MARKETING MARGINS AND PRICE TRANSMISSION ON THE HUNGARIAN BEEF MARKET•

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ABSTRACT

There is a wealth of literature on farm-retail price spread for different commodities and countries. However, with the exception of Bojnec, (Slovenian pork and beef) and Bakucs–Fertő, (Hungarian pork) none have studied price transmission and marketing margins in the transition economies. It is a common belief that because of the distorted markets inherited from the pre 1989 period, the deficiency of the price-discovery mechanisms, unpredictable policy interventions, price transmission is generally asymmetric. We apply the Gregory and Hansen procedure with recursively estimated breakpoints and ADF statistics, and found that the prices are cointegrated with a time trend and a structural break occurring in August 1997. Exogeneity tests reveal the causality running from producer to retail prices. Homogeneity is rejected, suggesting a mark-up pricing strategy. Price transmission modelling suggests that, despite the common belief, price transmission on the Hungarian beef meat market is symmetric on both long and short run.

Keywords: marketing margins, asymmetric price transmission, cointegration in the presence of structural breaks, error correction, beef market.

• 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.

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 and 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 beef 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.:

$$RP = PP + M \qquad (1)$$

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

$M = a + b*RP$, where $a \geq 0$ and $0 \leq b < 1$. (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:

$RP = \alpha_1 + \varepsilon_{PP}PP$ (3)

,where ε_{PP} is the price transmission elasticity from the producer price (PP) towards the consumer price (RP). If $\varepsilon_{FP} = 1$, we have perfect transmission, and thus the *mark - up* will be $(e^{\alpha 1} - 1)$. $0 < \varepsilon_{PP} < 1$ 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:

$PP = \alpha_2 + \varepsilon_{RP}RP$, (4)

where ε_{RP} is the elasticity of transmission between the consumer price (RP) and the producer price (PP). 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 and 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 and 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 and Hayenga, 2001, p. 554).

3.2 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 and 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 and Holt (1999), Goodwin and Harper (2000) and Ben-Kaabia, Gil, Boshnjaku (2002) studying the US beef sector, US pork sector and Spanish lamb sector, respectively. Goodwin and 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 and 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 and 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 and 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 and 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 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):

$y_t = \rho y_{t-1} + e_t$, 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:

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

and thus the test becomes:

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

Maddala and Kim (1998) argues, that because of the size distortions and poor power problems associated with the Augmented Dickey-Fuller unit root tests, it is preferable to use the DF-GLS unit root test, derived by Elliott, Rothenberg and Stock (1996).

Elliott, Rothenberg and Stock develop the asymptotic power envelope for point optimal autoregressive unit root tests, and propose several tests whose power functions are tangent to the power envelope and never too far below (Maddala and Kim, 1998). The proposed DF-GLS test works by testing the $a_0=0$ null hypothesis in regression (7):

$$\Delta y^d_t = a_0 y^d_{t-1} + a_1 \Delta y^d_{t-1} + \ldots + a_p \Delta^d_{t-p} + e_t \tag{7}$$

where y^d_t is the locally detrended y_t series that depends on whether a model with a drift or linear trend is considered. In case of a model with a linear trend, the following formula is used to obtain the detrended series y^d_t:

$$y^d_t = y_t - \hat{\beta}_0 - \hat{\beta}_t \, . \tag{8}$$

$\hat{\beta}_0$ and $\hat{\beta}_t$ are obtained by regressing $\bar{y}$ on $\bar{z}$, where:

$$\bar{y} = [y_1, (1-\bar{\alpha}L)y_2, \ldots, (1-\bar{\alpha}L)y_T] \tag{9}$$

$$\bar{z} = [z_1, (1-\bar{\alpha}L)z_2, \ldots, (1-\bar{\alpha}L)z_T] \, . \tag{10}$$

Elliott, Rothenberg and Stock argue that fixing $\bar{c} = -7$ in the drift model, and $\bar{c} = -13.5$ in the linear trend model, used in (11) and (12), the test is within 0.01 of the power envelope:

$$z_t = (1, t)' \tag{11}$$

$$\bar{\alpha} = 1 + \frac{\bar{c}}{T} \, . \tag{12}$$

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. Several unit root tests were developed to handle the problem. Depending on specification, the Perron (1989) test considers three models: with an exogenous break in the intercept (13), with an exogenous break in the trend (14), and with a break in both the intercept and trend (15).

$$y_t = \alpha_1 + \beta_1 t + (\alpha_2 - \alpha_1)DU_t + e_t, \quad t=1,2,\ldots.T \tag{13}$$

$$y_t = \alpha_1 + \beta_1 t + (\beta_2 - \beta_1)DT_t + e_t, \quad t=1,2,\ldots.T \tag{14}$$

$$y_t = \alpha_1 + \beta_1 t + (\alpha_2 - \alpha_1)DU_t + (\beta_2 - \beta_1)DT_t + e_t, \quad t=1,2,\ldots.T \tag{15}$$

where, $DT_t = \begin{cases} t & if \quad t > TB \\ 0 & else \end{cases}$

and $DU_t = \begin{cases} 1 & if \quad t > TB \\ 0 & else \end{cases}$.

The problem with the Perron test is that the breakpoint must be known *a priori*. Zivot and Andrews (1992) modified the Perron test, to endogenously search 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.

3.2 Cointegration analysis

The two most widely used cointegration tests are the Engle-Granger two-step method (Engle and Granger, 1987) and Johansen's multivariate approach (Johansen, 1988). Let's consider a simple relationship in the form of (16), used by several cointegration tests:

$$\Delta \mathbf{y}_t = \pi \mathbf{y}_{t-1} + \eta_t , \tag{16}$$

where $\mathbf{y}_t$ is an (n x 1) vector of nonstationary 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 (17):

$$y_{1t} = \mu_1 + \mu_2 y_{2t} + e_t\,, \tag{17}$$

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

The residuals from (17) 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 (17), 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 and 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:

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha^{\mathrm{T}} y_{2t} + e_t \;, \; \mathrm{t} = 1,\ldots,\mathrm{n}. \tag{18}$$

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

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^{\mathrm{T}} y_{2t} + e_t \;, \; \mathrm{t} = 1,\ldots,\mathrm{n}. \tag{19}$$

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

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t} \varphi_{t\tau} + e_t \quad \mathrm{t} = 1,\ldots,\mathrm{n}. \tag{20}$$

Because usually the time of the break in not known *a priori*, models (18) – (20) are estimated recursively allowing τ to vary between the middle 70% of the sample:

$$|0.15n| \le T \le |0.85n| \tag{21}$$

For each possible breakpoint, the ADF statistics corresponding to the residuals of models (18) – (20) 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 and Hansen (1996).

3.3 Asymmetrical error correction representation

Most asymmetry analysis uses the following Ward (1982) specification, based on the earlier Woffram (1971) and Houck (1977) specification:

$$\Delta RP_t = \alpha + \sum_{j=1}^{K} (\beta_j^{+} D^{+} \Delta FP_{\,t-j+1}) + \sum_{j=1}^{L} (\beta_j^{-} D^{-} \Delta FP_{\,t-j+1}) + \gamma_t \qquad (22)$$

Here, the first differences of the producer prices are split into increasing and decreasing phases by the D^- and D^+ dummy variables. Asymmetry is tested using a standard F-test to determine whether β_j^+ and β_j^- are significantly different.

These approaches do not pay attention to the time series properties of the data and many of them suffer serial autocorrelation that usually suggests spurious regression.

With the development of cointegration techniques, attempts were made to test asymmetry in a cointegration framework. Von Cramon-Taubadel (1998) demonstrated that the Wolffram-Houck type specifications are fundamentally inconsistent with cointegration and proposed an error correction model of the form:

$$\Delta RP_t = \alpha + \sum_{j=1}^{K} (\beta_j^{+} D^{+} \Delta FP_{t-j+1}) + \sum_{j=1}^{L} (\beta_j^{-} D^{-} \Delta FP_{t-j+1}) + \varphi^{+} ECT^{+}_{t-1} + \varphi^{-} ECT^{-}_{t-1} +$$

$$+ \sum_{j=1}^{P} \Delta RP_{t-j} + \gamma_t \qquad (23)$$

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

$$ECT_{t-1} = \mu_{t-1} = RP^{R}_{t-1} - \lambda_0 - \lambda_1 FP_{t-1}\ ;\ \lambda_0 \text{ and } \lambda_1 \text{ are coefficients.} \qquad (24)$$

and,

$$ECT_{t-1} = ECT^{+}_{t-1} + ECT^{-}_{t-1}. \qquad (25)$$

Using a VECM representation as in (23), 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 (23). Following

Boswijk and Urbain (1997) we test for the short-run exogeneity by estimating the marginal model (26), than perform a variable addition test of the fitted residuals $\hat{v}_t$ from (26) into the structural model, (23):

$$\Delta P^{P}_{t} = \psi_0 + \psi_1(L)\, \Delta P^{R}_{t-1} + \psi_2(L)\, \Delta P^{P}_{t-1} + v_t \qquad (26)$$

Long-run exogeneity is tested by the significance of the error correction terms in the equations (23), and (26).

4. DATA AND RESULTS

Our sample contains 99 monthly (January 1992 – March 2000) farm-gate and consumer prices. Farm-gate prices (FPB) are represented by the monthly producer purchase price of live cattle for slaughter, whilst consumer price (RPB) is defined as retail price of beef steak. The Hungarian Central Statistical Office supplied all price data. All data were deflated to January 1992 prices, using the monthly Hungarian Consumer Price Index (CPI). The 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 and Tiffin, 2000). The evolution of real farm and retail level prices is presented in Figure 1.

- insert Figure 1. here -

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, results are presented in table 1. The tests indicate that all price series contain unit roots with constant and/or trend. The unit root null hypothesis is rejected for the first differences of all price series. Therefore we conclude that all three series are integrated of order one.

- insert Table 1. here -

The Zivot – Andrews unit root tests in the presence of structural breaks (not shown here, but available on request) reinforce the results obtained with the DF-GLS test.

Using the Engle and Granger two step procedure, first an OLS regression in the form of (17) is run between the producer and retail price variables. The residuals (figure 2), are then tested for unit roots, using the DF – GLS test.

- insert Figure 2. here -

The test statistic with a constant only specification is – 1.543, with a constant and trend specification is – 1.785, none rejecting the no-cointegration null hypothesis.

Therefore we apply the Gregory - Hansen procedure next, to test for cointegration in the presence of structural breaks. Models 2 to 4 (equations 18 to 20) 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 and a linear trend (model 3). The recursively estimated ADF statistics for the different breakpoints are presented in figure 3. The min ADF statistic is – 5.252, - significant at 5% - corresponding to a break occurring in August 1997.

- insert Figure 3. here -

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

$$RPB = \underset{(15.59)}{1.843} - \underset{(-13.23)}{0.148}\varphi_t + \underset{(12.22)}{0.002}t + \underset{(28.49)}{0.807}FPB \qquad (27)$$

where $\varphi_t = \begin{cases} 0 & if \quad t < August \quad 1997 \\ 1 & if \quad t \geq August \quad 1997 \end{cases}$, and t is the time trend.

To ensure that the prices are indeed cointegrated, the residuals of (27), (figure 4) are tested for unit roots using the DF-GLS procedure. The test statistic with a constant only specification is – 4.093, with a constant and trend specification is – 4.544, rejecting the unit root null at 1%.

4.2 Price spread and price transmission analysis

To test the competitive transmission null hypothesis, we impose the $\beta_{RPB} = \beta_{FPB}$ restriction on (27). The F-statistic is 45.883, rejecting the null hypothesis at 1%. 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 $\varepsilon_{FPB} = 0.807$.

- insert Figure 4. here -

The residuals of (27) and 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: ΔFPBM, ΔFPBP. The transformed equation (23) 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 (26), 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 (23) and its symmetric counterpart, are presented in the bottom of table 2. The test statistics are not significant. As discussed in section 3.4, 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 26 is tested. Results (not presented here) show that none were significant. It therefore appears that both the short and long run causality runs from the producer to the retail prices.

- insert Table 2. here -

Table 2 presents the regression estimates of the asymmetrical and symmetrical representations, the symmetry and some diagnostic tests. The two models are well specified, there are no traces of serial autocorrelation of order 1, 4, and 12. The Ljung-Box Q statistic doesn't reject the null hypothesis of no serial correlation amongst the first 36 residuals. There is no evidence of heteroskedasticity, therefore heteroskedasticity. The residuals are nonnormal, 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). The variable addition tests indicate that the marginal equations' residuals are not significant in the models, therefore the null hypothesis that the retail prices are weakly exogenous with respect to the short-run parameters too, cannot be rejected. The error correction terms (ECT_{t-1}, $ECTM_{t-1}$ and $ECTP_{t-1}$) have the right (negative) sign, and $ECTM_{t-1}$ causes a slightly greater change in the retail price than $ECTM_{t-1}$. However, the F-test of long-run symmetry null hypotheses cannot be rejected, suggesting price transmission symmetry. The short-run symmetry hypotheses are then tested using an F-test. At 1% probability, the nulls of symmetry cannot be rejected in this case either. Results from Table 2 are used to estimate impulse response functions, that graphically depict the effect of a one unit negative and positive shock to the farm price on the retail price variable.

- insert Figure. 5 here –

The graph reinforces the symmetrical transmission results obtained. The magnitude of the response to a positive shock coincides with the response to the negative shock, and it takes the same period of time for the retail price to get back to its long run equilibrium.

5 CONCLUSIONS

With many empirical studies of livestock markets in developed countries, we have examined how retail price is formed and how price transmission works in a transition country's livestock market. We analysed the long-run relationship between two retail prices and the farm-gate price for beef meat in Hungary. Vertical price transmission was analysed in the cointegration framework, using relatively new cointegration technique that allows cointegration in the presence of structural breaks. Results indicate that the retail and farm gate prices in the Hungarian beef meat market move together in the long run, that is, they are cointegrated for the January 1992 to March 2000 period, with a structural break occurring in August 1997. 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 causality findings are in line with most empirical studies carried out on livestock markets (Von Cramon-Taubadel, 1988; Bojnec, 2002; Abdulai, 2002; Ben-Kaabia et al, 2002; just to name a few). Marketing analysis found that there is 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. These results suggest that the less developed markets in the transition economies cannot perform as competitive markets.

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 in either case. 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.

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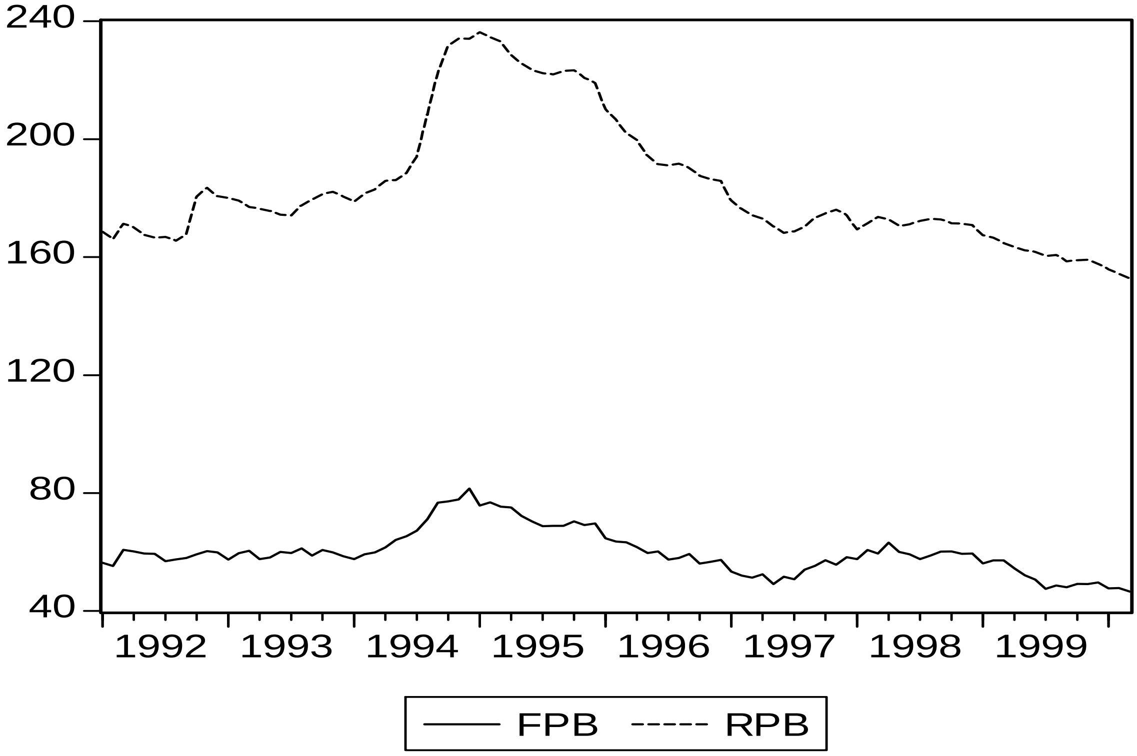
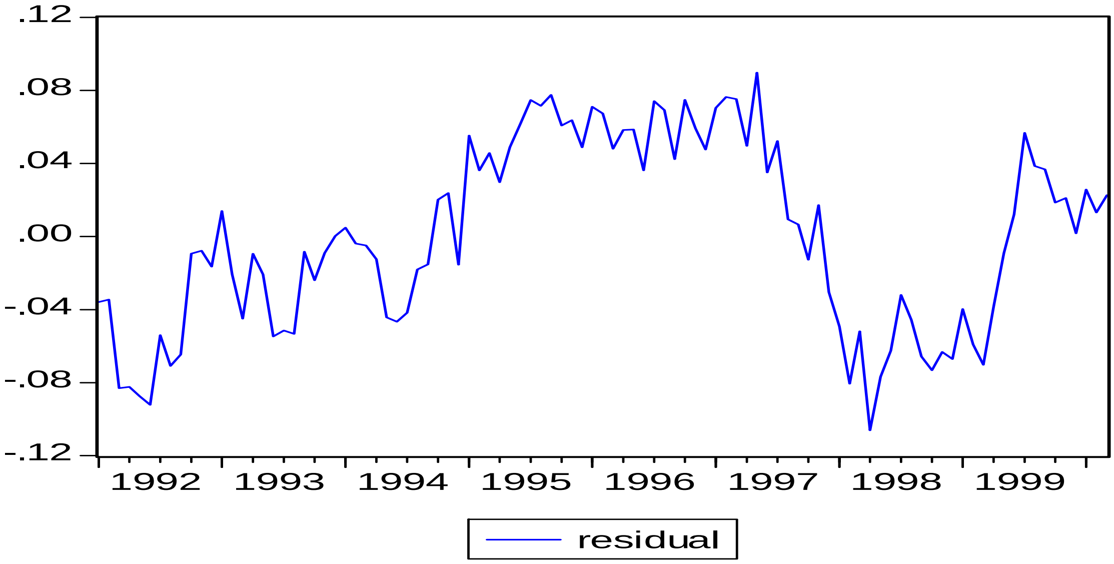
Figure 1. Monthly real farm- gate and retail prices in HUF/kg



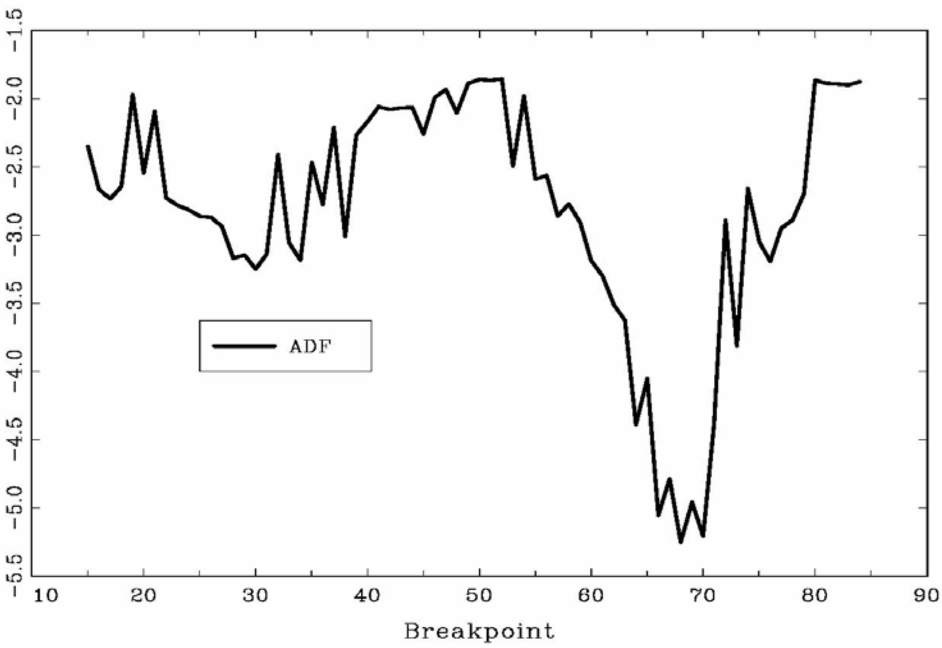
Source: Author’s own calculations, data supplied by the Central Statistical Institute

Figure 2. OLS residuals of the regression RPB on FPB



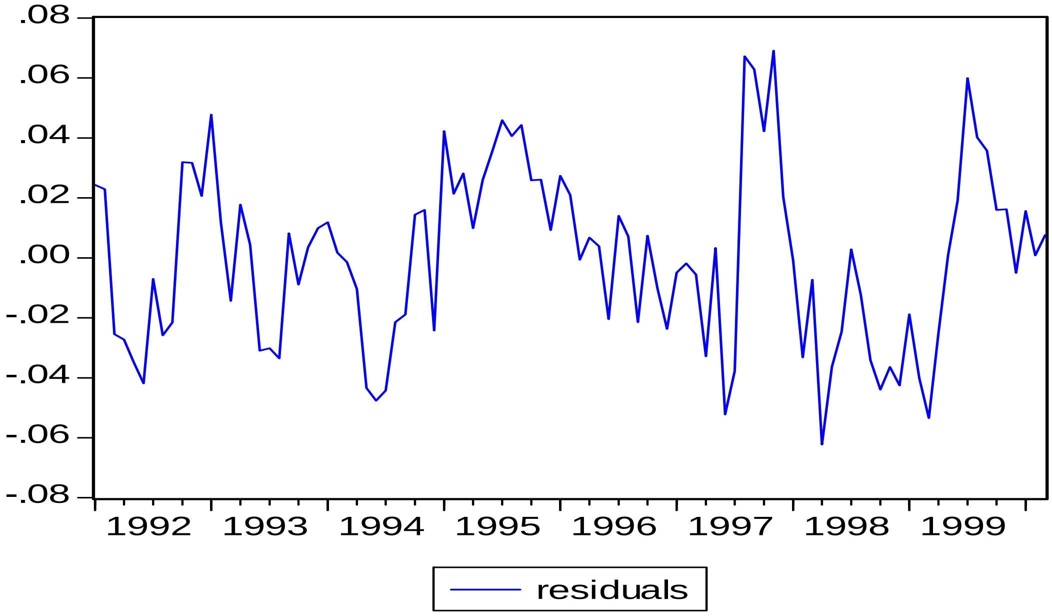
Source: Author's own calculations

Figure 3. Recursively estimated Gregory – Hansen statistics



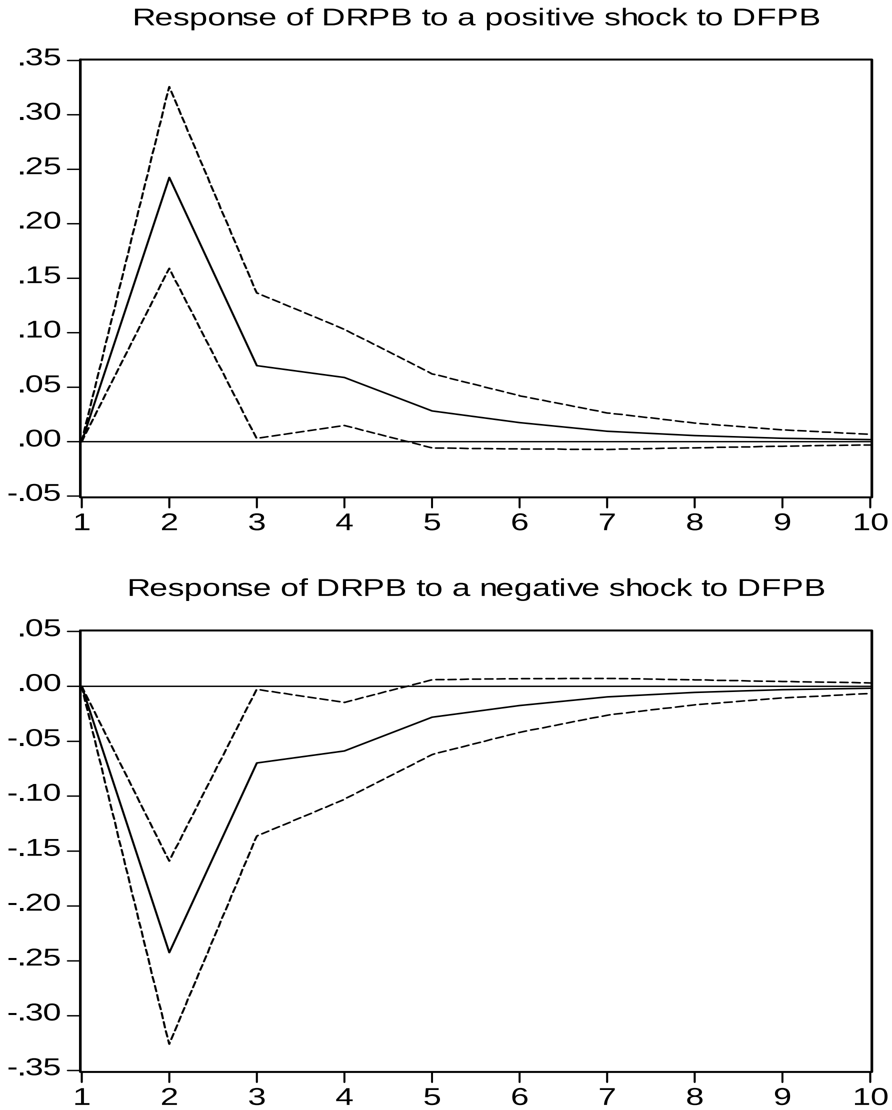
Source: Author's own calculations

Figure 4. The cointegrating residuals



Source: Author's own calculations

Figure 5. Symmetric response of RPB to unit positive and negative shocks to FPB



Source: Author's own calculations

Table 1. DF-GLS unit root test results

Variable	Specification	Lags	Test statistic
FPB	constant	6	- 1.630
	constant and trend	6	- 1.966
RPB	constant	1	- 1.122
	constant and trend	1	- 1.260
ΔFPB	constant	5	- 2.220
ΔRPB	constant	0	- 4.503

The 0.90 (0.95) confidence intervals critical values for the DF-GLS tests with constant are –1.614 (-1.944), with constant and trend are –2.749 (-3.039). The lag length was selected by the AIC criteria.

Table 2. Symmetric and asymmetric VECM models (dependent variable ΔRPB)

Independent variable	Symmetric representation (standard errors in brackets)	Asymmetric representation (standard errors in brackets)
ΔFPB_t	0.2407** (0.041)	-
$\Delta FPBP_t$	-	0.2595** (0.067)
$\Delta FPBM_t$	-	0.2264** (0.065)
ΔRPB_{t-1}	0.4234** (0.075)	0.4217** (0.076)
ECT_{t-1}	- 0.1633** (0.045)	-
$ECTP_{t-1}$	-	- 0.1423** (0.070)
$ECTM_{t-1}$	-	- 0.1821** (0.079)
Adjusted R^2	0.503	0.496
Autocorrelation LM(1)	0.051	0.024
Autocorrelation LM(4)	1.201	1.015
Autocorrelation LM(12)	1.038	1.009
Autocorrelation (Ljung – Box Q statistic)	Q(36) = 23.358	Q(36) = 23.113
Normality (Jarque–Bera)	86.38**†	90.06**†
Heteroskedasticity (White)	1.665	1.346
Variable addition test (v_t, marginal model residuals)	0.015 [~F(1,92)]	0.007 [~F(1,90)]
Long-run symmetry	-	0.112 [~F(1,92)]
Short-run symmetry	-	0.100 [~F(1,92)]

*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).