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Summary

In this paper we review the existing empirical literature on price asymmetries in commodities, providing a way to classify and compare different studies which are highly heterogeneous in terms of econometric models, type of asymmetries and empirical findings. Relative to the previous literature, this paper is novel in several respects. First, it presents a detailed and updated survey of the existing empirical contributions on the existence of price asymmetries in the transmission mechanism linking input prices to output prices. Second, this paper presents an extension of the traditional distinction between long-run and short-run asymmetries to new categories of asymmetries, such as: contemporaneous impact, distributed lag effect, cumulated impact, reaction time, equilibrium and momentum equilibrium adjustment path, regime effect, regime equilibrium adjustment path. Third, each empirical study is critically discussed in the light of this new classification of asymmetries. Fourth, this paper evaluates the relative merits of the most popular econometric models for price asymmetries, namely autoregressive distributed lags, partial adjustments, error correction models, regime switching and vector autoregressive models.

Keywords: Price asymmetries, Cointegration, Partial adjustment, Threshold regime switching

JEL Classification: C22, D40, Q40

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1 Introduction

Consumers are generally very concerned when retailers decide to increase the price of their products as a consequence of increases of wholesales prices but not to reduce the price as a consequence of a fall in wholesale prices. This sharp attention to product price variations is particularly addressed to those goods which significantly contribute to the consumers' daily expenditure.

One of the products whose price variations consumers are particularly sensible to is gasoline. Given the importance of individual mobility in modern societies, it is quite natural to think that a reduction in the price of gasoline makes consumers very happy just as a price rise makes them very upset. What is less obvious is whether this last statement could be considered an appropriate description of the real consumers' sentiment. Wouldn't be better to say that consumers are very happy if the price of gasoline decreases, while they are very, very upset in the price of gasoline increases?

As illustrated, among others, by Brown and Yucel (2000) for the US gasoline market, many consumers complain for the existence of price asymmetries, which they interpret as evidence of monopolistic behavior in the markets for oil and petroleum products.

The perception of price asymmetries in the mechanism of transmission which links input prices to output prices is not confined to the gasoline market, but it is typically extended to many agricultural products (e.g. vegetables, meat, dairy products, etc.) and financial markets (e.g. interest rates, bank deposits, etc.). In any case, the crucial question is whether output prices respond symmetrically to variations of input prices, or if prices behave as the consumers' common sense seems to suggest.

Economic theory offers a limited number of justifications for price asymmetries. For example, it is well known that a necessary condition for the presence of asymmetric price effects is price rigidity, which gives rise to sluggish adjustment of output prices to shocks in cost conditions. However, according to Peltzman (2000), "economic theory suggests no pervasive tendency for prices to respond faster to one kind of cost change than to another" (p. 467). Yet, there are a number of theoretical arguments that can be put forth on the issue of asymmetric adjustment. To begin with, profit maximizing behavior forces firms in competitive markets to adjust their prices to new cost conditions immediately, and presumably symmetrically. This hold when frictions and imperfections are absent. "Menu" costs, however, preclude instantaneous price adjustment even if firms have no market power. Similarly, accountancy rules and inventory valuation may be responsible for the sluggish adjustment of final prices with respect to increases or decreases in the value of major exogenous variables. For instance, when a historical criterion (first in first out or FIFO) is adopted to value inventories, the firm does not adjust its output immediately when costs change, but awaits until the stocks of inputs bought at the old price are depleted. When instead a replacement cost criterion (last in first out or LIFO) is applied, the firm adjusts its price very rapidly in response to changes in input costs. The accounting convention chosen by a firm can therefore have an influence on the speed of adjustment: application of a FIFO criterion results in longer lags than in the case of a LIFO principle. Market power is probably the main concern to those who observe that output prices respond more quickly to input price increases than they do to decreases. The standard argument goes as follows. Retailers allegedly try to maintain their "normal" profit margin when price rise, but they try to capture the larger margins that result, at least temporarily, when wholesale prices fall. In both cases, the situation is not to last because, for example, consumer search costs are present. When costly search is completed, profits go down and prices tend to competitive levels. A version of this story that emphasizes tacit collusion in oligopolistic markets notes that, when wholesale prices rise, each firm is quick to increase its selling price in order to signal its competitors that it is adhering to the tacit agreement; when wholesale prices fall, it is slow in adjusting its price because it does not want to run the risk of sending a signal that it is cutting its margins and breaking away from the agreement. Finally, another argument is that adjusting production is costly. When a cost shock occurs, profit maximizing competitive firms absorb part of the shock by depleting inventories when input prices fall and storing output when they rise. This leads prices not to immediately adjust to cost shocks even in competitive markets, although it is perfectly consistent with firms enjoying market power. According to Borenstein, Cameron and Gilbert (1997), asymmetry is consistent with the above story if the net marginal convenience yield (the change in net distribution costs resulting from a change in inventory levels) is convex. In this case cost increases will be accommodated more quickly.

If compared with theoretical economics, the empirical literature on price transmission asymmetries is very large and mixed. Studies generally differ in terms of analyzed goods, dependent and explanatory variables (Table 1), countries under scrutiny (Table 2), time frequencies, time periods, specifications of the models employed, and even type of journal (Table 3). As a consequence, the empirical findings are not always unique, making it difficult to determine whether prices do behave in an asymmetrical way or consumers are wrong.

In this paper we review the existing works about price asymmetries in commodities, providing a way to classify and compare different studies which are highly heterogeneous in terms of econometric models, type of asymmetries and empirical findings. In particular, we answer three fundamental questions, which have been systematically ignored by the available literature:

1. If the empirical evidence in favor of (or against) price asymmetries is, as one could expect, not model-invariant, what are the most popular models used in the literature to investigate the input-output price transmission mechanism?

2. Does the term "‘price asymmetries”’ define a homogeneous concept, or should alternative types of asymmetries be identified and introduced?

3. Would it be possible to classify the models currently used in the empirical work on price asymmetries in terms of their ability to describe specific types of asymmetries ?

Relative to the previous literature, this paper is novel in several respects (see among others Geweke (2004), von Cramon-Taubadel and Meyer (2004)). First, it presents a detailed and updated survey of the existing empirical contributions on the existence of price asymmetries in the transmission mechanism linking input prices to output prices. Second, this paper presents an extension of the traditional distinction between long-run and short-run asymmetries to new categories of asymmetries, such as: contemporaneous impact, distributed lag effect, cumulated impact, reaction time, equilibrium and momentum equilibrium adjustment path, regime effect, regime equilibrium adjustment path (Table 4). Third, each empirical study is critically discussed in the light of this new classification of asymmetries. Fourth, this paper evaluates the relative merits of the most popular econometric models for price asymmetries, namely autoregressive distributed lags, partial adjustments, error correction models, regime switching and vector autoregressive models (Table 5).

The plan of the paper is as follows. Section 2 discusses the traditional as well as the new definitions of asymmetry. Section 3 presents the empirical work based on early econometric models, namely autoregressive distributed lag specifications. Section 4 surveys is dedicated to the equilibrium correction approaches, that is partial adjustment, error correction and threshold autoregressive models. In Section 5 the more recent econometric models are illustrated, such as regime switching and vector autoregressive models. Some open questions in the analysis of price transmission asymmetries are addressed in Section 6. Section 7 concludes.

2 Price asymmetries

Prior to answering if the relationship between the price of an input and the price of one or more outputs is symmetric or asymmetric, it is crucial to understand that the word "asymmetry" does not have a unique meaning and, consequently, to distinguish among different types of asymmetries.
A widely used classification is between short-run (SR) and long-run (LR) asymmetries, since, in general, a SR analysis is more indicated to compare the intensity of output price variations to positive or negative changes in input prices, whereas a LR perspective is needed if the empirical investigation concentrates on the computation of reaction times, length of fluctuations, as well as speeds of adjustment toward an equilibrium level.

Specific econometric models focus on different aspects of the relation between input and output prices, or, equivalently, on different types of asymmetries. In this paper we identify five major classes of econometric models, namely the autoregressive distributed lag (ARDL) model, the partial adjustment model (PAM), the error (or equilibrium) correction model (ECM), the regime switching model (RSM) and, finally, their multivariate extensions. Since the different concepts of asymmetry which are defined in this section apply to both univariate and multivariate models, for sake of simplicity we discuss asymmetries using single-equation specifications. In particular we define eight types of asymmetries (A)/symmetries (S), namely: contemporaneous impact (COIA/COIS); distributed lag effect (DLEA/DLES); cumulated impact (CUIA/CUIS); reaction time (RTA/RTS); equilibrium adjustment path (EAPA/EAPS) and momentum equilibrium adjustment path (MEAPA/MEAPS); regime effect (REA/RES) and regime equilibrium adjustment path (REAPA/ REAPS).

In a ARDL, a variable y_t, $t = 1,...,n$, depends on its own lags (autoregressive part, or AR) and on a vector of variables X, both contemporaneous and lagged (distributed lag part, or DL).
If with x we indicate a single explanatory variable, i.e an element of X, a typical ARDL can be specified as:

$$y_t = \sum_{h=1}^{r} \phi_h y_{t-h} + \sum_{i=0}^{s} \alpha_i x_{t-i} + \epsilon_t \tag{1}$$

where ϵ_t is a white noise.
Model 1 can be generalized to incorporate asymmetries by assuming that x has a different impact on y, according to whether its sign is positive (+) or negative (−):

$$y_t = \sum_{h=1}^{r} \phi_h y_{t-h} + \sum_{i=0}^{s} \alpha_i^+ x_{t-i}^+ + \sum_{j=0}^{q} \alpha_j^- x_{t-j}^- + \epsilon_t, \tag{2}$$

Clearly, the above specification supports various types of asymmetries, which can be classified in four main categories.

First, a test of the null hypothesis: $\alpha_0^+ = \alpha_0^-$ provides information about the contemporaneous impact of x^+ and x^- on y, which is defined to be asymmetric (COIA) or symmetric (COIS) according to whether the null is rejected or not.

Second, it is easy to check whether the impact of x^+ and x^- is the same at any lag by testing the null hypothesis: $\alpha_i^+ = \alpha_j^-$, $i = 1, ..., s$, $j = 1, ..., q$, which, if rejected (not rejected), will denote an asymmetry (symmetry) due to a distributed lag effect (DLEA/DLES).

It is worth noting that $s \neq q$ implies a DLEA, while the vice versa is clearly false. Another kind of asymmetry which is linked to DLEA is the mean lag asymmetry. A mean lag is generally defined as a weighted average of the lags on x, with weights being the coefficients of the model. Mean lag asymmetry occurs when the mean lags for positive and negative variations of x are different.

If the mean lags are symmetric, we can not reach any reliable conclusion on the presence of DLEA, as the distributed lag effect can be either symmetric or asymmetric. Conversely, when the mean lags are asymmetric, then we can support a DLEA.

COIA and the DLEA are SR asymmetries, since they are comparing the impacts on y of x^+ and x^- at a given instant in time; however, it is evident that specification (2) can also incorporate LR asymmetries. A third possibility is whether the cumulated effect of x^+ and x^- at lag $t-k$ is symmetric. This can be checked by testing the null hypothesis: $\sum_{i=k}^{s} \alpha_i^+ = \sum_{j=k}^{q} \alpha_j^-$, with $k \in [0, \min(s, q)]$; hearafter we will denote this asymmetry as CUIA. Note that testing this hypothesis for all $k \in [0, \min(s, q)]$ is equivalent to jointly test the two hypothesis $\alpha_i^+ = \alpha_j^-$, $\alpha_0^+ = \alpha_0^-$ described above. Moreover the joint existence of DLES and COIS is a sufficient but not necessary condition for CUIS but the coexistence of DLEA and COIA does not imply either CUIA or CUIS.

Finally, impulse response or cumulative adjustment functions are used to compute the number of periods needed by the dependent variable to (re-)adjust to an equilibrium level once an asymmetric shock to x^+ and/or x^- has occurred. Clearly this kind of asymmetry is related to the persistence of the effects of x^+ and x^-, but, unlike a test on $s = q$, it also accounts for the impact of the other variables in the model. From now on, this asymmetric (symmetric) response will be referred to as reaction time asymmetry (RTA) (symmetry (RTS)).

A PAM, instead, assumes that there exists a target level for y (say y^*) and relates the actual value of y

to its value at time $t-1$ and to the deviation of y_{t-1} from the actual target level y_t^*:

$$y_t = \beta y_{t-1} + (1-\phi)(y_t^* - y_{t-1}) + \epsilon_t \tag{3}$$

Since ϕ expresses the speed of convergence of y to y^*, if $\phi = 0$ the adjustment to the equilibrium level is instantaneous, while $\phi = 1$ implies an infinite adjustment process.

This model can be generalized to incorporate asymmetries by assuming that the adjustment depends on whether y_{t-1} is above or below the equilibrium level:

$$y_t = \beta y_{t-1} + \phi^+ \varphi(y_{t-1}^* - y_{t-1})^+ + \phi^- \psi(y_{t-1}^* - y_{t-1})^- + \epsilon_t \tag{4}$$

where φ e ψ are functions of the disequilibrium, which may be equal to the identity function. Henceforth we will denote this kind of asymmetry as an equilibrium adjustment path asymmetry (EAPA).

If the series considered are stationary, then ARDL and PAM can be consistently estimated using OLS; but if the time series are nonstationary then, as shown by Granger and Newbold (1974), standard linear regression analysis can lead to spurious results, that is, the relation between two variables is only apparently significant. Economic variables are often integrated of order one variables (I(1)), meaning that they can be made stationary by first differencing; for this reason, a solution for nonstationarity widely used in the past is to estimate a model in first differences.

However, modern econometric analysis proposes a different framework for modeling nonstationary data.

In their seminal paper, in fact, Engle and Granger (1987) point out that, given a pair of $I(1)$ processes, if there is a stationary linear combination of them, then the two processes move together in the long run and are said to be cointegrated.

In order to exploit the concept of cointegration, Engle and Granger develop an equilibrium correction representation (ECM), which, given two $I(1)$ variables y, x which are cointegrated, with cointegrating vector $(1-\theta)$, can be written as follows:

$$\Delta y_t = \alpha \Delta x_t + \lambda(y_{t-1} - \theta x_{t-1}) + \epsilon_t \tag{5}$$

Lagged variables and autoregressive effects can be added to this model, which is also able to incorporate asymmetries as proposed by Granger and Lee (1989):

$$\begin{aligned} \Delta y_t = &\sum_{h=1}^{r} \beta_h \Delta y_{t-h} + \sum_{i=0}^{s} \alpha_i^+ \Delta x_{t-i}^+ + \sum_{j=0}^{q} \alpha_j^- \Delta x_{t-j}^- \\ &+ \lambda^+ ECT^+ + \lambda^- ECT^- + \epsilon_t \end{aligned} \tag{6}$$

where $ECT = (y_{t-1} - \theta x_{t-1})$.

Model (6) considers all the asymmetries which are testable within the ARDL specification and supports a test of equilibrium adjustment path symmetry too. As a matter of fact, if $\lambda^+ \neq \lambda^-$, the convergence process is different depending on the direction of the deviation from the equilibrium level.

Model (6) can also be extended as suggested by Enders and Granger (1998):

$$\Delta y_t = \sum_{i=0}^{s} \alpha_i \Delta x_{t-i} + \gamma^+ ECT_{t-1} I_t + \gamma^- ECT_{t-1}(1 - I_t) \tag{7}$$

where

$$I_t = \begin{cases} 1 & \text{if } \Delta ECT_{t-1} \geq 0 \\ 0 & \text{if } \Delta ECT_{t-1} < 0 \end{cases} \tag{8}$$

In this case, asymmetries arise depending on whether the deviation from the equilibrium is increasing or decreasing, rather then on the level of the shift. These asymmetries are known as momentum equilibrium adjustment path asymmetries (MEAPA).

The models described so far are based on the idea that some or all the explanatory variables X may have a non-linear impact on y.

However, the analysis can be extended to incorporate the possibility that the relationship between y and X as a whole depends on the state of a variable v, which can be one of the explanatory variables. Generally, it is said that the level of v, with respect to a threshold δ, describes different states of the world or regimes, hence the name regime switching models (RSM).

Specifically we can define two different models, depending on the nature of the threshold variable, the deterministic RSM and the stochastic RSM. In the first case we know which regime prevails in each instant of time; in the second case, instead, the shift from one regime to another is random.

According to the value of the threshold variable we can consider a further classification of RSM based on whether or not v belongs to X.

If v does not belong to X, then a general RSM, with $p+1$ states of the world, can be written as follows:

$$\begin{cases} y_t = f(X) & \text{if } v < \delta_1 \\ y_t = f^{'}(X) & \text{if } \delta_1 \leq v \leq \delta_2 \\ \ldots\ldots & \\ y_t = f^{*}(X) & \text{if } \delta_{p-1} \leq v \leq \delta_p \\ y_t = f^{*'}(X) & \text{if } v > \delta_p \end{cases} \tag{9}$$

where within each regime, the relationship between y and X can assume either an ARDL, PAM or ECM form.

If we consider a general ECM with DL effects, then model (9) becomes:

$$\begin{cases} \Delta y_t = \phi_0 + \sum_{i=1}^{r} \beta_i \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{-} \Delta x_{t-i}^{-} + \lambda^{+} ECT_{t-1}^{+} + \lambda^{-} ECT_{t-1}^{-} + u_t \\ \qquad \text{if } v < \delta_1 \\ \ldots\ldots \\ \Delta y_t = \phi_0^{*'} + \sum_{i=1}^{r} \beta_i^{*'} \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{*'+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{*'-} \Delta x_{t-i}^{-} + \lambda^{+*'} ECT_{t-1}^{+} + \lambda^{-*'} ECT_{t-1}^{-} + u_t^{*'} \\ \qquad \text{if } v > \delta_p \end{cases} \tag{10}$$

In this model, the five types of asymmetry we have described are still meaningful and their presence in each regime can be tested.

Keeping on considering the general case of an ECM, when v does belong to X, instead, the threshold can be set on either x or the error correction term ECT, leading to:

$$\begin{cases} \Delta y_t = \phi_0 + \sum_{i=1}^{r} \beta_i \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{-} \Delta x_{t-i}^{-} + \lambda^{+} ECT_{t-1}^{+} + \lambda^{-} ECT_{t-1}^{-} + u_t \\ \qquad \text{if } \Delta x_t < \delta_1 \\ \ldots\ldots \\ \Delta y_t = \phi_0^{*'} + \sum_{i=1}^{r} \beta_i^{*'} \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{*'+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{*'-} \Delta x_{t-i}^{-} + \lambda^{+*'} ECT_{t-1}^{+} + \lambda^{-*'} ECT_{t-1}^{-} + u_t^{*'} \\ \qquad \text{if } \Delta x_t > \delta_p \end{cases} \tag{11}$$

or

$$
\begin{cases}
\Delta y_t = \phi_0 + \sum_{i=1}^{r} \beta_i \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{-} \Delta x_{t-i}^{-} + \lambda^{+} ECT_{t-1}^{+} + \lambda^{-} ECT_{t-1}^{-} + u_t \\
\hfill \text{if } ECT_{t-1} < \delta_1 \\
\ldots\ldots \\
\Delta y_t = \phi_0^{*'} + \sum_{i=1}^{r} \beta_i^{*'} \Delta y_{t-i} + \sum_{i=0}^{s} \alpha_i^{*'+} \Delta x_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{*'-} \Delta x_{t-i}^{-} + \lambda^{+*'} ECT_{t-1}^{+} + \lambda^{-*'} ECT_{t-1}^{-} + u_t^{*'} \\
\hfill \text{if } ECT_{t-1} > \delta_p
\end{cases}
\tag{12}
$$

In this case, the values of Δx and ECT not only directly affect Δy (as in the ARDL, PAM and ECM), but also they indirectly influence Δy through their effect on the other explanatory variables.
Models (11) and (12) allow us to define two new concepts of asymmetry. Specifically, we define the regime effect asymmetry (REA) as the situation where the existence of more than one regime defined by the variable x is significant; if it is not, we have a regime effect symmetry (RES). When the threshold variable is given by the ECT, instead, we define a regime equilibrium adjustment path asymmetry (REAPA) or symmetry (REAPS).

Finally, the Vector models we are presenting in this work are multivariate extensions of the uniequational specification presented above, thus supporting the same asymmetries. In particular, we are describing Vector Auto Regressive (VAR), Vector Error Correction (VEC) and Vector Regime Switching (VRS) Models.

In what follows, the existing literature is classified according to the model used and to the type of asymmetries studied. For each work, the reference market is specified, in order to account for the fact that, generally, irrespective of the good under scrutiny, input price variations are not transmitted from the first element of the distribution chain directly to the final consumer. On the contrary, input price changes are usually filtered by the intermediate nodes of the distribution system. For example, a variation in the price of oil does not directly affect the price of gasoline at the pump, but first hits the wholesale price of gasoline. A reference to the relevant sample period and the data frequency complete the presentation of each contribution.

3 Early asymmetric models

3.1 Autoregressive distributed lags models

The empirical literature on asymmetric price transmissions goes back to Farrel (1952). This is the first attempt to investigate empirically the irreversibility behavior of the demand function of some habitual consumption goods. Farrel analyzes the demand (Y) for tobacco, beer, wines and spirits in response to income (Z) and price (X) variations, using a functional form of the type:

$$\frac{Y_t}{Y_{t-1}} = e^c \left(\frac{Z_t'}{Z_{t-1}}\right)^{a'} \left(\frac{Z_t''}{Z_{t-1}}\right)^{a''} \left(\frac{X_t'}{X_{t-1}}\right)^{b'} \left(\frac{X_t''}{X_{t-1}}\right)^{b''} \tag{13}$$

where

$$\begin{cases} \frac{Z_t'}{Z_{t-1}} = \frac{Z_t}{Z_{t-1}} \text{ if } \frac{Z_t}{Z_{t-1}} \geq 1 \\ \frac{Z_t'}{Z_{t-1}} = 1 \text{ otherwise} \end{cases} \qquad \begin{cases} \frac{Z_t''}{Z_{t-1}} = \frac{Z_t}{Z_{t-1}} \text{ if } \frac{Z_t}{Z_{t-1}} < 1 \\ \frac{Z_t''}{Z_{t-1}} = 1 \text{ otherwise} \end{cases} \tag{14}$$

with a, b and c indicating unknow parameters.

The hypothesis to be tested is whether an income increase, or a price decrease, leads to an increase in consumption of habitual goods, while an income decrease, or price increase, leaves the levels of consumption virtually unaltered. The empirical findings seem to be inconclusive, but, in Farrel's words, "they do suggest that irreversibility may be quite an important factor in the change of tastes" (p. 186).

It is worth noting that the model proposed by Farrell applies to the levels of demand, income and price. If, instead, we consider its logarithmic transformation, equation (13) takes the following form:

$$\Delta y_t = c + a^+ \Delta z_t^+ + a^- \Delta z_t^- + b^+ \Delta x_t^+ + b^- \Delta x_t^- + \epsilon_t \tag{15}$$

which is the typical asymmetric specification used in the literature.

At the end of the Sixties and during the Seventies most of the studies on asymmetric price transmission concentrate, not surprisingly, on agricultural goods. Tweeten and Quance (1968) investigate the relationship between the level of output (y) and the ratio between input and output prices (x) in the agricultural sector, using an indicator variable to discriminate between positive and negative variations of x.

Two different models are estimated. The first is a two-equation system, one referring to the years of price increases, the other to the years of decreasing prices:

$$y_t = \alpha_0 + \alpha^+ \widetilde{x}_t^+ + \epsilon_t \tag{16}$$

$$y_t = \alpha_0 + \alpha^- \widetilde{x}_t^- + \epsilon_t \tag{17}$$

where $\widetilde{x}^+$ is equal to x, if its value has increased over the last year and zero otherwise (viceversa for $\widetilde{x}^-$). Asymmetry is present if the null hypothesis of $\alpha^+ = \alpha^-$ is rejected.
The second model, instead, combines the effects of increasing and decreasing prices in a single equation:

$$y_t = \alpha_0 + \alpha^+ \widetilde{x}_t^+ + \alpha^- \widetilde{x}_t^- + \epsilon_t \tag{18}$$

Using annual data for the period 1921-1966, the authors find some evidence of asymmetry in the empirical results produced by the first model only.

Wolffram (1971) shows that the approach followed by Tweeten and Quance to distinguish between periods of expansion and periods of reduction of the input/output price ratio can lead to biased estimates. As a solution, Wolffram suggests to redefine the variables $\widetilde{x}_t^-$ and $\widetilde{x}_t^-$ as:

$$\begin{aligned}
\widetilde{\widetilde{x}}_1^+ &= x_1 \\
\widetilde{\widetilde{x}}_2' &= \widetilde{\widetilde{x}}_1^+ + D(x_2 - x_1) \\
\widetilde{\widetilde{x}}_3' &= \widetilde{\widetilde{x}}_2^+ + D(x_3 - x_2) \\
&\ldots\ldots
\end{aligned}$$

$$\begin{aligned}
\widetilde{\widetilde{x}}_1^- &= x_1 \\
\widetilde{\widetilde{x}}_2^- &= \widetilde{\widetilde{x}}_1^- + (1 - D)(x_2 - x_1) \\
\widetilde{\widetilde{x}}_3^- &= \widetilde{\widetilde{x}}_2^- + (1 - D)(x_3 - x_2) \\
&\ldots\ldots
\end{aligned}$$

where D is a dummy variable which takes the value of 1 if the variation of the input/output price ratio is positive, while it is equal to 0 otherwise.
The main difference between the Wolfram's approach and the Tweeten and Quance's model is that the former explicitly considers the effect of cumulative variations in the variable x, while the latter takes into account the direct impact of period-to-period variations. Thus, it can be useful to divide the literature into two broad categories, depending on whether the explanatory variable is defined according to Tweeten and Quance or to Wolfram.

3.1.1 Cumulative price variations

In 1977 Houck proposes a work on inventories and prices of milk and beans in US, which improves the Wolffram's approach in the following direction:

$$y_t - y_0 = \alpha_0 t + \alpha^+ \sum_{i=0}^{t} \Delta x^+_{t-i} + \alpha^- \sum_{i=1}^{t} \Delta x^-_{t-i} + \epsilon_t \tag{19}$$

It is immediate to notice that the dependent variable in model (19) is no longer y_t, as in the previous studies, but the deviation of y_t from its starting value y_0. Moreover, the model directly considers the impact of positive and negative variations of x on y, cumulated from the first period ($i = t$) up to the current period ($i = 0$).

Hereafter we define $C\Delta x^+_{t-k} = \sum_{i=k}^{t} \Delta x^+_i$, $C\Delta x^-_{t-k} = \sum_{i=k}^{t} \Delta x^-_i$ with $k \in [0, t]$, the cumulative sums of price variations used in Wolfram's approach models as the independent variable.

When testing the null hypothesis $\alpha^+ = \alpha^-$ using annual data, Houck finds that, for the milk market only, the variation of the level of inventories over the sample period asymmetrically depends on the contemporaneous impact of cumulative price changes (COIA).

Young (1980) analyzes the specification of the consumer's demand function in response to price changes. In particular, he proposes three different kinds of asymmetric behavior.

The first model is Hook's (1977):

$$y_t - y_0 = \alpha_0 + \alpha^+ \sum_{i=1}^{t} \Delta x^+_i + \alpha^- \sum_{i=1}^{t} \Delta x^-_i + \epsilon_t \tag{20}$$

where y now denotes quantity, while x price. Noting that a generic variable x can be expressed as the sum of its initial value (at time $t = 0$) and all the positive and negative deviations observed up to current time t, that is

$$x_t = x_0 + \sum_{i=1}^{t} \Delta x^+_i + \sum_{i=1}^{t} \Delta x^-_i$$

Young rewrites Hook's model in the following form:

$$y_t - y_0 = \alpha^*_0 + \alpha^+ x_t + \alpha^* \sum_{i=1}^{t} \Delta x^-_i + \epsilon_t \tag{21}$$

where $\alpha^* = \alpha^+ - \alpha^-$. Model (21) clearly assumes that the demand may react asymmetrically to the total impact of cumulative price increases and decreases. However, Young suggests that the demand curve would rather show an asymmetric behavior when prices are "at unprecedently low or high levels" (p. 178). The definition of xm_i and xM_i as the minimum and the maximum price at time i yields the two alternative models:

$$y_t - y_0 = \alpha_0^* + \alpha^+ x_t + \alpha^* \sum_{i=1}^{t} \Delta x m_i + \epsilon_t \tag{22}$$

$$y_t - y_0 = \alpha_0^* + \alpha^+ x_t + \alpha^* \sum_{i=1}^{t} \Delta x M_i + \epsilon_t \tag{23}$$

Models (22) and (23) represent the first attempt to describe asymmetry in terms of the response of y to the deviation of x above or below a given threshold, and not to the sign of x itself. This approach has become very popular in the empirical literature.

All studies discussed up to this point analyze the asymmetric effects of prices on the demand levels of different goods. Ward (1982) is the first contribution which shifts the attention to the core topic of this survey, that is the transmission mechanism between prices (Tables 6a, 6b).

Specifically, Ward models the impact of wholesales prices (ws) on retail (rt) and FOB prices (fb), using monthly data of different types of fresh vegetables in the US market, and dividing the sample in l seasons of t observations each.

With respect to Houck (1977), Ward does not simply express the current value of rt or fb as a function of the positive and negative sum of the variation of ws over the observation period, but he rather allows the effect of those variables to persist over time:

$$rt_{lt} = \phi + \sum_{j=0}^{3} \alpha_{jlt}^+ C\Delta ws_{lt-j}^+ D_{jlt} + \sum_{j=0}^{3} \alpha_{jlt}^- C\Delta ws_{lt-j}^- D_{jlt} + \epsilon_{lt} \tag{24}$$

where D_{jlt} are dummy variables used to identify the existence of $C\Delta w_{lt-J}$, as the required lagged quantities are not available for the first three observations of each season l.

Hence, equation (24) is a generalization of Houck's model, since it defines a new kind of asymmetry, namely the distributed lag effect of the cumulative variations.

Since equation (24) is defined over a very large parameter space, Ward actually analyzes a simplified version of it, along Young's line:

$$rt_{lt} = \phi + \sum_{j=0}^{3} \alpha_{jl}^+ (ws_{lt-j} - ws_{l0}) D_{jl} + \sum_{j=0}^{3} (\alpha_{jl}^- - \alpha_{jl}^+) C\Delta ws_{lt-j}^- D_{jl} + \epsilon_{lt} \tag{25}$$

The actual value of rt is now expressed as a function of the deviation of ws from its initial value and of the cumulative impact of its negative variations; the effect of positive deviations is not made explicit.

Starting from equation (25), the parameter space can be easily reduced if the impact of each variables depends on the time considered:

$$\begin{aligned} \alpha_{jl}^+ &= \gamma_1 + \gamma_2 \xi_j \\ \alpha_{jl}^- - \alpha_{jl}^+ &= \gamma_3 + \gamma_4 \xi_j \end{aligned} \tag{26}$$

where ξ_j is a known coefficient, equal to $\sqrt[3]{j}$, which forces the impact of cumulative prices ws to decrease with the time lag.

Substituting model (26) into equation (25), we eventually obtain Ward's specification:

$$\begin{aligned} rt_{lt} = \phi + \gamma_1 \sum_{j=0}^{3} (ws_{lt-j} - ws_0) D_{jl} + \gamma_2 \sum_{j=0}^{3} (ws_{lt-j} - ws_0) \xi_j D_{jl} + \gamma_3 \sum_{j=0}^{3} C\Delta ws^{-}_{lt-j} D_{jl} \\ + \gamma_4 \sum_{j=0}^{3} c\Delta ws^{-}_{lt-j} \xi_j D_{jl} + \epsilon_{lt} \end{aligned} \tag{27}$$

Model (27) depends on four parameters only; the significance of γ_3 and γ_4 is supportive of asymmetry in the distributed lag effect of cumulative price variations.

It is important to point out that the biunivocal correspondence between the coefficients α^{-}_{jl}, α^{+}_{jl}, γ_1, γ_2, γ_3 and γ_4 implies $\alpha^{-}_{jl} = \alpha^{+}_{jl}$ if and only if $\gamma_3 + \gamma_4 \xi_j = 0$; DLEA occurs when, for at least one $j \in [0, 3]$, $\gamma_3 + \gamma_4 \xi_j \neq 0$.

If just one of the two coefficients γ_3 and γ_4 is different from zero, DL asymmetry is clearly implied; but if γ_3 and γ_4 are both significant and have opposite sign, there can be a value of ξ_j such that $\gamma_3 + \gamma_4 \xi_j = 0$. However, this result can occur for just one particular value of ξ_j and, consequently, the significance of γ_3 and γ_4 always implies DLEA. Note that relationships (26) also imply that COIA occurs when $\gamma_3 \neq 0$.

To summarize, Ward finds out the existence of asymmetry in the contemporaneous and distributed lag effects of cumulative wholesale prices variations on both FOB and retail prices. Results about DLEA are also supported by the calculated mean lags, which are asymmetric in most of the cases.

Finally, Ward tests the hypothesis of symmetry in the cumulative impact of $c\Delta ws^{+}$ and $c\Delta ws^{-}$, finding that CUI asymmetry affects only the wholesale-retail transmission mechanism.

Ward's approach has been extended by several authors. For instance, in 1987 Kinnucan and Forker analyze the farm(fm)-retail price transmission for major dairy products in US, using monthly data over the period January 1971-December 1981. In order to make explicit the distributed lag impact of cumulative price variations, they choose not to use Ward's final specification (27), but to directly estimate equation (24), where they also introduce the role of marketing costs (mk):

$$rt_t - rt_0 = \phi t + \sum_{i=1}^{s} \alpha^{+}_i C\Delta fm^{+}_i + \sum_{j=1}^{q} \alpha^{-}_j C\Delta f^{-}_j + mk_t - mk_0 + \epsilon_t \tag{28}$$

Although the authors estimate the coefficients of the cumulative variables $C\Delta fm^{+}_j$ and $C\Delta fm^{-}_j$, they do not test the hypothesis $\alpha^{+}_i = \alpha^{-}_j$, $\forall i, j$; instead, they investigate the presence of DLEA by simply computing the mean lags, which appear to be different depending on the sign of farm price variations. However, this result shouldn't be interpreted as a reliable evidence in favor of asymmetry, especially given that no statistical test is provided. As a consequence, any consideration about the presence of distributed lag asymmetry is inconclusive, as the number of significant lags is symmetric, which does not exclude either DLA or DLS. Kinnucan and Forker formally test only the null hypothesis: $\sum_{i=1}^{s} \alpha_i^{+} = \sum_{j=1}^{q} \alpha_j^{-}$, which is rejected for all the considered goods in favor of CUI asymmetry.

The specifications proposed by Houck (1977), Ward (1982) and Kinnucan and Forker (1987) constitute a complete overview of the types of models proposed in the literature to evaluate the impact of cumulative variations of a price x on a price y. As a result, the empirical work based on Wolfram's approach can be related to one of these three models.

Griffith and Piggott (1994) use Kinnucan and Forker's specification to analyze the relationships between retail-wholesale prices, farm-wholesale prices and farm-retail prices for the Australian beef, lamb and pork markets, using monthly data since January 1971 to December 1988.
As Kinnucan and Forker, the authors test the null hypothesis: $\sum_{i=1}^{s} \alpha_i^{+} = \sum_{j=1}^{q} \alpha_j^{-}$, which is never rejected for the pork market, whereas it is not rejected only for the farm-wholesale price transmission and for the farm-retail price relationship in the beef and lamb markets, respectively. In the light of these findings, Griffith and Piggott suggest that "asymmetrical price response is a strategy used by beef and lamb retailers and wholesalers to adjust to changing input prices but not by pork retailers and wholesalers" (p. 307).
However, a closer analysis of these results leads to a different conclusion, namely that CUIA is not really relevant in the Australian markets for beef, lamb and pork, while DLEA is. As a matter of fact, except for the farm-wholesale price relationship in the lamb market, all models exhibit an asymmetric number of lags for the positive and the negative cumulative price variations, which imply an asymmetric behavior of their distributed lag effects.
Furthermore, the asymmetric lag structure leads to conclude that all three price relationships in the beef market, the farm-retail transmission in the markets for beef and pork, as well as the wholesale-retail price relation for the pork market are characterized by COIA.
Actually, in the first four cases only either the positive or the negative contemporaneous impact of upstream prices results to be statistically significant. In the latter case, the presence of CUIA, together with the structure of the significant lags, shows that the cumulative impact of the cumulative positive variations of farm prices up to time $t-2$ is equivalent to the contemporaneous impact of the negative variations, which consequently supports the presence of COIA.

Kinnucan and Forker's (1987) approach is also employed by Powers (1995), who studies the impact of

FOB prices on retail and wholesale prices, as well as the effect of wholesale prices on retail prices, together with the impact of hauling costs on both wholesale and retail prices, in twelve US cities. The analysis focuses on iceberg lettuce and relies on weekly data from March 9, 1986 to August 30, 1992.

The effect of hauling costs is tested for the presence of CUIS, which, as a whole, cannot be rejected in both markets.

As far as the price transmission is concerned, the author provides detailed information about three types of asymmetries.

First of all, he analyzes the cumulative impact of the cumulative price variations and finds out that CUIA generally characterizes the wholesale-retail and the FOB-retail transmission mechanisms, but not the FOB-wholesale price relation. Secondly, Powers looks for asymmetries in the reaction time by evaluating the median lags, which support RTA only in the wholesale-retail market. Finally, the author shows that the lag structure of the relationships retail-wholesale and retail-FOB is often asymmetric across the twelve cities considered, suggesting the existence of DLEA in both markets.

Zhang et al. (1995) apply Kinucan and Forker's model to the wholesale prices of peanuts and the price of peanuts butter in US, using monthly data over the period January 1984- July 1992. In such a market the hypothesis of CUI symmetry of the cumulative wholesale price variations cannot be rejected. As for the DL impact, the mean lags computed by the author look symmetric and do not provide any valid information about the presence or absence of price asymmetries, as shown in Section 2. However, the number of significant lags in the model differs between phases of rising and falling prices, which clearly implies DLEA.

Worth (2000) studies the relationship between FOB shipping point prices and retail prices of some fresh vegetables. In particular, he applies Kinnucan and Forker's model to monthly data from January 1980 to May 1999. It is notable that Worth chooses not to use "earlier data, though available, because of changes in agricultural markets since 1960" (p. 6), that might bias the analysis.

In this work the author tests only the CUIS hypothesis, which leads to conclude that "carrots and tomatoes are the only commodities which show evidence of price asymmetry" (p.8). In the light of our taxonomy of asymmetries, however, Word's results draw a different picture. CUIA, in fact, characterizes a subset of vegetables only (namely, carrots and tomatoes), while all goods under scrutiny are affected by DLEA, since, as the author points out, the number of significant lags for price increases and decreases is always asymmetric.

Parrott et al. (2001) analyze the transmission mechanism between retail and FOB shipping point prices, weighted by the volume of shipments, in the US fresh tomatoes market, using weekly data over the period June 1988-December 1993. By estimating the Kinnucan and Forker's model, the authors evaluate the presence of CUI, COI and DLE symmetry, all of which cannot be rejected.

The US fresh tomato market is also studied by Girapunthong et al. (2004), who, unlike Parrott et al., use Ward's specification. In particular, they focus on the producer-retail, producer-wholesale and wholesale-retail price relationships between May 1975 and February 1998, using monthly data.
The authors test for CUI symmetry and for both COI and DLE symmetry, which as we know, are biunivocally linked to the significance of γ_3 and γ_4 in equation (27).
Empirical findings suggest that no asymmetry occurs in the producer-retail market, while all types of asymmetries characterize the producer-wholesale price transmission. Finally, the different wholesale-retail price relationships exhibit both COIA and DLEA.
Note that the estimated mean lags confirm the existence of DLEA in the producer-retail market but appear to be symmetric at both the producer-wholesale and wholesale-retail stages. This result, however, is not surprising, since mean lag symmetry does not imply either DLEA or DLES.

Within the class of Wolframm's types of models, Mohanty's et al. (1995) forms a separate category, as it deals with spatial asymmetry. Up to this point we have analyzed the price transmission between the different levels of a market chain, but it is also possible to study the relationship between the same price in different countries.
The authors investigate the relationship between US FOB wheat prices and the correspondent prices in Canada, Australia, Argentina and European Union and, for each pair of countries, estimate a Kinnucan and Forker's (1987) type of model.
In particular, the authors, who consider monthly data from January 1980 to June 1990, propose two asymmetry tests. The first is directed to test CUIS and the other aims at testing the joint null hypothesis of COI and DLE symmetries, which, if refused, denotes the presence of at least one of them.
Results show that both hypothesis can be rejected, that is CUI and at least either DLE or COI are asymmetric.

3.1.2 Period-to-period price variations

Unlike Wolffram's approach, which evaluates the effect of cumulative price variations, Tweeten and Quance (1968) study the impact of period-to-period price increases and decreases. For this reason, the results on DL, CUI and COI asymmetries obtained by the papers which follow this approach cannot be directly compared with those described in Section 3.1.1.

The most effective way to introduce the Tweteen and Quance's (1968) model is to present the generalization proposed by Balke et at. (1998), which includes simple distributed lag effects.
In their work the authors study the transmission mechanism between prices at different levels of the gasoline distribution chain using a specification which, for the relationship spot(sp)-wholesale prices, can be written as:

$$ws_t = \phi + \sum_{i=0}^{s} \alpha_i sp_{t-i} + \sum_{i=0}^{s} \alpha_i^+ \widetilde{sp}_{t-i} + \sum_{j=1}^{r} \beta_i ws_{t-j} + \sum_{j=1}^{r} \beta_i^+ \widetilde{ws}_{t-j}^+ + \epsilon_t \tag{29}$$

where, as in Tweeten and Quance's model, $\widetilde{sp}$ is equal to sp if the value of sp has increased over the last period and it is zero otherwise (identical considerations apply to $\widetilde{ws}$).
In the context of equation (29) it is possible to test for symmetry by checking the significance of specific coefficients. For instance, the impact of negative spot price variations is described by the α_i coefficients, while the effect of positive spot price changes is measured by the sum $\alpha_i + \alpha_i^+$.
If the α_i^+ coefficients are statistically equal to zero, then the impact of spot price increases and decreases is the same (say, α); otherwise, the effect of spot prices depends on their sign through α_i^+.
The authors use US weekly data from January 1987 to August 1996 and consider all possible upstream-downstream price transmission mechanisms among crude oil and gasoline prices (namely, spot, wholesale and retail), with and without taxes. In particular, they test the joint null hypothesis: $\alpha_1^+ = 0, ..., \alpha_s^+ = 0$, which, if rejected, denotes the existence of, at least, either COI or DLE asymmetries.
Results show that crude prices have an asymmetric impact on retail prices (with and without taxes), spot prices asymmetrically affect wholesale and retail prices with taxes and, finally, asymmetry occurs also in the response of ex-taxes retail prices to wholesale costs.

Karrenbrock (1991) proposes a specification to test for symmetry in the impact of period-to-period wholesale price variations on retail prices:

$$\Delta rt_t = \phi + \sum_{i=1}^{s} \alpha_i^+ \Delta ws_{t-i}^+ + \sum_{j=1}^{q} \alpha_i^- \Delta ws_{t-j}^- + \epsilon_t \tag{30}$$

Using US monthly data over the period January 1983-December 1990, Karrenbrock tests the presence of CUI, COI and DL asymmetries in the market of premium gasoline, as well as regular leaded and unleaded gasoline.
The null hypothesis $\sum_{i=1}^{s} \alpha_i^+ = \sum_{j=1}^{q} \alpha_i^-$ is never rejected, that is the cumulative effect of wholesale price variation is symmetric. On the contrary, COI asymmetry characterizes the response of the premium gasoline price.

As for the distributed lag impact, Karrenbrock provides both a test of persistence and a direct test of equality among the coefficient at each time lag. In general Karrenbrock shows that, although the effects of wholesale price increases and decreases exhibit the same degree of persistence[1], their distributed lag effect is clearly asymmetric.

The distributed lag structure of the impact of a price on another price may also be estimated by a second model, described in a report of the General Accounting Office (GAO) (1993).

GAO investigates the crude-wholesale and the wholesale-retail price transmission in US on weekly data from January 1984 to March 1991. If we concentrate on the relationship between wholesale and retail prices, GAO proposes an extension of the Balke et al. (1998) model where the impact of positive and negative wholesale price variations is explicitly considered:

$$\Delta rt_t = \phi + \sum_{i=1}^{s} (\alpha_i \Delta ws_{t-i} + \alpha_i^+ \Delta ws_{t-j}^+) + \epsilon_t \tag{31}$$

The empirical findings suggest that both crude-wholesale and wholesale-retail markets do not exhibit any asymmetric behavior to upstream prices, irrespective of the type of impact (i.e. contemporaneous, lag distributed and cumulative), as none of the α_i^+ is statistically significant.

GAO exploits this result to argue that, even though the price transmission in the US gasoline market is, on average, symmetric, nevertheless it could well be asymmetric during periods of market shocks.

In the author's definition, a "market shock" is: "an event or rumor that substantially alters the actual or expected supply of and demand for crude oil or petroleum products" (p. 60). In this sense, price asymmetries may arise since, after a shock, existing inventories are generally sold at their stock-induced market values, independently of their acquisition costs. Consequently, GAO extends this model by including additional economic variables which can help explain the different price adjustment processes during periods characterized by the presence or the absence of economic shocks. A list of these variables includes the fuel stock, the petroleum refinery capacity utilization rate and a logistic time trend, which accounts for the increased flow of market informations over time. Clearly, all these variables affect the crude-wholesale price transmission, while only fuel stock has an impact on the wholesale-retail relationship.

Considering the wholesale-retail transmission and introducing a pair of dummy variables to account for

[1] The only exception is the leaded regular gasoline market, where persistence in the wholesale price decrease is slightly stronger than for price increases

shock-induced price increases and decreases:

$$D1_t = \begin{cases} 1 & \text{if} \quad \Delta ws_t > \delta \\ 0 & \text{otherwise} \end{cases}$$
$$D2_t = \begin{cases} 1 & \text{if} \quad |\Delta ws_t| > \delta \quad \text{and} \quad \Delta ws_t < 0 \\ 0 & \text{otherwise} \end{cases} \tag{32}$$

the modified model is:

$$\Delta rt_t = \phi + \sum_{i=1}^{s}(\alpha_i \Delta ws_{t-i} + \alpha_i^{+} \Delta ws_{t-i}^{+} + \gamma_{1i}(D1_t + D2_t) fs_{t-i} \Delta ws_{t-i} + \gamma_{2i}(D1_t) fs_{t-i} \Delta ws_{t-i}) + \epsilon_t \tag{33}$$

The empirical findings suggest that, after a market shock, wholesale prices respond asymmetrically to crude prices, while symmetry characterizes the wholesale-retail relation. In particular, the crude-wholesale market seems to be characterized by COIA.

A very interesting approach is proposed by Asplund et al. (2000). The authors analyze the Swedish gasoline market, focusing on how Rotterdam spot gasoline prices, exchange rates (er) and taxes (tx) are transmitted to the local retail prices of leaded premium gasoline. In all previous models we have considered the dependent variable as function of a set of explanatory variables which are split into positive and negative values; Asplund et al. reverse this approach by distinguishing between positive and negative values of the dependent variable. They also investigate the effects of using the cost of gasoline expressed in local currency as explanatory variable, rather then decomposing it into variations of the spot price (Δsp) and of the exchange rate (Δer).

To account for all these factors, the authors propose the following two equations, which are independently estimated for increasing and decreasing retail prices:

$$\Delta rt_t|_{\Delta rt_t >, <0} = \phi + \sum_{i=1}^{s} \alpha_i \Delta(sp_{t-i} er_{t-i}) + \gamma \Delta tx_t + \epsilon_t \tag{34}$$

$$\Delta rt_t|_{\Delta rt_t >, <0} = \phi + \sum_{i=1}^{s} \alpha_i \Delta sp_{t-i} er_{t-i} + \sum_{j=1}^{q} \alpha_j sp_{t-j} \Delta er_{t-j} + \gamma \Delta tx_t + \epsilon_t \tag{35}$$

Daily data from January 1980 to December 1996 highlight that the impact of cost variations is different depending on whether they are the consequence of spot price changes or exchange rate changes. With respect to asymmetry, results confirm that rising and falling retail prices respond differently to spot prices and exchange rate variations.

The studies by Balke et al., Karrenbrock, GAO and Asplund et al. provide a complete overview of the asymmetric ARDL models proposed to date; starting from these contributions, many other authors have studied the transmission mechanism between prices of different goods.

Balabanoff (1993) investigates the transmission mechanism between CIF crude oil and the composite barrel of retail prices in France, Germany, Italy, Japan, United Kingdom and US. Data are monthly and cover the period 1985-1992. The author uses Karrenbrock's model and tests for symmetry in the total cumulative effect of crude prices, which is never rejected. Conversely, the persistence of crude price variations is clearly asymmetric for all considered nations, and it implies the presence of DLA.
Even if the author tests directly only CUIA and DLA, the estimation results provide some indirect information about COIA. Actually, US and Germany show significant lags for crude price decreases only; as a consequence, the symmetry in the cumulative impact clearly suggests asymmetry in the contemporaneous effect of the crude price variations in both countries.

Shin (1994) applies Karrenbrock's (1991) model to the analysis of the US crude-wholesale price transmission, using monthly data over the period 1986-1992. In particular he starts with a distributed lag structure, but, since it is not statistically significant, he finally estimates a model with only contemporaneous effects. Results show that crude oil price variations have a symmetric impact on the wholesale market (i.e. they are characterized by COIS).

Another interesting work is due to Duffy-Deno (1996), who combines Karrenbrock's (1991) model with GAO's market shock analysis. In particular, the consequences on asymmetry detection of considering or ignoring the presence of a market shock are investigated.
The author focuses on the Salt Lake City gasoline market, using weekly data for the period 1989-1993 and shows that, when market shocks are not accounted for, COIA, DLEA and CUIA generally emerge. But when the model is extended to incorporate the effect of unusual price variations, wholesale price decreases exhibit a longer persistence and the cumulative impact of wholesale prices is now symmetric.

In 1999, the Energy Information Administration (EIA) investigates the upstream-downstream price transmission at the different levels of the gasoline distribution chain in the US Midwest, using weekly data for the period October 1992-June 1998. Though not considering the presence of market shocks, the authors apply GAO's asymmetry specification and, for each price relationship, they estimate a model in the form of equation (31). A joint and a single test for the significance of the α^+ coefficients provide strong evidence of asymmetries. Although the DLE of crude prices on each downstream market is always symmetric, COIA affects the response of pipeline and rack prices. Furthermore, COIA is also detectable in the response of rack prices to pipeline costs, while DLEA affects the response of retail prices to all upstream markets, with the exception of crude oil, and the Gulf Coast-Chicago pipeline transmission.

Aguiar and Santana (2002) follow Karrenbrock's (1991) approach to analyze the transmission mechanism between farm and retail prices of some agricultural products in Brazil. Monthly data over the period January 1987-June 1998 show that DLEA is a relevant issue in all markets, coffee excluded, where farm costs show only an instantaneous effect on retail prices.

Cumulated Impact Asymmetry (CUIA) is another relevant topic in the Brazilian agricultural market, since the null of CUIS is not rejected in two cases only (namely, onions and rice). Finally, COIA affects milk, rice and coffee markets, as the estimated coefficients do not support any reliable conclusion for the remaining products.

In a report by London Economics (2003) it is possible to find another application of Karrenbrock's model. The authors investigate the mutual relationship between retailer and producer prices of different types of vegetables, fruits, meat and dairy products in Austria, Denmark, France, Germany, Ireland, Netherlands, Spain and UK. In particular, the prices of apples, carrots, potatoes, beef, lamb, flour, eggs and chicken are investigated. Results show that asymmetry (either COIA or DLEA, at least) mainly occurs in the dairy market, where it affects the milk price transmission in UK, France and Denmark, the cheese prices in France and Denmark and, finally, the prices of butter in France and UK. However, asymmetry is also evident in Dutch beef, French bread and Danish chicken markets.

In 2003, Zachariasse and Bunte analyze how farm prices are transmitted to wholesale and retail prices in the Netherlands. In particular, they apply Karrenbrock's approach to evaluate the effects of positive and negative price shocks at the producer level in the markets of pork, beef, poultry, table potatoes and chips. Data from 1990 to 1997 show that RTA characterizes the farm-retail relationship in the pork and table potatoes markets, as well as the farm-wholesale transmission of poultry prices.

Finally, Punyawadee et al. (1991) propose a model for spatial asymmetry applied to Canadian data. This paper investigate whether the price of pork in Ontario has an asymmetric impact on the correspondent price in Alberta, over the period January 1965-December 1989. The authors use weekly data and consider six subperiods to capture potential changes in the price relationships.

The model proposed follows Karrenbrock's specification and allows to test for COIS, DLES and CUIS. Results clearly show that the cumulated impact of Ontario price variations on Alberta pork market is always symmetric. On the contrary, a joint test for the equivalence of the effects of price increases and decreases at each lag shows that COIS and DLES are present in all subsamples, except the time interval from January 1965 to October 1969.

Note that the asymmetry found in the period before the creation of the Alberta Board confirms the importance of structural breaks in the data, a topic which will be subject of separate discussion (see Section 6 below).

4 Equilibrium correction approaches

4.1 Partial adjustment models

During the 90s, the analysis of price asymmetries was enriched by the partial adjustment models (PAM), which describe the adjustment process of a price variable in response to deviations from a given target level (Tables 7a, 7b).

A widely quoted example is provided by Bacon (1991), who studies the relation among ex-refinery petroleum product prices (rf), retail gasoline prices net of taxes and exchange rate (er) using a quadratic PAM model:

$$rt_t = rt_{t-1} + \phi_1(rt^* - rt_{t-1})^2 + \phi_2(rt^* - rt_{t-1}) + \epsilon_t \tag{36}$$

In this model, if ϕ_1 and ϕ_2 are both positive, the adjustment is faster in response to cost increases than to cost decreases, while the converse is true if ϕ_1 is positive and ϕ_2 is negative. In particular, Bacon defines the equilibrium level by assuming the existence of a v-th lag between retail price variations and exchange rates or of a s-th lag between retail price variations and ex-refinery price changes:

$$rt_t^* = \gamma + \delta t + \frac{rf_{t-s}^{\mu}}{er_{t-v}^{\delta}} + \epsilon_t \tag{37}$$

where μ and δ are coefficients within the interval $[0, 1]$, which implies incomplete adjustment to cost variations.

Bacon finds that a statistically adequate specification can be achieved with $s = 1$ and $v = 2$, and proposes the following adjustment process:

$$rt_t = rt_{t-1} + \phi_1(\gamma + \delta t + \frac{rf_{t-1}}{er_{t-2}} - rt_{t-1})^2 + \phi_2(\gamma + \delta t + \frac{rf_{t-1}}{er_{t-2}} - rt_{t-1}) + \eta_t \tag{38}$$

which is jointly estimated with the equilibrium level (37). Using fortnightly UK data from 1982 to 1989, Bacon shows that retail prices rise much more slowly after an increase of ex refinery prices than they fall after a decrease. Moreover, Bacon also provides detailed information about the mean lags which are asymmetric and support the existence of DLEA.

Another interesting application is Salas (2002). Using weekly data for the period January 1999-February 2002, the author examines the relationship between retail and crude (cr) prices in the Philippine market and introduces a specification fairly different from Bacon's model:

$$\Delta rt_t = \phi^+(rt_t^* - rt_{t-1})^+ + \phi^-(rt_t^* - rt_{t-1})^- + \epsilon_t \tag{39}$$

The effect of positive and negative deviations from the target level is now made explicit, and the presence of asymmetry can be detected by testing the null hypothesis: $\phi^+ = \phi^-$. In the definition of the equilibrium level, r_t^*, Salas considers a eight-week lag between retail and crude price variations, which is translated into a moving average structure:

$$rt_t^* = \gamma + \alpha_1 cr_t + \epsilon_t + \theta_1\epsilon_{t-1} + ... + \theta_8\epsilon_{t-8} \tag{40}$$

By introducing the estimated target level into the adjustment equation (39), Salas shows that the adjustment speeds are asymmetric.

Additional applications of PAM to the analysis of price asymmetries are provided by Norman and Shin (1991) and by Shin (1994).

The former paper applies Bacon's model to the transmission mechanism between crude, wholesale and retail gasoline prices in US. Using two different samples of weekly data, one from January 1984 to March 1991, the other from January 1984 to July 1992, the authors find that retail prices respond symmetrically to wholesale and crude oil prices variations.

The latter contribution still employs Bacon's model and analyze the transmission mechanism between wholesale gasoline price and crude price variations in US. Monthly data over the period January 1986-May 1992 confirm the findings of the previous work, providing no evidence of EAP asymmetry.

4.2 Error correction models

The error correction model (ECM) proposed by Engle and Granger (1987) and subsequently modified by Granger and Lee (1989) allows to test all the asymmetries supported by both ARDL specifications and PAM.

Two different approaches have been proposed for ECM estimation, namely Engle and Granger (1987) and Stock and Watson (1993).

4.2.1 Engle and Granger's estimation method

Engle and Granger (1987) suggest a two-step procedure: i) estimate the equilibrium relation and test for cointegration; ii) estimate the ECM, that is a regression where all variables are expressed in first differences, apart from the stationary residuals from step i), the so-called error correction term (ECT), which can be interpreted as the deviations from the equilibrium level.

The application of the error correction approach to the analysis of price asymmetries goes back to Manning (1991), who studies the relationships between retail prices, excise duties (tx) and crude oil prices (cr) in UK, using monthly data over the period 1973-1988.

The analysis starts from a long-run relationship, where retail prices are expressed as a function of crude oil prices and taxes:

$$rt_t = \phi_0 + \phi_1 cr_t + \phi_2 tx_t + \epsilon_t \tag{41}$$

In presence of cointegration, the residuals from regression (41) are stationary and can be introduced as an additional regressor into the following asymmetric ECM:

$$\Delta rt_t = \phi + \varphi_t^+ + \sum_{i=0}^{s} \alpha_i \Delta cr_{t-i} + \sum_{j=0}^{q} \alpha_j^+ \Delta cr_{t-j}^+ + \sum_{h=1}^{r} \beta_h \Delta rt_{t-h} + \sum_{m=1}^{z} \theta_m \Delta tx_{t-m} + \lambda ECT_{t-1} + u_t \tag{42}$$

where ECT_{t-1} are the lagged residuals from equation (41), and φ_t^+ is an intercept dummy which is equal to one if $\Delta cr_t > 0$.

Potential asymmetries implied by variations of crude prices can be found by testing the significance of the coefficients α_j^+ and φ_t^+; actually, when α_j^+ and φ_t^+ are not statistically different from zero, the effect of positive and negative variations of cr is the same (i.e. α_i). Results support the existence of DLE asymmetry, while the cumulative adjustment function provides evidences of RT symmetry.

While Manning proposes a model to test for asymmetries in the direct impact of a price increase and decrease, Von Cramon-Taubadel (1998) focuses on asymmetries in the adjustment to the equilibrium. The author studies the relationship between retail and wholesale prices of pork in the German market, using weekly data over the period January 1990 - October 1993 and proposes the following specification:

$$\Delta rt_t = \phi + \sum_{i=0}^{k} \alpha_i \Delta ws_{t-i} + \sum_{h=1}^{r} \beta_h \Delta rt_{t-h} + \lambda^+ ECT_{t-1}^+ + \lambda^- ECT_{t-1}^- + u_t \tag{43}$$

where, even if a distributed lag structure is considered, the presence of asymmetry is tested only on the ECT. The estimated coefficients and the impulse response functions show the existence of both EAPA and RTA.

A different approach is proposed by Asplund et al. (2000), who we have already quoted in the context of ARDL models. The analysis focuses on the spot retail relationship in the Swedish gasoline market and relies on monthly data from January 1980 to December 1996. The authors propose two different adjustment equations, the first one where the impact of increases and decreases in the marginal costs, measured in local currency, is accounted for:

$$\Delta rt_t = \sum_{i=0}^{s} \alpha_i^+ \Delta(sp \times er)_{t-i}^+ + \sum_{j=0}^{q} \alpha_j^- \Delta(sp \times er)_{t-j}^- + \lambda ECT_{t-1} + u_t \tag{44}$$

while in the second equation the effects of increases and decreases in the spot price and in the exchange rate are made explicit:

$$\begin{aligned}\Delta rt_t = &\sum_{i=0}^{s} \alpha_i^+ sp_{t-i}(\Delta(er)_{t-i}^+) + \sum_{j=0}^{s} \alpha_j^{*+}(\Delta(sp)_{t-j}^+)(er)_{t-j} + \sum_{k=0}^{q} \beta_k^- sp_{t-k}(\Delta(er)_{t-k}^-)\\ &+ \sum_{l=0}^{q} \beta_l^{*-}(\Delta(sp)_{t-l}^+)(er)_{t-l} + \lambda ECT_{t-1} + u_t\end{aligned} \tag{45}$$

Data show that, when specification (44) is used, the presence of CUIS cannot be rejected, but the hypotheses of COI and DLE symmetries can. Specification (45) provides more details, since the adjustment to exchange rate variations is instantaneous and asymmetric, while the response to spot price increases and decreases is distributed over two periods and is affected by COIA and DLEA.

A slight modification of this model is proposed by Bettendorf et al. (2003), who examine the Dutch gasoline market using weekly data for the relationship between exchange rate, retail and spot prices for the period January 1996-December 2001. In this work, asymmetries in the response to exchange rate variations are not considered, leading to the following simplified form:

$$\Delta rt_t = \sum_{i=0}^{s} \alpha_i^+ \Delta sp_{t-i}^+ + \sum_{j=0}^{q} \alpha_j^- \Delta sp_{t-j}^- + \gamma \Delta er_t + \lambda ECT_{t-1} + u_t \tag{46}$$

Despite the availability of daily data, the authors find that a daily analysis is not significant and choose to estimate five distinct weekly models, each referred to a different day of the week. Unexpectedly, the results in terms of asymmetries turn out to depend on the selected day. The usual test for COIS and for DLES (i.e. $\alpha_i^+ = \alpha_j^-$) rejects the null hypothesis only for the Monday, Thursday and Friday models.
It is worth noting that these results contrast with the empirical evidence implied by the number of significant lags, which are different in all models. However, as the authors point out, the Akaike information criterion used to select the lag structure takes values which are almost identical for each pair of lags, meaning that the difference between s and q cannot be supportive of DL asymmetry. Since DLES and COIS imply CUIS, while the reverse is not necessarily true, we can infer that the models for Tuesday and Wednesday data also exhibit CUIS. Finally, the authors estimate the cumulative adjustment functions, which show RTA again only for Monday, Thursday and Friday.

The study by Bettendorf et al. (2003) demonstrates how data selection may influence the results in terms of presence or absence of asymmetries.

This issue is tackled by Galeotti et al. (2003) from two perspectives. First, the analysis concerns five different countries (Italy, France, Spain, Germany, UK); second, the asymmetries are explored not only between spot and retail prices(second stage), but also between crude and spot prices (first stage) and between crude and retail prices (single stage). The authors use monthly data for the period January 1985-June 2000 and focus on the market of leaded gasoline. An asymmetric ECM is estimated for each of the three relationships; for sake of brevity we report only the transmission model between retail and spot prices:

$$rt_t = \gamma_0 + \alpha_1 sp_t + \epsilon_t \tag{47}$$

$$\Delta rt_t = \gamma + \alpha^+ \Delta sp_t^+ + \alpha^- \Delta sp_t^- + \lambda^+ ECT_{t-1}^+ + \lambda^- ECT_{t-1}^- + u_t \tag{48}$$

The asymmetries specified in equation (48) clearly affect the actual response of retail prices to spot price variations and the adjustment to the equilibrium level. Furthermore, the authors calculate the number of weeks necessary to close the gap between current and desired levels of prices. The estimated models have a penchant toward the null hypothesis of symmetry. A result of this kind is not surprising, if we consider that the usual F-tests loose power when applied to an asymmetric ECM, as shown by Cook et al. (1999). By bootstrapping the F-statistics, the authors show that all countries under scrutiny are likely to support both COI and EAP asymmetries, even though there is no evidence of RTA.

In particular, for France and Germany both asymmetries arise in the first and in the single stage, while in Italy, Spain and UK asymmetries affect the spot-retail relationship. Furthermore for Italy also the crude-spot relationship rejects the null of EAPS, while in UK the single stage relationship seems to be characterized by COIA.

Consequently, though price asymmetry is a relevant issue in all countries considered by this study, it assumes different features depending on which market is analyzed.

Finally, it is worth mentioning that, within the crude-retail relationship, retail prices are more responsive to exchange rate increases than decreases.

Conforti et al. (2003) propose an extension of Manning's approach to study the link between world import prices (wp) and the domestic producer prices (dp) of wheat in Egypt. Their specification is:

$$\Delta wp_t = \phi + \sum_{i=0}^{s} \alpha_i \Delta dp_{t-i} + \sum_{h=1}^{r} \beta_h \Delta wp_{t-h} + \lambda ECT_{t-1} + \lambda^+ ECT_{t-1}^+ + u_t \tag{49}$$

In model (49) a dummy variable is used to account for asymmetry in the error correction process, but not in the direct impact of producer price increases and decreases. Monthly prices from January 1989 to May 2001 confirm the statistical significance of the λ^+ coefficient, that is the existence of EAPA.

The paper by Berardi et al. (2000) provides a very general model which encompasses many of the studies described above. Using weekly data from April 1996 to February 2000, the authors study the relationship between ex-refinery oil prices and wholesale prices in the Italian market of leaded and unleaded gasoline and diesel oil, looking for both short- and long-run asymmetries:

$$\begin{aligned}\Delta ws_t = \phi_0 + \sum_{i=0}^{s}(\alpha_i^+ \Delta rf_{t-1}^+ + \alpha_i^- \Delta rf_{t-1}^-) + \sum_{i=1}^{r}(\beta_i^+ \Delta ws_{t-1}^+ + \beta_i^- \Delta ws_{t-1}^-) \\ + \lambda^+ ECT_{t-1}^+ + \lambda^- ECT_{t-1}^- + u_t\end{aligned} \tag{50}$$

where the long-run equation takes the following form:

$$ws = \gamma_0 + \alpha_1 rf_t + \alpha_2 mk_t + \alpha_3 t + \epsilon_t \tag{51}$$

with mk a proxy of the marketing costs.

A test on the joint null hypothesis $\alpha_i^+ = \alpha_i^-$, $\forall i$, $\lambda^+ = \lambda^-$, as well as the behavior of the cumulative adjustment functions show that the three products are affected by the same asymmetries. In particular, RTA turns out to be a relevant issue in the Italian market, in addition to COIA and DLEA, though it is not possible to conclude in favor of both. Conversely, the adjustment path toward the equilibrium level is symmetric.

Model (51) is compared with a restricted form where the trend and the marketing costs are not considered. Results show that the restricted model leads to over-reject the hypothesis of symmetry. Since the general model is found to fit data better than the restricted specification, omitting relevant variables may lead to spurious results in terms of asymmetries.

Kaufmann and Laskowski (2005) propose a modified version of Von Cramon-Taubadel's (1998) model. The authors propose a different way of splitting ECT, which is based on the level of the explanatory price rather than on the sign of the deviations from the long-run equilibrium. The analysis focuses on the prices of heating oil and motor gasoline, it investigates both the crude-refinery and the refinery-retail relationships, and it takes into account the stock level (st) and the utilization rate (ur). If we consider, for instance, the refinery-retail transmission, the model looks as follows:

$$\begin{aligned}\Delta rt_t = \phi_0 + \sum_{i=0}^{s}\alpha_i \Delta rf_{t-i} + \sum_{i=1}^{r}\beta_i \Delta rt_{t-i} + \sum_{i=1}^{r}\gamma_i \Delta st_{t-i} + \sum_{i=1}^{r}\varphi_i \Delta ur_{t-i} \\ + \lambda^+ ECT_{t-1}^{up} + \lambda^+ ECT_{t-1}^{dw} + \sum_{i=1}^{11}\zeta_i D_i + u_t\end{aligned} \tag{52}$$

where D_i are monthly dummies and ECT^{dw} and ECT^{up} are defined as:

$$\begin{cases} ECT_t^{up} = ECT_t & \text{if } \Delta cr_t > 0 \\ ECT_t^{dw} = ECT_t & \text{if } \Delta cr_t \leq 0 \end{cases} \tag{53}$$

Monthly data since January 1986 to December 2002 are used for twelve US regions, showing that EAPA is not a relevant issue at the crude-refinery level. On he contrary, EAPA affects the heating oil market across US, while little evidence of asymmetry is found in the motor gasoline market, as the null hypothesis of symmetry can be rejected only for California, Lousiana and Idaho.

All the contributions so far described estimate the different ECMs with the Engle and Granger's method. However, the recent literature is far richer of applications of this type of models.
A representative selection of recent contributions should start with Salas (2002), who studies the Philippine retail gasoline market from January 1999-February 2002. In addition to a PAM, whose structure has been described in the previous section, he estimates an ECM along the lines of Asplund et al. (2000) to test the cumulative impact of crude price variations on retail price, which results to be asymmetric.

Conforti et al. (2003) also analyze the producer-import price relationship in the Ethiopian, Rwandan and Ugandan coffee markets over the period January 1990-December 2001, using a specification similar to Berardi et al. (2000) to test for COIS, DLES and for EAPS. Monthly data show that neither of these three countries exhibits any evidence of asymmetry[2].

Von Cramon-Taubadel and Meyer (2003) investigate the link between retail and wholesale prices of German lettuce and chicken, using weekly data for the the period May 1995-December 2000. In particular, two different set of data are used, the first based on individual store prices, while the second on average retail prices. Using the model proposed by Von Cramon-Taubadel (1998), the authors show that, for both products and when individual data are used, the null hypothesis of symmetry can be rejected, while aggregated data provide no evidence of asymmetric price behavior.

In a second paper, Conforti (2004) applies Conforti's et al. (2003) model to test the existence of EAPA in the adjustment of the local prices of different agricultural products in response to world prices variations, for a number of countries [3]. As a whole, annual data from 1969 to 2001 support the existence of asymmetry.

The National Department of Agriculture in South Africa (NDA)(2003) uses the model of Von Cramon-Taubadel (1998) to study the farm-retail transmission in the South African market of maize meal, bread, fresh and long life milk, cheddar cheese and cooking oil. Monthly data over the period January 2000-July 2003 are used to obtain the impulse response functions for farm price increases and decreases, which suggests the presence of RTA for all considered cases.

[2] The presence of COIS and DLES also implies CUIS [3] Namely, Argentina, Brazil, Chile, Costa Rica, Egypt, Ethiopia, Ghana, India, Indonesia, Mexico, Pakistan, Senegal, Thailand, Turkey, Uganda, Uruguay.

London Economics (2003) analyzes the mutual relationship between retailer and producer prices of a number of goods in Austria, Denmark, France, Germany, Ireland, Netherlands, Spain and UK (see Section (3.1.2)). In this study the authors also employ a variation of the Von Cramon-Taubadel (1997) ECM for the price series which turn out to be cointegrated. Empirical evidence supports the presence of EAPA in the producer-retail relationship for the markets of Danish carrots and of UK bread, and in the retail-producer transmission mechanism for the UK lamb market.

Bachmeier and Griffin (2003) analyze the UK gasoline market from February 1985 to November 1998. They compare a symmetric ECM with an asymmetric specification. Using daily data, they are able to find that the UK market is characterized by symmetries of the type COIS, DLES, CUIS, RTS and EAPS.

Krivonos (2004) investigates the transmission mechanism between local and world coffee prices before and after the structural reforms which affected the Sub-Saharan Africa and the Latin America coffee markets during the last 1980s and the first 1990s. Using monthly data from 1984 to 2003, the author finds no evidence of asymmetries in the pre-reform period, since local prices were driven by the local governments and were not directly influenced by world prices. In the post-reform period, local prices started to react to world price variations, and three of the twenty countries under study have shown an asymmetric price behavior. In particular, Kenya, Madagascar and Cameroon seem to be influenced by COIA and RTA. Note that while the former asymmetry is formally tested, the latter is only indirectly inferred from the degree of adjustment of domestic prices after six and twelve months.

Another interesting paper is proposed by Radchenko (2005a), who analyzes the link between oil price volatility and the asymmetric response of gasoline prices to oil price variations in the US market. In his work, Radchenko uses both a VAR model, which will be formally introduced in Section 5.2, and an error correction specification close to the model of Berardi et al. (2000). The author uses weekly data from March 1991 to February 2003 to compute the impulse response functions to crude price increases and decreases. Results show that RTA affects the response of retail prices.

In a different study, Radchenko (2005b) applies a similar model to the US gasoline market, using weekly data from March 1991 to February 2003. Impulse response function are used to investigate the crude-retail and the spot-retail price transmission mechanisms, both of which turn out to be affected by RTA.

Finally, Grasso and Manera (2005) analyze the three classical stage of the oil-gasoline price transmission mechanism, i.e. crude-spot, spot-retail and crude-retail, in France, Germany, Italy, Spain and UK. The authors follow Galeotti et al. (2003) in the choice of the sample period, which goes from January 1985 to March 2003 with a monthly frequency, and in the use of bootstrapped F-tests for the null hypothesis of symmetry. As for the model structure, they apply a specification similar to Berardi et al. (2000) specification,

though enriched with autoregressive asymmetrical effects and with asymmetric adjustments to exchange rate variations. The estimations results are then used to test for symmetries of the type EAPS, COIS, DLES.
When the crude-spot relationship is considered, all countries exhibit COIA in response to exchange rate variations. DLEA is found in all countries for the response to crude prices, with the exception of UK, while COIA affects only the Spanish market.
The spot-retail transmission mechanism is affected by EAPA in all considered countries; among those, Spain and UK are also affected by COIA and DLEA, while only DLEA is evident in the French market. Finally, in France, Italy and UK asymmetric autoregressive effects are also present.
With respect to the single stage analysis, autoregressive asymmetrical effects are found only in UK, which is also the only market where the hypothesis of DLES in the response to exchange rate variations can be rejected. The contemporaneous impact of the exchange rate is asymmetric not only in the UK market, but also in Germany and Italy. Finally, EAPA occurs in Italy and France, which has a behavior similar to UK and is also affected by COIA in the response to crude price variations.

4.2.2 Stock and Watson's method

The two-step Engle and Granger's approach is not the only alternative to estimate ECM. Stock and Watson (1993) propose a different method which essentially requires the simultaneous estimation of the long-run equilibrium and of the adjustment process.
Given the differences between the two approaches, an open question is whether they lead to the same results in terms of asymmetries or, if this is not the case, which is the more reliable.
In 1991 Manning published a paper which is one of the first applications of the Stock and Watson's procedure to the analysis of price asymmetries, and it provides a comparison with Engle and Granger's approach. Manning analyzes the UK gasoline market using monthly data over the period 1973-1988. The method of Stock and Watson, which avoids to impose a pre-determined equilibrium level on the relation among crude and retail gasoline prices and taxes, produces the following model:

$$\Delta rt_t = \phi + \varphi_t^+ + \sum_{i=0}^{s} \alpha_i \Delta cr_{t-i} + \sum_{j=0}^{q} \alpha_j^+ \Delta cr_{t-j}^+ + \sum_{h=1}^{r} \beta_h \Delta rt_{t-h} + \sum_{m=1}^{z} \theta_m \Delta tx_{t-m} + c_1 rt_{t-1} - c_2 cr_{t-1} + c_3 tx_{t-1} + u_t \tag{54}$$

The long-run equilibrium, which in the Engle and Granger's procedure is pre-estimated, can be identified from the estimation of the ECM. Given equation (54), the implied long-run solution is:

$$rt = \frac{\hat{\phi}}{\hat{c_1}} + \frac{\hat{c_2}}{\hat{c_1}} + \frac{\hat{c_3}}{\hat{c_1}} + \epsilon \tag{55}$$

In this case, results confirm the outcomes derived from Engle and Granger's procedure, that is the null hypothesis of RTS is not rejected, while the null hypothesis of DLES is rejected.

Arden et al. (1997) is an example of contrasting results from the two approaches. The analysis starts from the statement that it is inappropriate to look for asymmetries using a symmetric relationship. Actually, the first step of the Engle and Granger's procedure is based on a symmetric long-run relation, which may invalidate the identification of long-run asymmetries. Some authors have solved this problem using TAR (Threshold Auto-Regressive) models (see next section). Arden et al., instead, choose to jointly estimate the long-run equilibrium and the asymmetry parameters, as prescribed by the Stock and Watson's method.
The authors investigate how input prices x and labor costs lb affect output prices y in the UK manufacturing sector, during the period 1970-1996 (quarterly data). The production function is assumed to be Cobb-Douglas, and the ECM includes asymmetries in the adjustment to the equilibrium:

$$\begin{aligned}\Delta y_t = \phi_0 + \sum_{k=1}^{r} \Delta y_{t-i} + \alpha_1 \Delta x_t + \alpha_2 \Delta lb_t + \theta^+ (y_{t-1} - c_1 x_{t-1} - c_2 lb_{t-1} - \mu)^+ \\ + \theta^- (y_{t-1} - c_1 x_{t-1} - c_2 lb_{t-1} - \mu)^-\end{aligned} \tag{56}$$

For comparison, the authors choose to estimate this model using the two-step Engle and Granger's procedure. Testing the null hypothesis $\theta^+ = \theta^+$ with both techniques produces contrasting empirical findings. Specifically, only the Stock and Watson's method supports strong evidence of EAP asymmetry.

Among the contributions which rely on the simultaneous estimation of the long-run equilibrium and the adjustment process, a widely cited paper is Borenstein at al. (1997), which focuses on the US gasoline market for the period January 1986-December 1992.
Since the purpose of the article is to detect the presence of asymmetries at the different levels of the distribution chain, the authors use four different prices: crude spot prices, gasoline spot prices, wholesale prices and retail prices. All data are weekly, except retail prices which are biweekly.
The authors consider first crude and retail prices, and define the long-run relation as:

$$rt_t = \gamma_0 + \alpha_1 cr_t + \epsilon_t \tag{57}$$

Model (57) is then included in the ECM:

$$\Delta rt_t = -\theta_1\phi_0 + \sum_{i=0}^{p}(\alpha_i^+\Delta cr_{t-1}^+ + \alpha_i^-\Delta cr_{t-1}^-) + \sum_{i=1}^{p}(\beta_i^+\Delta rt_{t-1}^+ + \beta_i^-\Delta rt_{t-1}^-)$$
$$- \sum_{j=2}^{n}(\theta_1\eta_j D_{j,t}) + \theta_1 rt_{t-1} - \theta_1\phi_1 cr_{t-1} - \theta_1\phi_2 t + \epsilon_t \tag{58}$$

where D is a set of variables which denote seasonal effects. Each variable in D indicates the period of the year the data are referred to; consequently, n is equal to 24 or to 54 for biweekly or weekly data, respectively. This model captures both the contemporary and lagged impact of crude price variations and asymmetric autoregressive effects, but it imposes the presence of symmetry in the persistence of crude price variation, as the number of lags is constant and equal to p.

Since the contemporary variation of upstream prices can be considered endogenous, Borenstein et al. choose to use a Two Stage Least Square (TSLS) estimator, using the positive and the negative variations of the Brent and Forties crude spot and futures prices in England as instruments. The large difference between the estimated values of the coefficients α_0^+ and α_0^- confirms the presence of asymmetry in the contemporary effects of crude price variations, even though the authors do not provide any statistical test to support their conclusion. Then, in order to evaluate the presence of RTA, they estimate the cumulative adjustment functions B_j^+ and B_j^-, and test the significance of the quantity A_n defined as:

$$A_n = \int_{j=0}^{n}(B_j^+ - B_j^-)dj \tag{59}$$

The results confirm the existence of RT asymmetry. By extending their analysis to the intermediate levels of the distribution chain, Borenstein et al. find that almost all markets exhibit RT asymmetry; the only exception is the spot-wholesale transmission mechanism, which results to be symmetric.

Bachmeier and Griffin (2003) question whether Borenstein et al. (1997) have really proved the existence of asymmetry. In particular, they assert that the application of Stock and Watson's method to the model proposed by Borenstein et al. leads to over-reject the null hypothesis of symmetry. To demonstrate their claim, the authors analyze the same US crude refinery market analyzed by Borenstain et al., with a larger data sample (February 1985-November 1998).

Specifically, they estimate an ECM using both Engle and Granger's and Stock and Watson's techniques, and use an OLS estimator, rather than 2SLS, in both cases. In accordance to what they claim, the authors show that the null of RT symmetry is rejected only when the model is estimated using the Stock and Watson's approach. These results support Arden et al. (1997) conclusion that only the Stock and Watson's estimation method leads to strong asymmetries.

The debate on the empirical robustness of the two approaches is wide. A selection of contributions which use the Stock and Watson's method is reported below.

In 1996, Borenstein and Shepard analyze the US gasoline market from 1982 to 1991 and concentrate on retail, wholesale and crude oil prices. On the one hand, the authors analyze how the retail margin, which is given by the difference between retail and terminal prices, is affected by retail and terminal prices variations, as well as by the volume of gasoline consumption. On the other hand, they investigate the link between terminal and crude oil prices. Results confirm the presence of RTA in both relationships.

Eltony (1998) studies the instant response of gasoline prices at the pump to positive and negative variations of crude prices and exchange rate. They use monthly data for UK and US markets during the period January 1980-June 1996. Empirical evidence shows that both UK and Us gasoline prices exhibit COIA to crude price and exchange rate variations.

Reilly and Witt (1998) study the same problem analyzed by Eltony (1998) with the same model. Using monthly data for the UK market during the period January 1982-June 1995, they show that retail prices respond much more strongly to crude prices increases than decreases. Results suggest that retailers change their prices in response to exchange rate reductions but not to exchange rate increases.

Balke et al. (1998), who study the gasoline price transmission mechanism at the different levels of the distribution chain, use Manning's model. They find that DLE and COI asymmetries occur in almost all upstream-downstream relationships between crude, spot, wholesale and retail prices, with and without taxes. The only exception is the spot-retail ex-taxes price relationship, where the null hypothesis of symmetry cannot be rejected.

Another application of Manning's model is provided by Peltzman (2000), who offers a very detailed discussion of the issue of price transmission, by investigating how the output prices of 77 consumer and 165 producer goods in US react to cost variations. The analysis relies on monthly data over the period 1978-1996 and focuses on the cumulative impact of input prices. Results show that asymmetry is a relevant factor in all considered markets.

Eckert (2002) investigates the response of weekly retail gasoline prices to wholesale prices in Ontario, Canada, from November 1989 to September 1994. The model follows Borenstein et al. and confirms that the Ontario market is affected by both COIA and RTA.

In 2004, Contin et al. investigate the retail gasoline mark in Spain. They analyze the relationship between crude and retail prices before and after the abolition of the system of ceiling price regulation, which took place in 1998. Data are weekly and the authors provide both a statistical test for the equality of the contemporaneous and the distributed lag impact of positive and negative crude price variations and the cumulative adjustment functions, in response to crude price increases and decreases. The Spanish market results to be affected by COIA, DLEA and RTA, before (i.e. January 1993-September 1998) and after (i.e. October 1998-December 2002) the reform. However, it is worth noting that in the first period gasoline prices adjust faster in response to crude price decreases than to crude price increases, while the reverse is true for the second period.

Finally, a completely different solution to the problem of jointly estimating the long-run equilibrium and the error correction process is discussed in Verlinda (2004), who uses a Bayesian estimation procedure. A non-linear Bayes regression, with informative prior, is used to estimate the impact of spot gasoline prices on retail ones, which is represented as follows:

$$\Delta rt_t = \phi + \sum_{i=0}^{s} \alpha_i \Delta ws_{t-i} + \sum_{j=0}^{q} \alpha_j^+ \Delta ws_{t-j}^+ + \sum_{h=1}^{r} \beta_h \Delta rt_{t-h} + \sum_{n=1}^{r} \beta_r^+ \Delta rt_{t-n}^+ + \theta(rt_{t-1} - c_1 - c_2 ws_{t-1}) + u_t \quad (60)$$

Specification (60) is a mixture of Bettendorf's et al. (2003) model and Manning's (1991) approach. The author considers weekly station prices in Southern California from September 2002 to May 2003 to study the price transmission mechanism at the station level and at an aggregated level. Cumulative impulse response functions for spot price increases and decreases show that both average and local retail prices are affected by RTA.

So far we have illustrated a number of contributions which use the Engle and Granger's and the Stock and Watson's approaches to model cointegrated variables and long-run asymmetries. These studies in general suggest that the two approaches can often produce different outcomes in terms of asymmetries. Unfortunately, the available literature does not help the applied researcher to determinate which results are more reliable. However, a third approach to asymmetric cointegration is available in the literature and will be the subject of the next section.

4.2.3 TAR and M-TAR cointegration

As we have already pointed out, Engle and Granger's (1987) originates from a symmetric long-run relationship and allows to introduce asymmetries in the estimation of the adjustment process of a ECM. However, some authors assert that, if the true long-run relationship between two prices is asymmetric, a test for cointegration based on a symmetric long-run equilibrium may lead to misleading results (see, among others, Arden et al., 1997; Abdulai, 2002). As discussed in the previous section, Stock and Watson's estimation technique does not impose symmetric behavior when testing for cointegration; a different solution to this problem is proposed by Ender and Granger (1998), who introduce Threshold Auto-Regressive (TAR) and Momentum-TAR (M-TAR) cointegration.

In a Engle and Granger (1987) context, given the long-run relationship:

$$y_t = \phi_0 + \phi_1 x_t + \epsilon_t \tag{61}$$

x and y are cointegrated if the null hypothesis $\rho = 0$ is rejected in the following regression model:

$$\Delta\epsilon_t = \rho\epsilon_{t-1} + u_t \tag{62}$$

Enders and Granger (1998) suggest to use the alternative TAR specification, where the relation between $\Delta\epsilon_t$ and ϵ_{t-1} is supposed to vary between two regimes, depending on the value of ϵ_{t-1}:

$$\Delta\epsilon_t = I_t\rho_1\epsilon_{t-1} + (1 - I_t)\rho_2\epsilon_{t-1} + \epsilon_t \tag{63}$$

I_t is an indicator function defined as:

$$I_t = \begin{cases} 1 & \text{if } \epsilon_{t-1} \geq 0 \\ 0 & \text{if } \epsilon_{t-1} < 0 \end{cases} \tag{64}$$

If the null hypothesis $\rho_1 = \rho_2 = 0$ in equation (63) is rejected, then x and y are cointegrated and the asymmetric ECM which stems from the TAR specification is:

$$\Delta y_t = \sum_{i=0}^{s} \alpha_i \Delta x_{t-i} + \gamma^+\epsilon_{t-1}I_t + \gamma^-\epsilon_{t-1}(1 - I_t) \tag{65}$$

Model (65) is equivalent to the usual specification:

$$\Delta y_t = \sum_{i=0}^{s} \alpha_i \Delta x_{t-i} + \gamma^+ECT^+_{t-1} + \gamma^-ECT^-_{t-1} + \epsilon_t \tag{66}$$

and, consequently, a test on the null hypothesis $\gamma^+ = \gamma^-$ is a test for EAPA.

Enders and Granger propose a second model for cointegration, known as M-TAR. This name comes from the financial definition of "momentum", which describes the rate of acceleration of prices. As the authors assert, M-TAR models are especially valuable when adjustment is asymmetric in a way that the series exhibit more momentum in one direction than in the other.

In M-TAR models the threshold is placed on the variation of ϵ_{t-1}, rather than on ϵ_{t-1}, and the indicator function in equation (63) and (65) takes the form:

$$I_t = \begin{cases} 1 & \text{if } \Delta\epsilon_{t-1} \geq 0 \\ 0 & \text{if } \Delta\epsilon_{t-1} < 0 \end{cases} \tag{67}$$

In this case equation (66) is not suitable to test for the presence of EAPA anymore. As a matter of fact, the term $\gamma^{+}\epsilon_{t-1}I_t + \gamma^{-}\epsilon_{t-1}(1 - I_t)$ is no longer equivalent to the term $\gamma^{+}\epsilon^{+}_{t-1} + \gamma^{-}\epsilon^{-}_{t-1}$, since γ^{+} and γ^{-} are now the weights of ϵ_{t-1} when its variation with respect to the previous period is positive or negative, respectively. This model clearly introduces a new kind of asymmetry in the adjustment to the equilibrium, which we will call Momentum Equilibrium Adjustment Path Asymmetry (MEAPA).

TAR and M-TAR cointegration has become increasingly popular in the literature on asymmetric price transmission during the very last years.

One recent contribution in this direction is Abdulai (2000), who studies the Ganan market of maize during the period May 1980-October 1997. The major aim of this analysis is the transmission mechanism between the wholesale price of the central market of Techiman and the corresponding local prices formed in the Accra and Bolgatanga markets.

Abdulai tests for cointegration using both the linear Engle and Granger's model and the TAR and M-TAR specifications, finding the existence of a long-run relationship in both the Techiman-Accra and the Techiman-Bolgatonga markets. However, the M-TAR specification seems to fit data better than the others, thus leading the author to use model (63) with the indicator function given by equation (67).

The estimation results and the impulse response functions indicate that both markets are affected by MEAPA and RTA.

In 2001 Hassan and Simioni study the link between shipping point prices and retail prices of French tomatoes and chicory, using both TAR and M-TAR cointegration. Specifically, they investigate 22 relationships for the first market and 20 for second market, where each relation is identified by the type of tomato/chicory and the origin of production.

The empirical findings suggest that, even if almost half of the relationships under study for the two markets exhibit an asymmetric behavior, different products are affected by different kind of asymmetries, as tomatoes display MEAPA while chicory EAPA. This result confirms that M-TAR models fit well in presence of very quick adjustments of retail prices to wholesale prices (as justified by the highly perishable nature of tomatoes). TAR cointegration, instead, is appropriate when the less perishable nature of a good (e.g.

chicory) makes retail prices less affected by supply fluctuation (see Abdulai and Rieder, 1999).

In 2002, Abdulai proposes a different contribution, which investigates the mutual relationship between retail and wholesale prices of pork meat in the Swiss market, from January 1988 to September 1997, using monthly data. In this case, a symmetric cointegration test along the Engle and Granger's (1987) approach does not provide any evidence of cointegration, while both TAR and M-TAR models support cointegration. Standard AIC and SBC selection criteria indicate M-TAR as the best fitting specification. The estimated coefficients and the impulse response functions show that there is no impact of retail prices on producer prices, while the response of the retail market to production costs is affected by both MEAP and RT asymmetries.

Gonzales et al. (2003) apply TAR and M-TAR cointegration to the analysis of the French value chain of cod and salmon, using monthly data ranging from February 1988 to December 1999. The paper looks for EAPA, MEAPA and RTA in the transmission mechanism between upstream and downstream prices, but the estimated coefficients and the impulse response function provide no evidence of asymmetries .

Grasso and Manera (2005) analyze the gasoline market in Italy, France, Germany, Spain and UK and employ, among other models, a M-TAR specification. Results show that MEAP asymmetries occur in the crude retail price transmission for all the considered countries, as well as also in the crude-spot relationship for Italy and Spain.

5 Modern econometric models

5.1 Regime switching models

Regime switching models started to be applied to the analysis of price asymmetries in the middle 90s with the pioneering work of Powers (1995), which we have already described in the context of the ARDL models based on Wolffram's approach (Table 8).

As we have seen, this class of models applies to situations in which the relation between y and a set of variables X depends on the state of a variable v, which can belong to X but does not need to.

Powers analyzes the Californian market for lettuce and argues that the relationship between local retail prices and shipping point FOB prices (fb) may be different depending on the volume of locally grown lettuce. The argument goes as follows: since the FOB price of locally grown lettuce is influenced by both national and local market, the link between retail price and the FOB price of lettuce might weaken when the arrivals of locally grown lettuce increases.

In order to test this hypothesis, Power proposes the following model:

$$\begin{cases} rt_t - rt_0 = \sum_{i=0}^{s} \alpha_i^{+} C\Delta fb_{t-i}^{+} + \sum_{i=0}^{s} \alpha_i^{-} C\Delta fb_{t-i}^{-} + \beta_i^{+} C\Delta mk_{t-i}^{+} + \beta_i^{-} C\Delta mk_{t-i}^{-} + \epsilon_t & \text{if } LL_t \leq \delta_1 \\ rt_t - rt_0 = \sum_{i=0}^{s} \alpha_i^{+*} C\Delta fb_{t-i}^{+} + \sum_{i=0}^{s} \alpha_{i*}^{-} C\Delta fb_{t-i}^{-} + \beta_i^{+} C\Delta mk_{t-i}^{+} + \beta_i^{-} C\Delta mk_{t-i}^{-} + \epsilon_t & \text{if } LL_t > \delta_1 \end{cases} \tag{68}$$

where the impact of FOB prices depends on the actual level of local lettuce (LL_t) and within-regime asymmetries are allowed. Weekly data from March 1986 to August 1992 show that the levels of local lettuce do influence the retail-FOB relationship and that CUIA occurs in both states of the world.

Goodwin and Holt (1999) investigate the relationship between farm, wholesale and retail beef prices in US, using a three-regime error correction specification, with a threshold placed on the ECT, rather than on a variable external to the model as in Power's:

$$\begin{cases} \Delta rt_t = \phi_0 + \sum_{i=1}^{r} \beta_i \Delta rt_{t-i} + \sum_{i=0}^{s} \alpha_i \Delta ws_{t-i} + \sum_{i=0}^{p} \gamma_i \Delta fm_{t-i} + \lambda ECT_{t-1} + u_t & \text{if } ECT_{t-1} \leq \delta_1 \\ \Delta rt_t = \phi_0^{'} + \sum_{i=1}^{r} \beta_i^{'} \Delta rt_{t-i} + \sum_{i=0}^{s} \alpha_i^{'} \Delta ws_{t-i} + \sum_{i=0}^{p} \gamma i^{'} \Delta fm_{t-i} + \lambda^{'} ECT_{t-1} + u_t^{'} & \text{if } \delta_1 < ECT_{t-1} < \delta_2 \\ \Delta rt_t = \phi_0^{''} + \sum_{i=1}^{r} \beta_i^{''} \Delta rt_{t-i} + \sum_{i=0}^{s} \alpha_i^{''} \Delta ws_{t-i} + \sum_{i=0}^{p} \gamma i^{''} \Delta fm_{t-i} + \lambda^{''} ECT_{t-1} + u_t^{''} & \text{if } \delta_1 ECT_{t-1} \geq \delta_2 \end{cases} \tag{69}$$

Weekly data over the period January 1981 - March 1998 show that the existence of more than one regime cannot be rejected, which means that the dynamic relationship between farm, wholesale and retail prices as a whole is different according to the deviation from the long-run equilibrium, which we call REAPA.
The authors also look for RTA using impulse response functions, but in this case no evidence of asymmetry is found.

In 2001, Goodwin and Piggott apply model (69) to the daily price transmission between the central market and three local markets of corn and soybeans in North Carolina, from January 1992 to March 1999. Although the investigated market is not comparable with the one studied by Goodwin and Holt (1999), results still suggest the presence of REAPA and RTS.

A different specification is provided by Godby et al. (2000), who use a deterministic error correction regime switching model to investigate the existence of REA between crude and retail gasoline prices in thirteen major cities across Canada:

$$\begin{cases} \Delta rt_t = \phi_0 + \sum_{i=0}^{s} \alpha_i \Delta cr_{t-i} + \lambda ECT_{t-1} + u_t & \text{if } f(\Delta cr_t) \leq \delta_1 \\ \Delta rt_t = \phi_0^{'} + \sum_{i=0}^{s} \alpha_i^{'} \Delta cr_{t-i} + \lambda^{'} ECT_{t-1} + u_t^{'} & \text{if } f(\Delta cr_t) > \delta_1 \end{cases} \tag{70}$$

In this case the threshold is not placed on the ECT, but on a function $f(.)$ of crude prices; in particular, five different cases are considered, according to whether the threshold is imposed on the mean of crude price variations, calculated on the most recent eight lagged values (in this case $f(.)$ is the mean function), or directly on the lagged values at one, two, three and four lags respectively (in this case $f(.)$ is the lag operator).

Using weekly data over the time span January 1990 - December 1996, the authors test the equality of all coefficients between the two regimes and are not able to reject the null of symmetry under any circumstances.

Johnson (2002) uses a modified version of specification (70), where autoregressive effects are considered and a zero-threshold is posed on the input price variation.

In particular, the author analyzes the weekly transmission of wholesale price variation on retail prices in US, focusing on fifteen gasoline and diesel markets over the period July 1996 - June 1998.

In about half of the examined cases, the existence of two regimes determined by the value of wholesale price variations is supported by data, thus showing the presence of REA in both diesel and gasoline markets.

Johnson notes also that, for both fuels, the behavior of the cumulative impulse response functions depends on the sign of wholesale price shocks, which confirms the existence of RTA.

If the models described above consider either between-regime or within-regime asymmetries, Lewis (2004) proposes a specification which supports both:

$$\begin{cases} \Delta rt_t = \phi_0 + \sum_{i=1}^{r} \beta_i \Delta rt_{t-i} + \sum_{i=0}^{s} \alpha_i^{+} \Delta ws_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{-} \Delta ws_{t-i}^{-} + \lambda ECT_{t-1} + u_t & \text{if } ECT_{t-1} \leq \delta_1 \\ \Delta rt_t = \phi_0^{'} + \sum_{i=1}^{r} \beta_i^{'} \Delta rt_{t-i} + \sum_{i=0}^{s} \alpha_i^{'+} \Delta ws_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{'-} \Delta ws_{t-i}^{-} + \lambda^{'} ECT_{t-1} + u_t^{'} & \text{if } ECT_{t-1} > \delta_1 \end{cases} \tag{71}$$

In particular, the author studies the spot-retail gasoline price transmission in California, using weekly data over the period January 2000 - December 2001, and, in order to test for asymmetry, he compares the above model with the corresponding symmetric specification.

Results suggest that the Californian market is affected by REAPA and, within each regime, by COIA, CUIA and DLEA; furthermore, the cumulative adjustment functions for wholesale price increases and decreases, evidence the existence of RTA too.

Grasso and Manera (2005) use a model with between-regime asymmetries as in Johnson (2002). Monthly data over the period January 1985, March 2003 show that in the Italian, French, German, Spanish and English gasoline markets REA occurs at all levels of the gasoline distribution chain, with the exception of the crude-spot transmission in Italy .

Finally, Radchenko (2005b)proposes, along with the ECM described above, a very interesting RS specification, where the transition from a regime to the other is driven by a Markov chain, rather than by a deterministic process.

The author, who focuses on the US gasoline market, supposes that the relationship between crude and retail prices and between spot and retail prices can depend on whether a price shock in the upstream market is viewed as long-lasting or or short-lasting. He assumes that the consumers' beliefs about shocks can be described by an unobserved state variable S following a Markov process with transition probability matrix:

$$P\left[S_t = i | S_{t-1} = j\right] = p_{ij} \tag{72}$$

being i, j in $\{1, 2\}$, where indices i and j indicate a long-term and a short-term shock, respectively. Another interesting feature of Radchenko's model is that only the impact of upstream prices is allowed to vary between regimes, while the effect of all other variables is maintained fixed; for sake of brevity, we report only the crude-retail model, but all considerations are equally valid for the spot-retail specification:

$$\begin{cases} \Delta rt_t = \phi_0 + \sum_{i=1}^{r} \beta_i^{+} \Delta rt_{t-i}^{+} + \sum_{i=1}^{r} \beta_i^{-} \Delta rt_{t-i}^{-} + \sum_{i=0}^{s} \alpha_i^{+} \Delta cr_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{-} \Delta cr_{t-i}^{-} & \\ \qquad\qquad +\lambda^{+} ECT_{t-1}^{+} + \lambda^{-} ECT_{t-1}^{-} + u_t & \text{if } S_t = 1 \\ \Delta rt_t = \phi_0 + \sum_{i=1}^{r} \beta_i^{+} \Delta rt_{t-i}^{+} + \sum_{i=1}^{r} \beta_i^{-} \Delta rt_{t-i}^{-} + \sum_{i=0}^{s} \alpha_i^{'+} \Delta cr_{t-i}^{+} + \sum_{i=0}^{q} \alpha_i^{'-} \Delta cr_{t-i}^{-} & \\ \qquad\qquad +\lambda^{+} ECT_{t-1}^{+} + \lambda^{-} ECT_{t-1}^{-} + u_t & \text{if } S_t = 2 \end{cases} \tag{73}$$

A Bayesian regression with informative prior on weekly data from March 1991 to February 2003 reveals that retail prices do respond differently depending on the nature of crude and spot prices shocks. However, this does not provide any information about asymmetries, since, in this model, only within-regime asymmetries are considered. In particular, Radchenko focuses on RTS and tests for it using a cumulative response function for upstream price increases and decreases, which confirms the presence of asymmetry in both states of the world.

5.2 Systems of equations

The standard single equation analysis of price asymmetries has been generalized to systems of equations in order to take into account the potential interdependences among input and output prices and other exogenous variables.

In the literature, multivariate extensions of ARDL, EC and RS models have been proposed, which are respectively known as Vector Auto-Regressive (VAR), Vector Error Correction (VEC) and Vector Regime Switching (VRS) models (Table 9).

5.2.1 VAR models

VAR models are generally defined as:

$$Y_t = \Phi_1 Y_{t-1} + \Phi_2 Y_{t-2} + ... + \Phi_q Y_{t-q} + \epsilon_t \tag{74}$$

where Y_t is a $n \times 1$ vector of the variables of interest, ϵ_t is a $n \times 1$ vector of error terms, and $\Phi_i, .., \Phi_q$ are $n \times n$ matrices of coefficients.

In model (74), each element of vector Y, $\{y_1, ..., y_n\}$, is a symmetric function of its own lagged values and of the lagged values of all the other components of Y. However, VAR models can be generalized to incorporate asymmetries, along the same line of the corresponding univariate specifications.

A very general asymmetric version of model (74) is:

$$Y_t = \Phi_1^+ Y_{t-1}^+ + \Phi_1^- Y_{t-1}^- + ... + \Phi_q^+ Y_{t-q}^+ + \Phi_q^- Y_{t-q}^- + \epsilon_t \tag{75}$$

where all (or only some) elements of vector Y are now split into positive and negative values, according to their sign.

In 1993 Capps proposes a VAR specification for the wholesale-retail price transmission of fifteen meat products in the Houston market. The model is a restricted form of the VAR described above, where direct cross-price effects are not considered, but only contemporaneous correlation between errors in different equations is accounted for.

The mth product is specified according to Houck's (1977) model:

$$rt_{mt} - rt_{m0} = \phi_{m0} + \phi_{m1}^+ C\Delta ws_{mt}^+ + \phi_{m2}^- C\Delta ws_{mt}^- + \epsilon_{mt} \tag{76}$$

and the resulting system is estimated with seemingly unrelated regression (SUR). Weekly data, from September 1986 to November 1988, show asymmetry in the contemporaneous impact of cumulative wholesale price variations.

Another interesting example is provided by Willett et al. (1997), who analyze the transmission mechanism between wholesale, retail and shipping point prices of red delicious apples, in the West, the Northeast and the North Central regions of the US, over the period 1975-1991. A Granger causality test suggests that, in each of the three regions, the transmission mechanism does not simply flow from the upstream to the downstream market, but that feedback effects are also present. In particular, shipping point prices both lead and are led by wholesale and retail prices. On the contrary, retail prices are only driven by shipping point prices, while wholesale prices are also affected by retail costs.

Since the shipping point price of red delicious apples is set in the Western region, irrespective of the final market where this product is sold, it is appropriate to study the transmission mechanism between prices in each region separately, or, alternatively, to model the interdependences among regions; the choice to use a VAR goes into the second direction.

As a matter of fact, Willett et al. formulate three region-specific equations for both wholesale and retail prices, as well as a single equation for shipping point prices, and then use SUR to estimate the corresponding parameters; in this way, for example, each transmission mechanism in the North-West is affected by the behavior of the other regional markets.

As for the structure of the seven equations of the system, Willett et al. propose a modified version of Ward's (1982) model, where the dependent variable is no longer the actual price level but its first difference. That is, if we consider, for instance, the effect of wholesale on shipping point prices, Ward's specification (27) becomes:

$$
\begin{aligned}
sp_t - sp_{t-1} = \phi + \gamma_1 \sum_{j=0}^{3} \Delta ws_{t-j} D_j + \gamma_2 \sum_{j=0}^{3} \Delta ws_{t-j} \xi_j D_j + \gamma_3 \sum_{j=0}^{3} (C\Delta ws^-_{t-j} \\
- C\Delta ws^-_{t-j-1}) D_j + \gamma_4 \sum_{j=0}^{3} (C\Delta ws^-_{t-j} - C\Delta ws^-_{t-j-1}) \xi_j D_j + \epsilon_t.
\end{aligned}
\tag{77}
$$

Note that $C\Delta ws^-_{t-j} - C\Delta ws^-_{t-j-1} = \Delta ws^-_{t-j}$, and hence the two relations

$$
\begin{aligned}
\pi^+_{jlt} &= \gamma_0 + \gamma_1 \xi_j \\
\pi^-_{jlt} - \pi^+_{jlt} &= \gamma_3 + \gamma_4 \xi_j
\end{aligned}
$$

define the impact of a period-to-period variation in wholesale prices. Consequently, this model can be considered a vector extension of Tweeten and Quance's model, rather than of Wolframm's, as it allows to evaluate the effect of single price variations.

Using monthly data, Willett et al. find that shipping point prices respond symmetrically to retail and wholesale prices in the three regions. Wholesale and retail prices, instead, exhibit evidence of asymmetric behavior. In particular, in the West, the contemporaneous impact of retail prices asymmetrically affects wholesale prices. In the North Central Region, COIA and CUIA are observed in the transmission from shipping point prices to retail prices, while DLEA, instead, characterizes the shipping point-wholesale price relationship. In the NorthEast, wholesale prices respond asymmetrically to contemporaneous variations of both retail and shipping point prices. Finally, CUIA is present in the response of wholesale prices to shipping point prices.

A different approach is suggested by Miller and Hayenga (2001), who combine the investigation of price transmission with spectral analysis, in order to highlight whether asymmetries are linked to the frequency of price cycles. Their study focuses on farm, retail and wholesale prices of pork in the UK and it is based on weekly data from 1981 to 1995. Spectral analysis allows to convert a time series from time domain to frequency domain. In particular, since a stationary and invertible process is uniquely determined by its autocovariance function γ_k, where k is the lag, this function is translated, through the Fourier transformation, into a power spectral density function $hy(\omega)$, which allows to decompose the total variance into contributions from short- medium- and long-run frequencies:

$$hy(\omega) = \frac{1}{2\pi}(\gamma_0 + 2\sum_{k=1}^{\infty}\gamma_k cos(\omega k)) \tag{78}$$

where ω is the frequency varying between $[0, \pi]$.

After applying transformation (78), the authors select four frequency bands; [4] then for each band the VAR model (75) is estimated, after being appropriately adapted to the frequency domain.[5]

Results show that retail prices exhibit DLA in the response to low-frequency cycles in wholesale price, while farm-wholesale margins are affected by DLA at all frequencies. These results are even more interesting if we consider that, when testing the same hypothesis in the time domain, the authors are never able to reject the null of symmetry for the retail-wholesale margin. As the authors assert, this work show that "traditional time domain methods can mask underlying asymmetries that occur in subsets of the frequency domain" (p. 561). However, it is worth noting that these results can also be used to show that spectral analysis leads to over-reject the null of symmetry.

Another interesting example is provided by Shepherd (2004) who analyzes the world coffee market. This author uses a multivariate model to explain the relationship between producer and world prices in six exporting regions, namely Brazil, Colombia, Guatemala, India, Mexico and Uganda, together with the relationship between world and retail prices in US and Germany, which are consumer countries. The aim of this work is to assess whether there are any differences in the transmission mechanism of coffee prices prior and after the liberalization process, which took place around the year 1989.

If we indicate with pw the world price and with pr the producer price, the asymmetric VAR model used by Shepherd can be specified as follows:

$$\begin{cases} \Delta pw_t = \varphi^w + \phi_{w89}D_{89t} + \sum_{j=1}^{k}\phi_{wjw}pw_{t-j} + \sum_{j=1}^{k}\phi_{wjr}^{+}pr_{t-j}^{+} + \sum_{j=1}^{k}\phi_{wjr}^{-}pr_{t-j}^{-} + \epsilon_t^w \\ \Delta pr_t = \varphi^r + \Phi_{r89}D_{89t} + \sum_{j=1}^{k}\phi_{rjw}^{+}pw_{t-j}^{+} + \sum_{j=1}^{k}\phi_{rjw}^{-}pw_{t-j}^{-} + \sum_{j=1}^{k}\Phi_{rjr}pr_{t-j} + \epsilon_t^r \end{cases} \tag{79}$$

[4] $\omega \in [0, 0.2], [0.1, 0.3], [0.2, 0.4], [0.3, 0.5]$, which correspond to the week intervals $[0, 5], [3.33, 10], [2.5, 5], [2, 3.33]$ in the time domain. [5] A generic VAR $Y = X\Phi + \epsilon$ is transformed into $ZY = ZX\Phi + Z\epsilon$, where Z is the Fourier transform matrix. For details see Miller and Hayenga (2001).

where D_{89t} is an impulse dummy which is equal to one at the time of the liberalization and zero elsewhere. Using monthly data from January 1982 to December 2001, the author finds out that, before liberalisation, all countries, except Guatemala, were affected by DLEA in the transmission from the world to the local price; furthermore, in US, DLEA occurred both in the transmission from world to retail price and in the reverse direction.

After liberalisation, asymmetries disappeard in India and Mexico, but, in the remaining regions, DLEA characterizes both directions of the price transmission.

As anticipated in Section 4.2, Radchenko (2005a) uses a VAR model to test for RTA. The author studies the relationship between crude and retail gasoline prices in US, during the period March 1991 - February 2003, using a restricted form of equation (75), in which only crude price variations are divided into positive and negative. In accordance with the results derived from the EC specification, evidence of RTA is found.

The asymmetries supported by ARDL models can also be tested in a multivariate framework, as proposed by Radchenko and Tsurumi (2004). The authors present a model where the transmission mechanism between the crude price and the retail gasoline price is estimated together with gasoline consumption (G) per vehicle (S), inventory level (I) and production (Q):

$$\begin{cases} \Delta rt_t = \varphi_{11} + \phi_{12}^{+}\Delta cr_t^{+} + \phi_{12}^{-}\Delta cr_t^{-} + \phi_{13}\Delta I_t + \phi_{14}\Delta D_{st} + \phi_{15}\Delta D_{wt} + \epsilon_{1t} \\ \frac{G_t}{S_t} = \varphi_{21} + \phi_{22} rt_t + \phi_{23} Z_t + \phi_{23} D_{su} + \phi_{24} D_{wi} + \epsilon_{2t} \\ I_t = \varphi_{31} + \phi_{32} I_{t-1} + \phi_{33} D_{st} + \phi_{34} D_{wt} + \epsilon_{3t} \\ Q_t = G_t + I_t - I_{t-1} \end{cases} \tag{80}$$

where Z indicates income and D_s, D_w are seasonal dummy variables for summer and winter, respectively. Radchencko and Tsurumi apply a Bayesian Markov Chain Monte Carlo algorithm to estimate model (80), using the prior distribution defined in Zellner (1988), Tsurumi (1985) and Dreze (1976), among others. Monthly data for the US market, over the period January 1976 - December 1997, show that the contemporaneous impact of crude gasoline prices on retail prices is symmetric.

5.2.2 VEC Models

The second group of vector models, which have been widely used for the analysis of price transmission asymmetries, are the Vector Error Correction (VECM).

The general form of a VECM is:

$$\Delta Y_t = \Pi Y_{t-1} + \sum_{i=1}^{q} \Phi_i \Delta Y_{t-i} + \epsilon_t \tag{81}$$

where Y is a $n \times 1$ vector of observed variables, Γ and Π are two $n \times n$ matrices of coefficients. In particular, if the number of cointegrating vectors is equal to r, the long-run matrix Π can be decomposed into:

$$\Pi = \alpha_{n \times r} \beta'_{r \times n} \tag{82}$$

where β is the matrix of the r cointegrating vectors.

This model can be generalized to the asymmetric case along the same lines as the ECM described above.

In 1992, Kirchgassner and Kubler use a VECM to analyze the price transmission between German wholesale and retail prices of gasoline and light heating oil and the corresponding spot prices formed in the Rotterdam market.

The specification proposed is a vector extension of Stock and Watson's model, as the long-run equilibrium and the adjustment process are contemporaneously estimated. Considering, for example, the gasoline market, the model works as follows:

$$\begin{cases} \Delta ws_t = \varphi_1 + \sum_{j=1}^{k} \phi_{11j} \Delta ws_{t-j} + \sum_{i=0}^{k} \phi_{12i}^{+} \Delta sp_{t-i}^{+} + \sum_{i=0}^{k} \phi_{13i}^{-} \Delta sp_{t-i}^{-} + (c_1 ws_{t-1} + c_2 sp_{t-1}) + \epsilon_{1t} \\ \Delta rt_t = \varphi_2 + \sum_{j=1}^{k} \phi_{21j} \Delta rt_{t-j} + \sum_{i=0}^{k} \phi_{22i}^{+} \Delta sp_{t-i}^{+} + \sum_{i=0}^{k} \phi_{23i}^{-} \Delta sp_{t-i}^{-} + (c_3 rt_{t-1} + c_4 sp_{t-1}) + \epsilon_{2t} \end{cases} \tag{83}$$

The authors use monthly data from January 1972 to December 1989 and seek to verify whether there are any asymmetries and if there are any differences in the relationship between those prices prior and after 1980, since the number of spot contracts during the 1980s was much larger than during the 1970s. Results show that, during the 1970s, gasoline retail and wholesale prices and heating oil consumer prices exhibited RTA, DLEA and/or COIA in response to spot prices, while the wholesale heating oil market was affected only by either DLEA or COIA. During the second half of the sample period no evidence of asymmetries is found.

Another example of asymmetric VECM is offered by Chavas and Mehta (2002). The authors study wholesale and retail prices of US butter using monthly data over the period January 1980 - August 2001 and propose a regime switching VECM with seasonal effects (D):

$$\Delta Y_t = \Psi_0 + \Psi_1 t + \Pi Y_{t-1} + \sum_{k=1}^{r} \Omega_k D_{ts} + \sum_{i=1}^{q} \Phi_i^{+} \Delta Y_{t-i}^{+} + \sum_{i=1}^{q} \Phi_i^{-} \Delta Y_{t-i}^{-} + \epsilon_t \tag{84}$$

A test on the hypothesis $\Phi_i^+ = \Phi_i^-$, $\forall i$, clearly shows asymmetry in the distributed lag effect of both prices on each other; furthermore, impulse response functions suggest the existence of RTA. Besides, Chavas and Mehta propose to investigate a new kind of asymmetry, which captures the effects of expectations on the price transmission mechanism. In particular, the authors suggest to estimate the covariance between retail (rt) and wholesale (ws) prices using the specification:

$$s_{wr} = \omega_0 + \omega_r E(\Delta rt_t) + \omega_w E(\Delta ws_t) + \epsilon_t \tag{85}$$

Asymmetry occurs when $\omega_r \neq \omega_w$, that is when the covariance between ws and rt has an asymmetric response to expected price increases or decreases.

As ω_r results to be positive and ω_w negative, the authors conclude that "the contemporaneous linkage between retail and wholesale prices become weaker (stronger) when the wholesale (retail) price is expected to increase" (p. 11).

It is worth noting, however, that the asymmetry introduced by the authors does not provide any information about COIA and, consequently, it does not fit into the classification proposed in this paper.

In 2002 Gomez and Koerner published another interesting paper, where the transmission between international and retail coffee prices in US, France and Germany is studied using Kirchgassner and Kubler's 1992 model. The authors analyze the mutual relationship between the two prices, by taking into account the existence of an asymmetric effect of both the exchange rate on the local price and of the monthly rain precipitations on the international price.

Focusing attention on prices alone, monthly data over the period January 1990 to December 2000 show that in all countries both prices exhibit DLA with respect to their own lagged values and to the lagged value of the other price, in response to which they also display COIA.

5.2.3 VRS Models

Finally, VRSM are used to account for the existence of multiple regimes in a VAR or a VEC specification. A very interesting example is provided by Aguero (2003), who use a Vector Error Correction Regime Switching Model (VECRSM) to study the transmission between wholesale and retail prices of rice, tomatoes and potatoes in Peru from January 1995 to July 2001. For each product Aguero proposes the following model to test for REAPS:

$$\begin{cases} \Delta Y_t = \alpha^1 \beta' Y_{t-1} + \sum_{i=1}^q \Phi_i^1 \Delta Y_{t-i} + \epsilon_t & \text{if } \beta' Y_{t-1} < \delta \\ \Delta Y_t = \alpha^2 \beta' Y_{t-1} + \sum_{i=1}^q \Phi_i^2 \Delta Y_{t-i} + \epsilon_t & \text{if } \beta' Y_{t-1} \geq \delta \end{cases} \tag{86}$$

Model (86) is estimated using both Enders and Granger's (1998) and Hansen and Seo's (2002) approaches. As we have seen in Section (4.2.3), Enders and Granger's TAR and MTAR models allows to test for cointegration without imposing symmetry, by means of a two step procedure where the long-run equilibrium is estimated using a two-regime model with a known threshold.

Conversely, Hansen and Seo supply a method to deal with a regime switching cointegration model when the cointegrating vector β and the threshold δ are unknown and should be estimated. The existence of different states of the world can then be tested against the hypothesis of linear cointegration using an LM test, suitably modified to account for the fact that, under the null hypothesis, the value of δ is unknown.

Using daily data, the author finds out that the Enders and Granger's method is supportive of the null hypothesis of symmetry for the most perishable products (i.e tomatoes) but not for potatoes and rice, while the converse is true when Hansen and Seo's approach is used. Aguero has a preference towards the Hansen ans Seos's approach because this method does not to require a preliminary choice of the threshold value .

In 2003, Goodwin and Serra use a Threshold Vector Error Correction model (TVECM) to investigate how farm prices of raw milk are transmitted to the retail prices of blended cheese, cream caramel, pasteurized and sterilized milk in Spain. In particular, the authors propose a three-regime version of model (86) and estimate it using a sequential conditional iterated SUR in two steps, in order not to impose either a value on the threshold parameters or a cross-equation independence. Monthly observations over the period July 1994 - December 2000 highlight that only the sterilized milk model supports the presence of REAPA. RTA, instead, is a relevant issue for all the products considered, as shown by the impulse response function provided by the authors.

A different threshold multivariate model is proposed by Meyer (2003). The author observes that a single-threshold model, as the one by Aguero (2003), can be easily tested for the existence of a threshold, but it excludes the possibility of a "band" of non-adjustment. What the author complains is that, in a model like (86), "even very small deviations from the long term equilibrium will always lead to an adjustment process" (p. 2). Indeed, given for example the existence of transaction costs, there may be values of the error correction term for which no adjustment occurs. To solve this problem, Meyer proposes the alternative specification:

$$\begin{cases} \Delta Y_t = \alpha^1 + \sum_{i=1}^{q} \Phi_i^1 \Delta Y_{t-i} + \lambda^1 ECT_{t-1} + \epsilon_{1t} & \text{if } |ECT_{t-1}| \leq \delta \\ \Delta Y_t = \alpha^2 + \sum_{i=1}^{q} \Phi_i^2 \Delta Y_{t-i} + \lambda^2 ECT_{t-1} + \epsilon_{2t} & \text{if } |ECT_{t-1}| > \delta \end{cases} \tag{87}$$

where the adjustment is different depending on whether deviations from the equilibrium take extreme or intermediate levels.

Meyer applies model (87) to investigate the mutual relationship between German and Dutch pig farm prices, from June 1989 to March 2001, on a weekly basis. Results confirm that no significant adjustment occurs for small deviations from the long-run value, thus implying the presence of REAPA.

Finally, Luoma et al. (2004) reexamine the issue of not imposing a threshold value when estimating a VRSM, which they solve using a Bayesian estimation procedure. They study the transmission between producer and consumer prices in the Finnish beef and pork markets, over the period January 1981 - May 2003. The model adopted is a three regimes specification with no lags:

$$\begin{cases} \Delta Y_t = \alpha^1(\beta' Y_{t-1} - \delta_2) + \epsilon_t & \text{if } \beta' Y_{t-1} \leq \delta_1 \\ \Delta Y_t = \alpha^2(\beta' Y_{t-1} - \delta_2) + \epsilon_t & \text{if } \delta_1 < \beta' Y_{t-1} \leq \delta_2 \\ \Delta Y_t = \alpha^1(\beta' Y_{t-1} - \delta_2) + \epsilon_t & \text{if } \beta' Y_{t-1} \geq \delta_2 \end{cases} \tag{88}$$

which is used to test for REAPS and RTS on a monthly basis. For both beef and pork, the estimated coefficients and the impulse response functions provide no evidence of asymmetry.

6 Asymmetric price behavior: should we really care?

The presence of asymmetries in the price transmission mechanism has been investigated during the last twenty years throughout a wide variety of countries and commodities, even though the United States and the gasoline market seem to be the areas of greatest concern (Tables 10 and 11).

The existing literature seems to suggest that the presence of asymmetry is more than a murmur. Among the 69 papers considered in this survey, which provide a total of 83 estimated models, only 11 models show no evidence of any kind of asymmetry (Table 13). But should we really care of asymmetric price transmission? Are asymmetries really pervasive or, alternatively, are they even more relevant than it has been found? Indeed there are a number of issues which might affect the results of symmetry tests.

First, researchers have generally not accounted for all the possible sources of asymmetry. Table 12 shows that a relevant number of studies concentrates on COIS, DLES, CUIS and RTS, while RES and REAPS have scarcely been considered so far.

Second, various models and approaches have been proposed and the issue of which one is the most reliable is still open. For instance, since ARDL and ECM have been the most popular frameworks to investigate price asymmetries (Table 13), the debate on which of the two is the best model has been the focus of many studies.

It is well known that ECMs are generalizations of the ARDL specifications, used to account for the presence of cointegration. What happens if ARDL model are used for cointegrated series and, conversely, what are the consequences of using an ECM with non-cointegrated data?

Von Cramon-Taubadel (1998) shows that some kind of asymmetric ARDL models are incompatible with cointegration because, should they empirically support the presence of asymmetries, then either the series are not cointegrated or the results are spurious.

In order to demonstrate this point, Von Cramon-Taubadel considers two $I(1)$ processes, x and y, and the model:

$$y_t - y_0 = \alpha_0 + \alpha^+ \sum_{i=0}^{t} \Delta x^+_{t-i} + \alpha^- \sum_{j=1}^{t} \Delta x^-_{t-j} + \epsilon_t \qquad (89)$$

which, since any variable x can always be expressed as the sum of its starting value and all its variations up to time t (that is $x_t = x_0 + \sum_{i=0}^{t} \Delta x^+_{t-i} + \sum_{j=0}^{t} \Delta x^-_{t-j}$), can be rewritten in the form:

$$y_t = y_0 + \alpha_0 - \alpha^- x_t + \alpha^- x_0 + (\alpha^+ - \alpha^-) \sum_{i=0}^{t} \Delta x^+_{t-i} + \epsilon_t \qquad (90)$$

where $\sum_{i=0}^{t} \Delta x^+_{t-i}$ is clearly $I(1)$ because x is $I(1)$.

As Von Cramon-Taubadel asserts, the estimation of this model can provide three different results. The first case is when $(\alpha^+ - \alpha^-) \neq 0$ and ϵ_t is stationary: then x, y and $\sum \Delta x$ are cointegrated, thus precluding cointegration between x and y.

The second case is when $(\alpha^+ - \alpha^-) \neq 0$ and ϵ_t is $I(1)$ or when $(\alpha^+ - \alpha^-) = 0$ and ϵ_t is $I(1)$: this indicates a spurious regression with nonstationary variables.

Finally $(\alpha^+ - \alpha^-)$ can be equal to zero and ϵ_t stationary: that means that the dynamic is symmetric and x e y are cointegrated.

It is then proved that such a specification is inconsistent with the presence of both asymmetries and cointegration between x and y.

In this survey, 21 contributions dealing with asymmetric ARDL models for price transmission have been considered. Of these, only five provide information about the presence of cointegration, namely the studies by Shin (1994), Mohanty et al. (1995), Balke et al. (1998), EIA (1999) and Aguiar and Santana (2002). Within the class of the multivariate ARDL models, only two out of four papers investigate the existence of a long-run relationship, namely Miller and Hayenga (2001) and Shepherd (2004).

It is worth noting that, in the light of what was shown by Von Cramon-Taubadel, some of the results obtained with ARDL and VAR which support asymmetric price behaviors might actually be spurious.

As the use of the error correction models is concerned, some authors argue that they can be applied even to stationary data. An interesting debate on this issue can be found in six articles published in the review

Political Analysis between 1992 and 1993 (i.e. Beck, 1992; Beck, 1993; Durr, 1993a, 1993b; Smith, 1993 and Williams, 1993). This issue has been recently revisited by Keele (2005), who shows that the empirical properties of the ECM are maintained even when stationary data are used. As described in Sections 3.1.2 and 4.2.2, Balke et al. (1998) use stationary weekly data for the period 1987-1996 to study the linkage between crude, spot, wholesale and retail gasoline prices, with both an ARDL and an ECM. The empirical results suggest that the ECM, estimated using Stock and Watson's (1993) approach, provides much more evidence of asymmetry than the ARDL specification. In order to select between the two models, a more general specification, which nests both of them, is estimated:

$$\begin{aligned} ws_t = \phi + \sum_{i=0}^{s} \alpha_i sp_{t-i} + \sum_{i=1}^{s} \beta_i ws_{t-i} + \sum_{i=0}^{s} \alpha_i^+ \widetilde{sp}_{t-i} + \sum_{i=1}^{n} \beta_i^+ \widetilde{ws}_{t-i} \\ + \sum_{i=0}^{s} \xi_i \widetilde{sp}_{t-i-1} + \sum_{i=1}^{s} \upsilon_i \widetilde{ws}_{t-i-1} + \epsilon_t \end{aligned} \tag{91}$$

Results show that the ECM fits the data better than the ARDL. However, Balke et al. do not answer the question whether the use of an ECM with stationary data leads to misleading results, but they surely point out that different models can provide different results for the same data.

Another ECM with stationary data is estimated by Von Cramon-Taubadel et al. (2003). As described in Section 4.2.1, the authors study the German market for chicken and lettuce using weekly data for wholesale and retail prices over the period May 1995 - December 2000.

Although wholesale and retail chicken prices result to be cointegrated, while wholesale and retail lettuce prices are stationary, Von Cramon-Taubadel et al., for sake of comparison, choose to estimate an ECM for both goods.

In this case, the ECM produces the same outcomes for both chicken and lettuce prices, thus preventing any conclusion about the empirical performance of ECM in presence of stationary data.

To summarize, the literature analyzed in this paper does not help to understand whether ECMs can be applied independently of the presence of cointegration, but rather it shows that modeling non-cointegrated data with either ECM or ARDL specifications can affect the results of symmetry tests.

A third topic concerns the existence of structural breaks in the data. In the preceding sections we have reviewed several studies which accounts for structural breaks, namely Kinnucan and Forker (1987), Punyawadee et al. (1991), Kirchgassner and Kubler (1992), Worth (2000), Berardi et al. (2000), Eckert (2002), Abdulai (2002) Salas (2002), Contin et al. (2004), Krivonos (2004), Shepherd (2004) and clearly all the regime switching specifications discussed in Section 5.

A key question when looking for asymmetries is about the consequences of ignoring the possible presence of structural breaks in the data.

First of all, if we consider an ECM, the presence of unobserved structural breaks may lead to over-reject the cointegration hypothesis, as shown, among others, by Mosconi (1997). This can be easily solved by suitably modifying the cointegration tests (for more details see again Mosconi (1997)). However, in an asymmetric context an additional problem may arise, that is whether the presence of structural breaks in the cointegrating relationship can produce the false impression of asymmetry, as discussed by Von Cramon-Taubadel and Meyer (2003).

A structural break implies that, given two I(1) series x and y, their long-run equilibrium is on two regimes:

$$\begin{cases} y_t = \beta^1 x_t + \epsilon_t & \text{if } t \leq t_0 \\ y_t = \beta^2 x_t + \epsilon_t & \text{if } t > t_0 \end{cases} \tag{92}$$

$t = 1, ..., t_0, ..., T$, which clearly means that x and y are not cointegrated, even though cointegration is present on each of the two subsamples $[1, t_0]$ and $[t_0, T]$.

Von Cramon-Taubadel and Meyer's work relies on a Monte Carlo experiment to simulate two price series linked by a long-run relationship, with a structural break, and a symmetric adjustment process. These data are then used to evaluate in what percentage the null hypothesis of no cointegration and symmetry is rejected by an ARDL, an ECM estimated with the Engle and Granger's approach, an ECM with TAR cointegration and an ECM with MTAR cointegration.

Results show that all models, with the exception of the MTAR tend to over-reject the null of symmetry. MTAR models seem to be less vulnerable to detecting spurious asymmetries. Whereas with the other approaches small structural breaks lead to a rapid increase in the number of incorrect rejections of the null hypothesis of symmetry, this increase is much more gradual for the MTAR approach; over the range of parameters values for which the MTAR approach leads to a high proportion of rejection of the null hypothesis of symmetry, there is also a high probability that the cointegrations tests will indicate the absence of cointegration.

A fourth open question which should be considered in greater detail is the direction of causality (Table 14).

A typical maintained hypothesis is that price formation goes from upstream to downstream, or equivalently, input prices cause output prices; however, this causality direction is not necessarily true.

Adrangi et al. (2001), for instance, refer to the derived demand theory according to which prices should move from the downstream to the upstream market, as the price of an input should be formed by its contribution to the value of the final good.

In the considered literature, three tests have been used to check the direction of causality: Granger (1969) Sims (1972) and the variance decomposition as described by Sims (1980).

Granger's test relies on the idea that, given two series x, y and a set of relevant information z, which contains x, x causes y ($x \rightarrow y$) if the predictions of y based on z are more accurate than the predictions of y formed by exploiting the set of all relevant information x excluded.

Assuming, for example, that x and y are stationary [6], Granger proposes to estimate the following vector model:

$$\begin{aligned} x_t + b_0 y_t &= \sum_{i=1}^{m} a_i x_{t-i} + \sum_{j=1}^{m} b_j y_{t-j} + \epsilon_t \\ y_t + c_0 x_t &= \sum_{i=1}^{m} c_i x_{t-i} + \sum_{j=1}^{m} d_j y_{t-j} + \eta_t \end{aligned} \tag{93}$$

and then to consider the coefficient vectors b and c. If y causes x, then at least one element of b must be different from zero; that is, if $b_j \neq 0$, $j = 1, ...; m$, we say that there is instant causality. In the same way, d provides information about the effects of x on y. When $y \rightarrow x$ and $y \rightarrow x$, then x and y reciprocally cause each others ($x \leftrightarrow y$), a situation that is generally called "'feedback effect"'.

According to Sims (1972), instead, $x \rightarrow y$ if y is not influenced by the future values of x, but only by its past:

$$y_t = \sum_{i=0}^{m} a_i x_{t-i} + \sum_{j=1}^{m} b_j x_{t+j} \tag{94}$$

Naturally, if x causes y, then each b_j must be not statistically significant.

Finally, the variance decomposition described by Sims (1980), is based on the assumption that if a variable y is optimally forecasted from its own lagged values, then all its forecast error variance should be accounted for by its own disturbances. Conversely, if the variance of the forecast error is somehow explained by shocks to a different variable x, than we can conclude that x does Granger-cause y. Sims proposes to estimate a VAR model and to orthogonalize its residuals in order to isolate the impact of each shock, and then to check which percentage of the variance of each variable is explained by shocks in the others.

Of the 69 papers dealing with price transmission that we have considered in this survey, 19 studies test for causality. Specifically, 18 rely on Granger's (1969) test, while Sims' (1972) and Sims' (1980) approaches are used only by Aguiar and Santana (2002), who also employ Granager's method, and by Balke et al. (1998). It is also important to notice that evidence of "'from downward- to upward"' transmission is found in 7

[6] The test can be generalized to the non-stationary case

cases, a result which goes in the direction of Adrangi's et al. (1991).

Besides the choice of the proper specification, results can also be affected by the type of data used. This additional issue comprises both frequency and aggregation of the data.
Bachmeier and Griffin (2003) and Von Cramon-Taubadel (1997), among others, argue that the use of low frequency data can affect the results of symmetry tests, though it is not clear whether low frequency data would cause to the false impression of symmetry or asymmetry.
Bachmeier and Griffin, using an error correction specification, show that the use of daily data eliminates most of the asymmetries supported by a weekly analysis. Conversely, Von Cramon-Taubadel (2003) asserts that weekly data can hide asymmetries which occur within a day.
As a general conclusion, higher frequency data are likely to provide a more detailed view of the price transmission process, hence more reliable results. However, this is not always the case, as shown by Bettendorf et al. (2003). The authors, in fact, try to analyze the Dutch gasoline market using daily data but, since they do not find any significant result, they eventually choose to use weekly data.
It is worth noting that literature has focused on weekly and monthly data and that the number of cases where symmetry is found does not seem to differ significantly between the two categories (Table 15).

Finally the debate on data aggregation is still open. Among the very many contribution, we mention Peltzman (2000) and Von Cramon-Taubadel et al. (2003), whose analysis provide contrasting evidence.
As described in Section (4.2.2), Peltzman shows that asymmetry is very likely to occur in many US markets. However when he tests for symmetry using data on a single supermarket chain, rather than average data, no evidence of asymmetry is found. Conversely, Von Cramon-Taubadel et al. find evidence of asymmetry only when disaggregated data are used.
These results point out that neither aggregated nor disaggregated data are likely to bias symmetry tests, rather they show that researchers should be very careful when trying to infer from the behavior of a single agent the behavior of an economic system and viceversa.

7 Conclusions

Consumers often complain that retail prices rise more when input prices are rising than they decrease when costs are falling. In response to this sentiment, a wide range of empirical works have tried to clarify whether or not asymmetries do occur, proposing various definitions of asymmetries and using different econometric models.

In this paper, we have classified the existing empirical literature on price transmission according to the model selected and to the asymmetries tested. In particular, we have proposed an exhaustive classification of asymmetries into eight categories, namely contemporaneous impact, distributed lag effect, cumulated impact, reaction time, equilibrium and momentum equilibrium adjustment path, regime effect and regime equilibrium adjustment path. Moreover, this paper has evaluated the relative merits of the most popular econometric models for price asymmetries, namely autoregressive distributed lags, partial adjustments, error correction models, regime switching and vector autoregressive models.

Many are the issues which might affect the results presented by the empirical contributions on asymmetric price behavior surveyed in this paper. However, the literature we have critically discussed suggests that asymmetry, in all its forms, is very likely to occur in a wide range of markets. Among the 69 papers considered in this survey, which provide a total of 83 estimated models, only 11 models show no evidence of asymmetries of any kind.

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Table 1: List of variables used in the empirical literature

cr	Crude oil price	D	Dummy variable
dp	Domestic producer price	er	Exchange rate
fb	FOB price	fm	Farm price
G	Gasoline consumption	I	Inventory level
LL	Local Lettuce level	mk	Marketing costs
pr	Producer price	pw	World price
Q	Production	rf	Ex-refinery petroleum products price
rt	Retail price	S	Stock of vehicles
sp	Spot price	st	Stock level
tx	Taxes	ur	Utilization rate
wp	World import price	ws	Wholesale price
x	Generic variable	y	Generic variable
Z	Income		

Table 2: List of countries analyzed in the empirical literature

A	Argentina	Au	Australia	B	Brazil	C	Canada
D	Denmark	E	European Union	Fi	Finland	F	France
G	Germany	Gn	Ghana	Ir	Ireland	I	Italy
j	Japan	N	Netherlands	Pe	Peru	Ph	Philippines
Sp	Spain	S	Switzerland	SA	South Africa	Sw	Sweden
UK	United Kingdom	US	United States	World	Worldwide		

Table 3: List of journals and reviews

Ag	Agribusiness
AE	Applied Economics
AgE	Agricultural Economics
AJAE	American Journal of Agricultural Economics
CJAE	Canadian Journal of Agricultural Economics
CJE	Canadian Journal of Economics
ERAE	European Review of Agricultural Economics
EE	Energy Economics
EP	Energy Policy
IJER	International Journal of Energy Research
JA	Journal of Agribusiness
JDE	Journal of Development Economics
JPE	The Journal of Political Economy
OR	OPEC Review
PRE	Philippine Review of Economics
QJE	The Quarterly Journal of Economics
RES	The Review of Economics and Statistics
RIO	Review of Industrial Organization
SJE	Scandinavian Journal of Economics
WP	Working Paper

Table 4: Definition of asymmetries (symmetries)

COIA (COIS)	Contemporaneous impact
DLEA (DLES)	Distributed lag effect
CUIA (COIS)	Cumulated impact
RTA (RTS)	Reaction time
EAPA (EAPS)	Equilibrium adjustment path
MEAPA(MEAPS)	Momentum equilibrium adjustment path
REA (RES)	Regime effect
REAPA (RTS)	Regime equilibrium adjustment path

Table 5: Econometric models of asymmetric price transmission

ARDLpp	Autoregressive Distributed Lag Model based on period-to-period price variations
ARDLcu	Autoregressive Distributed Lag Model based on cumulative price variations
ECMeg	Error Correction Model estimated with Engle and Granger's method
ECMsw	Error Correction Model estimated with Stock and Watson's method
ECMth	Error Correction Model with threshold cointegration
PAM	Partial Adjustment Model
RSM	Regime Switching Model
DRS	Deterministic Regime Switching Model
SRS	Stochastic Regime Switching Model
VECM	Vector Error Correction Model
VAR	Vector Autoregressive Model
VARcu	Vector Autoregressive Model based on cumulative price variations
VRS	Vector Regime Switching Model
VRSeg	Vector Regime Switching Model estimated with Engle and Granger's method
VRShs	Vector Regime Switching Model estimated with Hansen and Seo's method

Table 6a: Summary of asymmetries by model - ARDL

Year	Authors	Journal	Model	Asymmetries					Country	Product	Sample	Data Frequency
				COI	DLE	CUI	RT	EAP				
1982	Ward	AJAE	ARDLwa	Y*	Y*	N*	-	-	USA	Agricultural	-	Monthly
1987	Kinnucan,Forker	AJAE	ARDLwa	-	Y	Y*	-	-	USA	Alimentary	1971-1981	Monthly
1991	Karrenbrock	WP	ARDLpp	Y*	Y*	N*	-	-	USA	Gasoline	1983-1990	Monthly
1991	Punnyawadee et al.	CJAE	ARDLpp	N*	N*	N*	-	-	C	Alimentary	1965-1989	Weekly
1991	Punnyawadee et al.	CJAE	ARDLpp	Y	Y	N*	-	-	C	Alimentary	1965-1969	Weekly
1993	Balabanoff	OR	ARDLpp	-	Y*	N*	-	-	F,I,J,UK	Gasoline	1985-1992	Monthly
1993	Balabanoff	OR	ARDLpp	Y*	Y*	N*	-	-	G,US	Gasoline	1985-1992	Monthly
1993	GAO	WP	ARDLpp	N*	N*	N*	-	-	USA	Gasoline	1984-1991	Weekly
1993	GAO	WP	ARDLpp	Y*	N*	N*	-	-	USA	Gasoline[a]	1984-1991	Weekly
1993	GAO	WP	ARDLpp	N*	N*	N*	-	-	USA	Gasoline[b]	1984-1991	Weekly
1994	Griffith,Piggott	AgE	ARDLwa	Y*	Y*	N*	-	-	USA	Alimentary[c]	1971-1988	Monthly
1994	Griffith,Piggott	AgE	ARDLwa	Y*	Y*	Y*	-	-	USA	Alimentary[d]	1971-1988	Monthly
1994	Shin	OR	ARDLpp	N*	-	-	-	-	USA	Gasoline	1986-1992	Monthly
1995	Mohanty et al.	CJAE	ARDLwa	Y	Y	Y*	-	-	C,A,Au,E	Agricultural	1980-1990	Monthly
1995	Powers	Ag	ARDLwa	-	Y*	Y*	Y*	-	USA	Agricultural[e]	1986-1992	Weekly
1995	Powers	Ag	ARDLwa	-	Y*	Y*	N*	-	USA	Agricultural[f]	1986-1992	Weekly
1995	Powers	Ag	ARDLwa	-	-	N*	N*	-	USA	Agricultural[g]	1986-1992	Weekly
1995	Zhang et al.	Ag	ARDLcu	-	Y*	N*	-	-	USA	Agricultural	1984-1992	Monthly
1996	Duffy-Deno	EE	ARDLpp	Y*	Y*	Y*	-	-	USA	Gasoline	1989-1993	Weekly
1996	Duffy-Deno	EE	ARDLpp	Y*	Y*	N*	-	-	USA	Gasoline[h]	1989-1993	Weekly
1998	Balke et al.	WP	ARDLpp	Y	Y	-	-	-	USA	Gasoline	1987-1996	Weekly
1999	EIA	WP	ARDLpp	Y*	Y*	-	-	-	USA	Gasoline	1989-1993	Weekly
2000	Worth	WP	ARDLcu	-	Y*	Y*	-	-	USA	Agricultural[i]	1980-1999	Monthly
2000	Worth	WP	ARDLcu	-	Y*	N*	-	-	USA	Agricultural[j]	1980-1999	Monthly
2001	Parrott	JA	ARDLcu	N*	N*	N*	-	-	USA	Agricultural	1988-1993	Weekly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] Crude wholesale relationship during a market shock [b] Wholesale retail relationship during a market shock [c] Pork [d] Beef and Lamb [e] Wholesale-retail
[f] FOB-retail [g] FOB-wholesale [h] During a market shock [i] Carrots,tomatoes [j] Celery,lettuce,onions,potatoes

Table 6b: Summary of asymmetries by model - ARDL

Year	**Authors**	**Journal**	**Model**	**Asymmetries**					**Country**	**Product**	**Sample**	**Data Frequency**
				COI	DLE	CUI	RT	EAP				
2002	Aguiar,Santana	Ag	ARDLpp	-	Y*	Y*	-	-	B	Agricultural[a]	1987-1998	Monthly
2002	Aguiar,Santana	Ag	ARDLpp	Y*	Y*	Y*	-	-	B	Agricultural[b]	1987-1998	Monthly
2002	Aguiar,Santana	Ag	ARDLpp	Y*	Y*	N*	-	-	B	Agricultural[c]	1987-1998	Monthly
2002	Aguiar,Santana	Ag	ARDLpp	-	Y*	N*	-	-	B	Agricultural[d]	1987-1998	Monthly
2002	Aguiar,Santana	Ag	ARDLpp	Y*	N*	Y*	-	-	B	Agricultural[e]	1987-1998	Monthly
2003	LondonEconomics	WP	ARDLpp	Y	Y	-	-	-	Au,D,F,G,Ir,N,Sp,UK	Alimentary	1985-2003	Monthly
2003	Bunte,Zachariasse	WP	ARDLpp	-	-	-	N*	-	N	Alimentary[f]	1990-1997	-
2003	Bunte,Zachariasse	WP	ARDLpp	-	-	-	Y*	-	N	Alimentary[g]	1990-1997	-
2004	Girapunthog et al.	WP	ARDLcu	N*	N*	N*	-	-	USA	Agricultural[h]	1975-1998	Monthly
2004	Girapunthog et al.	WP	ARDLcu	Y*	Y*	Y*	-	-	USA	Agricultural[i]	1975-1998	Monthly
2004	Girapunthog et al.	WP	ARDLcu	Y*	Y*	N*	-	-	USA	Agricultural[j]	1975-1998	Monthly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] Beans,tomatoes [b] Milk [c] Rice [d] Onions [e] Coffee [f] Farm retail transmission for beef, poultry and chips; farm wholesale transmission for pork, beef, potatoes and chips [g] Farm retail transmission for pork and potatoes; farm wholesale transmission for poultry [h] Producer-Retail [i] Producer-Wholesale [j] Wholesale-Retail

Table 7a: Summary of asymmetries by model - PAM and ECM

Year	Authors	Journal	Model	Asymmetries						Country	Product	Sample	Data Frequency
				COI	DLE	CUI	RT	EAP	MEAP				
1991	Bacon	EE	PAM	-	Y*	-	-	Y*		UK	Gasoline	1982-1989	Fortnightly
1991	Manning	AE	ECMeg	-	Y*	-	N*	-	-	UK	Gasoline	1973-1988	Monthly
1991	Manning	AE	ECMsw	-	Y*	-	N*	-	-	UK	Gasoline	1973-1988	Monthly
1991	Norman,Shin	WP	PAM	-	-	-	-	N*	-	US	Gasoline	1984-1992	Weekly
1994	Shin	OR	PAM	-	-	-	-	N*	-	US	Gasoline	1986-1992	Monthly
1996	Borenstein,Shepard	WP	ECMsw	-	-	-	Y*	-	-	USA	Gasoline	1982-1991	Monthly
1997	Arden et al.	WP	ECMeg	-	-	-	-	N*	-	UK	Manufacturing	1970-1996	Quarterly
1997	Arden et al.	WP	ECMsw	-	-	-	-	Y*	-	UK	Manufacturing	1970-1996	Quarterly
1997	Borenstain et al.	QJE	ECMsw	Y	-	-	Y*	-	-	USA	Gasoline[a]	1986-1992	(Bi/)Weekly
1997	Borenstain et al.	QJE	ECMsw	Y	-	-	N*	-		USA	Gasoline[b]	1986-1992	(Bi/)Weekly
1998	Balke et al.	WP	ECMsw	Y*	Y*	-	-	-	-	C	Gasoline	1987-1996	Weekly
1998	Eltony	IJER	ECMsw	Y*	-	-	-	-	-	UK,USA	Gasoline	1980-1996	Monthly
1998	Reilly,Witt	EE	ECMsw	Y*	-	-	-	-	-	UK	Gasoline	1982-1995	Monthly
1998	VonCramon-Taubadel	ERAE	ECMeg	-	-	-	Y*	Y*	-	G	Alimentary	1990-1993	Weekly
2000	Abdulai	JDE	ECMth	-	-	-	Y*	-	Y*	Gn	Agricultural	1980-1997	Monthly
2000	Asplund et al.	SJE	ECMeg	Y*	Y*	-	-	-	-	Sw	Gasoline	1980-1996	Monthly
2000	Berardi et al	WP	ECMeg	Y	Y	-	Y*	N*	-	I	Gasoline	1996-2000	Weekly
2000	Peltzman	JPE	ECMsw	-	-	Y*	-	-	-	USA	Various	1978-1996	Monthly
2001	Hassan,Simioni	WP	ECMth	-	-	-	-	Y*	-	F	Agricultural[c]	-	-
2001	Hassan,Simioni	WP	ECMth	-	-	-	Y*	-	Y*	F	Agricultural[d]	-	-
2002	Abdulai	AE	ECMth	-	-	-	Y*	-	Y*	S	Agricultural	1988-1997	Monthly
2002	Salas	PRE	PAM	-	-	-	-	Y*	-	Ph	Gasoline	1999-2002	Weekly
2002	Salas	PRE	ECMeg	-	-	Y*	-	-	-	Ph	Gasoline	1999-2002	Weekly
2002	Eckert	CJE	ECMsw	Y*	-	-	Y*	-	-	C	Gasoline	1989-1994	Weekly
2003	Bachmeier,Griffin	RES	ECMeg	N*	N*	N*	N*	N*	-	USA	Gasoline	1985-1998	Daily
2003	Bachmeier,Griffin	RES	ECMsw	-	-	-	Y*	-	-	USA	Gasoline	1985-1998	Weekly
2003	Bachmeier,Griffin	RES	ECMsw	-	-	-	Y*	-	-	USA	Gasoline	1985-1998	Daily
2003	Bettendorf et al.	EE	ECMeg	Y*	Y*	-	Y*	-	-	N	Gasoline[e]	1996-2001	Weekly
2003	Bettendorf et al.	EE	ECMeg	N*	N*	N*	N*	-	-	N	Gasoline[f]	1996-2001	Weekly
2003	Conforti et al.	WP	ECMeg	-	-	-	-	Y*	-	World	Agricultural	1989-2001	Monthly
2003	Conforti et al.	WP	ECMeg	N*	N*	N*	-	N*	-	World	Agricultural	1990-2001	Monthly
2003	Galeotti et al.	EE	ECMeg	Y*	-	-	N*	Y*	-	I,F,Sp,G,UK	Gasoline	1985-2000	Monthly
2003	Gonzales et al	WP	ECMth	-	-	-	N*	N*	N*	F	Alimentary	1988-1999	Monthly
2003	VonCramon-Taubadel,Meyer	WP	ECMeg	-	-	-	-	N*	-	G	Alimentary[g]	1995-2000	Weekly
2003	VonCramon-Taubadel,Meyer	WP	ECMeg	-	-	-	-	Y*	-	G	Alimentary[h]	1995-2000	Weekly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] $cr \rightarrow rt, cr \rightarrow sp, ws \rightarrow rt$ [b] $sp \rightarrow ws$ [c] Chicory [d] Tomatoes [e] Weekly data are referred to Monday, Thursday, Friday [f] Weekly data are referred to Tuesday,Wednesday [g] Aggregated Data [h] Non-Aggregated Data

Table 7b: Summary of asymmetries by model - PAM and ECM

Year	Authors	Journal	Model	Asymmetries						Country	Product	Sample	Data Frequency
				COI	DLE	CUI	RT	EAP	MEAP				
2003	LondonEconomics	WP	ECMeg	-	-	-	-	Y*	-	D,UK	Agricultural	1985-2003	Monthly
2003	NDA	WP	ECMeg	-	-	-	Y*	-	-	SA	Agricultural	2000-2003	Monthly
2004	Conforti	WP	ECMeg	-	-	-	-	Y*	-	World	Agricultural	1969-2001	Annual
2004	Contin et al.	WP	ECMsw	Y*	Y*	-	Y*	-	-	Sp	Gasoline	1993-2002	Weekly
2004	Krivonos	WP	ECMeg	N*	-	-	N	-	-	Africa	Agricultural	1984-1990	Monthly
2004	Krivonos	WP	ECMeg	Y*	-	-	Y	-	-	Africa	Agricultural	1990-2003	Monthly
2004	Verlinda	WP	ECMsw	-	-	-	Y*	-	-	USA	Gasoline	2002-2003	Weekly
2005	Grasso,Manera	WP	ECMeg	Y*	Y*	-	-	Y*	-	F,G,I,Sp,UK	Gasoline	1985-2003	Monthly
2005	Grasso,Manera	WP	ECMth	-	-	-	-	-	Y*	F,G,I,Sp,UK	Gasoline[a]	1985-2003	Monthly
2005	Grasso,Manera	WP	ECMth	-	-	-	-	-	N*	F,G,I,Sp,UK	Gasoline[b]	1985-2003	Monthly
2005	Kaufmann,Laskowski	EP	ECMeg	-	-	-	-	N*	-	USA	Gasoline[c]	1986-2002	Monthly
2005	Kaufmann,Laskowski	EP	ECMeg	-	-	-	-	Y*	-	USA	Gasoline[d]	1986-2002	Monthly
2005a	Radchenko	EE	ECMeg	-	-	-	Y*	-	-	USA	Gasoline	1991-2003	Weekly
2005b	Radchenko	EE	ECMeg	-	-	-	Y*	-	-	USA	Gasoline	1991-2003	Weekly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] Crude-retail transmission in F,G,I,Sp,UK, crude-spot relationship in I, Sp [b] Spot-retail transmission in F,G,I,Sp,UK, crude-spot relationship in F,G,Uk [c] Motor gasoline and heating oil crude-refinery transmission in US; motor gasoline refinery-retail transmission in US but California, Lousiana and Idaho [d] Heating oil refinery-retail transmission in US; motor gasoline refinery-retail transmission in California, Lousiana and Idaho

Table 8: Summary of asymmetries by model - RSM

Year	**Authors**	**Journal**	**Model**	**Asymmetries**							**Country**	**Product**	**Sample**	**Data Frequency**
				WithinRegime					**BetweenRegime**					
				COI	DLE	CUI	RT	EAP	REA	REAPA				
1995	Powers	Ag	DRS	-	-	Y*	-	-	-	-	USA	Agricultural[a]	1986-1992	Weekly
1999	Goodwin&Holt	AJAE	DRS	-	-	-	N*	-	-	Y*	USA	Alimentary	1981-1998	Weekly
2000	Godby et al.	EE	DRS	-	-	-	-	-	N*	-	C	Gasoline	1990-1996	Weekly
2001	Goodwin,Piggott	AJAE	DRS	-	-	-	N*	-	-	Y*	USA	Agricultural	1992-1999	Daily
2002	Johnson	RIO	DRS	-	-	-	Y*	-	Y*	-	USA	Gasoline	1996-1998	Weekly
2004	Lewis	WP	DRS	Y*	Y*	Y*	Y*	-	-	Y*	USA	Gasoline	2000-2001	Weekly
2005	Grasso,Manera	WP	DRS	-	-	-	-	-	Y*	-	F,G,I,Sp,UK	Gasoline	1985-2003	Monthly
2005	Grasso,Manera	WP	DRS	-	-	-	-	-	N*	-	F,G,I,Sp,UK	Gasoline[b]	1985-2003	Monthly
2005b	Radchenko	EE	SRS	-	-	-	Y*	-	-	-	USA	Gasoline	1991-2003	Weekly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] Wholesale-retail [b] Italy: crude-spot relationship

Table 9: Summary of asymmetries by model - VAR, VECM and VRSM

Year	Authors	Journal	Model	Asymmetries							Country	Produact	Sample	Data Frequency
				WithinRegime					BetweenRegime					
				COI	DLE	CUI	RT	EAP	REA	REAPA				
1992	Kirchgsser,Kbler	EE	VECM	Y	Y	-	Y*	-	-	-	G	Gasoline	1972-1980	Monthly
1992	Kirchgsser,Kbler	EE	VECM	N	N	-	N*	-	-	-	G	Gasoline	1980-1989	Monthly
1993	Capps	WP	VARcu	Y*	-	-	-	-	-	-	USA	Alimentary	1986-1988	Weekly
1997	Willett et al.	Ag	VAR	N*	N*	N*	-	-	-	-	USA	Agricultural[a]	1975-1991	Monthly
1997	Willett et al.	Ag	VAR	N*	Y*	N*	-	-	-	-	USA	Agricultural[b]	1975-1991	Monthly
1997	Willett et al.	Ag	VAR	Y*	N*	Y*	-	-	-	-	USA	Agricultural[c]	1975-1991	Monthly
1997	Willett et al.	Ag	VAR	Y*	N*	N*	-	-	-	-	USA	Agricultural[d]	1975-1991	Monthly
2001	Miller,Hayenga.	AJAE	VAR	-	Y*	-	-	-	-	-	UK	Alimentary	1981-1995	Weekly
2002	Chavas,Mehta	AJAE	VECM	-	Y*	-	Y*	-	-	-	USA	Alimentary	1980-2001	Monthly
2002	Gomez,Koerner	WP	VECM	Y*	Y*	-	-	-	-	-	USA,F,G	Alimentary	1990-2000	Monthly
2003	Aguero	WP	VRSeg	-	-	-	-	-	-	Y*	Pe	Agricultural[e]	1995-2001	Daily
2003	Aguero	WP	VRSeg	-	-	-	-	-	-	N*	Pe	Agricultural[f]	1995-2001	Daily
2003	Aguero	WP	VRShs	-	-	-	-	-	-	N*	Pe	Agricultural[g]	1995-2001	Daily
2003	Aguero	WP	VRShs	-	-	-	-	-	-	Y*	Pe	Agricultural[h]	1995-2001	Daily
2003	Goodwin,Serra	AE	VRS	-	-	-	Y*	-	-	N*	Sp	Alimentary[i]	1994-2000	Monthly
2003	Goodwin,Serra	AE	VRS	-	-	-	Y*	-	-	Y*	Sp	Alimentary[j]	1994-2000	Monthly
2003	Meyer	WP	VRS	-	-	-	-	-	-	Y*	G,N	Alimentary	1989-2001	Weekly
2004	Luoma et al.	WP	VRS	-	-	-	N*	-	-	N*	Fi	Alimentary[k]	1981-2003	Monthly
2004	Radchenko,Tsurumi	WP	VAR	N*	-	-	-	-	-	-	USA	Gasoline	1976-1997	Monthly
2004	Sheperd	WP	VAR	-	Y*	-	-	-	-	-	World	Agricultural	1982-2001	Monthly
2005a	Radchenko	EE	VAR	-	-	-	Y*	-	-	-	USA	Gasoline	1991-2003	Weekly

Notes: Y, N indicate the presence or absence of asymmetry, respectively; * denotes that the presence or absence of asymmetry is supported by a statistical test.

[a] $sp \rightarrow rt$ West,NorthEast; $sp \rightarrow ws$ West; $rt \rightarrow ws$ NorthCenter; $ws \rightarrow sp$, $rt \rightarrow sp$ West,NorthEast, NorthCenter [b] $sp \rightarrow ws$ NorthCenter [c] $sp \rightarrow rt$ NorthCenter,$sp \rightarrow ws$ NorthEast [d] $rt \rightarrow ws$ NorthEast, West [e] Potatoes, Rice [f] Tomatoes [g] Potatoes, Rice [h] Tomatoes [i] Blended cheese, cream caramel and pasteurized milk [j] Sterilized milk [k] Sterilized milk

Table 10: Number of studies which analyze a given country

Country	Number of studies
Argentina	1
Australia	1
Brazil	1
Canada	4
Denmark	1
European Union	1
Finland	1
France	6
Germany	7
Ghana	1
Ireland	1
Italy	4
Japan	1
Netherlands	3
Peru	1
Philippines	1
Spain	4
Switzerland	1
South Africa	1
Sweden	1
United Kingdom	9
United States	33
Worldwide countries	3

Table 11: Number of studies which analyze a given market

Market	Number of studies
Agricultural	18
Alimentary	16
Gasoline	33
Other	2

Table 12: Testing the null hypothesis of symmetry by type

Type of Asymmetry	Number of Tests	Frequency of no rejection
COIS	48	25%
DLES	54	25%
CUIS	39	59%
RTS	42	35%
EAPS	20	40%
MEAPS	6	33%
RES	4	50%
REAPS	11	36%

Table 13: Percentage of studies which do not support any kind of asymmetry by model

Model	Number of studies	Frequency of no rejection
ARDLpp	11	9%
ARDLcu	10	10%
ECMeg	20	15%
ECMsw	11	0%
ECMtar	5	20%
PAM	4	50%
RSM	8	25%
VAR	6	17%
VECM	3	0%
VRSM	5	20%
Total	83	13%

Table 14: Causality tests between input and output prices

Test	Number of studies	Percentage
Granger (1969)	17	24.5%
Granger (1969), Sims (1972)	1	1.5%
Sims (1980)	1	1.5%
No test	50	72.5%
Total	69	100%

Table 15: Percentage of studies which do not support any kind of asymmetry by data frequency

Data frequency	Number of studies	Frequency of no rejection
Daily	3	67%
Weekly	29	10%
Beweekly	1	0%
Monthly	40	15%
Quarterly	2	50%
Annual	1	0%

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