MARKETING AND PRICING DYNAMICS IN THE PRESENCE OF STRUCTURAL BREAKS - THE HUNGARIAN PORK MARKET

LAJOS ZOLTÁN BAKUCS[1] , IMRE FERTŐ[2]

Paper prepared for presentation at the 98th EAAE Seminar, June 29th – July 2nd, Chania, Greece

Research for this paper has been supported by a grant from the GDN and CERGE-EI foundation (grant no: RRC V-23), 'Price Transmission on the Hungarian Agri-Food Markets'. We thank CERGE-EI seminar participants for their useful comments.

ABSTRACT

The study of marketing margins and price transmission on various commodity markets has been a popular research topic of the past decades (see MEYER, VON CRAMON-TAUBADEL, 2004, for a recent survey), however with a few exceptions these studies focused on developed economies. In this paper we examine the above phenomena on the: Hungarian pork market. The Johansen (maximum likelihood) or Engle and Granger (two step) cointegration tests do not reject the no-cointegration null hypothesis between the Hungarian pork producer and retail price series. Therefore we apply the Gregory and Hansen procedure with recursively estimated breakpoints and ADF statistics, and found that the prices are cointegrated with a structural break occurring in April 1996. Exogeneity tests reveal the causality running from producer to retail prices both on long and short run. Homogeneity tests are rejected, suggesting a mark-up pricing strategy. Price transmission modelling suggests that, price transmission on the Hungarian pork meat market is symmetric on the long, but asymmetric on the short-run, i.e. processors, wholesalers or retailers might take temporary advantage should price changes occur.

Keywords: price transmission, marketing margin, pricing, structural breaks, Hungarian pork market

[1] Junior Research Fellow, Institute of Economics, Hungarian Academy of Sciences, Budapest (Hungary). bakucs@econ.core.hu

[2] Senior Research Fellow, Institute of Economics, Hungarian Academy of Science,, Budapest (Hungary). ferto@econ.core.hu

1 INTRODUCTION

Measuring the spread in vertical price relationships and analysing the nature of price transmission along the supply chain from the producer to consumer have evolved as widely used methods to gain insight into the functioning of, and degree of competition in food markets. Asymmetric price transmission has been studied by numerous authors using different econometric methods, from the classical WOLFFRAM (1971), and HOUCK (1977), specification to cointegration (VON CRAMON-TAUBADEL, 1998) and threshold autoregressive models (e.g. GOODWIN, HARPER, 2000). However none of these studies (except BOJNEC 2002, BAKUCS, FERTŐ 2005) focus on a transition economy. Because of the inherited pre-1989 distorted markets, low developed price-discovery mechanisms and often ad-hoc policy interventions, transitional economies could be expected to have generally larger marketing margins and more pronounced price transmission asymmetries.

The aim of this paper is to investigate the dynamics of the marketing margin on the Hungarian pork meat market. The paper is organised as follows. Section 2 reviews some of the theoretical literature concerning marketing margins and price transmission, while section 3 describes the empirical procedures we apply. Our data and results are reported and discussed in section 4, with a summary and some conclusions presented in section 5.

2 MARKETING MARGIN AND PRICE TRANSMISSION

2.1 Theoretical background

The marketing margin is the difference between the retail and the producer or farm gate price. It represents marketing costs such as transport, storage, processing, wholesaling, retailing, advertising, etc.:

$$RP = FP + M \tag{1}$$

M, the marketing margin, is composed of an absolute amount and a percentage or mark-up of the retail price:

$$M = a + bRP \text{, where } a \geq 0 \text{ and } 0 \leq b < 1. \tag{2}$$

With the use of logarithmic data, the long-run elasticity between the prices is readily available from the marketing margin model. If prices are determined at producer level, we use the mark-up model:

$$RP = \alpha_1 + \varepsilon_{FP} FP , \tag{3}$$

where ε_{FP} is the price transmission elasticity from the producer price (FP) towards the consumer price (RP). If $\varepsilon_{FP} = 1$, we have perfect transmission, and thus the mark - up will be ($e^{\alpha 1}$ - 1). $0 < \varepsilon_{FP} < 1$ implies that the transmission between the two prices is not perfect.

If however, prices are determined on consumer level, than the use of the mark-down model is appropriate:

$$FP = \alpha_2 + \varepsilon_{RP} RP, \tag{4}$$

where ε_{RP} is the elasticity of transmission between the consumer price (RP) and the producer price (FP). As before, there is perfect transmission, if $\varepsilon_{RP} = 1$, and the mark - down equals (1 - $e^{\alpha 2}$). Imperfect transmission results if $\varepsilon_{RP} > 1$.

A common perception is that responses to price increases differ from responses to price decreases. More exactly, retailers tend to pass more rapidly price increases to consumers, whilst it takes longer for consumer prices to adjust to producer prices if the latter decrease. There are several major explanations for the existence of price asymmetries. First, asymmetrical price transmission occurs when firms can take advantage of quickly changing prices. This is explained by the theory of the *search costs* (MILLER, HAYENGA, 2001). They occur in locally imperfect markets, where retailers can exercise their local market power. Although customers would have a finite number of choices, they might face difficulties in quickly gathering information about the pricing of the competing stores because of the search costs. Thus firms can quickly raise the retail price as the producer price rises, and reduce much slower retail prices when upstream prices decline. Second comes the problem of *perishable goods* (WARD, 1982), that withholds retailers from raising prices as producer prices rise. Wholesalers and retailers in possession of perishable goods may resist the temptation to increase the prices because they risk a lower demand and ultimately being left with the spoiled product. Third, the *adjustment costs or menu costs* (GOODWIN, HOLT, 1999) may underlie asymmetric price adjustments. Menu costs involve all the cost occurring with the re-pricing and the adoption of a new pricing strategy. As with perishable goods, menu costs also act against retailers changing prices. Finally, the exercise of *oligopoly power* can favour asymmetric price transmission. It appears in markets with highly inelastic demand and concentrated supply; many food chains have such market organisation characteristics. It also needs to be mentioned that such collusive behaviour is rather difficult to maintain in long run, because of the incentive for one firm to cheat the others (MILLER, HAYENGA, 2001, pp. 554).

2.2 Empirical evidence

There are a great number of empirical studies dealing with marketing margin and asymmetry problems in livestock markets. VON CRAMON-TAUBADEL (1998) finds asymmetrical price transmission on the German pork market. DAWSON, TIFFIN (2000) identify a long-run price relationship between UK lamb farm-retail prices, and study the seasonal and structural break properties of the series, concluding that the direction of Granger causality is from the retail to producer prices; thus lamb prices are set in the retail market. Threshold Autoregressive Models were developed by GOODWIN, HOLT (1999), GOODWIN, HARPER (2000) and BEN-KAABIA, GIL, BOSHNJAKU (2002) studying the US beef sector, US pork sector and Spanish lamb sector, respectively. GOODWIN, HOLT (1999) find that farm markets do adjust to wholesale market shocks, whilst the effect of the retail market shocks are largely confined to retail markets. GOODWIN, HARPER (2000) in their pork market study find a unidirectional price information flow from farm to wholesale and retail levels. Farm markets adjust to wholesale market shocks, but retail level shocks are not passed on to wholesale or farm levels. BEN-KAABIA, GIL, BOSHNJAKU (2002) establish a symmetric price transmission, concluding a long-run perfect price transmission, where any supply or demand shocks are fully transmitted through the system. They also observe that an increased horizontal concentration allows retailers to exercise market power.

ABDULAI (2002) uses a Momentum-Threshold Autoregressive Model (M-TAR) when studying the price transmission on the Swiss pork market. He also concludes that price transmission between producer and retailer market levels is asymmetric, i.e. increases in producer prices that would diminish the marketing margin are passed on more quickly than producer price decreases that widen marketing margins. MILLER, HAYENGA (2001) study the US pork market price transmission in conjunction with price cycles, concluding that wholesale prices adjust asymmetrically to changes in farm prices in all cycle frequencies. BOJNEC (2002) finds that both the Slovenian farm-gate beef and pork markets are weakly exogenous in the long run, with a mark-up long-run price strategy for beef and a competitive price strategy for the pork meat market. REZITIS (2003) applies a Generalised Autoregressive Conditional Heteroscedastic (GARCH) approach when studying causality, price transmission and volatility spillover effects in lamb, beef, pork and poultry markets in Greece. BAKUCS, FERTŐ (2005) use VECM to study the price transmission on the Hungarian pork meat market, and found competitive pricing and no evidence of price transmission asymmetries.

Most empirical results emphasise the presence of feedback between the different market levels, and support the imperfect price transmission between farm and retail markets in all meat categories studied. In short, most studies find asymmetrical price transmission in livestock markets, and they also establish a mostly unidirectional price information flow from farm to wholesale and finally retail levels.

3 EMPIRICAL PROCEDURE

Most macroeconomic time series are not stationary over time, i.e. they contain unit roots. That is, their mean and/or variance are not constant over time. Utilising the standard classical estimation methods (OLS) and statistical inference can result in biased estimates and/or spurious regressions.

Even though many individual time series contain stochastic trends (i.e. they are not stationary at levels), many of them tend to move together over the long run, suggesting the existence of a long-run equilibrium relationship. Two or more non-stationary variables are cointegrated if there exists one or more linear combinations of the variables that are stationary. This implies that the stochastic trends of the variables are linked over time, moving towards the same long-term equilibrium.

3.1 Testing for unit roots

Consider the first order autoregressive process, AR(1):

$x_t = \rho x_{t-1} + e_t$, t =…,-1,0,1,2,…, where e_t is white noise. (5)

The process is considered stationary if $|\rho| < 1$, thus testing for stationarity is equivalent with testing for unit roots ($\rho = 1$). (5) is rewritten to obtain:

$\Delta x_t = \delta x_{t-1} + e_t$ $\Delta y_t = \delta y_{t-1} + e_t$, where $\delta = 1 - \rho$ (6)

and thus the test becomes:

$H_0 : \delta = 0$ against the alternative $H_1 : \delta < 0$.

There are a large number of unit root tests in the literature (see MADDALA, KIM, 1998 for a comprehensive review), and due to their sensibility on the choice of the lag length and deterministic form it is a common practice to apply several tests. With structural breaks in the time series, the unit root tests often lead to the misleading conclusion of the presence of a unit root, when in fact the series are stationary with a break. There are however unit root tests that can handle the problem. Depending on specification, the PERRON (1997) test considers three models: with a break in the intercept, with a break in the trend, and with a break in both the intercept and trend. The test endogenously searches for the breakpoints. That is achieved by computing the t-statistics for all breakpoints, then choosing the breakpoint selected by the smallest t-statistic, that being the least favourable one for the null hypothesis of a unit root.

3.2 Cointegration analysis

The two most widely used cointegration tests are the Engle-Granger two-step method (ENGLE, GRANGER, 1987) and Johansen's multivariate approach (JOHANSEN, 1988). Let's consider a simple relationship in the form of (7), used by several cointegration tests:

$$\Delta y_t = \pi y_{t-1} + \eta_t , \tag{7}$$

where y_t is an (n x 1) vector of non-stationary variables, π is an (n x n) matrix, and η_t is a vector of possibly serially correlated normally distributed disturbances. The Johansen procedure is based on estimating π and its rank. Has the advantage that it allows for the existence of more than one cointegrating relationship (vector) and the speed of adjustment towards the long-term equilibrium is easily computed. The procedure is a Maximum Likelihood (ML) approach in a multivariate autoregressive framework with enough lags introduced to have a well-behaved disturbance term.

The Engle and Granger two step method uses an OLS regression to estimate the long-run relationship (8):

$$y_{1t} = \mu_1 + \mu_2 y_{2t} + e_t \, y_{1t}, \tag{8}$$

where y_{it} are non-stationary variables, μ are coefficients to be estimated, and e_t are disturbances.

The residuals from (8) are then tested for unit roots. The null hypothesis of unit roots is equivalent with the no cointegration hypothesis. If however the null hypothesis is rejected, the variables are considered to be cointegrated. If however, unlike (8), the true data generating process contains various regime shifts, then the Engle and Granger test is likely not to reject the no-cointegration null hypothesis.

GREGORY, HANSEN (1996) introduce a methodology to test for the null hypothesis of no-cointegration against the alternative of cointegration with structural breaks. 3 models are considered under the alternative. Model 2 with a change in the intercept:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^{\mathrm{T}} y_{2t} + e_t \ , \ \mathrm{t} = 1,\ldots,\mathrm{n}. \tag{9}$$

Model 3 is similar to model 2, only contains a time trend:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^{\mathrm{T}} y_{2t} + e_t \ , \ \mathrm{t} = 1,\ldots,\mathrm{n}. \tag{10}$$

Finally, model 4 allows a structural change both in the intercept and the slope:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t} \varphi_{t\tau} + e_t \quad \text{t = 1,…,n.} \tag{11}$$

Because usually the time of the break in not known a priori, models (9) – (11) are estimated recursively allowing T to vary between the middle 70% of the sample:

$|0.15n| \le T \le |0.85n|$

For each possible breakpoint, the ADF statistics corresponding to the residuals of models (9) – (11) are computed, then the smallest value is chosen as the test statistic (being the most favourable for the rejection of the null). Critical values are non-standard, and are tabulated in GREGORY, HANSEN (1996). JOHANSEN ET AL. (2000) generalised the JOHANSEN (1988) maximum likelihood approach in order to include up to two breaks. The procedure estimates the following model:

$$\Lambda Y_t = \alpha \begin{pmatrix} \beta \\ \mu \end{pmatrix}^{'} \begin{pmatrix} Y_{t-1} \\ tE_t \end{pmatrix} + \gamma E_t + \sum_{i=1}^{p-1} \Gamma_i \Delta Y_{t-i} + \sum_{i=1}^{p} \sum_{j=2}^{q} \kappa_{j,i} D_{j,t-i} + u_t \tag{12}$$

where Y_t is a vector of non-stationary variables, p is the lag number, $E_t = (E_{1t}\ E_{2t}...E_{qt})'$ is a vector of q dummy variables, where $E_{j,t} = 1$ if observation t belongs to the j^{th} period and 0 otherwise, $D_{j,t-i}$ is an impulse dummy that equals 1 if observation t is the i^{th} observation of the j^{th} period, meant to render the corresponding residuals to zero. Γ_i and $K_{j,i}$ are short run matrices, α is the speed of adjustment parameter matrix, β are the long run cointegration coefficients and μ are the long run drift parameters. The u_t residuals are supposed to be independently and identically distributed with zero mean and symmetric and positive definite variance-covariance matrix Ω. Restrictions on the model can be tested using likelihood ratio tests.

Notice the contradiction between the GREGORY, HANSEN (1996) and JOHANSEN (2000) procedures: the former searches endogenously for the break date, it only allows one structural break, whilst the latter allows for two, but pre-determined break dates.

3.3 Asymmetrical error correction representation

With the development of cointegration techniques, attempts were made to test asymmetry in a cointegration framework. VON CRAMON-TAUBADEL 1998, demonstrated that the earlier specifications are fundamentally inconsistent with cointegration and proposed an error correction model of the form:

$$\Delta RP_t = \alpha + \sum_{j=1}^{K}(\beta_j^+ D^+ \Delta FP_{t-j+1}) + \sum_{j=1}^{L}(\beta_j^- D^- \Delta FP_{t-j+1}) + \varphi^+ ECT_{t-1}^+ + \varphi^- ECT_{t-1} + \sum_{j=1}^{P}\Delta RP_{t-j} + \gamma_t \tag{13}$$

ECT^+_{t-1} and ECT^-_{t-1} are the segmented error correction terms resulting from the long-run (cointegration) relationship:

$ECT_{t-1} = \mu_{t-1} = RP_{t-1} - \lambda_0 - \lambda_1 FP_{t-1}$; λ_0 and λ_1 are coefficients. (14)

and,

$$ECT_{t-1} = ECT_{t-1}^+ + ECT_{t-1}^- . \tag{15}$$

Using a VECM representation as in (12), both the short-run and the long-run symmetry hypothesis can be tested, using standard tests. Valid inference requires one price to be weakly exogenous on both long and short run with respect to the parameters in (13). Following BOSWIJK, URBAIN 1997, we test for the short-run exogeneity by estimating the marginal model (16), than perform a variable addition test of the fitted residuals $\hat{\nu}_t$ from (16) into the structural model, (13):

$$\Delta FP_t = \psi_0 + \psi_1(L)\Delta FP_{t-1} + \psi_2(L)\Delta RP_{t-1} + \nu_t . \tag{16}$$

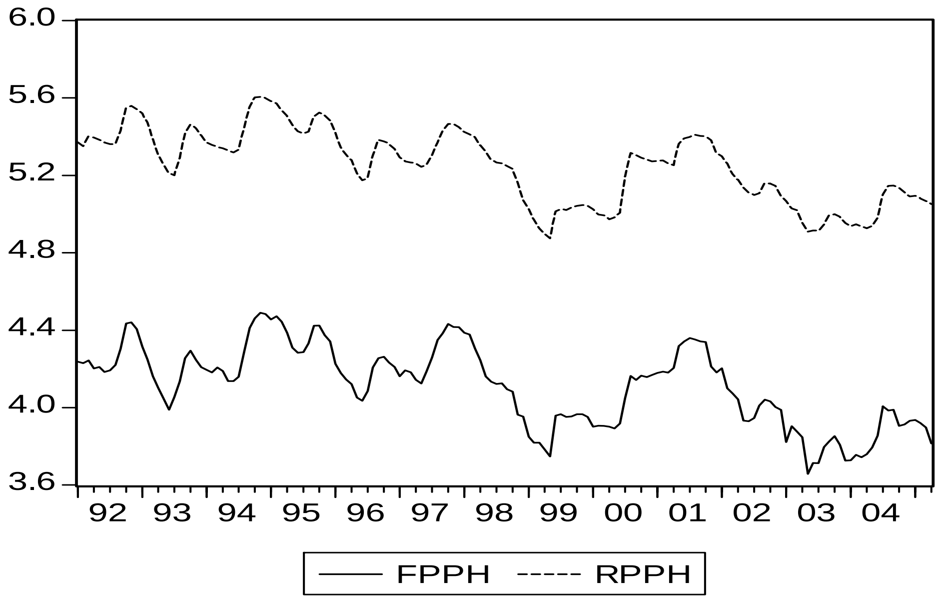
Long-run exogeneity is tested by the significance of the error correction terms in the equations (13), and (16).

4. DATA AND RESULTS

Our dataset consists of 160 monthly[3] (January 1992 – April 2005) farm-gate and consumer prices for Hungary. Hungarian farm-gate prices (FPPH) are represented by the monthly producer purchase price of live pigs for slaughter. Hungarian consumer price (RPPH) is defined as the average retail price of various meat cuts. The data was deflated to January 1992 prices, using the monthly Hungarian Consumer Price Index (CPI). All data was transformed in logarithms, because when analysing cointegrating relationships between variables, it is common to use logarithms, because otherwise, with trending data, the relative error might decline through time and this is inappropriate (DAWSON, TIFFIN 2000). The evolution of real farm and retail level prices is presented in Figure 1.

[3] With monthly data, seasonal effects might be present. Graphical analysis suggest that the time series exhibit seasonality. Therefore, following common practice, throughout this paper, monthly centred seasonal dummies were included in the VARs and regressions.

Figure 1: Log of real monthly producer and retail prices in Hungary



4.1 Stationarity and integration tests

First, we test unit roots in the logarithms of retail and farm gate prices and also their first differences using ADF (DICKEY, FULLER 1979, 1981), DF-GLS (ELLIOTT, ROTHENBERG, STOCK 1996), and PERRON tests[4] in the presence of structural breaks (PERRON 1997). As expected, results (not presented here, but available upon request) depend on the choice of lag length and deterministic assumptions. All tests however reject the null hypothesis of a unit root in the first differences, therefore we conclude that all three series are integrated of order one.

Both the ENGLE-GRANGER two step, and the JOHANSEN ML procedures accept the no-cointegration null hypothesis. Therefore next we apply the GREGORY-HANSEN procedure[5], to test for cointegration in the presence of structural breaks. Models 2 to 4 (equations 9 to 11) were subsequently estimated, starting with model 4 (models 2 and 3 are nested within 4). The null hypothesis of no-cointegration was rejected in the favour of the alternative of cointegration with a structural break in the intercept (model 2)[6]. The recursively estimated

[4] RATS code, and EVIEWS software was used to test the order of integration.

[5] The GREGORY-HANSEN cointegration tests in the presence of structural breaks were carried out using a GAUSS code.

[6] Results were substantiated using the JOHANSEN 2000 maximum likelihood cointegration procedure in the presence of structural breaks. MALCOLM code, in RATS programming language is available to test cointegration with up to 2 structural breaks.

ADF statistics for the different breakpoints are presented in figure 2. The min ADF statistic is – 5.864, - significant at 1% - corresponding to a break occurring in April 1996.

Figure 2: Recursively estimated Gregory – Hansen ADF statistics

The resulting cointegrating relationship (t - statistics in brackets) is:

$$RPPH = 2.000E_1 + 1.922E_2 + 0.802FPPH \qquad (17)$$

(28.41) (-10.42) (51.03)

where $E_1 = \begin{cases} 1 & if \quad t < April \quad 1996 \\ 0 & if \quad t \geq April \quad 1996 \end{cases}$, and

$$E_2 = \begin{cases} 0 & if \quad t < April \quad 1996 \\ 1 & if \quad t \geq April \quad 1996 \end{cases}.$$

To ensure that the prices are indeed cointegrated, the residuals of (17), are tested for unit roots using the DF-GLS procedure. The test rejects the unit root null at 1%.

4.2 Price spread and price transmission analysis

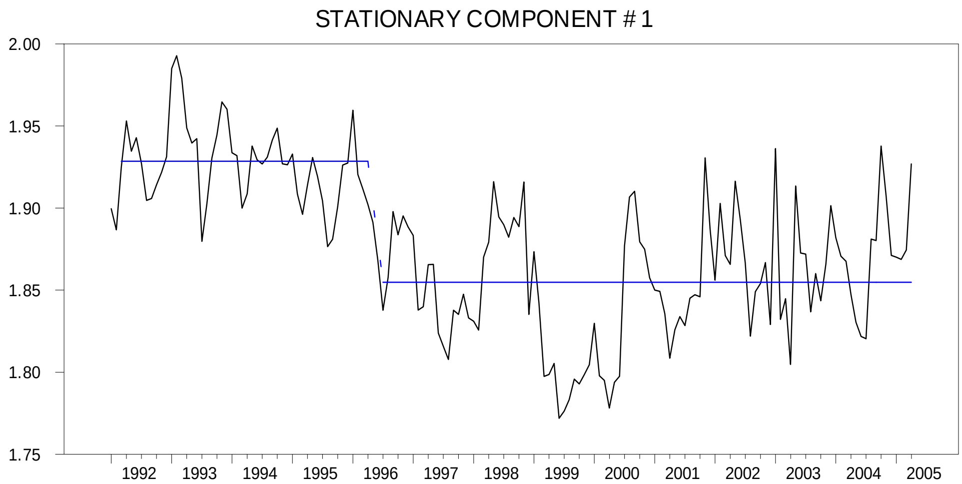
Long-run exogeneity tests cannot reject the null hypothesis of weakly exogenous producer prices ($0.459 \sim \chi^2$, p=0.38). It follows that the long run causality on the Hungarian pork meat market runs from the producer towards the consumer level. To test the competitive transmission null hypothesis, we impose the $\beta_{RPPH} = \beta_{FPPH}$ restriction on (17). The homogeneity restrictions, are rejected (12.43~ χ^2, p=0.00). The significance of the break point is also tested by variable exclusion tests, however the null hypothesis that the intercepts in the 2 sub-periods are equal is rejected (12.32~ χ^2, p=0.00). Re-estimating the model by imposing the exogeneity restrictions can improve its statistical properties. Equation (18) and figure 3 represents the re-estimated long-run relationship between the producer and retail prices on the Hungarian pork meat market:

$$RPPH = 1.928E_1 + 1.8542E_2 + 0.819FPPH \tag{18}$$

where $E_1 = \begin{cases} 1 & if \quad t < April \quad 1996 \\ 0 & if \quad t \geq April \quad 1996 \end{cases}$, and

$$E_2 = \begin{cases} 0 & if \quad t < April \quad 1996 \\ 1 & if \quad t \geq April \quad 1996 \end{cases}.$$

Figure 3: Cointegrating relationship with a structural break on the Hungarian pork meat market



With the use of logarithms, the long-run elasticity between the prices is readily available. Thus the Hungarian beef producer and retail prices are cointegrated with an imperfect transmission of $\varepsilon_{FPPH} = 0.819$. Equation 18, and figure 3 shows, that after the structural break in April 1996, the margin on the Hungarian pork market squeezed. The residuals of (18) are now saved and segmented into negative and positive phases. The first differences of the farm prices are also split into negative and positive sections as follows: $\Delta FPPHM_{t,}$ $\Delta FPPHP_t$.

Table 3: Symmetric and asymmetric VECM models (dependent variable ΔRPPH)

Independent variable	Symmetric representation (standard errors in brackets)	Asymmetric representation (standard errors in brackets)
$\Delta FPPH_t$	0.519^{**} (0.03)	-
$\Delta FPPH_{t-1}$	0.156^{**} (0.054)	-
$\Delta FPPHM_t$	-	0.175^{**} (0.039)
$\Delta FPPHM_{t-1}$	-	0.216^{**} (0.054)
$\Delta FPPHP_t$	-	0.831^{**} (0.036)
$\Delta FPPHP_{t-2}$	-	-0.171^{**} (0.06)
$\Delta RPPH_{t-1}$	0.105 (0.064)	0.227^{**} (0.038)
$\Delta RPPH_{t-2}$	-	0.102^{*} (0.052)
ECT_{t-1}	-0.277^{**} (0.056)	-
$ECTM_{t-1}$	-	-0.203^{**} (0.067)
$ECTP_{t-1}$	-	-0.198^{*} (0.093)
c	- 0.0303 (0.001)	-0.006^{*} (0.002)
Adjusted R^2	0.79	0.89
Autocorrelation LM(1)	1.254	0.03
Autocorrelation LM(4)	0.783	0.483
Autocorrelation LM(12)	0.565	0.721
Autocorrelation (Ljung – Box Q statistic)	Q(36) = 23.496	Q(36) = 35.187
Normality (Jarque–Bera)	$84.71^{**\dagger}$	$26.85^{**\dagger}$
Variable addition test (ν_t, marginal model residuals)	0.082 [~F(1,151)]	0.093 [~F(1,147)]
Long-run symmetry	-	0.001 [~F(1,148)]
Short-run symmetry	-	7.943^{**} [~F(1,148)]

*significant at 5%, ** significant at 1%

†Non-normality – implies that the test results must be interpreted with care, although asymptotic results do hold for a wider class of distributions (VON CRAMON-TAUBADEL 1998).

The transformed equation (13) was first estimated with 4 lags, and then reduced to more parsimonious models. Before proceeding to the price transmission analysis, the direction of the causality must be determined. The marginal models (16), not shown here, were also estimated, and the fitted residuals $\hat{\nu}_t$ saved. The variable addition test results of the saved $\hat{\nu}_t$ residuals into model (13) and its symmetric counterpart, are presented in the bottom of table 1. The $\hat{\nu}_t$ residuals are not significant in the structural equation therefore the short-run causality on the Hungarian pork meat market runs from the producer towards the consumer prices. As discussed in section 3.3, to test the long run causality, the significance of the error correction terms (ECT_{t-1}, $ECTM_{t-1}$, $ECTP_{t-1}$) in the marginal equation 16 is tested. Results (not presented here) show that none were significant, therefore the long-run causality on the Hungarian Pork meat market also runs from the producer to retail level.

The models appear to be well specified with a quite high coefficient of determination, there are no traces of serial autocorrelation of order 1, 4, 12, and the Ljung-Box Q statistic does not reject the null hypothesis of no serial correlation amongst the first 36 residuals. The residuals however are non-normal, which implies that the test results must be interpreted with care, although asymptotic results do hold for a wider class of distributions (VON CRAMON-TAUBADEL, 1998). Both $ECTM_{t-1}$ and $ECTP_{t-1}$ are highly significant, and of the right sign, $ECTM_{t-1}$ being slightly bigger (in absolute values) than $ECTP_{t-1}$. A formal test however cannot reject the null hypothesis the two correction terms being equal, suggesting that long-run price transmission is symmetric. The short-run symmetry null hypothesis is rejected, an increase in farm prices induces a bigger increase in retail prices (on short-run) than a decrease of farm prices.

5 CONCLUSIONS

With many empirical studies of livestock markets in developed countries, we have simultaneously examined how retail price is formed and how price transmission works in the livestock markets of a transition economy. We analysed the long-run relationship between retail prices and the farm-gate price for pork meat in Hungary. Vertical price transmission was analysed in the cointegration framework, using relatively new cointegration technique that also allows cointegration in the presence of structural breaks. Results indicate that the retail and farm gate prices move together in the long run, that is, they are cointegrated, with a structural break occurring in April 1996. The exogeneity tests found the farm prices were weakly exogenous on both long and short-run and established a unidirectional long-run Granger causality from producer to retail prices. Prices are set on the farm level market and transmitted up through the wholesale and processing level to the retailers. Our long run causality findings are in line with most empirical studies carried out on livestock markets

(VON CRAMON-TAUBADEL, 1998; BOJNEC, 2002; ABDULAI, 2002; BEN-KAABIA ET AL., 2002; to name just a few). Marketing analysis found that Hungary possess a non-competitive market structure, where processors and retailers charge a mark-up of the retail price plus a constant absolute margin that might suggest the exercise of market power. The existence of a mark-up pricing strategy, concur with BOJNEC (2002), who studied the Slovenian pork and beef meat market, and found competitive pork but non-competitive beef marketing margin formation processes.

We carried out both short and long run asymmetry tests, and contrary to popular belief, we found that the null of symmetrical price transmission cannot be rejected on the long-run. This result contradicts the findings of the studies set in developed markets that usually establish asymmetrical price transmission on livestock markets and a farm to wholesale to retail price information flow. Short-run price transmission however proved to be asymmetric, retailers tend to quickly pass increasing short-run producer price movements.

REFERENCES

ABDULAI, A. (2002): Using threshold cointegration to estimate asymmetric price transmission in the Swiss pork market, *Applied Economics, Vol. 34*, pp. 679-687.

BAKUCS, L.Z., FERTŐ, I. (2005): Marketing Margins and Price Transmission on the Hungarian Pork Meat Market, *Agribusiness, Vol. 21 (2)*, pp. 273-286.

BEN-KAABIA, M., GILL, JM., BOSHNJAKU, L. (2002): Price transmission asymmetries in the Spanish lamb sector. *Paper presented at the X. Congress of European Association of Agricultural Economists, 28-31 August*, Zaragoza, Spain.

BOJNEC, S. (2002): Price Transmission and Marketing Margins in the Slovenian Beef and Pork Markets During Transition. *Paper presented at the X. Congress of European Association of Agricultural Economists, 28-31 August*, Zaragoza, Spain.

BOSWIJK, H.P., URBAIN, J.P. (1997): Lagrange-multiplier tests for weak exogeneity: A synthesis. *Econometric Reviews, Vol. 16*, pp. 21-38.

VON CRAMON-TAUBADEL, S. (1998): Estimating asymmetric price transmission with the error correction representation: An application to the German pork market, *European Review of Agricultural Economics, Vol. 25,* pp. 1-18.

DAWSON, P.J. AND TIFFIN, R. (2000): Structural breaks, cointegration and the farm-retail price spread for lamb, *Applied Economics, Vol. 32,* pp. 1281-1286.

DICKEY, D.A. AND FULLER, W.A. (1979): Distributions of the Estimators For Autoregressive Time Series With a Unit Root, *Journal of the American Statistical Association, Vol. 75,* pp. 427- 431.

DICKEY, D.A., FULLER, W.A (1981): Likelihood Ratio Statistics for Autoregressive Time Series With a Unit Root, *Econometrica, Vol. 49,* pp. 1057-1072.

Elliott, G., Rothenberg, T.J., Stock, J.H. (1996): Efficient tests for an autoregressive unit root, *Econometrica*, *vol. 64*, PP.813-836.

ENGLE, R.F. GRANGER, C.W.J (1987): Cointegration and error correction: Representation, estimation and testing, *Econometrica, Vol. 55,* pp. 251-276.

GRANGER, C.W.J. (1969): Investigating Casual Relations by Econometric Methods and Cross-spectral Methods, *Econometrica, Vol. 37,* pp. 24-36.

GREGORY, A.W., HANSEN, B.E. (1996): Residual-Based Tests for Cointegration in Models with Regime Shifts, *Journal of Econometrics, Vol. 70,* pp. 99-126.

GOODWIN, B.K., HARPER, D.C. (2000): Price Transmission, Threshold Behaviour, and Asymmetric Adjustment in the U.S. Pork Sector, *Journal of Agricultural and Applied Economics, Vol. 32,* pp. 543 -553.

GOODWIN, B.K., HOLT, M.T. (1999): Price Transmission and Asymmetric Adjustment in the U.S. Beef Sector, *American Journal of Agricultural Economics, Vol. 81,* pp. 630-637.

HARRIS, R.I.D (1995): Using Cointegration Analysis in Econometric Modelling. Prentice Hall/Harvester Wheatsheaf.

HOUCK, J.P. (1977): An Approach to Specifying and Estimating Nonreversible Functions, *American Journal of Agricultural Economics, Vol. 59*, pp. 570-572.

JOHANSEN, S. (1988): Statistical Analysis of Cointegrating Vectors, *Journal of Economic Dynamics and Control, Vol. 12,* pp. 231-254.

MADDALA, G.S., KIM, IN-MOO (1998): Unit Roots, Cointegration, and Structural Change. Cambridge University Press.

MILLER, J. D., HAYENGA, M. L. (2001): Price Cycles and Asymmetric Price Transmission in the U.S. Pork Market, *American Journal of Agricultural Economics, Vol. 83,* pp. 551 – 561.

PERRON, P. (1997): Further evidence on breaking trend functions in macroeconomic variables, *Journal of Econometrics, Vol.80,* pp.355-385.

WARD, R.W. (1982): Asymmetry in Retail, Wholesale, and Shipping Point Pricing for Fresh Vegetables, *American Journal of Agricultural Economics, Vol. 64,* pp. 205-212.

WOLFFRAM, R. (1971): Positivistic Measures of Aggregate Supply Elasticities: Some New Approaches – some critical notes, *American Journal of Agricultural Economics, Vol. 31,* pp. 356 –359.