Using Threshold Cointegration to Estimate Asymmetric Price Transmission in the Swiss Pork Market.

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Using threshold cointegration to estimate asymmetric price transmission in the Swiss pork market

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This paper employs threshold cointegration tests that allow for asymmetric adjustment towards a long-run equilibrium relationship to examine the relationship between producer and retail pork prices in Switzerland. The short-run adjustments are also examined with asymmetric error correction models that are compared to the conventional symmetric error correction models. The results indicate that price transmission between the producer and retail levels is asymmetric, in the sense that increases in producer prices that lead to declines in marketing margins are passed on more quickly to retail prices than decreases in producer prices that result in increases in the marketing margins.

I. INTRODUCTION

The process of transmission of price changes arising in upstream stages through to final consumer prices in the food sector has long been one of the most studied areas in the agricultural economics literature for policy purposes (Palaskas, 1995). Given that price is the primary mechanism by which various levels of the market are linked, the extent of adjustment and speed, with which shocks are transmitted between producer and retail prices, is a significant factor showing the actions of participants at alternative market levels. Many observers have asserted that middlemen are more apt to increase than to lower the prices of food items. As a result, cost increases are completely and rapidly passed on to consumers, whilst there is a slower and less complete transmission of cost savings.

In examining the transmission of price changes arising from upstream to the final consumer in the food sector, a number of empirical studies have employed the Houck (1977) model. Allowing for price changes over more than one time period, Boyd and Brorsen (1988) employ the model to test for asymmetry in the US pork market, and find no evidence of asymmetric vertical price transmission. Hahn's (1990) later study on price transmission in pork and beef markets, however, suggested prices at all levels of the US beef and pork marketing chains are less sensitive to shocks from price declines than to those from price increases.

A major shortcoming of previous studies on asymmetric price transmission in the food marketing chain is that they fail to take into account the possibility of the presence of equilibrium relationship between any price series being examined (von Cramon-Taubadel, and Loy, 1997). Food retail prices and producer prices may drift apart in the short run due to policy changes or seasonal factors, but if they continue to be too far apart in the long run, economic forces, such as market mechanisms may bring them together (Palaskas, 1995; Enders, 1995).

The cointegration and its corresponding error correction model used in recent studies to account for the shortcomings outlined above implicitly assume that the tendency to move towards a long-run equilibrium is present every time. It is, however, possible that movement towards equilibrium need not occur in every period. Balke and Fombe (1997) point out that the presence of fixed costs of adjustment may prevent economic agents from adjusting continuously. Only when the deviation from the equilibrium exceeds a critical threshold do the benefits of adjustment exceed the costs and, hence economic agents act to move the system back towards the equilibrium. Threshold models of

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dynamic economic equilibria have therefore gained increasing significance in the analysis of price transmission asymmetries (e.g., Azzam, 1999).

The contribution of this paper is the application of a recent methodology developed by Enders and Granger (1998) that considers the time series properties of the variables, but does not presuppose a linear-symmetric adjustment to study the transmission of producer price changes to changes in retail prices in the Swiss pork market, over the 1988–1997 sample period. Given the widely held belief that prices are sticky downwards, the tests employed in this study have more power than the Engle and Granger (1987) and Johansen (1988) tests used in recent studies.

The remainder of the paper is organized as follows. The next section briefly discusses some competing theories of asymmetric adjustment. Section III explains the econometric approach used in the empirical analysis. In Section IV, the empirical results of the analysis are presented, while concluding remarks are given in the final section.

II. EXPLAINING ASYMMETRIC PRICE TRANSMISSION

A number of competing theories have been put forward to explain the existence of asymmetric farm-retail price transmission. Kovenock and Widdows (1998) point out that if input costs changes are perceived as temporary, then the menu costs likely to arise may serve as an incentive not to adjust prices when input costs decrease. These costs may result from the repricing of goods and informing sales people and customers, or may be generated by more sophisticated processes. For instance, firms may not be willing to signal their customers that market conditions have changed, since those customers could re-engage in search behaviour. Blinder (1982) also shows that firms normally alter their behaviour in response to sustained price movements that cause inventories to decline or increase over time, but allow inventories to build up or decline for short periods of time as prices change in a transitory fashion. He argues that the principal reason firms do not respond to transitory price changes are menu costs.

Azzam (1999) posits a model to demonstrate how asymmetry arises from inter-temporal optimizing behaviour of spatially competitive retailers. He shows that although retail prices may indeed rise more than they decline, the rise is smaller when competition is vigorous, and the decline is larger when competition is less vigorous. This suggests that retail competition does not necessarily translate into the larger retail price declines producers expect during periods of declining farm prices. He also shows that when retailers incur repricing costs there is a range of farm price changes over which retailers would choose not to reprice, resulting in less frequent repricing, or rigidity, as a response to upward as well as downward movements in the farm price.

Another frequently cited reason for asymmetric adjustment in the farm-retail price transmission is the presence of search costs in locally imperfect markets (Blinder *et al.*, 1998). In many areas, retail firms may enjoy local market power due to the lack of similar firms in the neighbourhood. As a result of search costs, customers of these firms may not be able to acquire complete information about prices offered by other firms, although they face a finite number of choices. Even if customers observe a price increase at a particular retail shop, they may be uncertain as to whether prices in other shops have increased. Under such conditions, firms can quickly raise prices as producer prices increase and can slowly reduce prices as producer prices fall.

The 'trigger price' model of oligopolistic co-ordination has also been used to explain asymmetric retail price adjustment (Borenstein *et al.*, 1997). The model considers the case in which a few dominant firms explicitly cooperate to coordinate prices in an industry. To ensure market power, the collusive firms can use trigger prices in such cases to identify firms that cheat. For example, a retailer may be punished by the other firms if it attempts to increase its market share by reducing prices below a trigger price set by the firms as a minimum. Retailers may therefore hesitate to lower prices too quickly in order to avoid punishment. However, as producer prices increase, retailers can quickly raise their prices to maintain profits without entertaining any fears of punishment from other firms.

III. MODELLING PRICE TRANSMISSION

Several analysts have employed a modification of the Houck (1977) model to examine the nature of vertical price transmission in the food industry. The reduced form of this static model is given as:

$$\sum_{\tau=1}^{t} \Delta PR_{\tau} = \alpha_0 + \alpha_1 \sum_{\tau=1}^{t} \Delta PPI_{\tau} + \alpha_2 \sum_{\tau=1}^{t} \Delta PPF_{\tau} + \varepsilon_t \quad (1)$$

where ΔPR_{τ} represents changes in the retail prices, ΔPPI_{τ} and ΔPPF_{τ} are the positive and negative changes in the producer prices for pork, respectively, α_0, α_1 and α_2 are coefficients to be estimated, t is current time, and ε_t is the random error term. The null hypothesis of symmetric adjustment is then tested by determining whether $\alpha_1 = \alpha_2$.

For reasons indicated in the previous section, some studies have estimated this specification employing cointegration techniques. However, von Cramon-Taubadel and Loy (1997) have recently demonstrated that the specification in Equation 1 is inconsistent with the concept of cointegration. More recently, Azzam (1999) shows that particularly in the presence of rigidity due to repricing costs, use of the conventional non-reversible functions (Houck, 1977) to test for symmetry is no longer appropriate. Given that the cointegration techniques used in previous studies assume symmetric adjustment, this paper employs a cointegration model that recognizes the fact that a shock may have to reach a critical level before a significant response is provoked.

Consider the simple relationship that is used as the basis for several cointegration analyses:

$$\Delta x_t = \pi x_{t-1} + \upsilon_t \tag{2}$$

where x_t is a vector of nonstationary random variables, π is an $n \times n$ matrix, and v_t is a vector of normally distributed disturbances. The Johansen's approach involves the estimation of π and determining its rank. The implicit assumption in this test is that if the rank (π) is not equal to zero, the system exhibits symmetric adjustment around $x_t = 0$ in that for any $x_t \neq 0$, Δx_{t+1}always equals πx_t. In this case, πx_t can be regarded as an attractor with its pull strictly proportional to $\|x_t\|$ (Enders and Siklos, 1998).[1]

The Engle–Granger (1987) two-step approach also tests for symmetric adjustment. The approach uses OLS to estimate the long-run equilibrium relationship as:

$$x_{1t} = \beta_0 + \beta_2 x_{2t} + \cdots + \beta_n x_{nt} + \mu_t \tag{3}$$

where x_{it} are non-stationary variables, β_i are parameters to be estimated, and μ_t is the disturbance term which may be serially correlated. OLS is then used to estimate ρ in the following relationship:

$$\Delta \mu_t = \rho \mu_{t-1} + \varepsilon_t \tag{4}$$

where ε_t is a white noise process. The rejection of the null hypothesis of no cointegration (i.e., accepting the alternative hypothesis of $-2 < p < 0$) implies that the residuals in Equation 3 are stationary with mean zero. Equation 3 can therefore be viewed as an attractor with its pull strictly proportional to the absolute value of μ_τ (Enders and Siklos, 1998). According to the Engle–Granger (1987) theorem, if $\rho \neq 0$, 3 and 4 jointly imply the existence of an error correction model that can be represented as:

$$\Delta x_{1t} = \delta_j(x_{1t-1} - \beta_0 - \beta_2 x_{2t-1} - \ldots - \beta_n x_{nt-1}) + \sum_{j=1}^{k} \beta_{2j} \Delta x_{2,t-j} + \ldots + \sum_{j=1}^{k} \beta_{nj} \Delta x_{n,t-j} + \upsilon_{1t} \tag{5}$$

where ν_{1t} is a white noise disturbance term and k is the lag length. The term inside the brackets provides the error correction mechanism. Enders and Granger (1998) argue that the cointegration tests from the Engle–Granger and Johansen frameworks are misspecified if adjustment is asymmetric. When these tests are employed in producer-retail price transmission analysis, the implicit assumption is that the price responses are symmetric in the sense that a shock to producer prices of a given magnitude would elicit the same response in retail prices, regardless of whether the shock reflected a price increase or price decrease.

Enders and Granger (1998) consider an alternative error correction specification called threshold autoregressive (TAR) model, in which Equation 4 is represented as:

$$\Delta \mu_t = \begin{cases} \rho_1 \mu_{t-1} + \varepsilon_t & \text{if} \quad \mu_{t-1} \geq 0 \\ \rho_2 \mu_{t-1} + \varepsilon_t & \text{if} \quad \mu_{t-1} < 0 \end{cases} \tag{6}$$

A necessary condition for $\{\mu_t\}$ to be stationary is: $-2 < (\rho_1, \rho_2) < 0$. Enders and Granger (1998) show that if the sequence is stationary, the least squares estimates of ρ_1 and ρ_2 have an asymptotic multivariate normal distribution. The adjustment process is then formally quantified as:

$$\Delta \mu_t = \mathrm{I}_t \rho_1 \mu_{t-1} + (1 - \mathrm{I}_t) \rho_2 \mu_{t-1} + \varepsilon_t \tag{7}$$

where I_t is the Heaviside indicator function such that:

$$\mathrm{I}_t = \begin{cases} 1 & \text{if} \quad \mu_{t-1} \geq 0 \\ 0 & \text{if} \quad \mu_{t-1} < 0 \end{cases} \tag{8}$$

where 0 represents a critical threshold. Models using 7 and 8 are referred to as threshold autoregression models (TAR), while the test for threshold behaviour of the equilibrium error is termed threshold cointegration test. Assuming the system is convergent, $\mu_t = 0$ can be considered as the long-run equilibrium value of the sequence. If μ_t is above its long-run equilibrium value, the adjustment is $\rho_1 \mu_t$ and if μ_t is below its long-run equilibrium, the adjustment is $\rho_2 \mu_t$. The equilibrium error therefore behaves like a threshold autoregression.

Given that adjustment is symmetric if $\rho_1 = \rho_2$, the Engle–Granger approach turns out to be a special case of Equations 7 and 8. The consistency of Equations 3, 7 and 8 with a wide variety of error correction models, allow an error correction representation for the system. Given the existence of a cointegrating vector in the form of Equations 3, the error correction representation as presented in (5) can be written as:

$$\Delta x_{1t} = \rho_{1.1} I_t \mu_{t-1} + \rho_{2.1}(1 - I_t)\mu_{t-1} + \sum_{j=1}^{k} \beta_{2j} \Delta x_{2,t-j} + \cdots + \sum_{j=1}^{k} \beta_{nj} \Delta x_{n,t-j} + \upsilon_{1t} \tag{9}$$

where $\rho_{1.1}$ and $\rho_{2.1}$ are the adjustment coefficients for positive and negative discrepancies, respectively.

[1] $\|x_t\|$ denotes the Euclidean norm of the vector x_t. The norm (or length) of a vector $x = (x_1, x_2, \ldots x_n)$ is defined as $\|x\| = \sqrt{x_1^2 + x_2^2 + \cdots + x_n^2}$ (see, for example, Berck and Sydsaeter, 1991 for detailed discussion on norms of vectors and matrices).

Enders and Granger (1998) show that Equation 7 can be augmented with lagged changes in the $\{\mu_t\}$ sequence such that it becomes a p-th order process:

$$\Delta\mu_t = \mathrm{I}_t\rho_1\mu_{t-1} + (1 - I_t)\rho_2\mu_{t-1} + \sum_{i=1}^{p-1}\gamma_i\Delta\mu_{t-i} + \varepsilon_t \quad (10)$$

Diagnostic checks of the residuals (such as the autocorrelogram of the residuals and Ljung–Box tests) and various model selection criteria (such as Akaike Information Criteria (AIC) or Schwartz Bayesian Criteria (SBC) are, however, required to determine the appropriate lag length, if one chooses to work with specifications such as Equation 10.

Instead of estimating Equation 7 with the Heaviside indicator 8 depending on the level of μ_{t-1}, the decay could also be allowed to depend on the previous period's change in μ_{t-1}. The Heaviside indicator could then be specified according to the following rule:

$$I_t = \begin{cases} 1 & \textit{if} \quad \Delta\mu_{t-1} \geq 0 \\ 0 & \textit{if} \quad \Delta\mu_{t-1} < 0 \end{cases} \quad (11)$$

According to Enders and Granger (1998), replacing 8 by 11 is especially valuable when adjustment is asymmetric such that the series exhibits more 'momentum' in one direction than the other. Models estimated using Equations 3, 7 and 11 are termed Momentum-Threshold Autoregression Models (M-TAR).

Thus, for the TAR model if, for example, $-2 < \rho 1 < \rho 2 < 0$, the negative phase of the $\{\mu_t\}$ sequence will tend to be more persistent than the positive phase. For the M-TAR model if, for example, $|\rho_1| < |\rho_2|$ the M-TAR model exhibits relatively less decay for positive values of $\Delta\mu_{t-1}$ than for negative values of $\Delta\mu_{t-1}$. Thus, while the TAR model allows the degree of autoregressive decay to depend on the state of the variable of interest, the M-TAR allows the degree of autoregressive decay to depend on the first differences of the variable.

The TAR model can capture asymmetrically 'deep' movements in a series, while the M-TAR model is particularly useful in capturing the possibility of asymmetrically sharp or 'steep' movements in a series (Enders and Granger, 1998). As demonstrated by Sichel (1993), negative deepness (i.e., $|\rho_1| < |\rho_2|$) of a time series implies the exhibition of negative skewness relative to the mean or trend. That is, it should have fewer observations below its mean or trend than above, but the average deviation of observations below the mean or trend should exceed the average deviation of observations above. A positive deepness suggests the opposite. In contrast, if a time series exhibits steepness, then its first differences should exhibit negative skewness. That is, the sharp decreases in the series should be larger, but less frequent, than the more moderate increases in the series. Depending on the magnitudes of $|\rho_1|$ and $|\rho_2|$, in the TAR and M-TAR estimations, the extent of deepness or steepness can be determined.

The test statistics for the null hypothesis ($\rho_1 = \rho_2 = 0$) using the TAR specification of 7 and 8, and the M-TAR specification of 7 and 11 are called Φ_μ and Φ^*_μ, respectively. Three main factors determine the distributions of Φ_μ and Φ^*_μ. These include the number of lags of μ_t in Equation 10, the number of variables and the type of deterministic elements included in the cointegrating relationship. The appropriate critical values for Φ_μ and Φ^*_μ are tabulated in Enders and Siklos (1998) and Enders and Granger (1998).

IV. AN APPLICATION TO THE SWISS PORK MARKET

The relatively high concentration in the Swiss pork processing sector makes it interesting to study. The packaging sector has a three-firm concentration ratio of over 80% (Koch, 1998). The approach described in the previous section is now used to estimate the relationship between producer prices and retail prices of pork in Switzerland. In line with previous studies in this area (e.g., Palaskas, 1995; von Cramon-Taubadel and Loy 1997), the cointegrating regression is specified as:

$$PR_t = \beta_0 + \beta_1 PP_t + \mu_t \quad (12)$$

where PR_t is retail price, PP_t is producer price, β_0 is a constant, and μ_t the residual series to be tested for stationarity to establish long-run relationship between PR_t and PP_t. Price transmission requires a long-run relationship between PR_t and PP_t after the disappearance of transient effects such as exogenous elements affecting production and consumption.

Data and stationarity test

The data used in this analysis is based on 117 monthly observations of producer and retail prices for pork in Switzerland from January 1988 to September 1997. The monthly retail prices are weighted prices of pork cuts in Swiss Francs per kg, obtained from the Swiss Statistical Services. The producer prices are represented by pig prices in Swiss francs per kg slaughterweight, obtained from the Swiss Farmers' Union.[2]

The hypothesis that the price series concerned are nonstationary, is tested using the augmented Dickey-Fuller (ADF) test. The AIC is used to determine the appropriate lag-length truncation, which is found to be two in both cases. The estimated values for the producer price series

[2] Details of the weighting mechanism can be found in Koch (1998).

are 2.15 and 5.92 for levels and first difference, respectively, while the corresponding values for the retail price series are 2.18 and 5.06. These compare with the critical value of 2.89, leading to the conclusion that both series are first difference stationary. To ensure that the series are first difference stationary, a second test by Kwiatkowski *et al.* (1992) was employed. The estimated values for the producer price series are 0.905 and 0.106 for levels and first difference, respectively, while the corresponding values for the retail price series are 0.793 and 0.128. The appropriate one-sided critical value is 0.463 (see Kwiatkowski *et al.*, 1992: p. 166). Taken together, these results suggest that both series are $I(1)$.

Threshold cointegration estimations

Following the cointegration analysis in the Engle–Granger (1987) sense, 12 is estimated by ordinary least squares (OLS). The estimated long-run equilibrium relationship (with t-statistics in parentheses) is:

$$PR_t = \underset{(27.56)}{1.863} + \underset{(16.47)}{0.581}\, PP_t + \hat{\mu}_t \tag{13}$$

As documented in Koch (1998), the BSE (Bovine Spongiform Encephalopathy) crisis initiated in March 1996 led to a decline in beef consumption and an increase in pork demand, exerting an upward pressure on pork prices in April 1996. The possibility that this BSE crisis may have initiated a structural break in the cointegrating vector in the first half of 1996 is therefore considered. The sequential Chow test with April 1996 as the breakpoint is employed to test the hypothesis of parameter stability. The null hypothesis of no structural break over the period cannot be rejected because the test yields $F(2, 113) = 1.05$ and the 5% significance level critical value equals 3.01, suggesting parameter stability.

Following the Engle–Granger procedure, the residuals of Equation 13 are used to estimate:

$$\Delta\hat{\mu}_t = \rho_1\hat{\mu}_{t-1} + \gamma_1\Delta\hat{\mu}_{t-1} + \varepsilon_t \tag{14}$$

As reported in Table 1, the estimated value of $\rho_1 = -0.2083$ and the t-statistic for the null hypothesis that $\rho_1 = 0$ is -3.02. The critical values for the Engle–Granger test are -3.03, -3.37 and -4.07 at the 10%, 5%, and 1% significance levels, respectively. Hence at conventional levels, the Engle–Granger test indicates that the two price series are not cointegrated.[3] The Ljung–Box Q-statistics reported in Table 1 indicate that the residuals of Equation 14 are not significantly correlated.

The TAR models are estimated next in the form of Equation 10 for different lag lengths. Here again, a model augmented by one lagged change in $\{\Delta\mu_t\}$ is selected by both AIC and SBC. The sample value of Φ_μ-statistic $= 6.30$ is greater than the critical value at the 5% level, which is 5.98. The null hypothesis of $\rho_1 = \rho_2 = 0$ can therefore be rejected, indicating that the series are cointe-

Table 1. *Estimates of price transmission in the Swiss pork market*

	Engle–Granger	Threshold	Momentum-threshold	Momentum-consistent[1]
ρ_1	-0.2083 (-3.06)[1]	-0.1157 (-1.354)	-0.0644 (-0.707)	-0.0808 (-0.971)
ρ_2	NA	-0.3391 (-3.375)[2]	-0.3552 (-3.863)	-0.3702 (-3.790)
SBC[3]	-263.28	-268.26	-270.55	-399.12
AIC	-269.85	-271.42	-278.71	-386.57
Φ_μ[4]	NA	6.303	7.56	11.25
$\rho_1 = \rho_2$[5]	NA	3.093 (0.081)	5.410 (0.023)	12.51 (0.000)
Q(4)[6]	0.62	0.67	0.69	0.73
Q(8)	0.78	0.81	0.86	0.89
Q(12)	0.86	0.88	0.84	0.83

Notes:
1. Entries in this row are the t-statistics for the null hypothesis $\rho_1 = 0$.
2. Entries in this row are the t-statistics for the null hypothesis $\rho_2 = 0$.
3. The *SBC* is calculated as: $T^*(SSR) + n^*\log(T)$; where: $T =$ number of usable observations, $SSR =$ sum of squared residuals and $n =$ number of regressors. The *AIC* is calculated as $T^*(SSR) + 2^*n$ where: $T =$ number of usable observations, $SSR =$ sum of squared residuals and $n =$ number of regressors (Enders, 1995).
4. Entries in this row are the sample values of Φ_μ and Φ_μ^*. Critical values for Φ_μ for a two variable case and one-lagged are 4.99, 5.98, and 8.21 for 10%, 5%, and 1%, respectively. The corresponding values for Φ_μ^* are 5.43, 6.45, and 8.75. Critical values are from Enders and Siklos (1998).
5. Entries in this row are the sample F-statistic for the null hypothesis that the adjustment coefficients are equal. Significance levels are in parentheses below.
6. Q(p) is the significance level of the Ljung–Box statistic that the first p of the residual autocorrelations are jointly equal to zero.
7. No intercept was included in the momentum-consistent model since the mean does not significantly differ from the attractor.

[3] The Johansen procedure is not able to detect a long-run equilibrium relationship between the two price series at conventional significance levels.

grated. Given that the price series are cointegrated, the null hypothesis of symmetric adjustment (i.e., $\rho_1 = \rho_2$) can be tested using a standard F-distribution (Enders and Granger, 1998). The sample value of $F = 3.09$ has a p-value of 0.08. The null hypothesis of symmetric adjustments is accepted only at the 10 per cent level of significance.

The M-TAR model is now estimated. Both the AIC and SBC select a model augmented by one lagged change in $\{\Delta\mu_t\}$. The sample value of the Φ^*_μ-statistic is 7.56, which is above the critical value of 6.45 at the 5% level of significance, indicating that the null hypothesis of $\rho_1 = \rho_2 = 0$ can be rejected. Hence, the two series are cointegrated, supporting the previous results. A test of symmetric adjustments ($\rho_1 = \rho_2$) using the F-distribution gives a sample value of 5.41 with a p-value of 0.013. In this case, it is possible to reject the null hypothesis of symmetric adjustment at the 5% level of significance. This finding suggests that adjustment is asymmetric in the sense that positive shocks to the marketing margin tend to persist but negative shocks revert quickly towards the attractor.

Since both the TAR and M-TAR models suggest asymmetric adjustment mechanism for the series, it would be interesting to ascertain whether adjustment follows a TAR or M-TAR process. For such a test, Enders and Granger (1998) suggest using the SBC or AIC test values to select the model with the best overall fit. As is evident in Table 1, the M-TAR model yields the lowest SBC and AIC and is therefore preferable to the TAR model for explaining asymmetric adjustment of the producer-retail margin.

Tong (1983) argues that if adjustment is asymmetric, the sample mean of the residuals is a biased estimate of the attractor. Enders and Granger (1998) show that in a TAR model such as 6 and 7, if $-2 < \rho_1 < \rho_2 < 0$, the $\{\mu_t\}$ sequence will exhibit relatively more persistence whenever $\mu_{t-1} < 0$. In such cases, the sample mean will exceed that of the attractor. Chan (1993) demonstrates that searching overall values of possible attractors to minimize the sum of squared errors from the fitted model yields a super-consistent estimate of the threshold. Chan's (1993) method was therefore employed to determine the consistent threshold estimate, which turned out to be -0.021.[4] Using this estimate, the M-TAR model estimate as reported in Table 1 (with t-statistics in parentheses) is:

$$\Delta\hat{\mu}_t = -\underset{(-0.971)}{0.0808}\, \mathrm{I}_t \hat{\mu}_{t-1} - \underset{(-3.786)}{0.3722}(1 - \mathrm{I}_t)\hat{\mu}_{t-1} + \underset{(2.0873)}{0.0163}\, \Delta\hat{\mu}_{t-1} + \hat{\varepsilon}_t \tag{15}$$

where

$$\mathrm{I}_t = \begin{cases} 1 & \text{if} \quad \Delta\mu_{t-1} \geq -0.021 \\ 0 & \text{if} \quad \Delta\mu_{t-1} < -0.021 \end{cases} \tag{16}$$

Diagnostic checking indicates there is no evidence of serial correlation in the estimated equation. The AIC and SBC confirmed a model augmented with one lagged change in $\{\Delta\mu_t\}$. The sample value Φ^*_μ-statistic for the test that ($\rho_1 = \rho_2 = 0$) is 11.25, with a critical value of 6.45 at the 5% level of significance. The F-value for symmetric adjustment ($\rho_1 = \rho_2$) is 12.51, which can be rejected at any conventional significance level, supporting the hypothesis of asymmetric adjustment. Based on the AIC and SBC shown in Table 1, it is apparent that the M-TAR model with the consistent estimate of the threshold fits the data better than the other models.

Asymmetric error correction models

The positive finding of asymmetric cointegration with the M-TAR adjustment implies that it is incorrect to examine the short-run dynamics with a symmetric error correction model. A symmetric error correction model would not reveal differential adjustments of the positive and negative changes (Enders and Granger, 1998). Hence, asymmetric error correction models are employed in the analysis. They can be represented as:

$$\Delta RP_t = \sum_{s=1}^{k} \alpha_s \Delta RP_{t-s} + \sum_{s=0}^{k} \beta_s \Delta PP_{t-s} + \gamma_1 Z\_plus_{t-1} + \gamma_2 Z\_\min us_{t-1} \tag{17}$$

where k is lag-length, Z_plus_{t-1} and Z_minus_{t-1} are the error correction terms from the threshold cointegration regressions, representing adjustments to positive and negative shocks to the marketing margin, respectively. They can be represented as:

$$Z\_plus_{t-1} = I_t(RP_{t-1} - 1.863 - 0.581PP_{t-1})$$

$$Z\_minus_{t-1} = (1 - I_t)(RP_{t-1} - 1.863 - 0.581PP_{t-1})$$

and I_t is the momentum Heaviside indicator function 16. Given that the response of retail prices to changes in producer prices is generally not instantaneous but rather distributed over time, the SBC is used to determine the lag-length.

Table 2 presents the results of the error correction models. For purposes of comparison, estimates of both asymmetric and symmetric error correction models are presented. Estimates for the asymmetric adjustment are

[4] The threshold estimate can be considered as the long-run equilibrium value of the sequence. Although the threshold estimate does not seem to be significantly different from the attractor (zero) used in estimating the M-TAR model, both the AIC and BIC suggest this consistent estimate is preferable to the zero attractor.

Table 2. *Estimates of the error correction models*[a]

	Asymmetric error correction		Symmetric error correction	
	ΔRP_t	ΔPP_t	ΔRP_t	ΔPP_t
Constant	0.127 (1.146)	0.108 (1.552)	0.082 (0.936)	0.094 (1.407)
ΔRP_{t-1}	0.134 (2.658)	0.328 (1.185)	0.114 (2.649)	0.297 (1.610)
ΔRP_{t-2}	0.099 (1.667)	0.203 (1.136)	0.102 (2.078)	0.226 (1.358)
ΔRP_{t-3}	0.074 (1.296)	−0.093 (1.665)	0.069 (1.249)	0.184 (0.973)
ΔRP_{t-4}	0.059 (1.802)	0. 117 (0.988)	0.083 (0.772)	0.062 (1.029)
ΔPP_t	0.360 (3.822)		0.314 (2.056)	
ΔPP_{t-1}	0.313 (2.973)	0.214 (2.684)	0.256 (1.514)	0.204 (2.272)
ΔPP_{t-2}	0.192 (1.886)	0.124 (1.913)	0.127 (1.176)	0.118 (1.443)
ΔPP_{t-3}	0.124 (1.178)	−0.072 (0.083)	0.080 (1.012)	0.067 (1.039)
ΔPP_{t-4}	0.069 (1.703)	0.105 (1.009)	0.098 (0.926)	0.042 (0.731)
Z_plus_{t-1}[b]	0.074 (1.692)	0.038 (1.308)		
Z_minus_{t-1}	0.221 (3.460)	0.146 (0. 983)		
ECT_{t-1}[c]			0.103 (1.672)	0.058 (1.423)
$Q(4)$[d]	1.81 (0.77)	1.82 (0.76)	0.89 (0.92)	0.89 (0.92)
ARCH(χ_1^2)[e]	0.27	1.53	2.28	1.96
Jarque-Bera (χ_2^2)[f]	1.15	3.58	2.97	4.86

a. t-statistics are in parentheses. Subscripts RP and PP are used to represent retail and producer prices, respectively.
b. $Z\_plus(t-1)$ and $Z\_minus(t-1)$ are error correction terms showing adjustments to increasing and decreasing deviations from the long-run, respectively.
c. ECT_{t-1} is the error correction term showing adjustments to deviations from the long-run in the symmetric model.
d. The Q-statistics denote the Ljung–Box statistic that the first four of the residual autocorrelations are jointly equal to zero. The significance level is in parentheses below.
e. Test for first order ARCH residuals.
f. Jacque–Bera test for normality in the residuals.

presented in the first two columns, followed by the estimates from the symmetric error correction models. The t-statistics for Z_plus_{t-1} and Z_minus_{t-1} in the column for retail prices indicate that retail prices respond strongly to negative shocks, but positive shocks in the margin are allowed to persist. However, the t-statistics for Z_plus_{t-1} and Z_minus_{t-1} in the column for producer prices suggest that producer prices do not seem to respond to negative or positive changes in the margin. Thus, retail prices adjust to correct long-run disequilibria in retail and producer prices, while producer prices do not significantly respond to long-run disequilibria.

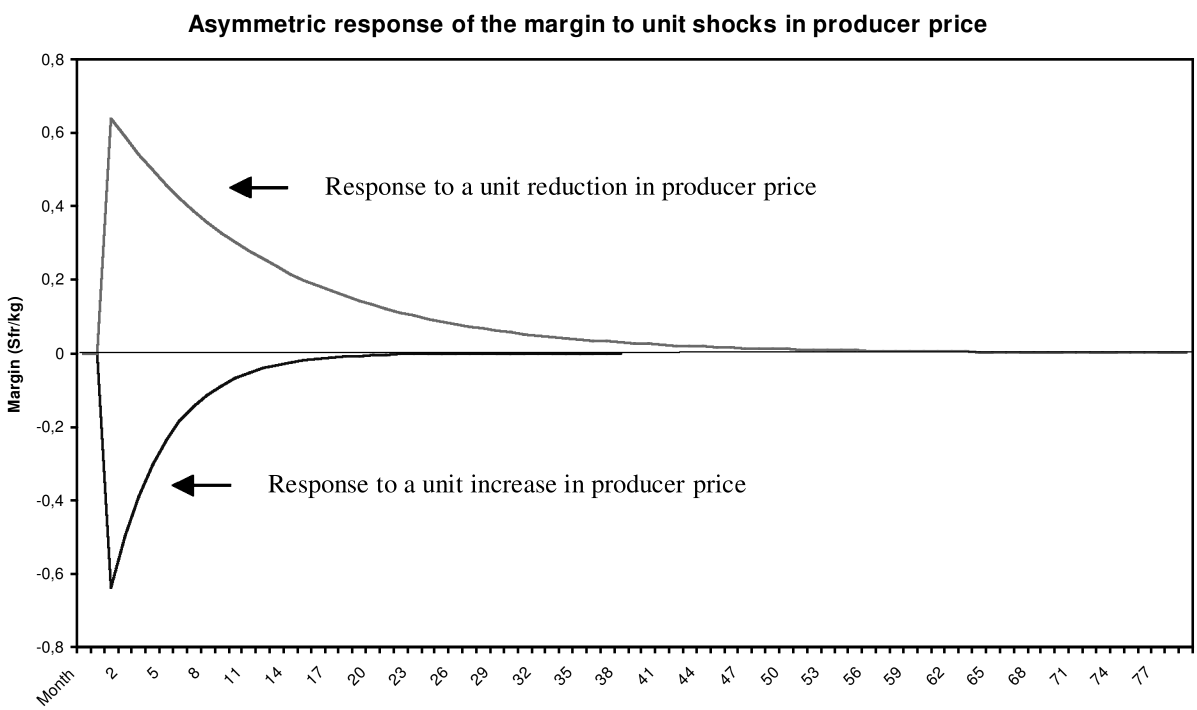
The results also show that both contemporaneous and lagged changes in producer prices induce statistically significant responses from retail prices. However, no significant response is realized by producer prices to shocks in retail prices. The $F_{(5,112)}$-statistic for the null hypothesis that current and lagged changes in producer prices do not affect retail prices is 3.86, while the $F_{(4,\ 113)}$-statistic for the null hypothesis that lagged changes in retail prices do not affect producer prices is 1.32, against a critical value of 3.08 at the 5% level of significance. These results suggest there is uni-directional causation from producer to retail prices. Because livestock are non-storable commodities subject to biological production lags with highly inelastic supply in the short run, producers are unable to adjust production in response to transitory price changes. By contrast, retailers can immediately respond to changes in producer prices by adjusting their prices.

Focusing on adjustments of retail prices to restore equilibrium, the point estimates of the adjustment coefficients given in Table 2 indicate that, within a month, retail prices adjust so as to eliminate approximately 22% of a unit negative change in the deviation from the equilibrium relationship created by changes in producer prices. On the other hand, retail prices adjust by only 7% of a positive change in deviation from the equilibrium created by changes in producer prices. These findings indicate that adjustments towards the long-run equilibrium relationship between producer and retail prices is faster when changes in deviations are negative (i.e., when producer prices rise to decrease the marketing margin) than when they are positive (i.e., when producer prices decline to increase the marketing margin).

The t-statistics from the symmetric error-correction specification indicate that the error-correction terms in the symmetric equations are not significant at conventional levels, implying that there is no convergence towards long-run equilibrium. An interesting point is that the multivariate values of AIC and SBC indicate that the asymmetric model is preferable to the symmetric model. The multivariate AIC and SBC are −453.74 and −426.59, respectively, for the asymmetric error correction system given in the second and third columns in Table 2. These values are lower than the corresponding figures of -424.27 and -411.78 for the linear error correction model system given by the fourth and fifth columns in Table 2, in spite of the additional coefficients appearing in the asymmetric error correction models.

The estimates of the asymmetric error correction model presented in Table 2 are used to develop impulse response functions to further examine the dynamic interrelationship among the producer and retail prices. As pointed out by Potter (1995), the response to a price shock in symmetric adjustment models is independent of the history of the time series and the sign and magnitude of the postulated shock. By contrast, asymmetric adjustment models produce impulse response functions that are themselves functions of the history of the price series and the sign and magnitude of the shock. The results of the asymmetric error correction model imply that a unit increase in producer price (negative shock to the marketing margin) leads to an increase of 0.36 units in retail price. This results in a decline of the marketing margin by 0.64 units, which is corrected asymptotically by a factor of about 0.221 per period in the following months as the producer price continues to grow. Figure 1 illustrates responses to one positive and negative shocks in the marketing margin. As shown in the lower panel of Fig. 1, the margin returns to its equilibrium level within 20 months. This is in contrast to a decline in producer price (positive shock to the marketing margin) that induces a reduction in retail price of about 0.36 units. The resulting excess margin of 0.64 units is corrected by a factor of 0.074 units per period in the ensuing months, taking almost 56 months to return to equilibrium level. These results demonstrate that there is a tendency for retail price level movements to reverse themselves with negative, but not positive shocks to the margin. This is consistent with the notion that the margin is corrected more rapidly when squeezed relative to its long-run level than when stretched. These findings are consistent with the findings of Hahn (1990) for the US pork market, but are in contrast with the results reported by Boyd and Brorsen (1988) for the US pork market.

The results generally suggest the presence of price transmission asymmetries in the Swiss pork market. As pointed out earlier, the three-firm concentration ratio in the Swiss pork market is over 80%. Very restrictive food regulations also serve as entry barriers for newcomers into this sector. In addition, the restriction of access to import quotas to few retailers in the sector tends to protect them from competition. However, the findings of this study do not imply market power or supernormal profits among retailers in the Swiss pork market. Although some of the hypotheses discussed earlier do suggest that temporary market power could explain the asymmetry, the other explanations are consistent with competitive markets. Thus, there is the need for further research to explain the underlying price transmission asymmetry.



***Source*: Author's calculations**

Fig. 1. *Asymmetric response of the margin to unit shocks in producer price*

V. CONCLUDING REMARKS

Recent studies of price transmission in the food marketing chain have suggested that middlemen use market power to pass on input price increases to consumers more rapidly and probably more completely, than input price reductions. This paper has employed recent statistical techniques to examine the long- and short-run relationship between producer and retail prices, giving special attention to the time series properties of the price data. Using Swiss pork price data over the 1988–1997 sample period, both the threshold autoregressive (TAR) and momentum-threshold autoregressive (M-TAR) models provide strong and clear evidence supporting asymmetric pricing behaviour on the part of pork retailers in Switzerland. Increases in producer prices of pork that result in the reduction of the marketing margin appear to be passed on to retail prices faster than reductions in producer prices that lead to increases in the marketing margin. Thus, the marketing margin is corrected more rapidly when squeezed relative to its long-run level, than when stretched.

The estimates of symmetric and asymmetric adjustment error-correction models were also compared to further examine the nature of the short-run adjustment process. The asymmetric error-correction model yielded a dynamic path consistent with adjustments to eliminate deviations from the long-run equilibrium. In contrast, the symmetric error-correction specification did not show any dynamic path that adjusts to eliminate deviations from the long-run equilibrium. An evaluation of an impulse response function also suggests that price transmission asymmetries do exist in the Swiss pork market. Further, Granger-causality tests indicate a uni-directional relationship from producer prices to retail prices with no evidence of the reverse causality feedback. This result substantiates the notion that retailers do adjust to shocks in producer prices, while effects of retail market shocks are largely confined to retail markets. Future empirical work in this area should strive to explain the reasons underlying observed asymmetric price transmission.

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