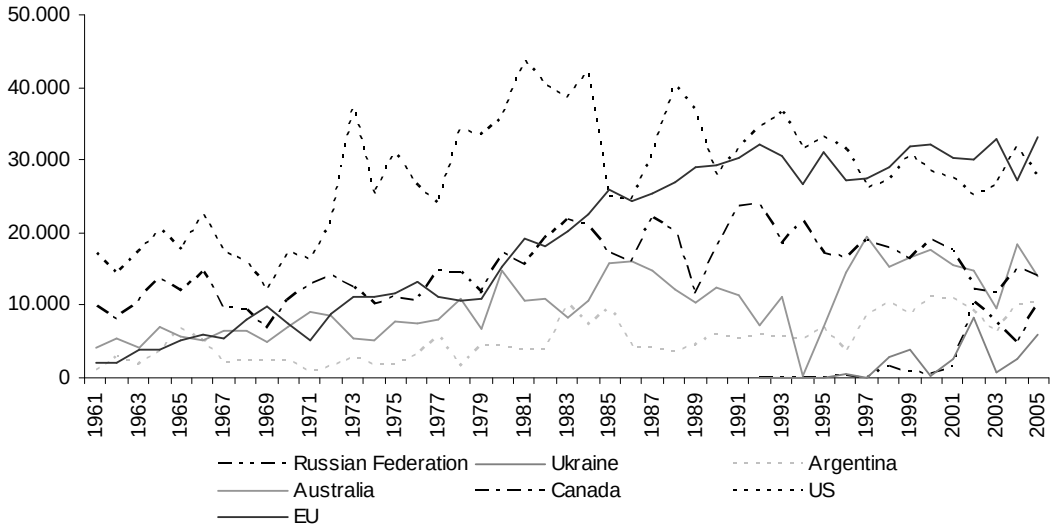
Wheat is a food product which is produced and/or consumed in almost every country in the world. It is a temperate product, grown both in the Northern hemisphere, where the harvest occurs in June to October, and in the Southern Hemisphere, where it is November to January (Ghoshray 2002, p.301).

Soft wheat is a heavily traded commodity; in fact, wheat dominates international cereals exports (Faostat 2004). The main exporting countries are Argentina, Australia, Canada, United States (US) and the European Union (EU); in most years, together they account for more than 80% of world wheat exports (Faostat 2007; see figure 4.1, 4.2 and 4.3). Algeria, Brazil, Egypt, Japan and the EU have been some of the biggest importers in the past years, although the import

demand is much less concentrated than the export supply (USDA 2005, p.10; see figure 4.4).

The EU is a major player, being amongst both the largest exporters and the largest importers; but 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 (Argentina, Australia, Canada, the EU and the US), Mohanty *et al.* (1999) show that there is no distinct leader[39]. However, in this work, the US price will be assumed to be representative of the world one. This seems a reasonable assumption; since 1975-76, the US share of global exports has fluctuated between 25 and 45 percent of the world total (USDA 2005, p.8; figure 4.2).

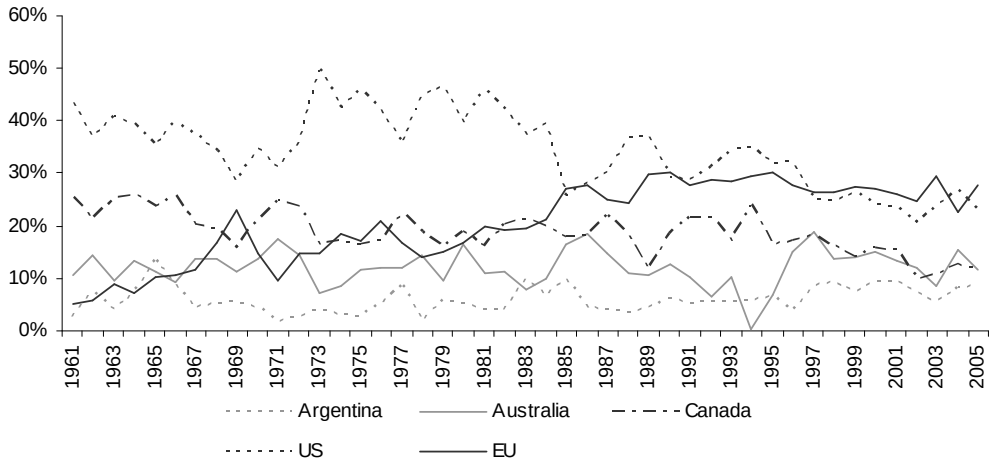
Figure 4.1 World wheat exports (1000 tonnes)



Source: Faostat 2007

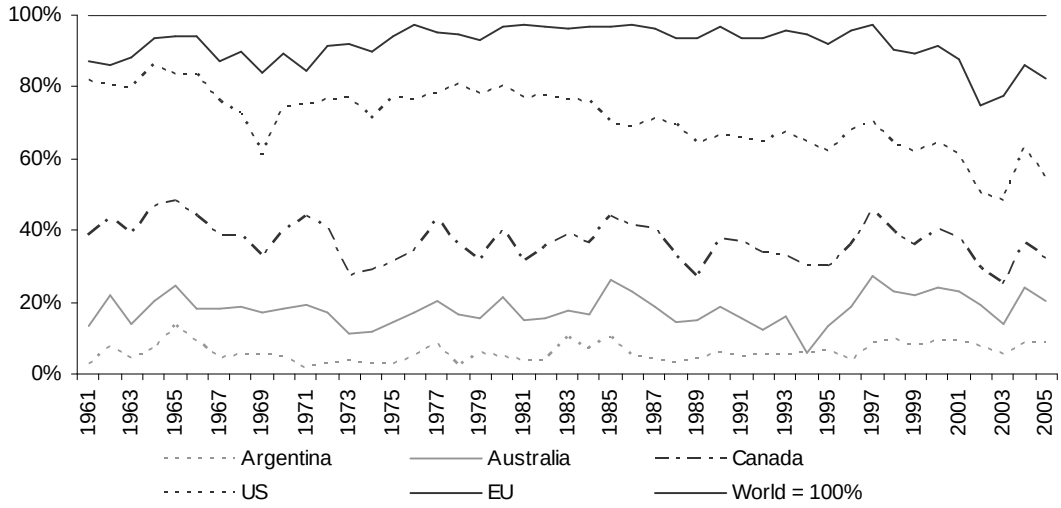
[39] They find out that, while the US price is affected by the Canadian and the Australian prices only, the EU responds to US and Canadian changes but doesn't have any influence on their prices. The EU responds to the US, but the reverse is not true. Previous studies (see Mohanty *et al.* 1999), some 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; they consider the US as a price leader in the international wheat market. Dawson and Sanjuan (2007) find that Canada is the price leader in the relation with the US price for hard wheats. In Ghoshray and Lyoid (2003) Canada is found to be the price leader in the North American market for hard wheat exports.

Figure 4.2 World wheat exports (share of the world total)



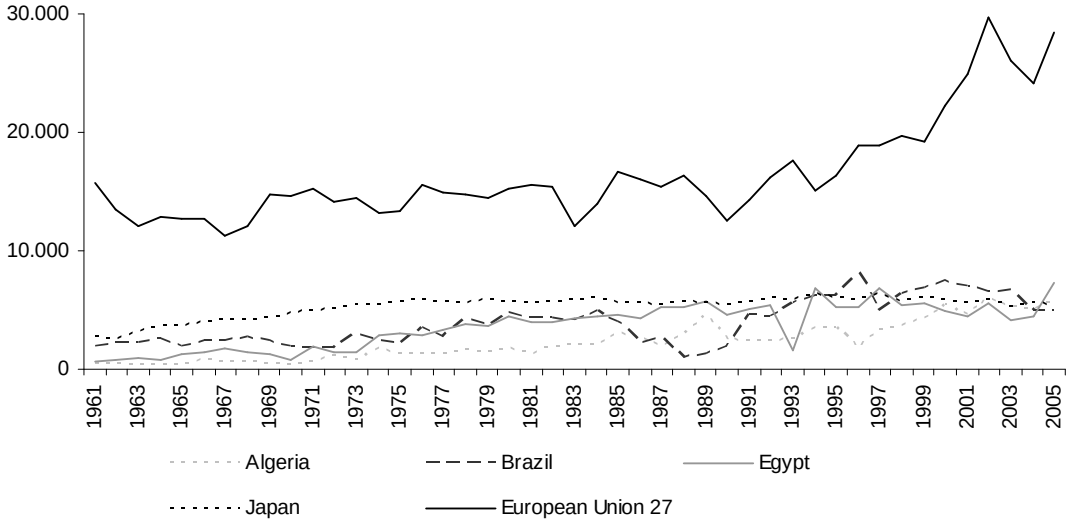
Source: Faostat 2007

Figure 4.3 World wheat exports (cumulated share of the world total)



Source: Faostat 2007

Figure 4.4 World wheat imports (1000 tonnes)



Source: Faostat 2007

In general, in the past century, we expect commodities' price linkages to have improved for a number of reasons. Bukenya and Labys (2005, p. 302) mention the gains in information technology which have improved the dissemination of demand and supply conditions across markets; in some cases (e.g. metals), the centralization of the commodity markets; the increased coordination between central bank activities; the greater congruency between national and international business cycles; the improved liberalization.

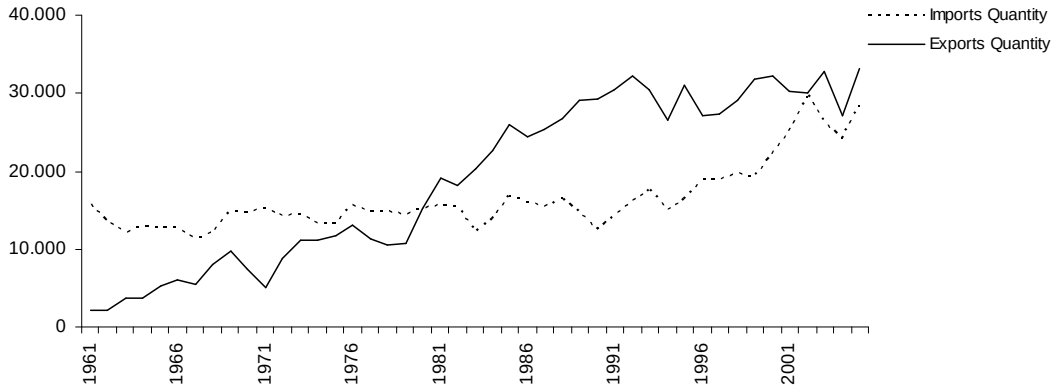
However, it is well known that agricultural markets are often subject to considerable policy intervention. This creates a gap between world and domestic prices, and generates cross-country variations (Mundlak and Larson 1992, p. 399).

In the case of soft wheat, Ghoshray *et al.* (2000, p.3) point out that policy regimes play a significant role in production and export shares. To this respect, the EU Common Agricultural Policy is a case in point, since thanks to its measures, during the 1980s, the EU emerged as the second largest exporter of wheat, having previously been a net importer (as it is shown in figures 4.5 and 4.6). In general, also other governments have a role in developing strategic responses[40] (Ghoshray 2002): for example, Canadian and Australian Wheat

[40] Due to product differentiation, market concentration (it is dominated by 5 major exporters) and government intervention, the world wheat market has also been considered as imperfectly competitive. For example, Ghoshray (2002) uses cointegration models allowing for asymmetric adjustments, assuming that prices will respond more quickly to others' prices increases than decreases.

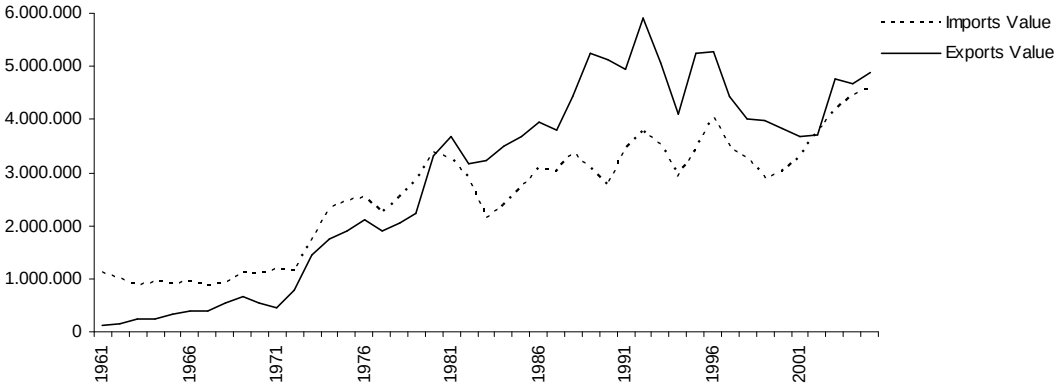
Boards, the marketing institutions of these countries, are responsible for most of their exports; US and EU policy actions, which will be examined later on in this paragraph, are very likely to have an influence on each other's agricultural policies as well as on world market prices (Barassi and Ghoshray 2007, p.79).

Figure 4.5 EU wheat imports and exports (.000 tonnes)



Source: Faostat 2007

Figure 4.6 EU wheat imports and exports (.000 dollars)



Source: Faostat 2007

4.2.2 European and international trade policies for soft wheat

In order to understand how the CAP is likely to have influenced international price transmission mechanisms, a short description of its main instruments, as well as an examination of the major changes it underwent in the past 30 years, are necessary. Indeed, the CAP has evolved considerably, moving from a heavy market regulation policy to the use of less distortive instruments, due to various forces which have contributed to make it go towards the direction most third countries desired (Anania 2007b, p. 1). These factors were concerns regarding its increasing financial costs and the competition with other sectors in the EU budget allocation, but also the growing exploitation of natural resources implied by supply-supporting policies, their inadequacy to the new needs of agricultural and rural areas (such as the protection of the environment), and the growing international pressure. All these reasons, over the years and especially in the past decade, caused an increasing share of the total expenditure to be removed from aids coupled to production, and from market intervention, and then allocated to decoupled aids, i.e. subsidies not requiring production but only eligible land, provided the respect of the environmental cross-compliance.

In this reform process, which is still underway, some major periods can be identified (Thompson 1999, p.5):

1971:01-1988:6 These years are characterized by the full functioning of the so-called Common Market Organizations (CMOs), that regulated the markets of various agricultural products. If necessary, to ensure that domestic prices never fell below the minimum guaranteed price on the domestic market (called "intervention price"), the European Commission, by means of private or public stockholding, would implement the withdrawal from the market of the excess supply. When not stocked, transformed, destroyed or even delivered through public nutrition programmes, the excess supply would be sold in international markets with "export refunds" paid to exporters, meant to cover at least the difference between the intervention and the world price[41]. To restrict imports (otherwise, intervention prices above the world price would result in large inflows of product and unsustainably high stocks levels), "variable levies" were set as the difference between the world price and the so called "entry price". The price on the internal market was then bounded between the intervention price and the entry price (figure 4.7). Since intervention prices were constantly set above what would have been the market equilibrium prices this system caused, in the 1980s,

[41] Actually, export subsidies are calculated as the difference between a current domestic EU price and the world one. However, to be effective, in theory, export refunds have to cover at least the difference between the intervention price -the minimum guaranteed price on the EU market- and the world one. For the sake of simplicity, this is what is assumed henceforth.

surpluses growth and budgetary costs escalation. On world markets, the EU emerged as a net exporter having previously been a net importer.

1988:07-1993:06 First CMO reforms concerning arable crops were put in place. New measures aimed at reducing the production surpluses and the budgetary costs were 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 prices would be reduced the following year), and voluntary set-aside for cereals.

1993:07-2000:06 Implemented in July, 1993, the MacSharry reform is the first structural change of the CAP. In fact, the small effects of the reforms already implemented made necessary the introduction of new policy measures with the objective of ensuring a progressive return to market mechanisms. The MacSharry reform represents the first move from a consumer financed (through high prices) to a taxpayers financed regime (through compensatory transfers; Thompson and Bohl 2002, p.7). The main concerns for EU policy makers concerning the grain sector included high growth in output relative to growth in demand, increase in export subsidies and stocks, and rising cost of the programmes (Mahe 1996, p. 1316). 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, p. 1317); 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.

With the MacSharry reform, for grains, substantial cuts in intervention prices were implemented (-30% over a three year period) to re-align them with the world prices; compensations to farmers through direct subsidies per hectare were put in place. Set aside requirements were established for producers of more than 92 tonnes[42].

Due to the decrease in intervention prices, which became closer to the world ones, both variable levies and export subsidies were reduced; however, this “old” system, on the other side, kept on insulating the EU market from the world one.

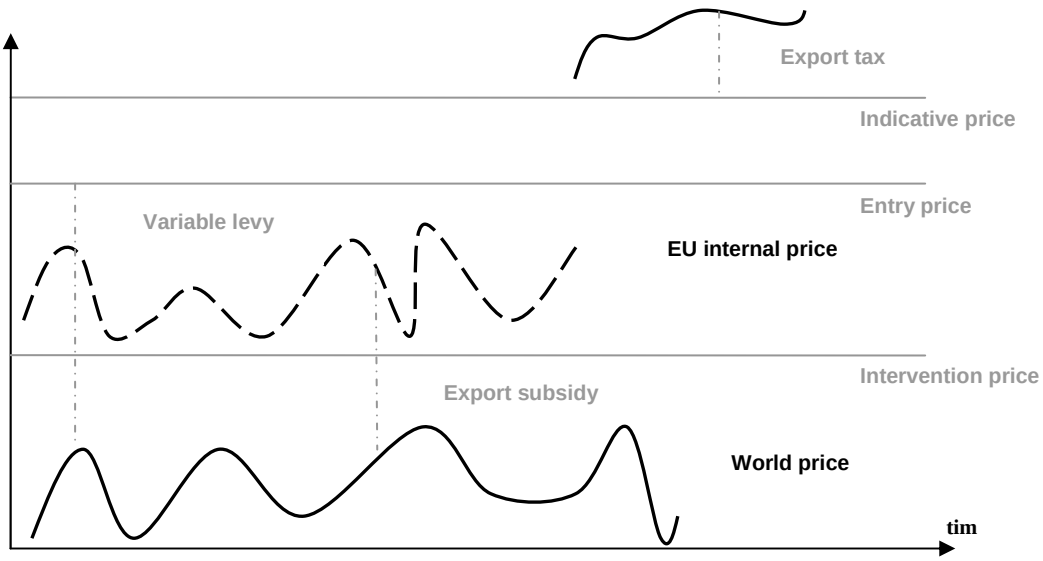
2000:06- … The Agenda 2000 reform continued along the path undertaken in 1992. Both a 15% reduction in two years of the intervention price for cereals and the introduction of decoupled payments were decided. The set-aside regime remained in force, too. In 2003, the Fischler reform, which was meant to be the “Mid Term Review” of Agenda 2000, turned out to be, instead, the most important single step[43] in the reform process of the CAP (Anania 2007b, p. 1). It

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

[43] Agenda 2000 and the Fischler Reform are put together despite the much bigger impact of the latter due to the lack of data in the most recent years.

introduced "fully decoupled" support for most of the CAP payments (this is expected to have a bigger impact in terms of the reduction of the distortionary effects of the CAP than in terms of the reduction in farm support; Anania 2007, p. 4). All sectors untouched by the reform have undergone, or are currently undergoing, similar changes. As far as market regulations for wheat are concerned, intervention prices were not further reduced, but the monthly seasonal adjustments applied to them were halved.

Figure 4.7 The Common Market Organization for cereals



EU trade policies for soft wheat did elicit responses from other countries. In 1985, the US retaliated to the EU heavy use of export refunds with the Export Enhancement Program (EEP), a targeted export subsidy program for wheat. But the EEP didn't alter substantially the EU-US price relationship, since, after its introduction, the EU might well have set its export subsidies in relation to US prices[44] (Mohanty *et al.* 1999, p.27). 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 (Barassi and Ghoshray 2007, p.79).

Moving to agricultural trade agreements, the most relevant event in the period examined has been the institution of the World Trade Organization (WTO) in

[44] Brooks *et al*. (1990) show that also 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.

1995, and the following implementation of the URAA, the first multilateral agreement explicitly referred to agriculture.

The URAA regulates both domestic support and export subsidies. For the first, the EU always kept a safe margin with respect to the maximum ceilings allowed for distorsive support; for the latter, the limits established for wheat were never binding for the EU (Anania 2007b).

As far as market access is concerned, the URAA abolished the possibility of keeping threshold prices and variable levies; with the "tariffication", all border measures were converted into import duties to be lowered in the following six years[45]. The EU has then consolidated the duties for cereals under the General Agreement on Tariffs and Trade (GATT) agreement.

However, some modifications were introduced in the so-called Blair House Agreement between the EU and the US. Due to the high levels of protections in the reference period chosen for the tariffication (the years 1986-1988), with regard to the main cereals, the tariff equivalents notified at the WTO were so high that a system for restricting duties was deemed necessary for the EU, in order to put a ceiling on the import entry price. The import entry price 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 (see table 4.1)

Table 4.1 Elements for calculating the duty on cereals

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.

There are five different wheat categories for which import duties are established. For all of these categories a US price is used to calculate the reference price. The duties are set every two weeks.

*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. The day's exchange rate is used for conversion.

Source: Gallezot 2007, p.16

[45] The reduction agreed was 36%, with a minimum reduction of 15% for each "8 digit" product of the Harmonized System codification; for agriculture, customs duties are for the most part specific duties, i.e., expressed in euro/t.

Both the changes in the CAP and those implied by the URAA implementation are likely to have deeply affected international price transmission for soft wheat.

Summing up, it is fairly clear that in the past thirty years the EU price has somehow been "isolated" from world markets, bounded from below by the intervention price (export subsidies[46] are meant to cover the difference between domestic prices and the world ones) and from above by a variable levy and, afterwards, by the "155% rule" established during the URAA. The most relevant EU trade policies for cereals are then export refunds and variable levies, with the objective of ensuring the maintenance of high prices on domestic markets.

In addition to the CAP reforms and the URAA, to the analysis that will follow, at least another event is worth of mention. Rather awkwardly, in 2002, unusual circumstances on world markets caused the EU price to be quite low despite the peak in US prices. This can be explained by the way EU import duties are calculated (Binfield 2002; Morgan 2003). In a nutshell, the shortages of milling quality wheat, due to poor harvests in US, Canada and Australia, had resulted in an increase of the US price for wheat. Since this price is used as the reference price to calculate EU import duties, the duties had fallen to zero. At the same time, large crops in Russia and the Ukraine increased the supply of feed quality wheat, that could enter the EU duty free. In practice, due to the lack of an import duty, and thanks to their proximity, Russian and Ukrainian wheat exports could enter into the EU at prices lower than domestic EU wheat prices, depressing them. The EU reckoned the situation as serious enough to change the EU protection system for cereals: on January 1, 2003, a Tariff Rate Quota (TRQ) of less than 3.0 Mt[47] was introduced for medium and low quality grains that are not traditionally imported in the EU from the US or Canada (the EU's major grain trading partners that are part of the WTO). The situation is then unlikely to occur again.

In the very last years, commodity world markets have been characterized by soaring food prices because of both transitory (like adverse meteorological conditions) and structural factors (like the increasing food demand from emerging economies; the growing energy costs; the biofuels demand which creates a pressure on crops production; the reduction in stocks due to policy reforms in developed countries; FAO 2008).

Even if their treatment goes beyond the scope of this work, it must be noticed that high food prices stimulated a renewed debate on the appropriateness of market intervention policies, already going on in the so-called CAP "Health Check" discussion in the EU (European Commission COM(207) 722, and COM(2008) 306/4), which should be concluded by the end of 2008 and will set the new rules to regulate the agricultural sector up to 2013. Substantial cutbacks

[46] 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.

[47] The in-quota tariff was set at euro 12/t; the after quota tariff was set at euro 95/t.

of the intervention system are currently being considered, which the Commission basically propose to function only as a "safety net" and only for soft wheat. Within this debate, an analysis of the effects that EU intervention policies have had on international price transmission is crucial. It is clear that their appropriate assessment requires adequate instruments of analysis.

At the same time, dramatic changes have occurred in international commercial policy: the World Bank and FAO estimates that almost 50 countries have put in place emergency measures to confront the rise in food prices[48].

For example, as a reaction to tight world markets and record high prices for cereals, the EU has decided to suspend import duties on cereals[49] for the whole 2007/2008 marketing year (Reg CE 1/2008); to date, it has been decided that the suspension will hold up to the end of the 2008/2009 campaign (June 30, 2009) unless market conditions justify their reintroduction before its end.

In turn, large exporters (such as Argentina, Russia, and Ukraine) have imposed taxes on wheat exports, with the same objective of controlling internal prices.

The evaluation of the impact of such measures on price transmission makes clear the need of an appropriate theoretical framework for the analysis.

However, how much of the rise in agricultural prices is due to short-run factors, or to speculative behaviour, or to a real structural break in their long-run downward trend, as well as which effect this is going to cause on international trade policies (whether a rise in protectionism, or, less likely, an increased liberalization), remain open questions.

4.3 Policy variables and price transmission: a theoretical framework

On the basis of the general overview given in paragraph 4.2.2, it is clear that, given that the policy actions have a significant role in the world soft wheat market, policy variables, and policy regimes, will need to be adequately considered while investigating price transmission.

Despite the fact that the role played by policy barriers in influencing price transmission is well recognized by the existing literature, as anticipated in chapter 3, in empirical estimates for agricultural commodities this is often not the case, whereas econometric models often make use of policy variables but, in turn, disregarding the time series properties of the data and making sometimes simplistic assumptions (AGMEMOD Partnership 2007a; 2007b).

Here, within the theoretical framework for price transmission presented in chapter 2, the explicit introduction of policy variables is proposed. The development of a sound and consistent theoretical framework will then be

[48] See http://www.fao.org/giews/english/policy/index.htm for a scheme of the various measures in place.

[49] All cereals expect oats, buckwheat and millet. Due to modest harvests in 2007, and low stocks, in 2007/2008 the EU has become a net importer.

completed by the elaboration of various empirical models, all with specific reference to cointegration models (which have been dealt with in paragraph 3.3), that will be used in the analysis which will follow in chapters 5 and 6.

Following the existing literature (for example, Thompson and Bohl 1999, Thompson et al. 2000, Thompson *et al.* 2002a and 2002b, Verga and Zuppiroli 2003), we will assume that the LOP hypothesis do apply on international soft wheat markets, and on this basis we will try and focus on the effects of policy intervention. In a certain way, our main hypothesis is then that the LOP holds but for policy intervention.

As explained in the previous paragraphs, analyzing how policy regimes affect international price transmission is a crucial issue. This analysis assumes particular relevance in light of the ongoing policy debate concerning the viability and the effectiveness of intervention policies on agricultural commodity markets.

The fundamental problem to tackle is the following one. Despite the fact that EU domestic and border policies aimed at insulating domestic markets from the world ones have, in practice, always been in place, we nonetheless aim at building a sensible framework to test for co-movement between EU domestic prices and the world ones. Indeed, because of trade barriers (namely, import variable levies and export restitutions), we expect domestic EU prices to be isolated from the world ones[50]; while, in turn, we suppose that EU export (i.e., freight on board, fob) prices are cointegrated with the world ones right because of the use of export restitutions (Barassi and Ghoshray 2007; Ghoshray 2002; Ghoshray *et al.* 2000). In fact, export subsidies are set with the objective of making EU exports competitive on world markets, bridging the gap between the internal (high) price and the world one.

Basically, different price transmission mechanisms will be in place according to the various combinations of the domestic and border policies described in the previous paragraph. A simple scheme is reported in table 4.2.

Intuitively, since in the period covered by the analysis the EU is a net exporter[51] (figures 4.5 ad 4.6) we expect that what occurs on the export side is more relevant to the analysis. We could think that the EU price will interact with the world one only if it is high enough to make export restitutions go to zero; i.e., at least higher than the intervention price.

[50] Van Meijil and van Tongeren (2002, p. 457) model the European market price as a weighted average of the entry and the intervention price, depending on the entity of net exports. The more the net exports, the closer the market price to the intervention one, and the other way round.

[51] This is true also for France, whose price will be assumed to be the representative price for the EU. Its self-sufficiency rate for soft wheat, between 1978 and 2000, was on average 218%, with a minimum of 168% and a maximum of 277% (Eurostat).

Table 4.2 Summary of the relationships between domestic EU and world prices.

Trade position		***Relevant price for EU price formation***
Import side		
Pre - URAA	world price < entry price	entry price
	world price > entry price	world price
Post - URAA	world price plus tariff < 155% intervention price (almost never the case)	world price plus tariff
	world price plus tariff > 155% intervention price	155% intervention price
	world price > 155% intervention price	world price
Export side		
world price < intervention price (export restitutions are paid)		intervention price
world price > intervention price		world price

Going back to table 4.2, on the import side, the domestic EU price will follow the world one only if the world price is high enough to make the variable levy go to zero, deactivating the border protection mechanism; otherwise, it will follow the entry price. After 1995, the functioning of the mechanisms is slightly more complicated, but analogous. If the world price plus tariff is below the 155% of the intervention price, then "pure tariffication" would apply, and the EU price would follow the world price plus tariff: but this was actually never the case (Gallezot 2007, p.16). Instead, what normally happens is that the world price plus tariff exceeds 155% of the intervention price; according to the "155% rule", the 155% of the intervention price acts as an entry price, and a variable levy covering the difference between the world price and the 155% of the intervention price is in place. We expect the EU price to be related to such an entry price. Should the world price be higher than 155% of the intervention price, then we would expect the variable levy to go to zero and the EU price to follow the world one.

On the export side, thanks to export subsidies[52], we can assume that the domestic EU price follows the highest between the intervention and the world price. The intervention price operates as a threshold below which the world price becomes inactive. If the world price is below the intervention price, indeed, export restitutions will prevent the EU price from falling; only if the world price is above the intervention price export restitutions will go to zero, and the EU price can be assumed to interact with the world market one.

[52] For the sake of simplicity, as they were never binding for the EU, we don't consider WTO limits on export subsidies.

Summing up, on the import side the entry price is the threshold above which there is co-movement between the EU and the world price, while on the export side we might think that it is the intervention price that performs as such[53]. Since the first condition is more restrictive[54], only if the world price is above both the intervention and the entry price we expect that it positively interact (or, analogously, is cointegrated) with the European one, because of the annulment of both export restitutions and import variable levies. Nevertheless, in this case we would end testing for the presence of co-movement between the EU price and the entry one, since, in practice, the world price was above the entry price only for a very short number of months[55]. In practice, the entry and the intervention price are very correlated, and proportional after 1995 because of the "155% rule". Testing for cointegration between the European and the entry price would then be equal, but for a constant term, to testing for cointegration between the European and the intervention price.

We might instead verify whether the EU price positively interact with the world one also when the latter is below the entry price but above the intervention one, a situation which occurred especially from the MacSharry reform onwards. The intervention price (which can be seen as the very basic instrument of EU agricultural market policy in the past[56]) is a threshold under which the world price has no influence on the intervention price.

In this way, we are implicitly identifying two regimes: if the world price is below the intervention one, then the intervention mechanism is in place, and we expect the EU price to follow the behaviour of the intervention one. If, in turn, the world price is above the intervention price, we might think that the EU price is free to follow the behaviour of the world one, provided export taxes will not eventually prevent him from rising too much.

In chapter 5, within this simple theoretical framework, different analytical models will be developed. They still are cointegration models aimed at testing for price co-movement (basically, the LOP validity), but at the same time provide alternative schemes for the analysis of price transmission in international markets accounting for policy regime changes.

[53] Actually, if both "thresholds" are binding (the word price is below them), the price transmission elasticity of the world to the EU price is zero; but the one from the intervention price to the world one is negative. In fact, the entry price and the export restitutions prevent the world price from affecting the EU price, but if the intervention (and then, the entry) price goes up, it will lower the world price through an increased export supply (Thompson 2000, p.720). We won't consider this case in this analysis, since our objective it to verify whether, under certain conditions, there is co-movement between the EU and the world price, or in other words, a positive price transmission elasticity exists.

[54] The entry price is by definition higher than the intervention price.

[55] In this analysis, 14 months over 25 years.

[56] In fact, it is basically from the intervention price that other domestic and commercial policy measures (export restitutions, entry prices and then import variable levies) are established.

To the author's knowledge, to date no such attempts have been made. Policy regimes have been accounted for in the revised literature by inserting dummy variables in linear regressions, which allow the intercept of the price equation or the transmission parameter to vary (Thompson *et al.* 2000; Thompson *et al.* 2002b; Vollrath and Hallahan 2006). In cointegration models, either dummies accounting for structural breaks are put outside the cointegration vector (Hui-Shung and Griffith 1998) or, most commonly, the analysis is repeated on the sample split into sub-samples according to policy regime changes[57] (Barassi and Ghoshray 2007; Thompson and Bohl 1999; see annex A for further references).

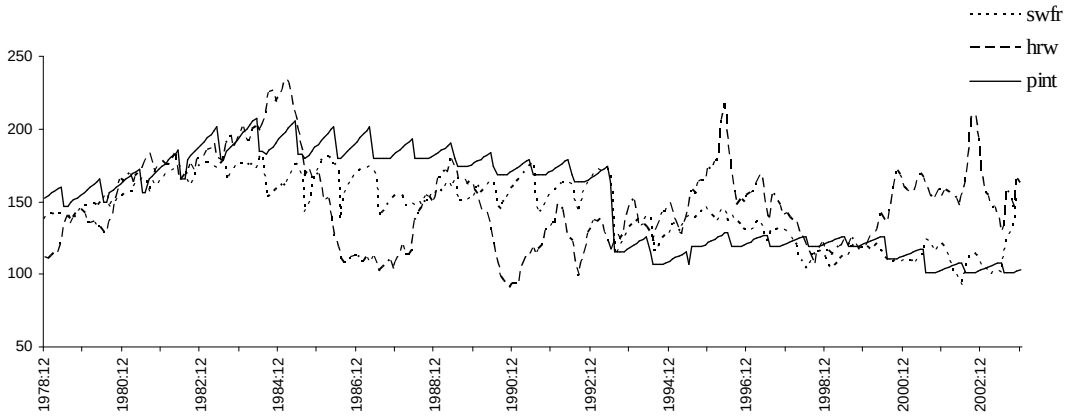
Cointegration models with structural breaks accounting for policy regime changes, which represent another method of considering policy regimes while testing for price transmission, (Dawson *et al.* 2006; Dawson and Sanjuan 2006) will instead be dealt with in chapter 6.

4.4 The data

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) are used (figure 4.8). Soft wheat French prices (*swfr*) are available on-line from the Eurostat database; unfortunately, while the Eurostat database allows to go back in time, no more recent observations were accessible. Data concerning US Gulf fob Hard Red Winter (HRW) wheat 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 European cif prices (*hrw*). 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[58; 59]. Eurostat monthly bilateral exchange rates have been used to convert in euro prices and freight rates expressed in US dollars[60]. 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 (see paragraph 3.4).

[57] Alternatively, Thompson *et al*. 2002 propose to include additional observations not to reduce the power of the tests.
[58] The monthly adjustments for 1985 have been calculated as arithmetical average.
[59] 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.
[60] Though the use of US dollars is more common in studying international agricultural markets, euro were preferred because of the analysis of EU policies which will be performed.

Figure 4.8 Soft wheat French price, US HRW price, intervention price (euro/t)



Source: Eurostat, International Grains Council, European Commission regulations

The three price series have quite a different behaviour. The US price presents a cyclical pattern, which is quite common for agricultural commodities, and is mostly due to the inherent characters of agricultural production (Hallett 1968).

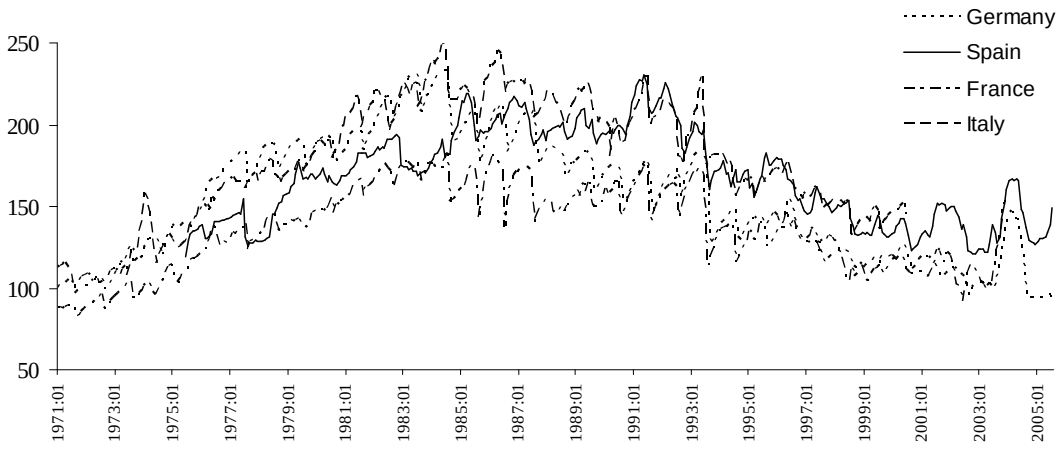
The French price appears to track closely the intervention one up to 1992, when its fluctuations seem to follow more the ones of the US price.

What emerges even from this simple visual inspection, is that the MacSharry reform clearly marks a turning point in the behaviour of the series. The intervention price values were constantly set at higher levels during the 1980s and up to 1992 when, in correspondence to the implementation of the MacSharry reform, a 30% reduction of their level was agreed (we can notice their strong seasonality, as seasonal adjustments are established monthly to cover the increase in storage costs; figure 4.8). Indeed, intervention prices switch from being constantly above to constantly below the world ones.

The French price is assumed to be the representative EU price, considering the fact that France is the biggest European wheat producer (Eurostat). In figure 4.9, moreover, we can see how the European prices from the four major agricultural producer countries move together over the past 30 years[61]; once again, we can notice how, during the eighties, high intervention prices allowed to keep high the prices in the European markets.

[61] As far as Italy and France (amongst the most important EU countries in this respect) are concerned, Verga and Zuppiroli (2003, p.27) find out that the strong relationships existing between them are not affected by the MacSharry Reform, which would imply that, besides sharing the same political context, the two countries are tied by strong commercial linkages. It must be said that older studies (Zanias 1993; Gjolberg *et al.* 1996) find evidence of non-integrated wheat markets, unless prices are corrected by Monetary Compensatory Amounts, which operated as export subsidies/taxes in the intra-EU trade as the "green exchange rate" was fixed and might diverge from the actual market exchange rate.

Figure 4.9 Soft wheat prices in Germany, Spain, France and Italy (euro/t)



Source: Eurostat

The selection of an appropriate "world" price for European markets is not straightforward.

In this work, the US Hard Red Winter wheat is assumed to be the relevant representative world price for EU markets, as it is normally the case in econometric models (AGMEMOD 2007a, 2007b). Moreover, it is assumed as a reference price also in practice. Indeed, its quotations are used to determine the import duties by the European Commission for medium quality EU soft wheat (Gallezot 2007, p.18).

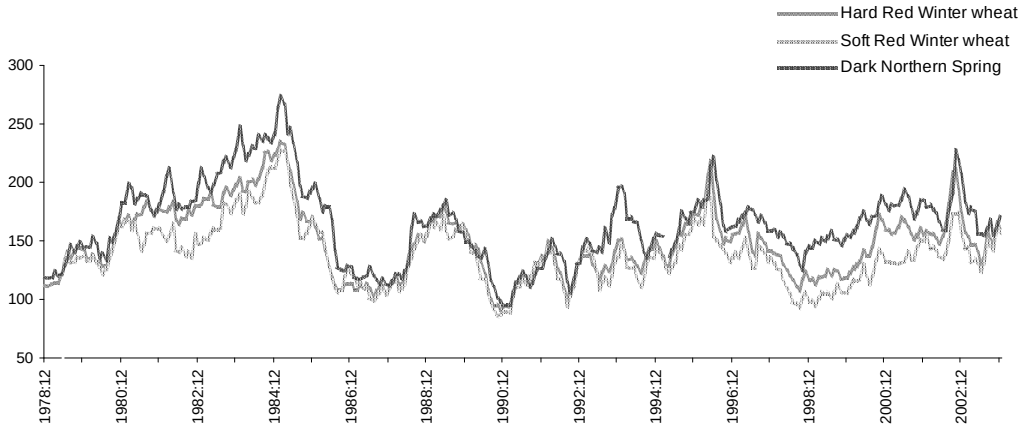
However, it might be argued that wheat is a non homogeneous product. This is a well known issue (for a table of the various classes of wheat see Ghoshray 2000, p. 5; Ghoshray and Lyoid 2003, p.28; Ghoshray 2002, p.303): Larue already (1991), shows that wheat is differentiated by end use and by country of origin, and that wheat protein content has a significant influence on prices. Quality differences, namely protein content[62], could influence international price linkages by making varieties of wheat imperfect substitutes of one another (Mohanty *et al.* 1999, p. 113; Ghoshray and Lyoid 2003); nevertheless, 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.

[62] Wheat type depends on hardness, a milling characteristic, (we have hard wheats, characterised by a high protein content, and soft wheats), and on dough strength, a baking quality (it determines the end use of wheat; strong wheats can produce bread with a large loaf volume and a good texture, differently from weak wheats). Canadian and Australian wheats are often considered to be of better quality than US wheats, which in turn are considered to be of higher quality than EU and Argentinean wheats; anyway, the substitution between the different classes of wheats has been made easier by technical innovations (Ghoshray 2002, p. 302).

While the EU soft wheat is a medium protein wheat, weak, the Hard Red Winter wheat has a higher protein content and is classified as medium strong. We mentioned already that the Hard Red Winter wheat has commonly been used to represent the world price, but it has also been argued that the Soft Red Winter wheat, which has a lower protein content, would indeed be a closer substitute for European soft wheat (Barassi and Ghoshray 2007; Verga and Zuppiroli 2003).

However, remembering the considerations mentioned above, and noting that, when the behaviours of different US wheat prices are compared, they seem to be very close but for a constant term (figure 4.10; the Dark Northern spring, used for the analysis in Thompson and Bohl (1999) has an even higher protein content than the Hard Red Winter Wheat), the use of the Hard Red Winter wheat as world representative seems a reasonable assumption.

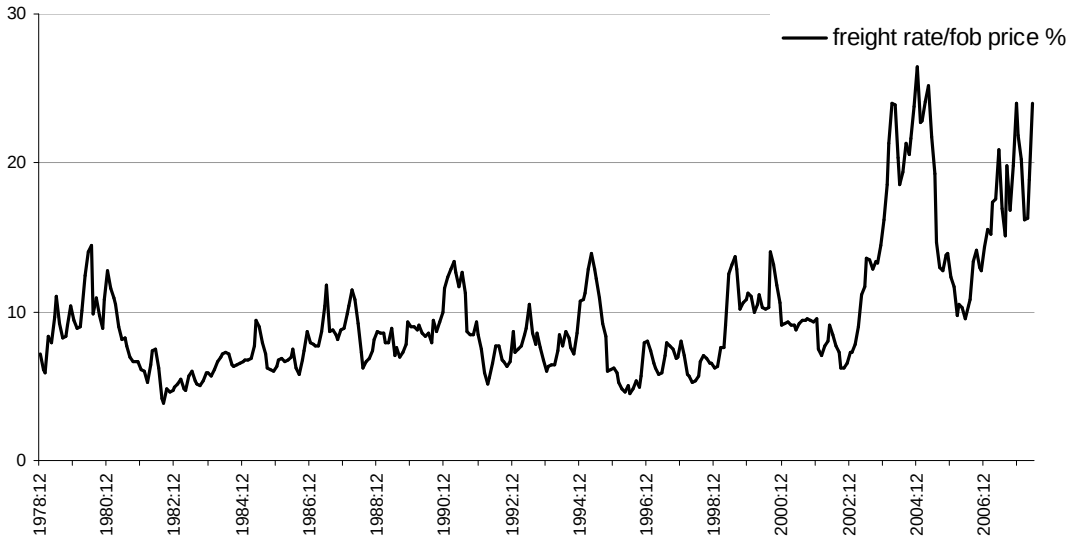
Figure 4.10 Various US wheat prices (euro/t)



Source: International Grains Council

Quality differences to a higher level of detail are not explicitly considered and, together with transport costs for which data are not accessible, the strong assumption for them to be stationary around a constant proportion of prices is made (see paragraphs 2.3, 3.2.1 and 3.4). Although very strict, this assumption seems not to be too far from reality in the period considered in the analysis. Indeed, the ratio between freight rates and fob US prices oscillates in a quite limited range, before starting a dramatic increase in 2004 (figure 4.11). This could be taken as evidence of a substantial stability of the proportion of transportation costs on prices in the years considered in the analysis.

Figure 4.11 Freight rates expressed as a percentage of fob prices



Source: International Grains Council

Policy regime changes will be taken into account through the various schemes which will be developed; all other components of possible market deviations from the LOP are not available, as it happens in most of the literature revised. As Thompson and Bohl (2002, p.1045) note, the analysis can be considered within what Barrett (1996, p. 825) defines as "level 2" of market analysis, which "more closely resembles spatial equilibrium theory".

Some descriptive statistics of the data, according to the major CAP changes and the URAA implementation, are provided in table 4.3.

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 decreases over time; *swfr*'s variability, which is always smaller than the *hrw's* one, decreases, too, but only 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 argue that *swfr's* variability was kept low thanks to protectionist EU agricultural policies, and then rose after 1993, when substantial policy reforms were introduced (Thompson 1999, p. 8). Verga and Zuppiroli (2003, p.28) 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, p. 10).

Table 4.3 Descriptive statistics of the price time series (euro/t)

	Time	Average	Standard deviation	Skewness	Kurtosis
Soft wheat French price	**1978:12-2003:12**	**144.16**	**22.67**	**-0.29**	**-1.02**
	1978:12- 1988:06	161.18	12.89	-0.23	-1.24
	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
	1978:12-1995:06	156.51	15.16	-0.54	-0.17
	1995:07-2003:12	120.08	13.81	0.414	-0.68
Intervention price	**1978:12-2003:12**	**151.49**	**32.81**	**-0.13**	**-1.53**
	1978:12- 1988:06	179.14	15.92	-0.34	-0.91
	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
	1978:12-1995:06	169.82	24.40	0.50	-0.42
	1995:07-2003:12	115.74	8.85	-0.47	-1.24
HRW US price	**1978:12-2003:12**	**148.23**	**29.75**	**0.48**	**-0.13**
	1978:12- 1988:06	157.51	36.16	0.13	-1.03
	1988:07-1993:06	133.30	24.32	-0.01	1
	1993:07- 2000:06	140.73	21.25	1.09	1.51
	2000:07- 2003:12	159.16	16.39	1.32	2.57
	1978:12-1995:06	147.59	32.77	0.50	0.89
	1995:07-2003:12	149.49	22.82	0.46	0.38

4.5 Concluding remarks

As anticipated, a general overview of trade flows and price data on international wheat markets, and of the main policies implemented by some of the biggest exporters and in multilateral trade agreements, makes clear that domestic and border policies play a relevant role in influencing price transmission on soft wheat markets.

In this work, the French and the US prices will be studied over the years 1978-2003; they are assumed to represent, respectively, the EU and the US price.

Though, after 1992, the CAP changed considerably, moving from more market oriented mechanisms towards the use of less distortive support, in practice, and although intervention prices were progressively reduced, the same scheme of protection, based on export subsidies and variable levies, kept being in place.

Under certain conditions, this should hinder the world prices from influencing the domestic ones.

In international trade policy, in 1995, the URAA of the WTO aimed at achieved an increased market liberalization; consequently, with an effective "tariffication", international price transmission elasticities should have increased. However, also in this respect, the so-called "155% rule" might well have reduced its effects on soft wheat markets.

In this chapter, all these policy changes have been described in detail, to assess their effects on price transmission in order to develop a consistent theoretical framework. This has the objective of understanding under which conditions the EU and the world prices are expected to interact, provided the presence of the EU domestic and border policies and of the implementation of the URAA of the WTO.

A thorough examination of the various combinations of domestic and international market conditions, considering the various policy measures, suggests that the EU and the US price are expected to be positively correlated only if the latter is above both the entry and the intervention price.

However, this has never been the case. What is argued is then that it is the intervention price which acts as a threshold under which the world price doesn't affect the European one. The European price is expected to interact with the world one only when the latter is above the intervention price; otherwise, it will follow the intervention price.

The relative position of the US and intervention prices allows the identification of different policy regimes, in which the price transmission relations are different depending on the effectiveness of policy intervention, represented through the position of the intervention price.

This theoretical framework represents a development of the one presented in chapter 2 that explicitly addresses the need for taking policy regimes into account while investigating price transmission, and will be used for the development of different empirical models in the next chapters.

The descriptive analysis of the dataset concludes the chapter. Some key facts that emerge are the higher volatility of the world price, if compared to the European one; and the fact EU prices show a constant decrease over time, while the world ones increase again after 1993.

5 EMPIRICAL ANALYSIS: COINTEGRATION MODELS ACCOUNTING FOR POLICY REGIME CHANGES

5.1 Introduction

The aim of this chapter is to introduce policy regimes in the analysis of international price transmission mechanisms.

The focus is on the soft wheat market between the EU and the US, in the years 1978-2003. As explained in the previous chapter, the French price is assumed to be representative of the EU price, whereas the US one of the world price.

Different econometric models will be derived on the basis of the considerations made in paragraph 4.3, where a theoretical framework has been developed with the objective of including policy regimes in the analysis, by combining policy and price data.

It has been demonstrated that the EU domestic price is expected to follow the world one only if it is above the intervention price; otherwise, it is expected to follow the latter one. This claim will here be explored in different ways.

The objective of the analysis is twofold: on the one side, suggesting different econometrical models accounting for policy regime changes, based on the theoretical framework elaborated. To the author's knowledge, to date not such attempts have been made.

On the other, shedding light on a controversial empirical evidence, the analysis aims at verifying if there is co-movement between the EU and the world prices, and how policy changes are likely to have affected this relation.

In fact, in the existing literature either the EU and US prices are found not to interact (Verga and Zuppiroli 2003), or their price transmission elasticity to be positive, although rather small (Thompson and Bohl 1999; Thompson *et al.* 2002a; 2002b), and in this case either influenced (Thompson *et al.* 2002a) or not by international policy regime changes (Thompson *et al.* 2002b). However, in

these works attempts to take policy regimes into account are limited to the inclusion of dummy variables, or the sub-division of the sample.

The chapter is structured as follows. In paragraph 5.2, the empirical models are presented. They are classified in two groups: in paragraph 5.2.1, a composite variable is introduced; in paragraph 5.2.2, three different *ad hoc* "threshold" cointegration models are presented, in which adjustment coefficients and cointegration parameters are assumed to vary depending on the policy regime in place. Paragraph 5.3 concludes.

5.2 The empirical models developed

The objective of this paragraph is to find out innovative ways of considering policy regimes while testing for price transmission and, once this has been made, to determine whether a long-run relationship exists between the French soft wheat price (*swfr*) and the US cif price (*hrw*).

In this paragraph, different cointegration models are developed. Indeed, cointegration models allow to explore both short and lung run dynamics. They all stem from the theoretical framework presented in paragraph 4.3: the basic idea is that the intervention price acts as a threshold above which the EU and the US price can interact.

Different ways of modeling such a relation can be figured out.

At first, a composite variable, equal to the maximum between the intervention and the US price, is introduced in a cointegration model and its relation with the EU price is studied. This is a straightforward representation of the simple assumption that "the EU price is expected to follow the maximum between the intervention and the US price". The creation of this composite variable is a very simple way of modeling the policy regime switches: indeed, also in some econometric models, even if by means of Ordinary Least Squares (OLS) estimates, the relevant European price is assumed to depend on either the US or the intervention price depending on which of them is higher.

In a second step, other models are estimated. This time, through the appropriate use of dummy variables, either the adjustment coefficients or the parameters of the cointegrating vector, or both, are allowed to vary according to the policy regime in place. In other words, these parameters are allowed to take different values if the US or the intervention price is the significant reference price for the EU.

Firstly, a particular cointegration model is estimated, with different adjustments coefficients depending on the observable policy regimes in place; the LOP is imposed between the French price and the highest between the US and the intervention price. Secondly, the price transmission elasticity is allowed to change according to which price the French one is related to (always, the highest between

the intervention and the US price). Finally, both adjustment coefficients and the cointegrating vector parameters are assumed to vary.

The cointegration models presented, though over-simplified, represent an attempt of combining policy and price data. To the author's knowledge, so far no such attempts have been made.

The "composite variable" model, Model 1, is presented in paragraph 5.2.1; whereas the other empirical models, Model 2, Model 3 and Model 4, are presented in paragraph 5.2.2.

5.2.1 Model 1: the use of a composite variable

As anticipated, to take into account the role played by EU policies for soft wheat, some alternative schemes can be proposed.

The first one is the creation of a composite variable, the "EU external reference price" ($wref_t$). Indeed, assuming that the fundamental 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 US (world) price; $wref_t = \max(hrw_t, pint_t)$.

What is argued is that the EU price follows the US price, when it is above the intervention price, and *vice versa*. The intervention price acts as an implicit downward threshold for the US price.

In this paragraph, the empirical study will be carried out both on the whole sample (301 observations, 1978:12 to 2003:12), and on two sub-samples, determined on the basis of the policy considerations made in the previous chapter, as described in table 5.1 [63].

Table 5.1 The two sub-samples used in the cointegration analysis

1978:12- 1993:06	175 observations	Precedent to substantial CAP reform (Regular CMO functioning, first reforms of 1988)
1993:07- 2003:12	126 observations	Following substantial CAP reform (Mac Sharry reform, Agenda 2000, Fischler reform)

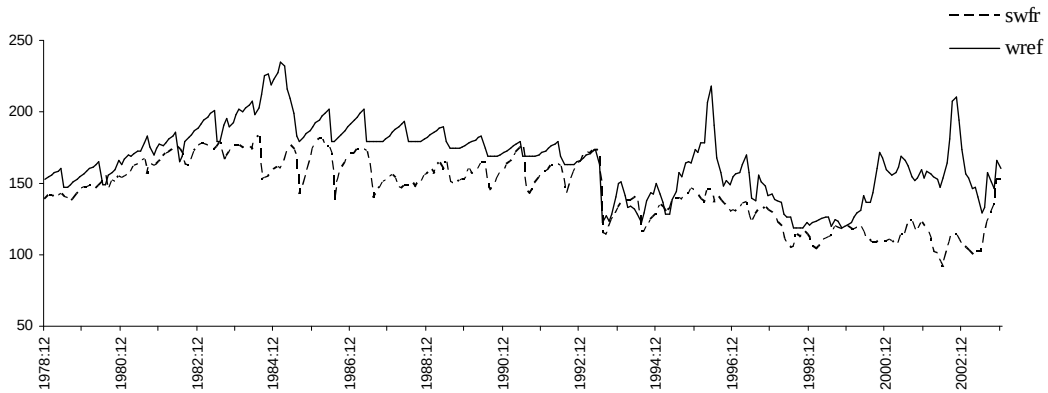
Indeed, the MacSharry reform of 1993 represents the first fundamental change in the CAP of the EU, since, for the first time, substantial cuts of the intervention prices (-30%) were implemented. In fact, as expected, the US price tends to be

[63] 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 paragraph 4.2.2. For this reason, and considering its political relevance, the MacSharry reform has been chosen as the fundamental break.

higher than the intervention price from the MacSharry reform onwards (see also figure 4.8).

Looking at the behaviour of the prices, we can also notice that the French price seems to follow more the intervention price before the implementation of the MacSharry reform, i.e. when it is normally higher than the US price, and the US price afterwards, when the reverse holds (figure 4.8, and 5.1).

Figure 5.1 Soft wheat French price and world reference price (euro/t)



Source: Eurostat; International Grains Council; European Commission regulations

Throughout the whole analysis that will follow, prices are expressed in logs.

Unit root tests have been repeated both in the whole period and in the two sub-periods identified above[64]. Augmented Dickey Fuller (ADF) tests have been run by adding the last significant lag up to a maximum of 18, due to the monthly nature of the data. Phillips-Perron (PP) tests were run with a number of lags determined by minimizing the Schwartz's Bayesian Information Criterion (SBIC) in a number of autoregressive specifications up to a maximum of 18 lags. The choice of the lag length was generally unaffected by the deterministic trend. Considering the monthly nature of tha data, 24 lags were used for the Kwiatkowski, Phillips, Schmidt e Shin (KPSS) tests. Finally, 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[65].

[64] Both STATA® and GRETL (www.gretl.sourceforge.net) softwares were used for the econometric estimations.

[65] With monthly data, both conventional and seasonal unit roots might be present; but the latter are less plausible in economic time series since they imply an evolving seasonal pattern (Dawson and Sanjuan 2006, p. 104). For this reason, we don't take them into account.

While on logarithms the null hypothesis of unit roots cannot be rejected, for log-differences it is possible to do so. This is generally true for all the series considered, both in the whole sample and in the two sub-samples identified as before and after the MacSharry reform. Results then overall confirm that the processes are I(1) (i.e., integrated of order 1, or difference-stationary), as it is the case in all the literature revised, despite the fact that they are nominal prices (Fanelli and Bacchiocchi 2005), and are reported in annex B.

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 Akaike Information Criterion (AIC) up to a maximum of 18 lags[66]. The Johansen and Juselius procedure has been used to estimate the rank of the cointegrating matrix (where not indicated differently, the null hypothesis has been rejected at 5% significance level) and the VECM in the "restricted constant case[67]. Autocorrelation was tested with a Lagrange Multiplier test with the null hypothesis of no autocorrelation up to the 18th lag (the null hypothesis has been rejected at 5% significance level). Dummy variables have been introduced outside the cointegrating vector to take account of the months in which export taxes have been imposed by the EU with the objective of preventing prices from raising too much, and also of the Russian and Ukrainian grain inflow of 2001/2002.

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. As expected from visual inspection of the two time series and from policy considerations, the rank of the cointegration matrix is zero (Johansen's tests are reported in annex C). However, by carrying out unit root tests on the French-US price spread in the overall sample and in different sub-samples, some evidence of stationarity emerges, which seems to be stronger after 1992 (see annex D).

Although these results are partly contrasting, we could conclude that the each of the two prices follows its own pattern and that the influence of the market policies is so strong that they are not related to each other. In other words, if policy variables and policy regimes are simply ignored in the analysis, and if such a long period of time is considered, the two prices wander following each its own path.

[66] The number of lags selected by the SBIC did not normally allow to remove autocorrelation in the equations.

[67] Visual inspection of the data shows that there is no trend in the series (figure 4.8), and both theory and visual inspection of the data imply the presence of a constant in the long-run relationship, accounting for all elements contributing to price differentials not explicitly modelled in the price transmission equation. This means that, even if there are no linear time trends in the level of the data, the cointegrating relation has a constant mean.

This should not come as a surprise, provided the very existence of the European border policies (namely, variable levies and export subsidies); this protection mechanisms, despite the decrease in the intervention price, has always been in place.

Other studies (Ghoshray *et al.* 2000; Ghoshray 2002; Barassi and Ghoshray 2007) show instead the presence of cointegration between the EU and the US prices. These different results can simply be 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: the very objective of export refunds is to cover the distance existing between the high domestic EU prices and the world ones, in order to allow the European excess supply to be sold in world markets at a competitive price. So, there should be no surprise also in their findings that the EU is a price follower in world markets, right because it applies export refunds on the basis of the world market prices (Ghoshray *et al.* 2000; Barassi and Ghoshray 2007). Instead, Thompson and Bohl (1999) use producer's level German prices and US fob Dark Northern Spring prices and find that they are cointegrated. They use a threshold cointegrating technique, which could partly explain this difference in results. On the other side, the selection of the Dark Northern Spring price (which has a even higher protein content than the Hard Red Winter wheat) was not deemed appropriate for our analysis. Finally, Thompson *et al.* (2002a), do not estimate but impose the cointegration relationship and, with a Seemingly Unrelated Regression Error Correction Model in which prices of different EU countries are related to the US price, find that price differentials are stationary and that adjustment coefficients are significant.

The cointegration analysis was repeated for *swfr* and *wref*, which contains 162 times the intervention price over 301 months (54% of the total number of observations) and is I(1), as well. The rank of the cointegration matrix turned out to be one (Johansen tests are reported in annex C). The estimates of Model 1 are reported in table 5.2[68].

For all the tables that will follow, the symbol α indicates the adjustment coefficients, while, for the sake of simplicity, for the coefficients of the cointegration vector, *cost* is the constant term and β_{wref} and β_{hrw} are the price transmission elasticities between *swfr* and either *wref* or *hrw*. In order to test the restrictions imposed on the cointegration vectors a likelihood ratio test is used (χ^2 distribution).

[68] The VECM was estimated with 13 lags (optimal number according to the AIC both with and without seasonal dummies). Monthly dummies for June, July and August only have been inserted (until 1985, the marketing campaign would begin in August and end in July the following year, since 1986/1987, it would begin in July and end in June); this model presented better information criteria that those estimated with all or no monthly dummies.

Table 5.2 Model 1 estimates for swfr and wref, 1978:12-2003:12

	Δswfr	**Δwref**	**Cointegration vector**	
α	-0.093*** (0.030)	0.064** (0.030)	$swfr_t = -\underset{(0.670)}{0.397} + \underset{(0.131)}{1.054}^{***} wref_t$	
LM test	0.207 (p-value 0.999)	0.201 (p-value 0.999)	$cost = 0$ $\beta_{wref} = 1$	$\chi^2 = 0.333$ (p-value 0.564) $\chi^2 = 0.161$ (p-value 0.688)
ARCH(13)	2.371 (p-value 0.999)	3.402 (p-value 0.996)	$\beta_{wref} = 0$	$\chi^2 = 18.886$ (p-value 0)

Standard errors are reported in parenthesis
LM test: Lagrange Multiplier test with the null hypothesis of no-autocorrelation
*significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients (#, ## and ### respectively if for β_{wref} the null hypothesis of equal to one is rejected)

The cointegration relation is $swfr_t$ = -0.397 + 1.054 $wref_t$. The imposition of a 1 price transmission elasticity coefficient is not rejected (χ^2 = 0.161; p-value 0.688). The constant term is not significantly different from zero (χ^2 = 0.333; p-value 0.564). Adjustment coefficients have the expected sign and are both significant. The residuals from the cointegration relationship show a stationary behaviour (figure 5.2). We can see that three major deviations from the long run relation, corresponding to major peaks of the error correction term, occurred in 1985, 1996 and 2002; in these months the French price was low, as explained before, because of the use of export taxes or of the specific situation of 2001/2002, when large inflows of products depressed domestic EU prices. Since these are accounted for in the short-run dynamics of the system but not inside the cointegration vector[69], what we observe is an increase in the disequilibrium term. To this point, the empirical findings confirm our expectations: the French price is linked to a series constituted by the maximum between the intervention and the US price, and the price transmission elasticity is very close to one.

To explore more in depth how this relation has possibly changed in time, the analysis has first been repeated in the sub sample 1978:12-1993:06. Here, *wref* is almost always constituted by the intervention price, which is higher than the US price 145 months over a total of 175 (83% of the total number of observations). Indeed, up to the MacSharry reform intervention price were normally set above the world price.

swfr is cointegrated with *hrw*, but the price transmission elasticity coefficient in the cointegration vector is very low (see table 5.3, Model 1a, where "a" indicates that estimates refer to the first sub-sample)[70].

[69] This could be done by cointegration techniques accounting for structural breaks (see chapter 6).
[70] The VECM was estimated with 5 lags (optimal number according to the AIC when monthly dummies were included) and with the monthly dummies for June, July, and August only, which allowed getting better

Figure 5.2 Residuals from the cointegration relationship of Model 1 for swfr and wref 1978:12-2003:12

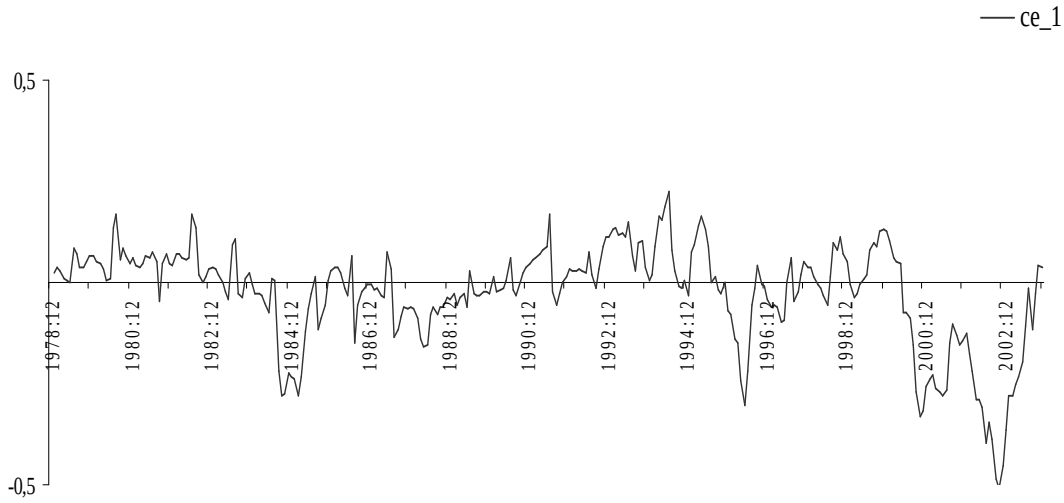


Table 5.3 Model 1a estimates for swfr and hrw, 1978:12-1993:06

	Δswfr	**Δhrw**	**Cointegration vector**
α	-0.172*** (0.041)	-0.007 (0.059)	$swfr_t = \underset{(0.341)}{3.968}^{**} + \underset{(0.069)}{0.226}^{\#\#\#:} {}^{***} hrw_t$
LM test	1.629 (p-value 0.063)	0.750 (p-value 0.753)	$cost = 0$ $\chi^2 = 5.940$ (p-value 0.015) $\beta_{hrw} = 1$ $\chi^2 = 6.901$ (p-value 0.009)
ARCH(13)	11.100 (p-value 0.049)	4.463 (p-value 0.485)	$\beta_{hrw} = 0$ $\chi^2 = 7.008$ (p-value 0.008)

Standard errors are reported in parenthesis
LM test: Lagrange Multiplier test with the null hypothesis of no-autocorrelation
*significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients (#, ## and ### respectively if for βhrw the null hypothesis of equal to one is rejected)

In the cointegration model, the transmission elasticity is statistically different from one and the constant is very high and significantly different from zero. Only the French price adjusts to disequilibria, whereas the US adjustment coefficient has not even the expected sign and is very small.

information criteria than when either all dummies were included or the model was estimated with 13 lags (optimal number when monthly dummies were not considered) and no dummies.

Model 1a was then estimated for *swfr* e *wref*, that are once again cointegrated: estimates are reported in table 5.4[71].

Table 5.4 Model 1a estimates for swfr and wref , 1978:12-1993:06

	Δswfr	**Δwref**	**Cointegration vector**
α	-0.211*** (0.074)	0.210*** (0.052)	$swfr_t = \underset{(0.643)}{0.679} + \underset{(0.124)}{0.848}^{***} wref_t$
LM test	0.430 (p-value 0.978)	0.293 (p-value 0.998)	cost = 0 χ2 = 0.758 (p-value 0.384) βwref = 1 χ2 = 0.990 (p-value 0.320)
ARCH(13)	34.030 (p-value 0.001)	25.881 (p-value 0.018)	βwref = 0 χ2 = 18.738 (p-value 0)

Standard errors are reported in parenthesis
LM test: Lagrange Multiplier test with the null hypothesis of no-autocorrelation
*significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients (#, ## and ### respectively if for β_{wref} the null hypothesis of equal to one is rejected)

The long run relationship is $swfr_t = 0.679 + 0.848\ wref_t$. Like in the overall sample, both adjustment coefficients have the expected sign and are significant, though here they are higher in magnitude. Restrictions imposing both a perfect price transmission and a zero-constant are not rejected (respectively, $\chi^2 = 0.990$; p-value 0.320; $\chi^2 = 0.758$; p-value 0.384).

In the first sub-sample, the relation between the French price and the combination between the US and the intervention price presents a stronger adherence to the LOP than when the US price enters the equation as such. This is somehow expected, provided the predominance of the intervention price in the *wref* series.

Finally, the analysis was repeated in the sub-sample 1993:07-2003:12 (Model 1b). Differently from the observations preceding the MacSharry reform, as expected *wref* is this time constituted almost exclusively by the US price, which is below the intervention price only 17 months over 126 (13% of the total number of observations). This basically happens because the MacSharry reform implemented a substantial reduction of the intervention prices.

For the French and the US price, there is evidence of cointegration among the series[72]. The Model 1b estimates are reported in table 5.5.

[71] The optimal number of lags selected by the AIC was 13 lags; 1 if monthly dummies were considered. One lag only did not allow to remove autocorrelation; the VECM was then estimated with 13 lags.

[72] The cointegration rank was 1 when 23 lags and seasonal dummies were considered. Since 23 lags seems however an overestimated number, the cointegration rank was checked also with 13 lags. In this case, the series are cointegrated if the null hypothesis is rejected at 10% significance, always with the inclusion of all the monthly dummies; the latter estimates are reported. See annex C.

Table 5.5 Model 1b estimates for swfr and hrw, 1993:07-2003:12

	Δswfr	**Δhrw**	**Cointegration vector**
α	-0.078** (0.034)	0.063 (0.060)	$swfr_t = \underset{(1.505)}{0.664} + \underset{(0.305)}{0.842}^{***} hrw_t$
LM test	0.452 (p-value 0.963)	0.221 (p-value 0.999)	$cost = 0$ $\chi^2 = 0.136$ (p-value 0.712) $\beta_{hwr} = 1$ $\chi^2 = 0.180$ (p-value 0.672)
ARCH(13)	16.849 (p-value 0.206)	17.552 (p-value 0.175)	$\beta_{hwr} = 0$ $\chi^2 = 8.549$ (p-value 0.003)

Standard errors are reported in parenthesis
LM test: Lagrange Multiplier test with the null hypothesis of no-autocorrelation
*significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients (#, ## and ### respectively if for β_{hrw} the null hypothesis of equal to one is rejected)

In Model 1b, the hypothesis of perfect price transmission in the cointegration vector is not rejected. Adjustments coefficients have the right sign. The US price performs as weakly exogenous.

wref and *swfr* are cointegrated, as well[73]; the corresponding VECM is estimated (Model 1b, table 5.6).

Table 5.6 Model 1b estimates for swfr and wref, 1993:07-2003:12

	Δswfr	**Δwref**	**Cointegration vector**
α	-0.073** (0.032)	0.063 (0.052)	$swfr_t = \underset{(1.624)}{0.072} + \underset{(0.328)}{0.990}^{***} wref_t$
LM test	0.346 (p-value 0.991)	0.155 (p-value 0.999)	$cost = 0$ $\chi^2 = 0.001$ (p-value 0.970) $\beta_{wref} = 1$ $\chi^2 = 0.0006$ (p-value 0.981)
ARCH(13)	13.086 (p-value 0.441)	10.644 (p-value 0.641)	$\beta_{wref} = 0$ $\chi^2 = 9.201$ (p-value 0.002)

Standard errors are reported in parenthesis
LM test: Lagrange Multiplier test with the null hypothesis of no-autocorrelation
*significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients (#, ## and ### respectively if for β_{wref} the null hypothesis of equal to one is rejected)

The long run relationship is $swfr_t = 0.072 + 0.990 wref_t$. Adjustment coefficients, that are very close to the one presented in table 5.5, have the expected sign. The *wref* one is not significant, which could suggest that it is

[73] The two series are not cointegrated when 2 lags are included (which is the optimum number of lags in all cases). When, in analogy with the models estimated in the overall sample and in the first subsample, 13 lags are included, evidence of cointegration emerges if the null hypothesis is rejected at 10% significance. See annex C. The insertion of the 3 summer dummies allowed to get better information criteria.

weakly exogenous. Once again, restrictions imposing both a perfect price transmission and a zero-constant are not rejected (respectively, χ^2 = 0.0006, p-value 0.981; χ^2 = 0.001; p-value 0.970).

Differently from what happens for the first sub-sample, this time the two relations estimated are much closer; the explanation lies of course in the predominance of the US price series in the composition of *wref*.

In the second subsample, nonetheless, the evidence of cointegration for *swfr* and *wref* is weaker than in the first and in the overall ones (the null hypothesis of a zero cointegration rank is in many cases rejected only at 10% significance; see annex C). This might be due to the fact that what happened in 2002 (when the Ukrainian or Russian price, which was much lower than the US price, was probably the true "world price" for the EU), is not explicitly considered inside the cointegration vector. This would get more importance as the number of observations is reduced, which is what happens when moving from all the observations to the second sub-sample.

Summing up, the composite variable used allows to consider in the cointegration model the presence of different policy regimes. When theoretical considerations are translated into empirical models, in general, the long run relation between *swfr* and *wref* is very close to the LOP both when all observations are considered, and in the two subsamples. This means that, basically after the MacSharry reform, the US price was above the intervention price and could interact with the EU price despite the same border policies scheme kept being in place. It was the reduction of the intervention price, and not the (non-) changes in the system of policy barriers[74], which increased price interdependency.

Nevertheless, this very simple model presents a number of shortcomings. If it has to be used for projections, unless the intervention price is considered as an exogenous, pre-determined and known threshold also for future months, we cannot actually identify what the *wref* series actually is constituted by, if the US or the intervention price. Moreover, whether the relation found is instead a linkage between the French and the intervention price, introduced as a threshold below which the US price has no influence on the EU one, poses some interpretative problems and requires further research[75].

[74] As a consequence of the lowering of intervention prices, both export subsidies and variable levies were reduced, but the same protection system kept being in place.

[75] Verga and Zuppiroli (2003), by using weekly data, find that the US Soft Red Winter wheat Rotterdam cif price is never cointegrated with domestic European prices, and that the intervention price is cointegrated with them in the period 1990-2002 but not in the sub-period 1995-2002. 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, evidence of cointegration emerges. They suggest that EU quotations could be linked to an "average" of the two prices, but also that this is a spurious vector not appropriate for the analysis (Verga and Zuppiroli 2003 p. 19).

5.2.2 Cointegration models accounting for policy regime changes

Based on the theoretical framework outlined in paragraph 4.3, there are other ways, in addition to the use of a composite variable presented in the previous paragraph, to further develop it into empirical applications.

Firstly, a particular cointegration model will be estimated, with different adjustments coefficients depending on the observable policy regimes (Model 2; paragraph 5.2.2.1). In this first model, the LOP will be imposed to hold between the French price and the highest between the US and the intervention price. Secondly, the price transmission elasticity will be allowed to change according to which price the French one is related to: always, the highest between the intervention and the US price (Model 3; paragraph 5.2.2.2). Finally, both adjustment coefficients and the cointegrating vector parameters will be assumed to vary (Model 4; paragraph 5.2.2.3).

5.2.2.1 Model 2: a threshold model allowing for policy regime changes

All models which follow still stem from the theoretical considerations outlined in the previous chapter: the French price is assumed to be linked to either the US or the intervention price depending on which of them is higher.

These models are derived by introducing *ad hoc* modifications in a general VECM as presented in equation 3.31 in chapter 3, in order to consider the influence of policy regimes. If a simple VECM would hold between the US and the French price, then the LOP would be expected to be valid at least in the long run, despite prices being allowed to diverge from it in the short run. But such a straightforward model is clearly not appropriate, as the two variables turn out not to be cointegrated (see annex C and annex D). In paragraph 5.2.1, we anticipated that this might due to overlooking the policy regime issue.

For this reason, in Model 2, we try and introduce appropriate modifications to take policy regimes in due account. We assume that the LOP holds between the French price and the US price only when the latter is above the intervention price; otherwise, the LOP will hold between the intervention and the French price[76]. This is the same as assuming that there are two distinct regimes; this model can be interpreted as a particular cointegration model in which the adjustment coefficients take different but non-zero values according to the observable policy regime the observations belong to.

Technically, this requires the creation of a regime dummy variable, which we call henceforth reg_t; if $hrw_t > pint_t$, $reg_t = 1$, and if $hrw_t < pint_t$, $reg_t = 0$.

We then estimate the following model:

[76] Price spread stationarity (i.e., the validity of the LOP in the long run) is here imposed rather than estimated.

$$\mathbf{\Delta p_t} = \boldsymbol{\alpha}_1 reg_{t-1} z_{1,t-1} + \boldsymbol{\alpha}_2 (1 - reg_{t-1}) z_{2,t-1} + \sum_{i=1}^{k-1} \mathbf{\Gamma_i \Delta p_{t-i}} + \boldsymbol{\varepsilon_t} \tag{5.1}$$

where $z_{1,t-1}$=($swfr_{t-1}$ - hrw_{t-1}) and $z_{2,t-1}$ = ($swfr_{t-1}$ - $pint_{t-1}$).

If $hrw_{t-1} > pint_{t-1}$, and then $reg_{t-1} = 1$, only $z_{1,t-1}$ is "active"; if $hrw_{t-1} < pint_{t-1}$, and then $reg_{t-1} = 0$, only $z_{2,t-1}$ is "active". The LOP will hold between the French and the highest between the US and the intervention price; adjustment coefficients are allowed to vary in either case.

Estimates of Model 2 are reported in table 5.7[77].

Table 5.7 Model 2 estimates

	Δswfr	**Δhrw**
α_1	-0.013 (0.015)	0.022 (0.017)
α_2	-0.099*** (0.028)	-0.025 (0.033)
LM test	0.303 (p-value 0.998)	0.543 (p-value 0.935)
ARCH (12)	2.265 (p-value 0.999)	25.818 (p-value 0.018)
OV test $z_{1,t-1}$, $z_{2,t-1}$	15.043 (p-value 0.005)	

Standard errors are reported in parenthesis
LM: LM test with null hypothesis of no-autocorrelation
OV: Omitted Variable test (χ^2)
* 10% significance; ** 5% significance; *** 1% significance

The omitted variable test for $reg_{t-1} z_{1,t-1}$ and $(1 - reg_{t-1}) z_{2,t-1}$ rejects the null hypothesis with a p-value of 0.005. In the French equation, the adjustment coefficients have the right sign, though only the one to the LOP holding with the intervention price is significant. As expected, for the French price $|\alpha_1| < |\alpha_2|$, i.e. it responds more quickly to the LOP holding with the intervention price than with the US price. In the US equation, α_1 has the right sign but is not significant, and α_2 has a negative sign, which means that it doesn't correctly adjusts to the disequilibria of the LOP holding between the French and the intervention price.

Actually, the interpretation of the US adjustment coefficient brings some problems. In fact, the intervention price is assumed to be a threshold for the US price; a positive, "correct" (in the conventional VECM interpretation) sign for the US adjustment coefficient would mean that, when the disequilibrium from the

[77] Autocorrelation was removed by adding the 12th differentiated lag. Monthly dummies were selected with specification tests.

LOP holding between the French and the intervention price is positive (i.e., the French price is above the intervention price), then the US price is expected to raise, and the other way round.

In reality, what we observe from the estimates is that the sign of the US adjustment coefficient is negative. This means that the US price doesn't adjust to the disequilibria of the LOP holding with the intervention price. This could just indicate a non response of the US price to the disequilibria from the LOP holding between the French and the intervention price.

But another possible, alternative explanation can be figured out. Recalling that the LOP is holding between the French and the intervention price right because the latter is above the US price (in other words, when the CAP market regulation is somehow "binding"), we might think that what happens is that when the domestic French price is above the intervention price, this is likely to depress the world's price through an increase in the export supply (Thompson 2000, p. 720), and that, alternatively, when the French price is below the intervention price, the world price would be allowed to increase.

5.2.2.2 Model 3: a changing cointegration vector

As it has been previously explained, the appeal that lead to such a widespread use of cointegration models in price transmission testing is basically the possibility of separating long and short run market dynamics.

In this respect, Model 3 is logically complementary to Model 2. Here, indeed, we assume that the observable regime changes don't affect the adjustment parameters, i.e. the short run relations, but the cointegration vector itself.

The French price is still expected to be linked to either the intervention or the US price according to which of them is higher; however, an attempt at discerning to which extent the two long-run relationships are different is made. In fact, the very nature of the relationship holding between the French and the intervention price is likely to be different from the one with the US price. Obviously, we suspect that the elasticity of transmission between the French and the intervention price is higher than the one between the French and the US price.

The cointegration vector is then assumed to be the following one (equation 5.2):

$$swfr_t - \beta_0 - \beta_1 wref_t - \beta_2 reg_t wref_t = z_t \text{ where } z_t \approx I(0) \tag{5.2}$$

which means that, if $hrw_t > pint_t$, and then $reg_t = 1$, it becomes

$$swfr_t = \beta_0 + (\beta_1 + \beta_2) hrw_t + z_t \tag{5.3}$$

and if $hrw_t < pint_t$, and $reg_t = 0$, the cointegration relation is instead

$$swfr_t = \beta_0 + \beta_1 p\text{int}_t + z_t \tag{5.4}$$

By combining the composite variable *wref* with the policy regime dummy, we obtain a particular cointegration vector whose parameters depend on the policy regime in place.

The cointegration vector was estimated with an OLS regression; its lagged residuals (for which the null hypothesis of non-stationarity yields an ADF statistic of -3.453, p-value 0.100) have been inserted in the VECM. Its estimates are reported in table 5.8[78]. As expected, the elasticity of transmission is lower with the US price (0.707-0.022) than with the intervention price (0.707), but, interestingly, the difference doesn't seem of a big magnitude. Moreover, such a value is consistent with the ones estimated for Model 1.

Table 5.8 Model 3 estimates

	Δswfr	**Δhrw**	**Cointegration vector**
α_3	-0.073*** (0.023)	0.002 (0.028)	
LM test	0.361 (p-value 0.993)	0.583 (p-value 0.910)	$swfr_t = \underset{(0.205)}{1.400}*** + \underset{(0.040)}{0.707}\, wref_t *** - \underset{(0.002)}{0.022}*** reg_t wref_t$
OV test z_{t-1}	10.214 (p-value 0.006)		
ARCH (12)	2.027 (p-value 0.999)	25.395 (p-value 0.013)	

Standard errors are reported in parenthesis
LM: LM test with null hypothesis of no-autocorrelation
OV: Omitted Variable test (χ^2)
* 10% significance; ** 5% significance; *** 1% significance

The omitted variable test for z_t rejected the null hypothesis with a p-value of 0.006. Both adjustments coefficients have the right sign; only the French one is significant. This means that the French price adjusts at a speed of about 7% per month to the disequilibria from the long run relationship between the French price and either the US or the intervention price, according to which of them is higher. The US price seems instead to perform as weakly exogenous even if it can't be excluded that this is due to the fact that the intervention price intervenes in the cointegration vector, and in some periods the US price is then expected to respond

[78] Autocorrelation was removed by adding the 12th differentiated lag. Monthly dummies were selected with specification tests.

to the disequilibria between the French and the intervention price (see previous paragraph).

5.2.2.3 Model 4: both adjustment coefficients and cointegration parameters can vary

Finally, in Model 4, both adjustment coefficients and the cointegrating relationship were allowed to differ according to the policy regime in place. In equation 5.5, the disequilibrium term z_t has been calculated as in Model 3, and then has been multiplied by the policy regime dummy like in Model 2.

$$\mathbf{\Delta p_t} = \boldsymbol{\alpha}_4 * \mathbf{reg_{t-1}} * \mathbf{z_{t-1}} + \boldsymbol{\alpha}_5 * (1 - \mathbf{reg_{t-1}}) * \mathbf{z_{t-1}} + \sum_{i=1}^{k-1} \Gamma_i \mathbf{\Delta p_{t-i}} + \boldsymbol{\varepsilon}_\mathbf{t} \qquad (5.5)$$

What happens is that, if reg_{t-1} = 1, the adjustment coefficients to the LOP holding with the US price are given by α_4; if reg_{t-1} = 0, then the adjustment coefficients to the LOP holding with the intervention price are given by α_5.

The empirical estimates reported in table 5.9[79] show that the adjustment coefficients of both the US and the EU prices behave consistently with Model 2, though slightly higher in absolute value.

Table 5.9 Model 4 estimates

	Δswfr	**Δhrw**	**Cointegration vector**
α_4	-0.023 (0.026)	0.021 (0.031)	
α_5	-0.260*** (0.049)	-0.069 (0.060)	
			$swfr_t = \underset{(0.205)}{1.400}*** + \underset{(0.040)}{0.707}***\, wref_t - \underset{(0.002)}{0.022}***\, reg_t\, wref_t$
LM test	0.233 (p-value 0.999)	0.537 (p-value 0.938)	
OV test $z_{1,t-1}$, $z_{2,t-1}$	28.815 (p- value 0)		
ARCH (12)	1.814 (p-value 0.999)	28.096 (p-value 0.005)	

Standard errors are reported in parenthesis
LM: LM test with null hypothesis of no-autocorrelation
OV: Omitted Variable test (χ^2)
* 10% significance; ** 5% significance; *** 1% significance

[79] Autocorrelation was removed by adding the 12th differentiated lag. Monthly dummies were selected with specification tests.

Interestingly, for the French price, the adjustment coefficients of Model 1 and 3, in which the LOP holds between a combination of US and intervention prices, are in between the two of Model 2 and 4, where they are allowed to vary according to which price the LOP holds with.

This is valid also for Model 3 for the US price (in Model 1, the US price is not present as such but in combination with the intervention price in *wref*).

The coefficients of the French price have always the expected sign, though they are significant only when the intervention price is included in the cointegration equations. US coefficients are instead significant only in Model 1. To which extent the US price is weakly exogenous (Model 3) or the cointegrating relationship is mainly driven by the intervention price (Model 2 and 4) requires further research. Answering to this question poses some interpretative problems, first of all because the intervention price is introduced as a threshold for the US price but in theory, when below such threshold, the two might be expected to have a diverging behaviour; and then also because the US price is assumed to be the relevant world price, whereas in some years (namely, the campaign 2001/2002), this might not have been the case.

5.3 Concluding remarks

In this chapter, the relation between soft wheat monthly prices for the US and France in the years 1978-2003 has been studied by developing various empirical models following the theoretical considerations made in chapter 4.

The econometric models presented, though simplified, are a first attempt of combining policy and price data; the EU intervention price has then been used in different ways to take into account the role played by policy regime changes. The basic idea is that the intervention price acts as a threshold for the US price.

Although empirical evidence is mixed in this respect, in our data domestic French prices turn out not to be cointegrated with the US ones. Indeed, what we argue is that this confirms that the presence of EU domestic and border policies prevents the two prices from sharing the same pattern, and need to be adequately considered when testing for price transmission.

First of all, the regime switch has been modelled *via* the creation of a composite variable, the “EU external reference price”, constituted by the maximum between the US and the intervention price. Evidence of cointegration is found with the French price. This means that, basically after the MacSharry reform of 1993, which reduced the intervention price allowing the US one to be much more often above it, the US prices interacted more with the EU domestic ones. We might argue that the reduction in intervention prices, despite a non-changing system of policy barriers, was what actually increased the degree of

interaction between the EU internal prices and the US ones. By lowering intervention prices, both export subsidies and variable levies were in fact reduced.

Secondly, a specific cointegration model has been estimated between the EU and the US prices, with different adjustments coefficients depending on the observable policy regime observations belong to. The LOP has been imposed between the French price and the highest between the US and the intervention price. The French price responds more to the LOP holding with the intervention price, but also the adjustment coefficient to the LOP holding with the US price has the correct sign. In turn, the interpretation of the coefficients of the US price is somehow problematic.

Thirdly, it is the price transmission elasticity which has been allowed to change, depending on the policy regime in place (once again, assuming that the French price is linked to either the US or the intervention price according to which of them is higher). The price transmission elasticity between the French and the intervention price is found to be stronger than the one with the US price. The French price adjusts to the disequilibria from the long run relation between the French price and either the US price or the intervention price, according to which of them is higher. The US price is weakly exogenous.

Finally, in Model 4, both adjustment coefficients and the cointegrating vector parameters are allowed to vary. Results are consistent with the previous ones.

The models presented here, though over-simplified, represent an attempt of explicitly combining policy and price data. The observable policy regime change has been simulated with the construction of an *ad hoc* composite variable and other econometric devices.

The consistent behaviour of EU adjustment coefficients suggests that the role of the US price might be understood only in light of adequate consideration for policy regimes.

The response of the French price is consistent throughout all the models: this price responds correctly to the disequilibria from the various long run relationships in which the intervention price has been introduced in different ways. The adjustment coefficients always have the correct sign, although they tend to be significant and higher in magnitude when the intervention price is included in the cointegration vector. Indeed, they are bounded between -0.013 (response to the LOP holding with the US price; model 2) and -0.260 (response to the LOP holding with the intervention price with a changing cointegration vector; model 4). However, they tend to be around 0.100 (see all model 1 estimates; model 3) when the *wref* series is included in the cointegration vector.

We could say that, once the EU regulatory framework has been taken into account, a positive albeit small response to the behaviour of the US price emerges.

The response of the US price leaves instead some questions open. US adjustment coefficients are usually much smaller in value and not significant

(when they have the expected sign, i.e., positive, they are bounded between 0.002, model 3, and 0.063, model 1b; when they are negative, they are bounded between -0.069, model 4, and -0.007, model 1).

This could be due to its exogeneity in the system, as model 3 would imply; but, also, the cointegration relation might be mostly driven by the intervention price (as can be noted in model 2 and 4), which in turns brings some interpretative problems. Indeed, the introduction of this threshold could explain why the US price has a negative adjustment coefficient (see paragraph 5.2.2.1).

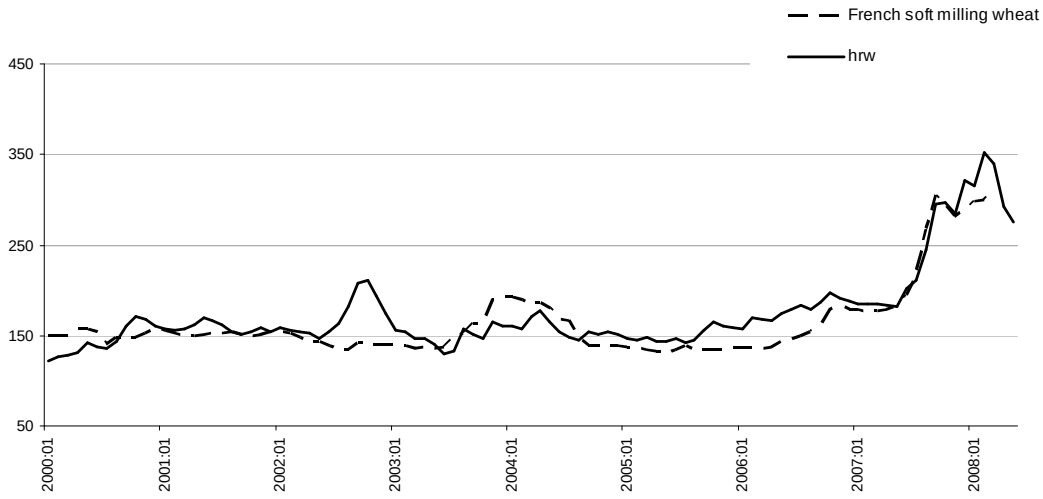
A final consideration concerns the reference period which has been used for the analysis.

In this work, in order to study in depth the role of domestic and border policies on price transmission, verifying the effects of the major changes they underwent in the past years, the use of longer time series (Eurostat data), which would allow to go back in time, has been preferred.

Unfortunately, this in turn implies the impossibility of using more recent data, and of considering what has recently happened on world markets.

The current rise in food prices (figure 5.3), which started in the summer of 2007, brought renewed interest in the debate on the appropriateness of agricultural domestic and border policies. At the same time, these issues are also currently discussed in the so-called "Health Check" of the CAP (European Commission, COM(2007) 722, and COM(2008) 306/4). The Health Check debate has the objective of completing the Fischler reform of 2003 and of establishing the new set of rules which will hold up to 2013.

Figure 5.3 French and US prices in the last years



Source: ISMEA; International Grains Council

In addition to this, as explained in the introduction of this chapter, soaring food prices immediately caused major changes in the world trade environment.

As high prices signal food scarcity, importers have removed import duties (such as the elimination of the EU tariffs for cereals up to 2009), while major exporters like Ukraine and Argentina have implemented, in turn, export taxes. These measures are of course expected to affect price transmission dynamics.

For these reasons, the study of the most recent dynamics on world markets, which goes well beyond the use of more up-to date data, probably constitutes the most interesting development of the model presented.

6 EMPIRICAL ANALYSIS: COINTEGRATION MODELS, STRUCTURAL BREAKS AND POLICY REFORM

6.1 Introduction

In this last chapter, a further investigation of the relationship between the EU and the US prices for soft wheat is proposed.

Differently from chapter 5, where a policy variable, the intervention price, was introduced in the empirical models to consider policy regimes, here policy regime changes will instead be assumed to be directly responsible for structural breaks in the cointegration vector which links, in the long run, EU and US prices.

Our research question is as follows: did the CAP reforms and the URAA implementation affect the relation between the two prices considered? We expect that the first, by reducing intervention prices, should have reduced the distance between EU and world prices; the second, by disciplining policy intervention, should have increased the price transmission elasticity.

This means that no other variables are introduced in addition to the domestic EU prices and the US ones; instead, the cointegration relationship is allowed to vary according to the most relevant policy regime changes.

This represents an alternative development of the econometric estimates proposed in chapter 5, that, focusing on the relationship between just the two prices, allows to avoid some of the interpretative problems which have been presented there. The underlying assumption here is that the two prices have always been interacting despite the presence of domestic and border policies, and that the policy regime switches had an impact on an existing relation.

The chapter is structured as follows. In paragraph 6.2, a short introduction to the theoretical framework used is proposed, explaining what is the rationale behind the use of cointegration models allowing for structural breaks in the deterministic trend. In paragraph 6.3, unit root tests accounting for structural

breaks are carried out and commented. The cointegration analysis is presented in paragraph 6.4, while paragraph 6.5 presents some concluding remarks.

6.2 Policy regimes, structural breaks and price transmission

According to the regulatory framework presented in chapter 4, since the EU market is isolated from the world one by variable levies and export subsidies, there should be, in theory, zero price transmission between the world and the EU price. Only if the world price is above the entry price (and consequently, above the intervention price, which is by definition below the entry price), we expect a positive price transmission elasticity with the domestic EU price.

Moreover, if the EU policy measures are binding, not only is the transmission from the world to the EU price zero, but the one from the intervention price to the world one should be negative, since increases in the first will depress the latter through the increase in the EU net export supply (Thompson 2000).

In chapter 5, various models have been adopted in order to allow for policy regime changes to be considered in cointegration models. In a nutshell, the introduction in the models of a policy variable, the intervention price, allows to distinguish between different policy regimes, in which we expect different price transmission relations to be in place.

In this chapter, the relation between the world (US) and the EU (French) price is instead studied by considering structural breaks. This allows to avoid a number of interpretative problems which arise when policy variables are introduced in the model (see paragraph 5.2).

We have proved already that the US and the EU prices turn out not to be cointegrated when standard Johansen tests are used (see the analysis reported in paragraph 5.2.1, and annex C). However, evidence of stationarity emerges when conventional unit root tests are repeated on the French-US price spread both in the overall sample and in the different subsamples (see annex D).

We also noticed that empirical evidence is mixed, since in some works co-movement between internal EU prices and the US prices is found (Thompson and Bohl 1999; Thompson *et al.* 2000; Thompson *et al.* 2002a and 2002b; see paragraph 5.2.1 and annex A).

Thompson (2000, p.723) argues that finding a low but nonzero price transmission elasticity also before the tariffication introduced by the URAA might be due to the fact that, in fixing domestic prices, the EU legislator responds somehow to the behaviour of world prices: increasing the intervention price relieves the budget in a situation of overproduction, and EU institutions will do this probably more with high than with low world prices (see also Thompson *et al.* 2000, p.723). The intervention and, in turn, the French price, could follow the world ones.

In addition to this, more interestingly, also market-oriented policy reforms could have played a role. We expect the MacSharry reform to have lowered EU domestic prices, and possibly reduced their distance from the world ones, and the URAA, depending on the effectiveness of tariffication, to have increased price transmission elasticities (Thompson *et al.* 2000, p.722).

Also in this respect, empirical evidence is mixed. Thompson and Bohl (1999), by using threshold cointegration models, find low price transmission elasticities, but increasing (from 0.18 to 0.30) after the liberalization reforms. Thompson *et al.* (2000), by ordinary OLS and 3SLS (3 Stages Least Squares) regressions, show a negative impact of the MacSharry reform on the level of the European prices, but also that the implementation of the URAA did not change the elasticity of price transmission between world and domestic prices, whose value is about 0.11. The same results apply to Thompson *et al.* (2002b). They find a price transmission elasticity of 0.18; the MacSharry reform is found to have caused a significant downward shift in domestic EU prices (its coefficient is -0.23), but an interaction term is not significant, which means that no increases in the price transmission elasticity took place. Thompson *et al.* (2002a), instead, by estimating a SURECM in different sub-samples according to the policy regimes in place, find evidence of increased price co-movement after the URAA implementation, although their price transmission coefficient is still small (0.183). The objective of this paragraph is then to analyze the existence and the evolution of a long run relationship between US and EU prices. This analysis can represent an interesting tool for the analysis of broader policy issues. Studying the implementation of the URAA can provide insights into the effects of more general liberalization reforms; analyzing the effects of the reduction of the intervention prices, in turn, is of a crucial importance in the current "Health Check" debate of the CAP, in which the elimination of this instrument is currently being considered. By using cointegration techniques allowing for structural breaks in the deterministic trend, it is possible to take into account observable policy regime changes. These are assumed to influence both the constant and the price transmission elasticity of the cointegration relation. This allows, on one side, to consider the time series properties of the data and, on the other, to take the entire number of observations available as a whole, studying the effects of the policy changes over time.

6.3 Unit root tests and structural breaks

Structural breaks due to policy reforms are likely to have affected both the pattern of EU prices and their long run relationship with the US ones.

As far as European wheat prices are concerned, we expect that the MacSharry reform, by dramatically lowering intervention prices, did indeed constitute a structural break which should not be ignored while testing for unit roots. As

explained in paragraph 3.3.1, the risk is that, without considering the presence of structural breaks, prices could result I(1) processes while indeed they are stationary around a broken level or trend. Although the standard unit root tests carried out in the two sub-samples identified as before and after the MacSharry reform (reported in annex B), that find all the series I(1), are an indirect confirmation of the appropriateness of this reasoning, also specific unit root tests which allow for the presence of one or more endogenous structural breaks are carried out.

The presence of unit roots allowing for structural breaks has been tested with the Perron and Vogelsang (1992), both additional outlier (AO) and innovational outlier (IO) tests, and with Clemente *et al.* (1998) tests (for a description, see paragraph 3.3.1). The dataset is the same as the one described in paragraph 4.4 and used for the analysis in chapter 5. All prices have been used in logarithmic form. Results are reported in tables 6.1 and 6.2. In order to better understand the behaviour of the French price, the intervention price time series has been examined, as well.

Table 6.1 Perron and Vogelsang (1992) tests

	AO model				**IO model**			
	Lags	**Time of the break**	**t stat on (ρ-1)**	**5% critical value**	**Lags**	**Time of the break**	**t stat on (ρ-1)**	**5% critical value**
$swfr_t$	12	1993:08***	-3.082	-3.560	12	1993:05***	-4.549**	-4.270
$pint_t$	14	1993:08***	-1.814	-3.560	13	1993:04***	-5.810**	-4.270
hrw_t	1	1986:05***	-3.998**	-3.560	1	1985:01***	-4.191	-4.270

***indicates rejection of the null hypothesis at 1%; ** at 5%; * at 10% significance

Table 6.2 Clemente et al. (1998) tests

	AO model				**IO model**			
	Lags	**Time of the break**	**t stat on (ρ-1)**	**5% critical value**	**Lags**	**Time of the break**	**t stat on (ρ-1)**	**5% critical value**
$swfr_t$	12	1993:08***; 1998:01 ***	-3.719	-5.490	12	1993:05***; 1997:12***	-5.934**	-5.490
$pint_t$	14	1993:08***; 2001:08 ***	-2.092	-5.490	13	1993:04***; 2001:05***	-6.838**	-5.490
hrw_t	17	1986:09***; 1996:02***	-3.968	-5.490	1	1985:01***; 2000:06**	-4.562	-5.490

***indicates rejection of the null hypothesis at 1%; ** at 5%; * at 10% significance

As expected, the behaviour of the French price follows the one of the intervention price, due of course to the strong regulatory framework in place.

With the additional outlier model, both for the French and the intervention price it is not possible to reject the null hypothesis of a unit root with structural break, occurred in August 1993. If two breaks are allowed, the situation is much less clear-cut; the 2001:08 break in the intervention price series coincides with the full implementation of Agenda 2000, but doesn't correspond to the one in the French price. Provided that the change caused by the MacSharry reform is very likely to have had a strong immediate effect rather than a gradual one, as confirmed by visual inspection of the series, for these two prices the innovational outlier model, both allowing for one and two breaks, (according to which the French and the intervention price are both I(0) processes) seems less appropriate, (and indeed, the time of the break moves backwards to a date which has no political sense[80]). For the US price it is possible to reject the unit root null hypothesis only in the AO model allowing for one break, only. This time the interpretation of the timing of the structural breaks is not immediate. The Zivot and Andrews test has been performed, as well (table 6.3). Both the French and the intervention prices turn out to be I(0) around a broken intercept. Nevertheless, remembering that the null hypothesis of unit root in this test excludes the presence of a structural break, and on the basis of the results of standard unit root tests carried out in the two subsamples identified as before and after the structural break constituted by the MacSharry reform (which, as recalled before, find evidence of nonstationarity; annex B), we conclude that the use of this test is not appropriate for the French and intervention price. For the US price it is not possible to reject the null hypothesis of a unit root process without structural break.

Table 6.3 Zivot and Andrews (1992) tests

	AIC lagmethod				**TTest lag method**			
	Lags	**Time of the break**	**t stat**	**5% critical value**	**Lags**	**Time of the break**	**t stat**	**5% critical value**
$swfr_t$	1	1993:07	-5.702**	-4.80	1	1993:07	-5.702**	-4.80
$pint_t$	1	1993:06	-6.042**	-4.80	0	1993:06	-5.795**	-4.80
hrw_t	1	1986:01	-4.686	-4.80	1	1986:01	-4.686	-4.80

AIC/TTest lag method: AIC criterion/5% significance of the t-test used for lag selection
***indicates rejection of the null hypothesis at 1%; ** at 5%; * at 10% significance

In order to test for the presence of a unit root in the French price differentiated series, provided the structural break identified in the French price evidently

[80] While, for the French price, this might suggest the hypothesis of some "rational expectations" behaviour of market operators whose actions anticipate the following reduction in the intervention price, this is actually not the case since the same occurs for the intervention price, which is the *very* policy variable.

corresponds to the MacSharry reform, reference can be made to standard unit root tests which have been repeated on the log-differentiated series in the two sub-samples identified as before and after the MacSharry reform (see annex B). For the US price, which most plausibly seems to be a I(1) process with no relevant structural breaks, evidence of its first-difference stationarity is confirmed by the conventional unit root tests reported in annex B.

6.4 A cointegration model allowing for structural breaks

At this point, the Johansen, Mosconi and Nielsen (2000) model has been tested on the data[81].

The model presented in paragraph 3.3.2.3, that is

$$\Delta \mathbf{p}_t = \boldsymbol{\alpha}\begin{bmatrix}\boldsymbol{\beta}\\ \boldsymbol{\mu}\end{bmatrix}'\begin{bmatrix}\mathbf{p}_{t-1}\\ \mathbf{tE}_{t-1}\end{bmatrix} + \boldsymbol{\gamma}\mathbf{E}_t + \sum_{i=1}^{k-1}\boldsymbol{\Gamma}_1\Delta\mathbf{p}_{t-i} + \sum_{i=1}^{k}\sum_{j=2}^{q}\mathbf{k}_{i,j}\mathbf{D}_{j,t+k-i} + \boldsymbol{\varepsilon}_t \qquad (6.1)$$

has been estimated restricting the broken level to the cointegration space[82], the only difference being the insertion of a constant term and the corresponding elimination of one of the *q* dummy variables used. This means that the coefficients of the dummies indicating the policy regimes will have to be interpreted as relative to the constant valid in the overall period. Throughout this chapter, the underlying assumption is that the rank of the cointegration matrix between the French and the US price series, once the various structural breaks have been taken into account, is one. And this because in the estimated models more than 2 structural breaks will be introduced, differently from the Johansen, Mosconi and Nielsen (2000) procedure, which doesn't allow to test for the cointegration rank with more than 2 structural breaks. According to conventional Johansen's test, the EU and the US price are not cointegrated (when tested in the restricted constant option with 5 lags[83], the null hypothesis of a zero rank in the cointegration matrix yield a trace and λ-max statistics of 16.086, p-value 0.173, and 12.538, p value 0.161, respectively). Nonetheless, standard ADF tests have been repeated on all the long-run residuals of the models estimated and confirm that they are I(0) at 5% level of significance.

[81] The prices have been de-seasonalized with the TRAMO/SEATS method developed by V. Gomez, A. Maravall (http://www.bde.es/servicio/software/econome.htm), in order to remove autocorrelation in the residuals. Unit root tests have been repeated on the de-seasonalized series and are reported in annex E.

[82] I.e., $t = \gamma = 0$. Indeed, the series don't show any linear trend in levels, but the presence of a constant in the cointegration space is needed to account for all those factors contributing to price differentials not explicitly modelled.

[83] The optimum number of lags selected for the underlying VAR is 5 according to the SBIC.

In the first VECM estimates, based on policy considerations (see paragraph 4.2.2), three basic structural breaks are imposed:

- the implementation of the MacSharry reform in 1993:07, which we expect had the effect of lowering the French price, due to the reduction of intervention prices (dummy variable $reg1_t = 1$ if $t \geq 1993{:}07$, and 0 otherwise);
- the implementation of the URAA in 1995:07, which, if tariffication was effectively in place, should have increased price transmission elasticities. For this reason, both the dummy $reg2_t = 1$ if $t \geq 1995{:}07$, and 0 otherwise, and the interaction term $reg2hrw_t = hrw_t * reg2_t$ are introduced in the cointegration vector;
- Agenda 2000 reform, that also set the lowering of intervention prices ($reg3_t = 1$ if $t \geq 2000{:}07$, and 0 otherwise).

By inserting these regime dummies and the interaction term, the cointegration vector becomes (assuming it is normalized with respect to the French price):

$$\left[swfr_t \quad \beta_0 \quad \beta_1 hrw_t \quad \beta_2 reg1_t \quad \beta_3 reg2_t \; \beta_4 reg2hrw_t \quad \beta_5 reg3_t\right] \tag{6.2}$$

Based on the effects that the various policy regime changes are likely to have had, we can make some assumptions concerning the expected sign of the parameters; we expect:

- $\beta_0 < 0$, i.e., a positive constant term in the cointegration relationship;
- $\beta_1 < 0$ and, presumably, $-1 \leq \beta_1 < 0$, i.e., overall, a positive price transmission elasticity between the US and the EU price, which might be at the most equal to 1 (perfect price transmission);
- β_2 , $\beta_5 > 0$, since both the implementation of the MacSharry reform and of Agenda 2000 are expected to have lowered the French price;
- $\beta_4 < 0$ (and, possibly, $|\beta_4 + \beta_1| \leq 1$), since we expect the URAA to have increased the value of the price transmission elasticity, at maximum up to the value of 1 (perfect price transmission) after the URAA.

Estimates are reported in table 6.4.

In the cointegration vector $\beta_1 = -0.250$ and $\beta_4 = -0.314$; this means that, after the URAA implementation, the price transmission elasticity between the two prices has increased up to the value of 0.564. As expected, β_2, β_3, $\beta_5 > 0$; this means that the constant of the cointegration relation between the French and the US price decreased (most likely, as the French price decreased following the implementation of the MacSharry and of Agenda 2000 reforms), from 3.835 to 3.671 (i.e., 3.835-0.164) after the Mac Sharry reform, to 2.009 (3.835-0.164-1.662) after the URAA implementation, and to 1.918 (3.835-0.164-1.662-0.091) after Agenda 2000.

Table 6.4 VECM with structural breaks: EU and WTO policy regimes

	$\Delta swfr_t$	Δhrw_t		Cointegration vector
α	-0.044***	0.142***		
	(0.014)	(0.047)		
$\Delta swfr_{t-1}$	-0.109*	0.080	β_0	-3.835***
	(0.060)	(0.202)		(0.347)
$\Delta swfr_{t-2}$	0.308***	0.255	β_1	-0.250***
	(0.060)	(0.203)		(0.069)
$\Delta swfr_{t-3}$	0.281***	-0.518**	β_2	0.164**
	(0.062)	(0.208)		(0.049)
$\Delta swfr_{t-4}$	0.195***	-0.064	β_3	1.662
	(0.064)	(0.216)		(0.819)
Δhrw_{t-1}	0.023	0.334***	β_4	-0.314
	(0.018)	(0.062)		(0.166)
Δhrw_{t-2}	-0.053***	-0.038	β_5	0.091
	(0.019)	(0.065)		(0.049)
Δhrw_{t-3}	-0.004	0.057		
	(0.019)	(0.065)		
Δhrw_{t-4}	-0.066***	0.047		
	(0.019)	(0.063)		
LM	2.249 (p-value 0.003)	0.717 (p-value 0.792)	$\beta_1 + \beta_4 = -1$	$\chi^2 = 2.318$ (p-value 0.128)
ARCH(5)	11.628 (p-value 0.040)	12.381(p-value 0.030)	$\beta_2,\beta_3,\beta_4,\beta_5 = 0$	$\chi^2 = 14.525$ (p-value 0.006)
SBIC	-8.146	HQN -8.595		

Standard errors in parenthesis. *significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients
LM: LM test with null hypothesis of no-autocorrelation up to the 18th lag. ARCH(5): LM test with the null hypothesis of no ARCH effects up to the 5th lag. SBIC : Schwartz's Bayesian Information criterion. HQN: Hannan-Quinn Information criterion

Both the French and the US adjustment coefficients (α) are significant and have the right sign: both prices respond to the disequilibria from the long-run relation where structural breaks have been taken into account. Interestingly, the US coefficient is higher in absolute value. This means that, being both adjustment coefficients significant, although none of the price is weakly exogenous, the US price responds more quickly to the disequilibria from the long run relation than the French price. However, the French price equation still displays autocorrelation, which indicates evidence of misspecification in the model.Indeed, within the simple model which has been presented, some additional factors can be considered (see paragraph 4.2.2), which also define specific policy regimes, although shorter in time, that have affected price transmission. First of all, we introduce export taxes, that have been imposed for very short periods of time ($t1_t$ =1 if 1984:08 $\leq$ t $\leq$ 1985:04 and zero otherwise; $t2_t$ =1 if 1995:12 $\leq$ t $\leq$ 1996:09 and zero otherwise), and the inflow of Russian and Ukrainian grains (u_t = 1 if 2001:07 $\leq$ t $\leq$ 2002:12 and zero otherwise). Introducing these additional factors, the cointegration vector becomes:

$$[swfr_t \quad \beta_0 \quad \beta_1 hrw_t \quad \beta_2 reg1_t \quad \beta_3 reg2_t \quad \beta_4 reg2hrw_t \quad \beta_5 reg3_t \quad \beta_6 t1_t \quad \beta_7 t2_t \quad \beta_8 u_t] \tag{6.3}$$

As explained in chapter 3, all these factors are expected to have lowered the French prices with respect to the US ones, that is, β_6, β_7, $\beta_8 > 0$, without having a direct impact on price transmission elasticities. Estimates are reported in table 6.5.

Table 6.5 VECM with structural breaks: EU and WTO policy regimes, export taxes, 2001/02 campaign

	$\Delta swfr_t$	Δhrw_t		Cointegration vector
α	-0.066*** (0.010)	0.012 (0.040)	β_0	-3.810*** (0.367)
$\Delta swfr_{t-1}$	-0.175*** (0.060)	0.297 (0.227)	β_1	-0.257*** (0.074)
$\Delta swfr_{t-2}$	0.231*** (0.060)	0.282 (0.228)	β_2	0.164** (0.045)
$\Delta swfr_{t-3}$	0.256*** (0.060)	-0.488** (0.228)	β_3	-0.664 (1.061)
$\Delta swfr_{t-4}$	0.197*** (0.062)	0.147 (0.235)	β_4	0.170 (0.218)
Δhrw_{t-1}	0.031* (0.017)	0.324*** (0.066)	β_5	-0.236*** (0.074)
Δhrw_{t-2}	-0.055*** (0.018)	-0.118* (0.069)	β_6	-0.013 (0.100)
Δhrw_{t-3}	0.008 (0.018)	0.062 (0.069)	β_7	-0.179* (0.118)
Δhrw_{t-4}	-0.079*** (0.018)	0.0002 (0.068)	β_8	0.372*** (0.082)
LM	1.161 (p-value 0.296)	0.547 (p-value 0.932)	$\beta 1 + \beta 4 = -1$	$\chi 2 = 9.213$ (p-value 0.002)
ARCH (5)	24.667(p-value .0002)	11.540 (p-value 0.042)		$\beta 2, \beta 3, \beta 4, \beta 5, \beta 6, \beta 7, \beta 8 = 0$ $\chi 2 = 34.668$ (p-value 0)
SBIC	-7.404	HQN -8.301		

Standard errors in parenthesis. *significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients
LM: LM test with null hypothesis of no-autocorrelation up to the 18th lag. ARCH(5): LM test with the null hypothesis of no ARCH effects up to the 5th lag. SBIC : Schwartz's Bayesian Information criterion. HQN: Hannan-Quinn Information criterion

The autocorrelation in the French equation has been removed. This time, even if both the French and the US adjustment coefficients have the expected sign, the US one is not significant, which may suggest its weak exogeneity. In the cointegration vector, although β_0, β_1 and β_2 have very similar values to those reported in table 6.4 where the additional factors were not considered, β_4, β_6 and β_7 have not the expected sign. For the latter two, this might well depend on the short number of observations available for these specific regimes, since export taxes were imposed for a very few months. Instead, $\beta_8 > 0$ as expected; moreover,

as the two dummies largely overlap, this might explain why β_5 , the coefficient of the dummy associated to the MacSharry reform, gets a negative sign. For this reason, the model is estimated without export taxes (table 6.6). This time, all the coefficients of the cointegration vector have the expected sign (for β_5, the same consideration made for the previous model holds). Both adjustment coefficients have the expected sign. Once again, the US price performs as weakly exogenous.

Table 6.6 VECM with structural breaks: EU and WTO policy regimes, 2001/02 campaign

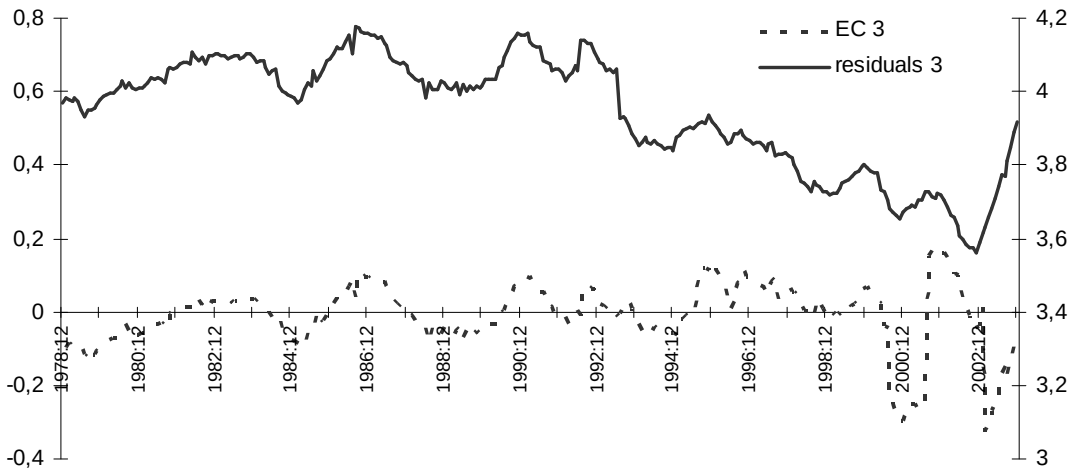
	$\Delta swfr_t$	Δhrw_t		Cointegration vector
α	-0.067***	0.048		
	(0.012)	(0.043)		
$\Delta swfr_{t-1}$	-0.185***	0.198	β_0	-4.066***
	(0.058)	(0.213)		(0.310)
$\Delta swfr_{t-2}$	0.248***	0.334	β_1	-0.203***
	(0.060)	(0.219)		(0.062)
$\Delta swfr_{t-3}$	0.255***	-0.416*	β_2	0.163**
	(0.060)	(0.218)		(0.044)
$\Delta swfr_{t-4}$	0.175***	0.010	β_3	1.072
	(0.062)	(0.225)		(0.740)
Δhrw_{t-1}	0.025	0.311***	β_4	-0.190
	(0.018)	(0.064)		(0.150)
Δhrw_{t-2}	-0.046**	-0.078	β_5	-0.154**
	(0.018)	(0.067)		(0.062)
Δhrw_{t-3}	0.001	0.042	β_8	0.380***
	(0.018)	(0.067)		(0.079)
Δhrw_{t-4}	-0.066***	0.025		
	(0.018)	(0.065)		
LM	1.554 (p value 0.074)	0.564 (p value 0.922)		$\beta_1 + \beta_4 = -1$ $\chi^2 = 15.920$ (p-value 0)
ARCH (5)	13.525 (pvalue 0.019)	13.519 (p value 0.019)		$\beta_2, \beta_3, \beta_4, \beta_5, \beta_6, \beta_7 = 0$ $\chi^2 = 33.485$ (p-value 0)
SBIC	-7.910	HQN -8.508		

Standard errors in parenthesis. *significant at 10%; **significant at 5%; *** significant at 1% for the null hypothesis of zero coefficients
LM: LM test with null hypothesis of no-autocorrelation up to the 18th lag. ARCH(5): LM test with the null hypothesis of no ARCH effects up to the 5th lag. SBIC : Schwartz's Bayesian Information criterion. HQN: Hannan-Quinn Information criterion

Since this third model presents better information criteria than the one presented in table 6.5, this last one is preferred. The null hypothesis of a perfect price transmission elasticity after the URAA implementation cannot be accepted. The long run residuals of the model (figure 6.1; *EC 1* dotted line, left-hand scale) show a stationary behaviour around zero. These long run residuals from the cointegration relation in which structural breaks are included are compared with the series called *residuals* = ($swfr_t$ - 0.203 * hrw_t) (figure 6.1; *residuals*; unbroken line, right-hand scale), obtained just subtracting from the French price the US one multiplied by its overall transmission parameter. This series is of course higher in magnitude, as the constant term is missing, and is reported to

show how the residuals from the overall cointegration relation have a constant which changes over time, in correspondence to structural breaks. In particular, the two series have a common patter up to 1993:06, after which the structural breaks are considered in the error correction term of the VECM estimated by allowing for a lower constant and an increased price transmission elasticity.

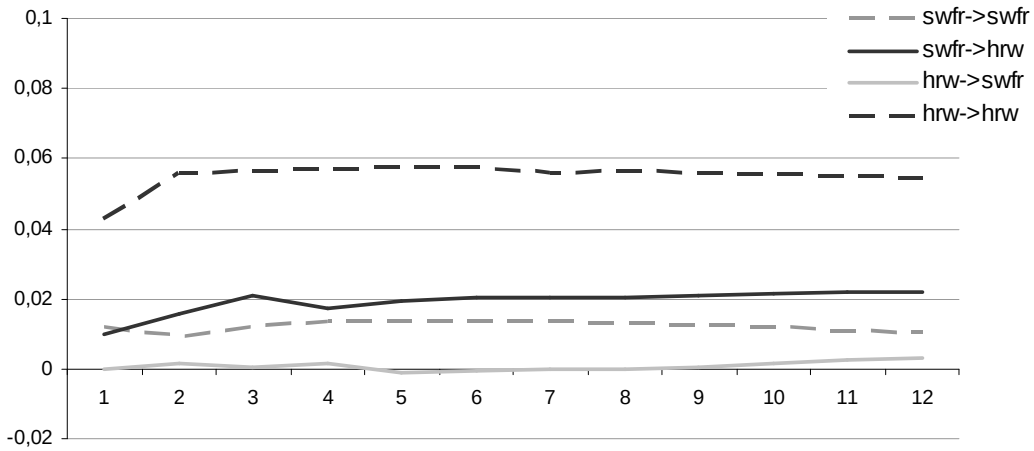
Figure 6.1 Long run residuals of the VECM with structural breaks (EU and WTO policy regimes, 2001/02 campaign)



When cointegration is present, we expect IRFs not to always die out: their effect can be permanent, as the whole system will move to a new long run equilibrium. Both transitory and permanent shocks can be present (Goodwin *et al.* 1999, p.168). To calculate orthogonal IRFs, the ordering of the Choleski decomposition for the VAR reformulated in levels is $swfr_t$, hrw_t. This means that, at time *t*, the French price depends on only one contemporaneous structural shock, whereas the US price depends on the two of them. We could interpret the first one as a shock occurring in the EU market, which affects contemporaneously both the French and the US price, and the second one as a shock happening on world markets, which at time *t* affects the US price only. In other words, we assume that the US price has no contemporaneous effects on the French price, but lagged values of the US price have an effect on the French price. The French price has, instead, a contemporaneous effect on the US one. We see that all IRFs, for a 12 months time horizon, are generally low but different from zero (figure 6.2), with the exception of the French response to a shock in the US price. Both the French and the US price response to a shock in themselves tend to be persistent, whereas the response of each price to a shock in the other price is generally lower and

shows an upward pattern. In the long run, the ratio between the effects of a shock to the French price on *swfr* and *hrw* is 0.473, the long run transmission elasticity.

Figure 6.2 IRFs of the VECM with structural breaks (EU and WTO policy regimes, 2001/02 campaign)



6.4.1 Assessing the robustness of the cointegration vector

To assess the robustness of the parameters of the cointegration vector, the estimates have been repeated also using an alternative estimator for cointegrated systems, the Dynamic OLS estimator (or DOLS) developed by Stock and Watson (1993). Indeed, the models presented in tables 6.4, 6.5 and 6.6 have been re-estimated by inserting the relevant policy regime dummies in equations like 6.4:

$$\mathbf{swfr_t} = \mathbf{hrw_t} + \sum_{\mathbf{i=-m}}^{\mathbf{+m}} \mathbf{\Delta hrw_{t-i}} + \mathbf{X_t} + \boldsymbol{\varepsilon}_\mathbf{t} \tag{6.4}$$

in which the French price was regressed on the US price and on his leads and lags differences (selected according to the SBIC), plus the regime dummies concerned (here indicated, for the sake of simplicity, with $\mathbf{X_t}$). Using the DOLS allows focusing on the long run relations between the prices[84].In general, all estimates yield consistent results, which confirms the robustness of the parameters of the cointegration vector. To give an example, the estimates of the model in

[84] Indeed, a possible objection to the estimates which have been presented in paragraph 6.4 is that also the short run parameters might be influenced by the structural breaks. This would require estimating different models for the different sub-samples, which is not advisable given the limited number of observations available.

which just EU and WTO policy regime changes are considered are presented in table 6.7.

Table 6.7 DOLS estimates: EU and WTO policy regimes

Variable	**Coefficient**
const	4.279***
	(0.143)
hrw_t	0.160***
	(0.028)
$reg1_t$	-0.191***
	(0.012)
$reg2_t$	-2.019***
	(0.290)
$reg2us_t$	0.396***
	(0.059)
$reg3_t$	-0.188***
	(0.019)
Δhrw_{t-6}	-0.177***
	(0.047)
Δhrw_{t-5}	-0.126***
	(0.035)
Δhrw_{t-4}	-0.177***
	(0.040)
Δhrw_{t-3}	-0.158***
	(0.043)
Δhrw_{t-2}	-0.137***
	(0.046)
Δhrw_{t-1}	-0.153***
	(0.046)
Δhrw_t	-0.149***
	(0.046)
Δhrw_{t+1}	0.103**
	(0.050)
Δhrw_{t+2}	0.144***
	(0.048)
Δhrw_{t+3}	0.073
	(0.048)
Δhrw_{t+4}	0.101**
	(0.046)
Δhrw_{t+5}	0.104**
	(0.046)
Δhrw_{t+6}	0.111**
	(0.051)
SBIC	-933.188

Standard errors in parenthesis. *significant at 10%; **significant at 5%;
*** significant at 1% for the null hypothesis of zero coefficients
SBIC : Schwartz's Bayesian Information criterion

It can be noticed that the long run transmission elasticity (0.160) and its increase after the URAA implementation (0.396) show values which are consistent with those presented in table 6.4.

6.5 Concluding remarks

This chapter provides a further examination of short and long-run co-movement relations between US and internal EU prices for soft wheat by considering policy regime changes.

The major EU and international market policy changes have here been taken into account by using cointegration techniques allowing for structural breaks in the deterministic trend of the cointegrating vector. The model proposed by Johansen, Mosconi and Nielsen (2000) has been modified by considering more than one structural breaks.

Also in light of the analysis carried out in the previous chapter, it is confirmed that, in the years 1978-2003, the US and the EU price common path emerges only when structural breaks, i.e. policy regime changes, are appropriately considered.

Indeed, the empirical analysis shows that both prices respond to deviations from the long run relation, which has been affected by policy regime changes: namely, the MacSharry and Agenda 2000 reform of the CAP, and the URAA implementation. The introduction of an additional dummy accounting for the peculiar tariff situation of the 2001/02 campaign allows to find a better model specification.

As expected, the MacSharry reform and Agenda 2000 are responsible for the reduction of domestic EU prices, while the URAA causes an increase in the price transmission elasticity (from about 0.20 before the implementation of the URAA to about 0.40 afterwards; and this in addition to a further reduction of prices).

Even if, remarkably, evidence of co-movement between the two prices is confirmed, still full price transmission is not in place.

Concerning the adjustment coefficients, interestingly, we note that the French ones are always significant and within the range that emerged in the previous chapter (see paragraph 5.3). Namely, they are close to the values shown when both the intervention and the US prices appear in the cointegration vector. This can be seen as a confirmation that the technique presented here is a consistent alternative to the ones presented in the previous chapter: the introduction of structural breaks in the deterministic term of the cointegration vector elicits the same response from the French price than when policy regime changes are represented by introducing the intervention price in the models.

Also the response of the US price is consistent: in fact, its adjustment coefficients are always positive and bounded between 0.012 and 0.142. In this respect, it would seem that the model presented in this chapter really solves some

of the interpretative problems which came out in the previous one (paragraph 5.2.2.1 and 5.3). Consistently with chapter 5, the US price behaves as weakly exogenous.

To sum up, our findings confirm and contribute to shed light on a complex and controversial empirical literature, bridging the gap between studies not finding any evidence of co-movement and those that, conversely, find evidence of increased price transmission elasticities following the URAA.

What we find is that co-movement between the two series exists, albeit the price transmission elasticity is small, and has been influenced by domestic and liberalizing policy reforms.

The analysis show that the implementation of the URAA almost doubled the elasticity of transmission between European and US prices. This means that the markets could benefit from an increased efficiency.

This exercise provides an alternative way of analyzing US and EU price patterns in the past 30 years by explicitly considering different policy regimes.

In order to verify the effects of the major changes of domestic and border policies on price transmission, the use of time series allowing to go back in time was preferred (Eurostat data). As noted already, this in turn implied that, unfortunately, more recent phenomena could not be investigated.

In the latest months, in particular, commodity world markets have been characterized by soaring prices whose causes have both a transitory and a structural nature. This provoked remarkable changes in the commercial policy measures in place in a number of countries, and in the EU. In our analysis this implies that, presumably, other structural breaks could be introduced in the cointegration vector, although their timing might not be so clear cut. In the EU, for example, the elimination of import duties, which was already included in the previous regulations to be applied in case of high world prices, keeps a transitory nature.

7 CONCLUDING REMARKS

This work aims at providing a theoretical framework to consider policy regimes while testing for price transmission, and at developing different models for the empirical analysis.

The case study is constituted by international soft wheat markets in the years 1978-2003. The French and the US prices are assumed to represent, respectively, the EU and the world ones.

This work tries to shed light on a mixed empirical evidence. Indeed, the Law of One Price is often found not to hold in practice. Many factors are supposed to prevent prices from convergence: amongst them, policy measures are expected to play a relevant role.

This is true in particular for the agricultural sector, traditionally characterized by high levels of policy intervention. The recent rise in food prices on international commodity markets drew new attention on price transmission mechanisms and, specifically, on the effects of policy intervention.

Theoretical considerations on the functioning of domestic and border policies on international wheat markets (paragraph 4.3) lead to the basic assumption that the domestic EU price and the world one are expected to interact only when the latter is above the EU intervention price. The relative position of this policy variable and of the US price allows to define different policy regimes; alternatively, the intervention price can be considered as a threshold for the US price to interact with the European one.

Accordingly to these theoretical considerations, different empirical models are derived.

At first (chapter 5) the intervention price is used in different ways. Either a composite variable is built, constituted by the maximum between the intervention and the US price, or the LOP is imposed to hold between the French price and either the US or the intervention price depending on the policy regime in place, i.e. on which of the two prices is higher. In this case, either the adjustment coefficients, or the long run transmission elasticities, or, finally, both, are allowed to change.

These models represent an attempt of explicitly combining policy and price data. The observable policy regime change has been simulated with the construction of an *ad hoc* composite variable and other econometric devices.

The consistent behaviour of EU adjustment coefficients suggests that the role of the US price might be understood only in light of adequate consideration for policy regimes. The adjustment coefficients of the French price always display the correct sign, although they tend to be significant and higher in magnitude only when the intervention price appears in the cointegration vector.

The response of the US price leaves instead some questions open: US adjustment coefficients are usually much smaller in value and not significant. This could be due to its exogeneity, but, also, to the fact that the cointegration relation is mostly driven by the intervention price, which in turns brings some interpretative problems.

The model presented in chapter 6, instead, directly explores the dynamic of a system constituted by the US and the French prices, only. In this case, the policy regime changes are assumed to influence the parameters of the long run relation linking the two prices. The policy regime changes are then expected to have an impact on an existing relation, as some previous empirical evidence would suggest (Thompson and Bohl 1999, Thompson *et al.* 2002a and 2002b). A cointegration model allowing for structural breaks in the deterministic term of the cointegration vector is tested. The Johansen, Mosconi and Nielsen (2000) model is partially modified by introducing more than one break.

Once again, the French response to the disequilibria in the long run relation is found to be consistent and significant; moreover, also the response of the US price is consistent, which solves some of the interpretative problems of chapter 5.

The theoretical assumptions are confirmed: indeed, the MacSharry and Agenda 2000 reforms, by lowering intervention prices, did reduce the distance of the EU prices from the world ones; while the URAA implementation lead to an increased price transmission elasticity, although perfect price transmission is still not in place.

Summing up, our empirical findings, which derive from the specific theoretical framework developed, confirm and contribute to shed light on a complex and controversial literature, bridging the gap between studies not finding any evidence of co-movement between the prices considered and those that, conversely, find evidence of positive and increased price transmission elasticities following the URAA.

In addition to this, these findings are of a more general interest for the study of the dynamics occurring in global markets which are increasingly complex and interconnected.

Indeed, the study of the outcomes of the implementation of the URAA provides insights about the effects of all liberalization reforms aimed at increasing market access.

Moreover, analyzing the effects of reduction of the intervention prices is of a crucial importance within the "Health Check" debate of the CAP, since a substantial reduction of this mechanism is currently being proposed by the European Commission.

However, a number of issues are left open for future research.

These basically stem from some of the simplistic assumption that necessarily had to be made in the empirical applications proposed: the way in which unknown transport costs are modelled; the presence of market power; other possible exogenous drivers for the price dynamics; a more detailed study of the homogeneity issue. In addition to this, a deeper analysis of the relation between US and intervention prices, and especially of its evolution in time, can contribute to better understand the French and US price dynamics.

Finally, it should be recalled that, in this work, the use of data which would allow to go back in time has been preferred although this, in turn, implied the impossibility of studying the latest years. In this respect, some additional considerations emerge considering the most recent market dynamics: the dramatic rise in energy prices (and in freight rates) could increase the complexity of the assumptions made on transaction costs; moreover, soaring food prices caused remarkable changes in the institutional settings of international trade.

The effects of more recent policy and market developments should be the object of further research; considering the evolving international context goes then well beyond the use of more recent data, and probably represents the most interesting possibility of developing the framework presented.

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ANNEXES

Annex A – Empirical analysis on price transmission: a review

METHODOLOGY	DATA USED	MAJOR FINDINGS
CORRELATION		
Isard, 1977		
By visual inspection and by ordinary regression (Cochrane-Orcutt procedure), the behaviour of relative prices for various industries, is examined.	Monthly time series of US, Germany and Japanese prices for a variety of industries collected at different frequencies for the years 1968-1975.	Substantial changes in exchange rates have substantial and persistent effects on the relative common currency prices for close substitute products of different countries.
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 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		
He applies both Ravallion's radial dynamic model of market integration and cointegration techniques. Cointegration relations are analyzed within the same market between commodities that can be considered close substitutes. His objective is to verify whether, in developing countries, the connections existing among markets of different agricultural commodities allow governments to use policy measures addressed to only one of them.	Deflated monthly wholesale maize, sorghum and millet prices (levels) from three markets in Ghana. 1977-1990.	Prices of different commodities are cointegrated. Often, the past prices of one product can be used to predict the prices of another one. Functional efficiency in Ghana's coarse grain markets is found.
Mundlak and Larson, 1992		
They investigate world markets to find out what proportion of world prices is transmitted to the domestic ones (coefficients of the regressions) and what proportion of the variation in domestic prices can be attributed to the variations in world prices (R^2 of the regressions). They analyze different panel models (all commodities for each country, nominal and real values; all countries pooled together; insertion of yearly dummies; between and within commodity regressions).	In a first analysis, the prices refer to 58 counties and 60 products, in the years1968-1978; then they examine the prices in EU countries, for 25 years.	This study finds very high transmission elasticities for all countries (in the range of 0.74-1.24) and very high values of R^2 (in the range of 0.66 and 0.96). The results are however likely to be biased by complete disregarding the time series properties of the data (autocorrelation tests are not reported).

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: RAVALLION'S CRITERIA		
Ravallion, 1986		
In his model of 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 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 really zero under the old CAP, and how it has been affected by policy regime changes. Germany is assumed to be a small country for which the world price is exogenous. An econometric model is also built. They analyze the effect of the change in the variability of prices on consumers' welfare.	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 the CPI of their respective country. 1976-1998.	The MacSharry reform caused a reduction in the levels of EU prices, but the URAA didn't have any significant effect on price transmission elasticities, that have an average value of 0.11 (which means that tariffication was not effective). The existence of low but not zero price transmission elasticities before the URAA suggests 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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: RAVALLION'S CRITERIA		
Thompson *et al.*, 2002b		
They develop an aggregate structural model of the EU wheat economy, linked to the rest of the world. Amongst the structural equations estimated with 3SLS techniques, the price transmission equation regresses the EU price on a constant term, the world price, a time trend, and a dummy variable accounting for the MacSharry reform.	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 the CPI of their respective country. 1976-2000.	They find a price transmission elasticity of 0.18, as the one obtained by Thompson *et al.* (2002a). The MacSharry reform is found to have caused a significant downward shift in prices; the addition of an interaction term is not significant.
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL (GC TESTS, IRFS, FEV)		
Goodwin and Schroeder, 1991a		
They estimate a VAR in levels between the prices. 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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL (GC TESTS, IRFS, FEV)		
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 markets. Week 1:1977 - week 50:1980	They find that markets are price-efficient, since the tests show that they are interdependent.
Myers *et al.*, 1990		
VAR methods are used to analyze the contribution of supply, demand and policy shocks to unpredictable fluctuations in the market for Australian wool. They develop a structural VAR in which the restrictions (which define contemporaneous interactions among variables) do not impose a recursive ordering of the variables. They perform IRF and FEV.	Quarterly prices for Australian wool, quantity purchased by private traders, total quantity of wool sold (in logs). 1971:1 to 1988:2.	Demand shocks are the dominant source of wool market fluctuations but intervention has blunted their effects.
Vollrath and Hallahan, 2006		
They investigate market integration for meat and livestock in the US and Canada, making use of three models: "streamlined" (prices expressed in a common currency) and "detailed" LOP models (to quantify both foreign price and exchange rate transmission), VAR models (IRF). They consider trade policy barriers and exchange rate changes, and evaluate the CUSTA (Canadian US Free Trade Agreement, since 1989) and its extension, the NAFTA (North American Free Trade Agreement, since 1994)	US and Canada wholesale monthly and weekly prices for slaughter (steers, hogs, whole chicken), two cuts of beef and two pork products (in logs); US/Canada exchange rate. 1988 - 2002	The market of whole chicken is segmented; hog and pork product markets are more integrated than steer and beef product markets. Changes in the Canadian-US exchange rate inhibits cross border integration.
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Ardeni, 1989		
For the first time, cointegration is proposed to analyze the LOP, arguing that previous tests are flawed for disregarding the time series properties of the data.	Quarterly export prices (in logs) for wheat, wool, beef, sugar, tea, tin and zinc for Australia, Canada, the UK, the US. Data are adjusted for exchange rates. January 1957-January 1986 (from 79 to 117 obs., depending on the commodity).	The LOP holds only in a small number of cases (but is nonetheless valid for US-Australian wheat, US-Canadian wheat); deviations from the pattern are permanent.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Asche *et al.*, (1999)		
The authors analyze the relation between market integration and product aggregation. In the deterministic case, if the Composite Commodity Theorem holds, the LOP holds, as well, and vice versa: in fact, for both conditions to hold, prices must move proportionally over time. In the stochastic case, the results also apply hold for relationships with stationary deviations.	fob Norwegian and US salmon prices for four different Pacific species. January 1986 – December 1996	The analysis indicates that the species compete in the same market, as their prices are cointegrated; the LOP holds; the generalized composite theorem holds.
Baek and Koo, 2005		
This study examines price dynamics in the U.S. and Canadian hard red spring and durum wheat markets by using a vector error-correction model. The presence of a structural break is taken into account by splitting the sample in two.	Monthly fob prices for the US and Canadian durum and hard red spring prices, already converted in dollars. July 1979 to June 2002	The Hard Red Wheat exporting industry and Canada have been the price leader in North American wheat markets.
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.	Same as Ardeni (1989), 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. The analysis is repeated on the two sub-samples identified.	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 MacSharry CAP reform was implemented. Before that, the only cointegrating relation existing is between the two US wheats. After that, they find a cointegrating vector between the US Hard Red Winter wheat and the US Soft Red Winter wheat, and another one between the US Soft Red Wheat and the EU. The EU acts as a price follower, as it would be expected since it makes use of export subsidies.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Brooks *et al.* (2005)		
They study the transmission of world to domestic prices in Brazil for a number of commodities. Standard cointegration procedures are compared in the overall sample and in two sub samples identified by the presence of policy regime changes, and then with the threshold model of Bailey, Balcombe and Brooks (BBB; 2006).	Brazilian monthly prices for wheat, maize, rice, dry beans, soybeans, coffee, poultry, pig meat and beef (logs). January 1989 to October 2003.	Cointegration tests for Brazilian prices support the hypothesis of discontinuous price transmission and the presence of different regimes following policy reforms in the mid-1990s (with differences amongst the different markets). Once threshold effects are taken into account, prices show a much faster rate of adjustment outside the bound. The thresholds identified by BBB are sometimes difficult to reconcile with data on trade flows.
Bukenya and Labys, 2005		
They examine the degree to which commodity prices have converged on world commodity markets by using correlation, regression, VAR and cointegration analysis.	Annual data for tin, copper, lead, coffee, cotton, wheat (in US dollars). 1930-1998	The empirical results don't support price convergence, with the exception of lead and wheat.
Conforti, 2004		
Both spatial and vertical price relations are investigated through the use of Autoregressive Distributed lag Models from which the corresponding Error Correction specification is derived. Annual price information is analyzed by testing for a long run equilibrium between the price series; monthly information by paying more attention to the dynamics of the relation between prices, to their causality (GC tests in both directions for the ARDL), and to the asymmetry of price transmission.	Both annual and monthly prices for a number of basic food commodities collected from various sources in 16 countries (Argentina, Brazil, Chile, Costa Rica, Egypt, Ethiopia, Ghana, India, Indonesia, Mexico, Pakistan, Senegal, Thailand, Turkey, Uganda, Uruguay). The study used import prices, producer prices, wholesale prices and retail prices (in logs).	Three main findings emerge: geographical regularity (African countries show lower degree of price transmission than Asian and Latin America ones); vertical transmission appears to be higher than the transmission of changes in the world reference prices; cereals show higher and faster price transmission, followed by oilseeds, while price transmission for livestock is poorer.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Dawson and Sanjuan and White, 2006		
They apply Johansen's procedure allowing for structural breaks. The time of occurrence of the structural break is known. Co-movement between futures prices can arise when commodities are substitutes. They investigate the relation between feed barley and wheat prices on the London International Financial Futures Exchange.	London International Financial Futures Exchange weekly data for feed wheat and barley future prices (in logs). 1996-2002	They find a significant break in October 2000 (due to Common Agricultural Policy intervention price reductions and poor sowing conditions). The barley–wheat futures market is perfectly integrated; the barley price Granger-causes the wheat price. Modelling structural breaks in price relationships appears of big importance.
Dawson and Sanjuan, 2006		
They apply Johansen's procedure allowing for structural breaks to investigate the dynamics of North America wheat markets and verify the effects of the Export Enhancement Program. 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. The EEP has caused the US/Canada price ratio to fall.
Dercon and Campenhout, 1999		
They apply a threshold error correction model, the Band-TAR model. Such a model, like switching regime models (they explicitly refer to Baulch's parity bound model), allows for trade discontinuities and trade flow reversals, but uses more reasonable distributional assumptions and is dynamic, explicitly considering the process of arbitrage in the form of non-linear correction.	Same as Baulch, (1997).	They find that in most markets full arbitrage is achieved with trade taking place; but, differently from Baulch, that in others arbitrage is slow and rather inefficient. This might be due to imperfect competition. This is explained by the fact that, in Baulch's model, even in the efficient arbitrage regime, the transfer costs are assumed to be measured with errors (but dynamic properties of the errors are not considered), whereas here the price margin is taken as such. On the other side, they probably overestimate the times in which the efficient arbitrage condition is violated.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
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. All wheats used for human consumption show a common trend, while feed wheats have another one.
Ghoshray, 2002		
He examines price differentials in the international wheat market using threshold autoregressive (TAR) and momentum threshold autoregressive (M-TAR) models, that allow for asymmetric adjustment.	Monthly average export price quotations on 8 series of Australian, Canadian, US and EU hard medium and soft wheat, analyzed in pairs (in logs). July 1980 to December 1985	The world wheat market is highly integrated and little evidence for asymmetry exists. Where asymmetry is found, it might be due to the quality of wheat (Canadian and Australian prices for higher quality wheats may respond easier to price increases than to price decreases).
Gjoelberg *et al.* (1996)		
A set of conditions to hold is proposed to have market integration: not only must prices be cointegrated in a dynamic specification and adjust rapidly towards equilibrium, but also they must not systematically fluctuate outside transaction costs ranges, they must react simultaneously to new information (i.e., have a common behaviour at the harvest time), display similar paths within the marketing year, and tend towards "one price".	Monthly producer prices for soft wheat (Eurostat) for Belgium, Italy, germany, the United Kingdom and France January 1984-December 1994	Evidence of no integration is found for EU wheat markets; only limited evidence of convergence towards a common price can be found. This can be due to distorsive effect of the CAP, due for example to the Monetary Compensatory Amounts, functioning as subsidies/taxes on intra-EU trade.
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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Goodwin *et al.*, 1999		
They evaluate spatial price dynamics and integration for selected Russian food markets by using conventional time series tests of spatial integration with an analysis of dynamic responses to price shocks (GC and IRF).	Weekly eggs, milk, vegetable oil and potatoes prices in five Russian cities in both retail and "gray" (free) markets (in logs). Results were repeated with deflated prices and were found similar. June 1993-December 1994.	Results provide tempered support for spatial integration, especially in retail market. Retail markets have bigger efficiency advantages and economies of scale, which allow a better price transmission than for "gray" markets. Because of gradual adjustments to price shocks, integration may occur mainly in the long run.
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 (the smaller markets exhibit a smaller degree of spatial dependence than the 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.
Hui Shung and Griffith, 1998		
They investigate short and long run dynamics in the Australian beef market between farm, wholesale and retail prices, with cointegration and IRF.	Monthly beef farm, wholesale and retail prices in Australia. January 1971 to September 1994	The three prices considered are cointegrated; the wholesale price is found to be weakly exogenous, which could be an indication of market inefficiency due in part to price levelling.
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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Mohanty and Langley, 2003		
They examine the co-movement 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.	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 misspecified for disregarding the time series properties of the data.
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 one.
Sharma, 2002		
The Engle-Granger procedure is used to test for price co-movement in Asian wheat markets.	Monthly wholesale or retail prices for cereals for eight Asian markets. January 1990-December 1999	Both the review of past estimates and their findings confirm that price transmission elasticities tend to be very low (0.2-0.4 in the short run, and higher in the long run, but still less than unity in the import case) for most developing countries.
Thompson and Bohl, 1999		
A cointegration analysis is performed to check how policy regime changes affected international price transmission elasticities. A threshold cointegrating technique is used.	Domestic prices for Germany are from Eurostat (in 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 any statistical difference. All prices in logs. June 1976 to December 1998.	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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Thompson *et al.*, 2002a		
Their focus is on the estimation technique and on the extent to which liberalizing policy reforms contribute to market spatial efficiency. They use Seemingly Unrelated Regression Augmented Dickey Fuller tests and error-correction models to account for cross-equation correlations among markets.	Soft wheat quarterly prices for Germany, France and the UK (Eurostat; domestic prices), and US cif Rotterdam. All prices are nominal and converted in dollars with IMF exchange rates (in logs; the analysis on deflated prices yields no significant differences). 1976:3-1999:4	The adjustment coefficients for EU countries (the same speed of convergence is imposed amongst all countries) increase after 1993 and 1995, but nonetheless keep low due to the imperfect tariffication realized under the URAA.
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 EU prices are cointegrated with the institutional prices (but they claim this relation is unstable and thus distorted) but not in the subsample 1995-2002. 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 Eurostat, and are all converted to ecu/euros. January 1975-December 2004.	The prices are highly cointegrated after the three countries joined EU. Beyond the presence of arbitrage, results could be explained by the fact that all products studied (except of potatoes) are under CAP support. In the pre-EU period, cointegration among prices was discovered for rye, soft wheat and barley between Austria and Germany (neighbouring countries with strong transportation ties).

METHODOLOGY	DATA USED	MAJOR FINDINGS
DYNAMIC REGRESSION MODELS BASED ON A POINT LOCATION MODEL: COINTEGRATION		
Wang and Tomek, 2007		
They investigate what seems to be inconsistency between price theory, which suggests that commodity prices should be stationary series, and tests for unit roots, which rather frequently imply that these prices are not stationary. Unit root tests are applied to nominal, deflated, log nominal, log deflated prices. Tests are run with and without a trend, and general to specific method for lag length is used.	Monthly prices for corn, soybeans, milk, barrow, gilts, weekly prices for corn and soybeans. All in Illinois. CPI is used for deflating prices 1960-2005	Results of unit root tests vary with their specification. Evidence favouring unit roots in commodity prices is not strong, even without accounting for structural breaks. .
Williams and Bewley, 1995		
They investigate the presence of arbitrage opportunities between four Australian cattle markets.	Weekly cattle prices for 4 Australian markets (in logs) March 1986 - August 1989.	Arbitrage opportunities are usually short lived and long run equilibrium is restored, except for one market which is identified as a satellite market.
Zanias, 1993		
For four products (for which different levels of political support exist), they run cointegration tests amongst all possible pairs of countries for which data are available. Both the unrestricted and the restricted version of the LOP (where perfect price transmission is I posed) are tested.	Monthly prices from 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. For soft wheat markets, price co-movement is found after prices have been corrected by Monetary Compensatory Amounts. The existence of a minimum intervention price is probably fundamental, but efficient arbitrage is also necessary.
SWITCHING REGIME-THRESHOLD MODELS		
Balcombe *et al.*, 2007		
They use a threshold cointegration model which introduces a generalization on existing threshold models: the prices could be attracted to either the edge of the threshold interval or the to some point within the interval. They are estimated with a Bayesian approach.	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 dollars.	Evidence of thresholds is found in three out of the five commodity pairs investigated. The size of the thresholds depends upon the commodities considered.

METHODOLOGY	DATA USED	MAJOR FINDINGS
SWITCHING REGIME-THRESHOLD MODELS		
Barrett and Li, 2002		
They introduce a new methodology based on maximum likelihood estimation of a mixture distribution model incorporating price, transfer costs, and trade flow data. Four different regimes are developed according to the presence or absence of market efficiency and integration.	Monthly data on prices, trade flows, and estimated inter-market transaction costs for soybean meal over the years 1990-1996 for Canada, Japan, Taiwan and the United States.	The methodology proposed, which is particularly useful when trade is discontinuous or bilateral and transfer costs are variable, is an attempt to bridge price and quantity based approaches to the study of price transmission, which allows to distinguish between market integration and efficiency. Competitive tradability and equilibrium conditions prevail in the market examined.
Baulch, 1997		
He develops the Parity Bound Model (PBM), an alternative methodology in which information on transfer costs is used in addition to prices. According to the size of price spreads and of transfer costs, three regimes are derived (within, at and outside the arbitrage band).	Monthly wholesale price (in levels) for Philippine rice coming from eight different 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 detects efficient arbitrage when other tests do not.
Goodwin and Piggott, 2001		
They evaluate corn and soybean markets in North Carolina; they use threshold autoregression and cointegration models to account for a neutral band representing transaction costs. Regime switching is triggered when deviations in prices exceed the "neutral band" represented by transport costs. The prices are analyzed in pairs.	Daily corn and soybean prices in 4 North Carolina terminal markets. (in logs). 2 January 1992-4 March 1999.	Results confirm the presence of thresholds and indicate support for market integration; threshold models suggest faster adjustments to disequilibrium than when threshold are ignored. Thresholds turn out to be bigger when there is a bigger distance between markets.
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 to see whether changes in policies have increased or not spatial efficiency.	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.

METHODOLOGY	DATA USED	MAJOR FINDINGS
SWITCHING REGIME-THRESHOLD MODELS		
Septhon, 2003		
They expand Goodwin and Piggott (2001) studies in three directions: they use the direct multivariate test of Hansen and Seo instead of a single equation search to look for confirmatory evidence; they use Hansen tests to determine the number of thresholds; they test for the presence of non-linear cointegration.	Same as Goodwin and Piggott, (2001).	Their results indicate the presence of one threshold in most of the bivariate commodity pairings, and are supported by more general tests of nonlinear cointegration.
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 and 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 market competitiveness.
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); nonparametric analysis of price parity.	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 B – Unit root tests

Table B.1 – Unit root tests, ADF statistics -18 (1978:12-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**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)
c+t	*12(11)*	-3.256 (0.074)	-3.606 (0.029)
c+s	*12(11)*	-1.737 (0.412)	-3.761 (0.003)
c+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)
c+t	*1(0)*	-3.325 (0.062)	-12.511 (0)
c+s	*1(0)*	-3.107 (0.026)	-12.477 (0)
c+t+s	*1(0)*	-3.183 (0.088)	-12.461 (0)

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

Table B.2 – Unit root tests, ADF statistics -18 (1978:12-1993:06)

Variable	**n° of lags** *logs (diff. logs)*	**Logs**	**Differentiated logs**
swfr			
n	*12(11)*	0.668 (0.860)	-3.679 (0.0001)
c	*17(11)*	-2.720 (0.070)	-3.737 (0.003)
c+t	*17(11)*	-2.882 (0.168)	-3.740 (0.019)
c+s	*12(11)*	-2.502 (0.115)	-3.025 (0.032)
c+t+s	*12(11)*	-2.566 (0.296)	-2.988 (0.135)
hrw			
n	*1(0)*	-0.072 (0.659)	-9.223 (0)
c	*11(0)*	-2.328 (0.163)	-9.196 (0)
c+t	*11(0)*	-2.984 (0.136)	-9.25 (0)
c+s	*4(0)*	-2.319 (0.166)	-8.957 (0)
c+t+s	*4(0)*	-3.061 (0.116)	-9.027 (0)

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

Table B.3 – Unit root tests, ADF statistics -18 (1993:07-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**Logs**	**Differentiated logs**
swfr			
n	*12(11)*	0.546 (0.8343)	-1.926 (0.0516)
c	*12(11)*	-1.807 (0.377)	-1.808 (0.3772)
c+t	*12(11)*	-2.245 (0.464)	-1.798 (0.706)
c+s	*12(0)*	-1.574 (0.496)	-8.327 (0)
c+t+s	*12(0)*	-2.113 (0.537)	-8.308 (0)
hrw			
n	*2(1)*	0.272 (0.765)	-7.950 (0)
c	*1(1)*	-3.081 (0.028)	-7.928 (0)
c+t	*1(1)*	-3.102 (0.106)	-7.9023 (0)
c+s	*1(1)*	-2.843 (0.052)	-7.269 (0)
c+t+s	*1(1)*	-2.853 (0.178)	-7.246 (0)

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

Table B.4 – Unit root tests, PP statistics (1978:12-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**5% critical values**	**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
c+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
c+t		-21.339	-15.387	-209.646
		-3.428	-2.871	-12.550

n= no constant; c= constant; c+ t=constant and trend

Table B.5 – Unit root tests, PP statistics (1978:12-1993:06)

Variable	**n° of lags** *logs (diff. logs)*	**5% critical values**	**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
c+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
c+t		-20.996	-8.467	-114.695
		-3.440	-2.338	-9.218

n= no constant; c= constant; c+ t=constant and trend

Table B.6 – Unit root tests, PP statistics (1993:07-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**5% critical values**	**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
c+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
c+t		-20.800	-13.508	-91.024
		-3.447	-2.646	-8.207

n= no constant; c= constant; c+ t=constant and trend

Table B.7 – Unit root tests, KPSS statistics (1978:12-2003:12)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	1.003	0.129
hrw	*24(23)*			
c		*0.463*	0.144	0.060

c= constant

Table B.8 – Unit root tests, KPSS statistics (1978:12-1993:06)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	0.132	0.184
hrw	*24(23)*			
c		*0.463*	0.289	0.145

c= constant

Table B.9 – Unit root tests, KPSS statistics (1993:07-2003:12)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	0.408	0.351
hrw	*24(23)*			
c		*0.463*	0.105	0.102

c= constant

Annex C – Cointegration tests

Table C.1 – Cointegration tests, Johansen tests for *swfr* and *hrw* (1978:12-2003:12)

$H_0(H_1)$	16 lags (nd)		16 lags (d)		16 lags (d6, d7, d8)	
	trace	λmax	trace	λmax	trace	λmax
r = 0 (r = 2)	18.457(0.086)♦	14.939(0.069)*	17.694(0.109)	14.159(0.092)	17.245(0.125)	13.870(0.102)
r = 1 (r = 2)	3.518(0.500)	3.518(0.500)	3.535(0.497)	3.535(0.497)	3.376(0.524)	3.376(0.524)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

♦ In the hypothesis of a 1 cointegration rank, the corresponding VECM was estimated, but the value of the transmission elasticity parameters were implausible. When 13 lags were used, in analogy with *wref*, the cointegration rank was always 0

Table C.2 – Cointegration tests, Johansen tests for *swfr* and *hrw* (1978:12-1993:06)

$H_0(H_1)$	13 lags (nd)		5 lags (d)		5 lags (d6, d7, d8)	
	trace	λmax	trace	λmax	trace	λmax
r = 0 (r = 2)	29.112(0.002)	25.585(0.0007)	24.885(0.009)	19.065(0.013)	25.323(0.008)	18.459(0.017)
r = 1 (r = 2)	3.527(0.498)	3.527(0.498)	5.820(0.212)	5.820(0.212)	6.863(0.137)	6.863(0.137)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

Table C.3 – Cointegration tests, Johansen tests for *swfr* and *hrw* (1993:07-2003:12)

$H_0(H_1)$	2 lags (nd)		23 lags (d)		2 lags (d6, d7, d8)	
	trace	λmax	trace	λmax	trace	λmax
r = 0 (r = 2)	18.853(0.0764)	12.100(0.186)	30.663(0.0009)	28.726(0.0002)	18.279(0.091)	11.352(0.235)
r = 1 (r = 2)	6.753(0.144)	6.753(0.144)	2.837(0.619)	2.8373(0.6176)	6.928(0.134)	6.928(0.134)
$H_0(H_1)$	13 lags (nd)		13 lags (d)		13 lags (d6, d7, d8)	
	trace	λmax	trace	λmax	trace	λmax
r = 0 (r = 2)	17.232(0.125)	13.467(0.118)	18.707(0.080)	15.647(0.053)	17.781(0.106)	14.588(0.079)
r = 1 (r = 2)	3.765(0.460)	3.765(0.459)	3.060(0.579)	3.060(0.577)	3.193(0.555)	3.193(0.554)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

Table C.4 – Cointegration tests, Johansen tests for *swfr* and *wref* (1978:12-2003:12)

H0(H1)	13 lags (nd)		13 lags (d)		13 lags (d6, d7, d8)	
	trace	λmax	trace	λmax	trace	λmax
r = 0 (r = 2)	32.852(0.0004)	29.137(0.0001)	28.183(0.003)	24.798(0.001)	27.621(0.003)	24.237(0.0013)
r = 1 (r = 2)	3.751(0.468)	3.751(0.468)	3.385(0.522)	3.385(0.522)	3.385(0.522)	3.385(0.522)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

Table C.5 – Cointegration tests, Johansen tests for *swfr* and *wref* (1978:12-1993:06)

$H_0(H_1)$	13 lags (nd)		1 lag (d)		1 lag (d6, d7, d8)	
	trace	**λmax**	**trace**	**λmax**	**trace**	**λmax**
r = 0 (r = 2)	31.470(0.0007)	26.264(0.0005)	37.889(0)	33.123(0)	45.083(0)	37.962(0)
r = 1 (r = 2)	5.206(0.271)	5.206(0.271)	4.766(0.321)	4.766(0.321)	7.1213(0.123)	7.1213(0.123)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

Table C.6 – Cointegration tests, Johansen tests for *swfr* and *wref* (1993:07-2003:12)

$H_0(H_1)$	2 lags (nd)		2 lags (d)		2 lags (d6, d7, d8)	
	trace	**λmax**	**trace**	**λmax**	**trace**	**λmax**
r = 0 (r = 2)	20.435(0.046)	13.769(0.1058)	19.993(0.059)	12.490(0.164)	19.548(0.061)	12.869(0.153)
r = 1 (r = 2)	6.665(0.150)	6.6655(0.150)	7.503(0.104)	7.503(0.104)	6.585(0.138)	6.585(0.138)
$H_0(H_1)$	**13 lags (nd)**		**13 lags (d)**		**13 lags (d6, d7, d8)**	
	trace	**λmax**	**trace**	**λmax**	**trace**	**λmax**
r = 0 (r = 2)	19.604(0.060)	15.642(0.053)	20.234(0.049)	17.200(0.029)	19.619(0.060)	16.469(0.038)
r = 1 (r = 2)	3.962(0.430)	3.962(0.430)	3.034(0.583)	3.034(0.583)	3.149(0.563)	3.149(0.563)

nd = no monthly dummies considered; d = eleven monthly dummies were included; d6,d7,d8 = only dummies for June, July and August were considered; trace = Johansen trace test; λmax = Johansen λmax test; p-values are reported in parenthesis

Annex D – Unit root tests on the price spread

Table D.1 – Unit root tests on the price spread, ADF statistics (1978:12-2003:12)

Spread	**n° of lags**	**ADF statistics**
n	*2*	-2.464 (0.132)
c	*2*	-2.495 (0.117)
c+s	*2*	-2.416 (0.137)

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

Table D.2 – Unit root tests on the price spread, ADF statistics (1978:12-1993:06)

Spread	**n° of lags**	**ADF statistics**
n	*1*	-1.615 (0.100)
c	*1*	-1.829 (0.366)
c+s	*1*	-1.654 (0.455)

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

Table D.3 – Unit root tests on the price spread, ADF statistics (1993:07-2003:12)

Spread	**n° of lags**	**ADF statistics**
n	*2*	-1.934 (0.051)
c	*2*	-3.658 (0.005)
c+s	*2*	-3.449 (0.009)

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

Table D.4 – Unit root tests on the price spread, PP statistics (1978:12-2003:12)

Spread	**n° of lags**	**5%critical values**	**PP statistics**
n	2	-8.000 -1.950	-11.519 -2.455
c	2	-14.000 -2.878	-11.704 -2.482

n= no constant; c= constant

Table D.5 – Unit root tests on the price spread, PP statistics (1978:12-1993:06)

Spread	**n° of lags**	**5%critical values**	**PP statistics**
n	*1*	-7.949 -1.950	-4.671 -1.403
c	*1*	-13.848 -2.885	-5.916 -1.625

n= no constant; c= constant

Table D.6 – Unit root tests on the price spread, PP statistics (1993:07-2003:12)

Spread	**n° of lags**	**5%critical values**	**PP statistics**
n	*2*	-7.917 -1.950	-3.280 -1.388
c	*2*	-13.750 -2.888	-9.194 -2.143

n= no constant; c= constant

Table D.7 – Unit root tests on the price spread, KPSS statistics (1978:12-2003:12)

Spread	**n°of lags**	**5% critical value**	**KPSS statistics**
c	*24*	*0.463*	0.455

c= constant

Table D.8 – Unit root tests on the price spread, KPSS statistics (1978:12-1993:06)

Spread	**n°of lags**	**5% critical value**	**KPSS statistics**
c	*24*	*0.463*	0.346

c= constant

Table D.9 – Unit root tests on the price spread, KPSS statistics (1993:07-2003:12)

Spread	**n°of lags**	**5% critical value**	**KPSS statistics**
c	*24*	*0.463*	0.373

c= constant

Annex E – Unit root tests on de-seasonalized time series

Table E.1 – Unit root tests on de-seasonalized time series, ADF statistics -18 (1978:12-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**Logs**	**Differentiated logs**
swfr			
n	*12(11)*	-0.179 (0.622)	-4.956 (0)
c	*12(11)*	-1.172 (0.696)	-4.929 (0)
c+t	*12(11)*	-2.552 (0.302)	-4.795 (0)
hrw			
n	*1(0)*	0.182 (0.739)	-12.729 (0)
c	*1(0)*	-3.164 (0.022)	-12.713 (0)
c+t	*1(0)*	-3.241 (0.076)	-12.699 (0)

n= no constant; c= constant; c+ t=constant and trend; p-values are based on MacKinnon (1996) and reported in parenthesis

Table E.2 – Unit root tests on de-seasonalized time series, ADF statistics -18 (1978:12-1993:06)

Variable	**n° of lags** *logs (diff. logs)*	**Logs**	**Differentiated logs**
swfr			
n	*12(11)*	0.072 (0.706)	-3.486 (0.0004)
c	*12(11)*	-2.118 (0.238)	-3.429 (0.010)
c+t	*12(11)*	-2.554 (0.32)	-3.843 (0.014)
hrw			
n	*1(0)*	-0.018 (0.677)	-9.451 (0)
c	*4(0)*	-2.390 (0.145)	-9.423 (0)
c+t	*4(0)*	-3.132 (0.238)	-9.487 (0)

n= no constant; c= constant; c+ t=constant and trend; p-values are based on MacKinnon (1996) and reported in parenthesis

Table E.3 – Unit root tests on de-seasonalized time series, ADF statistics -18 (1993:07-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**Logs**	**Differentiated logs**
swfr			
n	*12(11)*	-0.409 (0.536)	-3.502 (0.0004)
c	*12(11)*	-1.925 (0.321)	-3.410 (0.011)
c+t	*12(11)*	-0.702 (0.972)	-3.823 (0.015)
hrw			
n	*2(1)*	0.314 (0.776)	-8.037 (0)
c	*1(1)*	-3.138 (0.024)	-8.019 (0)
c+t	*1(1)*	-3.119 (0.102)	-7.999 (0)

n= no constant; c= constant; c+ t=constant and trend; p-values are based on MacKinnon (1996) and reported in parenthesis

Table E.4 – Unit root tests on de-seasonalized time series, PP statistics (1978:12-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**5%critical values**	**Logs**	**Differentiated logs**
swfr	*13(12)*			
n		-8.000 -1.950	0.001 0.006	-510.862 -19.107
c		-14.000 -2.878	-4.044 -1.424	-511.001 -19.090
c+t		-21.339 -3.428	-12.956 -2.682	-509.967 .-19.060
hrw	*2(1)*			
n		-8.000 -1.950	0.050 0.242	-212.827 -12.757
c		-14.000 -2.878	-14.323 -2.777	-212.971 -12.742
c+t		-21.339 -3.428	-14.794 -2.843	-213.103 -12.727

n= no constant; c= constant; c+ t=constant and trend

Table E.5 – Unit root tests on de-seasonalized time series, PP statistics (1978:12-1993:06)

Variable	**n° of lags** *logs (diff. logs)*	**5%critical values**	**Logs**	**Differentiated logs**
swfr	*2(1)*			
n		-7.949	0.020	-238.904
		-1.950	0.637	-19.058
c		-13.848	-7.048	-239.236
		-2.885	-2.541	-19.049
c+t		-20.996	-6.521	-242.493
		-3.440	-2.416	-19.498
hrw	*2(1)*			
n		-7.949	0.002	-116.846
		-1.950	0.012	-9.414
c		-13.848	-5.694	-116.862
		-2.885	-1.737	-9.386
c+t		-20.996	-8.240	-118.347
		-3.440	-2.357	-9.454

n= no constant; c= constant; c+ t=constant and trend

Table E.6 – Unit root tests on de-seasonalized time series, PP statistics (1993:07-2003:12)

Variable	**n° of lags** *logs (diff. logs)*	**5%critical values**	**Logs**	**Differentiated logs**
swfr	*3(2)*			
n		-7.917	0.006	-69.704
		-1.950	0.111	-6.928
c		-13.750	-3.165	-69.703
		-2.888	-1.167	-6.900
c+t		-20.800	2.095	-74.979
		-3.447	-0.529	-7.253
hrw	*2(1)*			
n		-7.917	0.033	-91.308
		-1.950	0.235	-8.314
c		-13.750	-12.482	-91.386
		-2.888	-2.546	-8.285
c+t		-20.800	-12.782	-91.389
		-3.447	-2.546	-8.252

n= no constant; c= constant; c+ t=constant and trend

Table E.7 – Unit root tests on de-seasonalized time series, KPSS statistics (1978:12-2003:12)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	1.007	0.335
hrw	*24(23)*			
c		*0.463*	0.144	0.098

c= constant

Table E.8 – Unit root tests on de-seasonalized time series, KPSS statistics (1978:12-1993:06)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	0.138	0.289
hrw	*24(23)*			
c		*0.463*	0.288	0.147

c= constant

Table E.9 – Unit root tests on de-seasonalized time series, KPSS statistics (1993:07-2003:12)

Variable	**n°of lags** *logs (diff. logs)*	**5% critical value**	**Logs**	**Differentiated logs**
swfr	*24(23)*			
c		*0.463*	0.454	0.335
hrw	*24(23)*			
c		*0.463*	0.105	0.098

c= constant

The notion of “price transmission” refers to the co-movement shown by prices of the same good in different locations. In this respect, many empirical works essentially aim at verifying whether the Law of One Price holds. Cointegration models especially have shown to have an “intuitive appeal”, as they allow disentangling short and long run dynamics. However, despite the use of increasingly sophisticated techniques, empirical findings are extremely mixed. Indeed, a number of factors are expected to prevent prices from convergence.

This work focuses on how international price transmission is affected by domestic and border policies. The case study is constituted by price transmission between EU and US soft wheat markets in the years 1978-2003.

The objective of this analysis is twofold. On the one side, to develop a consistent theoretical framework to understand under which conditions co-movement between EU and US prices is expected, after considering the presence of domestic and border policies: namely, the European Common Agricultural Policy and the implementation of the Uruguay Round Agreement on Agriculture of the WTO. On the other, to derive different econometric models in order to test whether such a co-movement exists, and to which extent it has been affected by policy reforms.

The work is organized as follows. After a general introduction to the study of price transmission and to empirical tests for spatial price analysis (chapters 2 and 3), international soft wheat markets are described (chapter 4). This allows to develop a consistent theoretical framework for the analysis, according to which different empirical models are developed and tested (chapters 5 and 6). They represent an original attempt of explicitly combining policy and price data. Findings contribute to shed light on a complex and controversial empirical evidence and, in addition, to provide analytical tools which are of a more general interest for the study of international price dynamics.

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