Considering threshold effects in the long-run equilibrium in a vector error correction model: An application to the German apple market

Linde Goetz[1] and Stephan von Cramon-Taubadel[2]

[1] Institute of Agricultural Development in Middle and Eastern Europe (IAMO), Halle/Saale, Germany
[2] Department of Agricultural Economics and Rural Development, University of Goettingen, Germany

Abstract — **We propose a three-step procedure to estimate a regime-dependent vector error correction model (VECM). In this model, not only the short-run adjustment process towards equilibrium is non-linear, as in threshold VECM and Markov switching VECM frameworks, but the long-run equilibrium relationship itself can also display threshold-type non-linearity. The proposed approach is unique in explicitly testing the null hypothesis of linear cointegration against the alternative of threshold cointegration based on the Gonzalo AND PITARAKIS (2006) test. The model is applied to apple price data on wholesale markets in Hamburg and Munich, using the share of domestic apples in total wholesale trade as the threshold variable. We identify four price transmission regimes characterized by different equilibrium relationships and short-run adjustment processes. This proposed approach is particularly suitable for capturing irregular seasonal threshold effects in price transmission typical for fresh fruits and vegetables.**

Keywords— **threshold cointegration, spatial price transmission, vector error correction model**

I. INTRODUCTION

Applications of the threshold vector error correction model (TVECM) to analyze price transmission assume that prices are linked by a constant long-run equilibrium relationship, while allowing for threshold or switching effects in the short-run adjustment process towards this equilibrium. The TVECM (e.g. GOODWIN AND PIGGOTT, 2001; MEYER, 2004; SERRA, GIL AND GOODWIN, 2006; BALCOMBE, BAILEY AND BROOKS, 2007) distinguishes between regimes depending on whether the deviation of prices from their long-run equilibrium, in other words the error correction term (ECT), is above or below a threshold value. For example, if the ECT exceeds a specific threshold which is determined by the size of the transaction costs, then more rapid adjustment to the constant long-run equilibrium is expected than if the ECT is smaller than the threshold value, in which case adjustment might even cease altogether. In the Markov-switching VECM (e.g. BRUEMMER ET AL., 2008), shifts between different adjustment regimes are triggered by unobservable state variables. Both models maintain the hypothesis of a linear long-run equilibrium relationship. This may not always be justifiable. For example, if product qualities or the direction of trade between two markets changes, then the long-run relationship between the prices on these markets may change as well. Failing to account for non-linearity in the long-run relationship can lead to misleading estimates of this relationship and the adjustment processes that lead to it.

In this paper we propose a three-step procedure to estimate a regime-dependent VECM. In this model, not only the short-run adjustment process towards equilibrium, but also the long-run equilibrium relationship itself can display threshold-type non-linearity, as a function of the size of a stationary variable with respect to a threshold value. The proposed approach is unique in explicitly testing the null hypothesis of linear cointegration against the alternative of threshold cointegration based on a test proposed by GONZALO AND PITARAKIS (2006). As GONZALO AND PITARAKIS (2006) point out, the use of the term 'threshold cointegration' in connection with threshold VECMs is misleading because in a threshold VECM it is actually the adjustment or error correction that is subject to threshold effects, while the cointegration itself (i.e. the long-run relationship) is assumed to be constant and linear.

We apply this procedure to data on daily apple prices on wholesale markets in Hamburg and Munich. Due to substantial seasonal variation in supply quantities, prices and price differences, we hypothesize that the equilibrium relationship between prices in Hamburg and Munich is subject to threshold effects, with the share of German as opposed to imported apples in total wholesale trade acting as the threshold variable.

We proceed as follows. Chapter 2 contains a literature review; chapter 3 presents the GONZALO AND PITARAKIS (2006) test and a three step procedure based on this test to study threshold cointegration in a regime-specific VECM. Chapter 4 describes the seasonal characteristics of supply and price determination on wholesale apple markets in Hamburg and Munich. Estimation and results are presented in chapter 5, and chapter 6 concludes.

II. METHODS

A. Test on threshold effects in cointegration

GONZALO AND PITARAKIS (2006) propose a test of the null hypothesis of linear cointegration:

(1) $y_t = \boldsymbol{\beta}'\mathbf{x}_t + u_t$

against the alternative hypothesis of cointegration with threshold effects:

(2) $y_t = \boldsymbol{\beta}'\mathbf{x}_t + \boldsymbol{\lambda}'\mathbf{x}_t I(q_{t-d} > \gamma) + u_t$

with $\mathbf{x}_t = \mathbf{x}_{t-1} + \mathbf{v}_t$, where u_t and v_t are scalar and *p*-vector valued stationary disturbance terms respectively, q_{t-d} with $d \geq 1$ is a stationary threshold variable lagged by d periods, and $I(q_{t-d} > \gamma)$ is an indicator function that equals one if $q_{t-d} > \gamma$, and zero otherwise.

GONZALO AND PITARAKIS (2006) propose a supLM test based on the following statistic:

(3) $LM_T(\gamma) = \frac{1}{\tilde{\sigma}_0^2} \mathbf{u}'\mathbf{M}\mathbf{X}_\gamma (\mathbf{X}_\gamma' \mathbf{M}\mathbf{X}_\gamma)^{-1} \mathbf{X}_\gamma' \mathbf{M}\mathbf{u}$

where $\mathbf{M} = \mathbf{I} - \mathbf{X}(\mathbf{X}'\mathbf{X})^{-1}\mathbf{X}'$, $\mathbf{X}$ stacks all values of $\mathbf{x}_t$ in the linear model (1), and $\mathbf{X}_\gamma$ stacks the values of $\mathbf{x}_t$ corresponding to the criterion $q_t > \gamma$ in the non-linear model (2). T is the length of the full sample, u is the residual, and $\tilde{\sigma}_0^2$ is the residual variance of the linear model (1).

The LM test statistic $LM_T(\gamma)$ is calculated for all possible values of the threshold variable q_t. A trimming parameter is employed to ensure a minimum number of observations on each side of the threshold. The supLM test statistic is given by

(4) $\sup LM = \sup_{\gamma \in \Gamma} LM_T(\lambda)$.

Critical values for this test statistic are taken from ANDREWS (1993).

B. A three-step procedure for estimating the threshold cointegration model

We propose the following three-step procedure to estimate a regime-specific VECM which includes non-linearities not only in the short-run and equilibrium adjustment process but also in the long-run equilibrium relationship between prices in question.

First, since the test for threshold cointegration by GONZALO AND PITARAKIS (2006) requires that the time series data be integrated of order 1, we determine the order of integration of the data series by conducting unit root tests.

Second, we test the null hypothesis of linear cointegration

(5) $y_t = \alpha_0 + \alpha_1 * x_t + u_t$

against the alternative hypothesis of threshold cointegration:

(6) $y_t = (\alpha_0 + \alpha_1 * x_t) + (\lambda_0 + \lambda_1 * x_t) I(q_{t-d} > \gamma) + u_t$

utilizing the supLM test proposed by GONZALO AND PITARAKIS (2006).

Third, we estimate an unrestricted, regime-specific ECM by including dummy variables defined by the indicator function $I(q_{t-d} > \gamma)$ corresponding to the threshold determined by the supLM test. This ECM takes the form:

(7)

$$\Delta y_t = \beta_0 + \delta_0 * I(q_{t-d} > \gamma) + \sum_{m=1}^{K} (\beta_{1m} \Delta x_{t-m+1} + \delta_{1m} \Delta x_{t-m+1} * I(q_{t-d} > \gamma))$$

$$+ \sum_{n=1}^{L} (\beta_{2n} \Delta y_{t-n} + \delta_{2m} * \Delta y_{t-n} * I(q_{t-d} > \gamma))$$

$$+\beta_3 * y_{t-1} + \delta_3 * y_{t-1} * I(q_{t-d} > \gamma) + \beta_4 * x_{t-1} + \delta_4 * x_{t-1} * I(q_{t-d} > \gamma) + \varepsilon_t$$

The regime-dependent cointegration vector can be retrieved from equation (7) as:

(8) $\alpha_0 = -(\beta_0 + \delta_0 * I(q_{t-d} > \gamma)) / (\beta_3 + \delta_3 * I(q_{t-d} > \gamma)$

and

(9) $\alpha_1 = -(\beta_4 + \delta_4 * I(q_{t-d} > \gamma)) / (\beta_3 + \delta_3 * I(q_{t-d} > \gamma)$.

Accounting for this type of non-linearity in a cointegration regression allows the long-run relationship to move back and forth between regimes as a function of a threshold variable, rather than hypothesising a one-off break in this relationship. This is an appealing model in settings, such as the one explored below (spatial trade in apples), in which price transmission is hypothesised to be seasonal, but the timing and duration of seasons differs from year to year depending on weather and harvests in the regions that are linked by trade. In such settings, the use of seasonal dummy variables to account for seasonal variation in the equilibrium relationship (e.g. CHAVAS AND MEHTA, 2004) might not be sufficiently flexible.

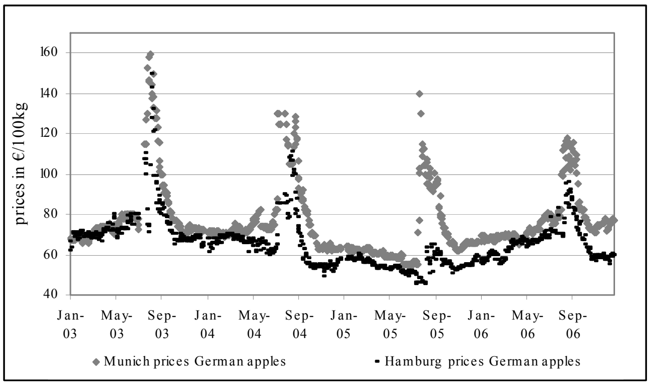
III. APPLICATION AND DATA: THE GERMAN WHOLESALE MARKETS FOR APPLES

The procedure outlined above is applied to spatial price transmission between wholesale prices for apples in Hamburg and Munich, Germany. The wholesale apple markets in Hamburg and Munich are the largest in Germany, together accounting for 42% of all apples traded on the five largest German wholesale markets between 2003 and 2006. We utilize 942 daily prices for German apples on wholesale markets in Munich and Hamburg between 2003 and 2006 (Figure 1).

About 60% of the apples produced in Germany are grown in the two largest apple growing areas: Niederelbe (8,840 ha), which is close (roughly 50 km) to the wholesale market in Hamburg; and Bodenseegebiet (7,000 ha), which is somewhat less proximate to the wholesale market in Munich (roughly 250 km). German apples are stored in large warehouses and can be supplied year-round by growers to wholesale markets. However, the supply of German apples on the wholesale markets in Hamburg and Munich is characterized by substantial seasonal variation in 1) quantities, 2) prices and 3) price differences.

First, the daily share of German apples in all apples traded on the wholesale markets in Hamburg and Munich varies seasonally between 1% and 60% (Figure 2). In addition to German apples, imports from Italy provide another ‚domestic' (i.e. intra-EU) source of apples that is roughly synchronised with German supply and accounts for up to 66% of all the apples traded on wholesale markets. Apple supply from these northern hemisphere countries is continuously high during the winter months, and decreases in spring until summer. At this time, the supply of apples originating in southern hemisphere countries (Argentina, Brazil, Chile, New Zealand, South Africa and Uruguay) increases, peaking at up to 90% of daily wholesale apple trade in early summer. When newly harvested German and Italian apples enter the market in late summer, the share of German apples traded increases steadily and the share of southern hemisphere apples drops sharply until they are driven out of the market in the fall. In the course of an average year, apples grown in Germany, Italy and southern hemisphere countries account for roughly 90% of all apples traded on wholesale markets in Hamburg and Munich.

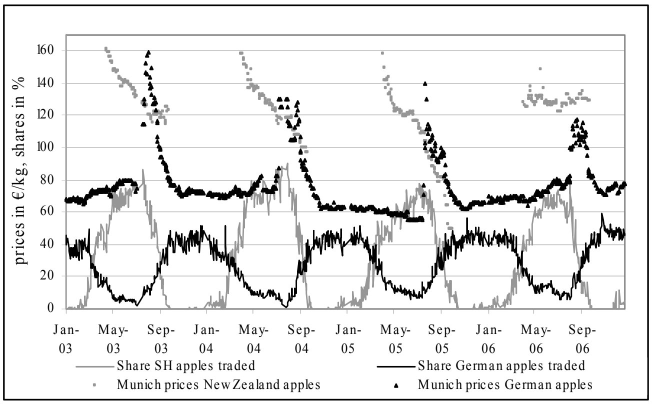
Figure 1: Prices of German apples on wholesale markets in Hamburg and Munich, 2003-2006



Data source: BLE

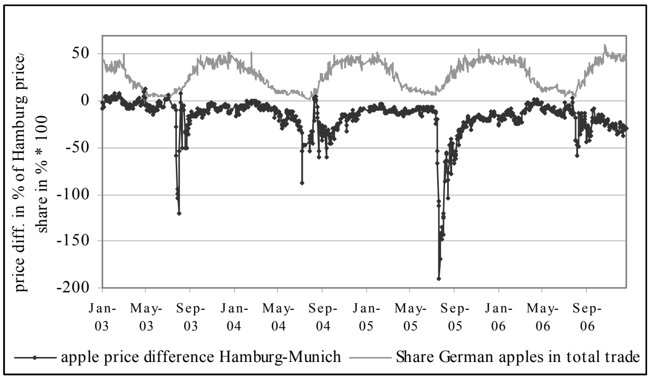
Second, the price of German apples is highest when newly harvested apples become available in late summer. Thereafter, prices drop continuously during the fall harvest, and remain almost constant during winter until spring when apples are sold from storage. In late spring, when apples from southern hemisphere countries reach a substantial market share, German apple prices slightly increase or decrease depending on the quantity and quality of southern hemisphere apples supplied. Similarly, prices of southern hemisphere apples are highest at the beginning of their season in spring, when they enter the market as newly harvested produce. At this time the price of southern

Figure 2: Prices and share of apples originating in Germany and southern hemisphere (SH) countries represented by New Zealand of all apples traded in Hamburg and Munich



Data source: BLE

Figure 3: Price difference between apple prices in Hamburg and Munich (in % of Hamburg price) and share of German apples in total daily trade



Data source: BLE; own calculations

hemisphere apples may exceed the price of German stored apples by up to 100%. Thereafter, prices of southern hemisphere apples drop continuously until the end of the season in fall.

Third, prices differ between wholesale markets in Germany. Figure 1 illustrates that the price level is higher in Munich than in Hamburg. The average difference amounts to 14%, but it varies, being relatively low and stable in the winter/spring months, and higher and more variable in late summer, when the share of German apples traded is low and newly harvested apples enter the market (Figure 3). Traders report that the Munich market demands higher quality than Hamburg, which explains the higher average prices in Munich. This is especially apparent in August/September when the first new-harvest domestic apples appear on the market and command premium prices. Furthermore, the closer proximity of the Hamburg market to the nearest production region in Germany leads to a lower transport cost component in apples prices there. Traders report that transport costs from the growing area to the wholesale market account for between 4%-7% of the wholesale price in Hamburg compared with 6%-9% in Munich.

Based on this description of the markets, we hypothesise that price transmission between the wholesale apples markets in Hamburg and Munich will be seasonally regime-dependent depending on whether these markets are mainly supplied from domestic or imported sources. However, an important characteristic of the seasonal pattern of apple prices and quantities is that it is irregular, caused by random variations in weather and the timing and quality of harvests in Germany and elsewhere. This irregular seasonality is typical for fresh fruits and vegetables markets (see e.g. RODRÍGUEZ AND HERNÀNDEZ, 2005). For example, the German apple season (defined as the date on which the share of German apples in total trade increases to over 10% for the first time in a year) started as early as July 22 in 2003 and July 14 in 2005, and as late as August 17 in 2004 and August 7 in 2006. Similarly, the beginning of the southern hemisphere apple season varies between January and March. Related to these fluctuations, the variety and quality composition of the domestic and imported apples traded in Hamburg and Munich can vary considerably from year to year.

For this reason, a modelling approach based on seasonal dummy variables would be too inflexible. Instead, we hypothesize that the equilibrium price relationship between wholesale prices in Hamburg and Munich is subject to threshold effects, with the share of German apples in total wholesale trade acting as the threshold variable. This specification allows for seasonal regime shifts to occur at different times from year to year, depending on the timing and volume of the German harvest.

V. EMPIRICAL RESULTS

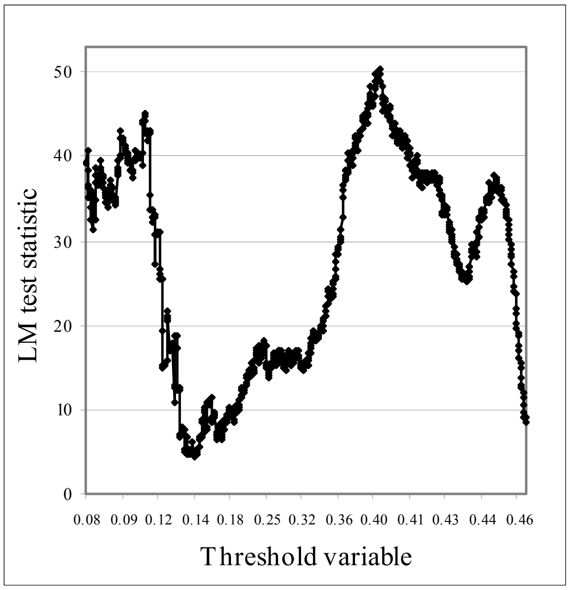
The results of the ADF test (DICKEY-FULLER, 1981) and the KPSS test (KWIATOWSKI ET AL., 1992) suggest that the wholesale apple prices in Hamburg (p^H_{GER}) and Munich (p^M_{GER}) (about 930 observations each) are I(1). Also, the Johansen test and residual based test on cointegration indicate that there is cointegration over the whole sample.

Next, we conduct the GONZALO AND PITARAKIS (2006) test for threshold cointegration between p^H_{GER} and p^M_{GER} for model (I) $p^H_{GER} = f(p^M_{GER})$ with p^H_{GER} as the dependent and p^M_{GER} as the independent variable and model (II) $p^M_{GER} = f(p^H_{GER})$ with the converse structure.

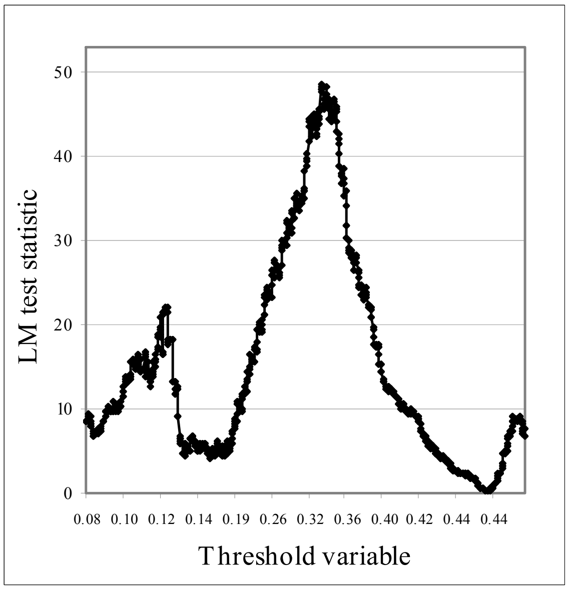
The daily share of apples produced in Germany in total wholesale trade in Hamburg and Munich is used as the threshold variable. Since this variable fluctuates from day to day, we smooth it by calculating the central moving average of the nearest 12 observations for each observation (see figure 5). In this way we avoid repeated, 'back and forth' regime changes that would otherwise occur in periods in which the variable is close to its threshold value. The LM-test statistic in (3) is estimated for all observed values of the threshold variable, with the trimming parameter is set to 0.08 to ensure that each regime contains at least 8% of all observations. Figure 4 presents the estimated value of the LM-test statistic and the corresponding value of the threshold variable for models I (panel a) and II (panel b). For model I, the value of the LM-test statistic is highest for the threshold values 0.105 and 0.399, corresponding to LM=45.20 (p-value<0.01) and LM=50.41 (p-value<0.01), respectively. For model II, the LM-test statistic is highest for the threshold values 0.121 and 0.335, corresponding to LM=48.69 (p-value<0.01) and LM=21.99 (p-value<0.01), respectively.

Figure 4: Values of the Gonzalo-Pitarakis test statistic

a) Model I ($p^H = f(p^M)$)



b) Model II ($p^M_{GER} = f(p^H_{GER})$)



Source: own calculations

Since the LM statistic displays two distinct peaks, we assume that there are two significant thresholds and correspondingly three long-run equilibria in the cointegration relationship between apple prices in Munich and Hamburg. To this end, three dummy variables are included in the model: SMALLER is defined by the indicator function $I(q_{t-d} < 0.121)$; BETWEEN is redefined by $I(0.121 < q_{t-d} < 0.335)$; and LARGER is defined by $I(0.335 < q_{t-d})$.

Furthermore, we note that the BETWEEN regime occurs twice each year; once during the transition from LARGER to SMALLER (from spring to early summer) as the share of German apples is falling; and once during the transition from SMALLER to LARGER (from late summer to fall) as the share of German apples is increasing. These two transitions represent very different market conditions: In the former, newer southern hemisphere apples progressively replace older stored apples from the last domestic crop; in the latter, the new domestic crop replaces imported southern hemisphere apples. To account for this, we divide BETWEEN into BETWEEN1 (spring) and BETWEEN2 (fall) with appropriate dummy variables. Figure 5 illustrates the mapping of observations into the four resulting regimes.

In the following, to shorten the presentation, we carry out the subsequent analysis using threshold values for model II (0.121 and 0.335) alone; results based on threshold values from model I (available on request) are qualitatively similar.

The presence of cointegration between the Hamburg and Munich prices in each of these four regimes is tested using both a residual-based (ADF) test and the Johansen trace-test (Table 1).

The results unambiguously point to cointegration in all regimes except BETWEEN1 (spring), where the results of the cointegrating ADF test only points to cointegration when Munich prices are regressed on Hamburg prices. Table 1 also shows unweighted mean and standard error of the price difference between the Munich and the Hamburg market for each regime.

The price difference and standard error are by far highest in regime BETWEEN2 (fall), followed by the regime SMALLER. The price difference and standard error is lowest for the regimes LARGER and BETWEEN1 (spring).

In the next step we estimate the unrestricted regime-dependent ECM with four regimes according to (7) in models I and II.

Our hypothesis of a regime-dependent model in which the long-run relationship (and correspondingly the adjustment process) displays threshold behaviour is supported by the results of the likelihood-ratio test. In this test, the value of the log-likelihood function of the regime-specific ECM with 4 regimes according to (7) (unrestricted model) is compared to that of an ECM over all observations without distinguishing between regimes (restricted model). The null hypothesis that the restricted model is superior to the unrestricted model is clearly rejected at low p-values in model frameworks I as well as II (Table 2).

Table 3 presents the estimates for the long-run price transmission elasticity and the speed of adjustment for the regimes SMALLER, BETWEEN2 and LARGER, for which the data series were identified as cointegrated. The t-values account for autocorrelation and heteroscedasticity in model I and for heteroscedasticity only in model II.

Results indicate that the price transmission elasticity varies significantly between regimes and model frameworks. The coefficient corresponding to the speed of adjustment to the long-run equilibrium has the correct negative sign and is statistically significant in all cases with the exception of the regime SMALLER in model framework II. This indicates that the price relationship is unidirectional in the regime SMALLER, with the Hamburg price error correcting, whereas the Munich market is dominating the price. In the other cases the price relationship is bi-directional and error correcting behaviour is identified for both markets. Furthermore, results for both models suggest that deviations from the long-run equilibrium are corrected fasted in regime BETWEEN2 in fall and slowest in regime LARGER during winter. The speed of adjustment is also quite high in the regime SMALLER in model I. Results obtained for the full data set (COMPLETE) suggest that the price relationship is unidirectional and that the Hamburg price only error corrects whereas the Munich market dominates the price.

Figure 5: Attribution of the observations of the Munich apple price to the four regimes based on thresholds retrieved from the GONZALO-PITARAKIS (2006)

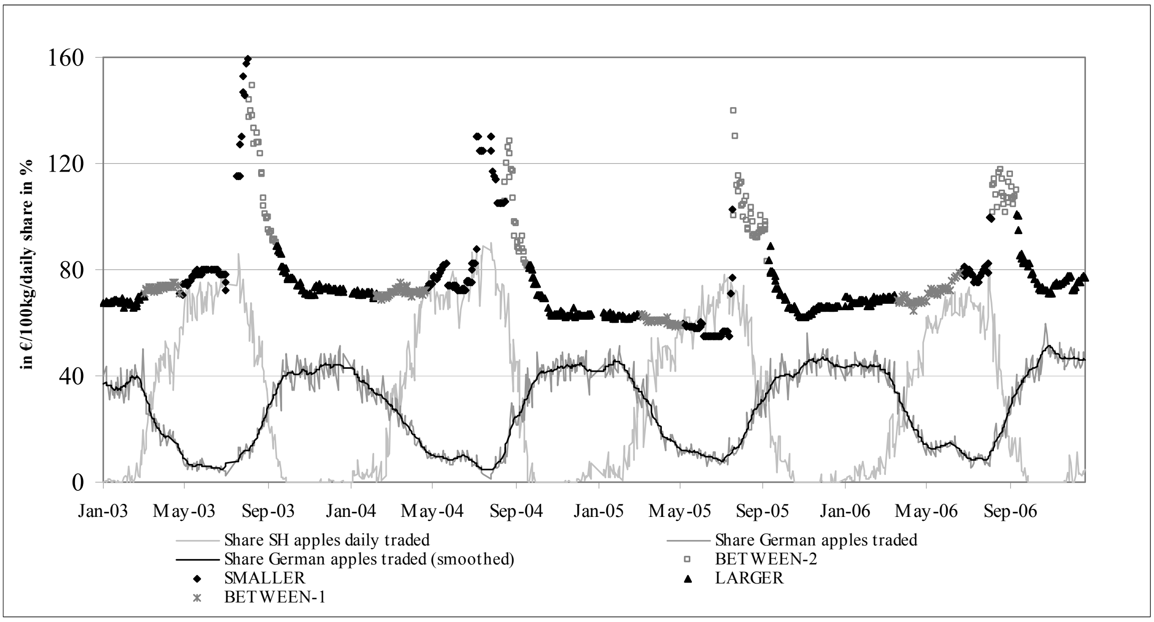


Table 1: Results test on cointegration and price volatility of the 4 regimes (Residual based test without intercept and trend), model II limits

Dummy variable Number of obsv. Threshold limits	**BETWEEN1** (spring) 185 obsv. $0.121 \le q_t \le 0.335$	**SMALLER** (summer) 214 obsv. $q_t < 0.121$	**BETWEEN2** (fall) 111 obsv. $0.121 \le q_t \le 0.335$	**LARGER** (winter) 424 obsv. $q_t > 0.335$
Test on cointegration				
a) Residual based test ADF t-value, [number of lags], (p-value) Model I $\ln(p^H_{GER}) = f(\ln(p^M_{GER}$ Model II $\ln(p^M_{GER}) = f(\ln(p^H_{GER}$	-2.34, [2 lags] (>0.1) -3.892, [0 lag] (<0.01)	4.999, [0 lag] (<0.01) -3.892, [0 lag] (<0.01)	2.998, [0 lag] (<0.05) 3.949, [1 lag] (<0.01)	-3.265, [1 lag] (<0.05) -4.557, [0 lag] (<0.01)
b) Johansen trace t. LR^{trace} zero versus one coint. rel., lags (p-value)	11.62, [2 lags] (>0.1)	25.39, [2 lags] (<0.01)	24.76, [2 lags] (<0.01)	25.63, [2 lags] (<0.05)
Price difference				
Mean (standard error) $\mid p^M_{GER} - p^H_{GER} \mid$	4.266 (2.445)	12.266 (14.435)	30.105 (16.309)	9.175 (5.341)

Table 2: Likelihood ratio test results

	Model I ($p^{H} = f(p^{M})$)	**Model II ($p^{M}_{GER} = f(p^{H}_{GER})$)**
LR-test statistic	42.524	56.835
Degrees of freedom	15	15
p-value	0.0002	<0.0001

Source: own calculations

Table 3: Estimates for the long-run price transmission elasticity and the speed of adjustment for the 4 regimes and the full sample

		Model I ($\ln(p^{H}_{GER}) = f(\ln(p^{M}_{GER}))$)		**Model II ($\ln(p^{M}_{GER}) = f(\ln(p^{H}_{GER}))$)**	
Dummy variable	**Parameter**	**Estimate**	**t-value**	**Estimate**	**t-value**
SMALLER (summer)	Price transmission elasticity	0.911		1.087	
	Speed of adjustment	-0.118	-2.077	0.034	0.870
BETWEEN2 (fall)	Price transmission elasticity	1.340		0.225	
	Speed of adjustment	-0.123	-2.510	-0.117	-2.985
LARGER (winter)	Price transmission elasticity	0.056		0.348	
	Speed of adjustment	-0.044	-3.645	-0.080	-4.054
COMPLETE	Price transmission elasticity	0.638		0.495	
	Speed of adjustment	-0.072	-3.131	-0.023	-1.470

Source: own calculations

Taking into account the time period, market condition and error-correcting behaviour, the four regimes can be characterized as follows:

Regime SMALLER (summer): Corresponds to the market conditions in May/June-July/August, when the remainder of the stored apples of the previous harvest and the first apples of the new harvest are supplied to the market. This is the only regime in which the price relationship is unidirectional with the Munich price not error correcting and thus dominating the Hamburg price. In contrast, the Hamburg price error corrects at relatively high speed. This may be attributed to the harvest season starting earlier in the southern parts of Germany implying that new apples are first sold on the Munich market. Thus, the initial price level for the new harvest is set on the Munich market and is transmitted to the Hamburg market. Price differences between the Hamburg and the Munich market are relatively high giving leeway to profitable arbitrage opportunities implying strong market integration.

Regime BETWEEN2 (fall): This regime matches with the time period July/August-September, when the daily traded share of newly harvested, apples grown in Germany increases implying that prices of German apples and the apple price level in general declines, inducing apple supply of southern hemisphere countries to vanish. Deviations from its long-run equilibrium are corrected fastest in this regime in both models compared to the other regimes.

The intense integration of markets with the compared to the other regimes highest speed of adjustment to the long-run equilibrium may be traced back to the highest mean difference between prices in Munich and Hamburg in this regime.

Traders confirm that large amounts of apples traded between northern and southern Germany in this time period. For example, substantial amounts of special varieties of apples (Boskop, Cox Orange), which are particularly grown in the northern part of Germany, are sold to the market in southern Germany to be stored in warehouses. Also, if the harvest is good in one and bad in the other area, e.g. due to hail or bad weather during bloom, producers the area with the bad harvest will buy apples from the other region to fill warehouses.

Regime LARGER (winter): Relates to the market conditions prevailing during September to March, when almost exclusively German and Italian apples stored in the regional warehouses are supplied to the wholesale markets. Prices in both markets do error correct, but the speed of adjustment is lowest compared to the other regimes. Yet, the speed of adjustment of the market in Munich is higher than of the market in Hamburg. This low speed of adjustment to the long-run equilibrium might be attributed to the relatively low mean price difference between the wholesale market in Hamburg and Munich limiting profitable interregional trade and implying a low degree of market integration.

Regime BETWEEN1 (spring): Is in accordance with the time period March-April/June, when the share of stored German apples sold declines and apple warehouses are cleared, whereas the share of newly harvested apples grown in southern hemisphere countries increases. Cointegration between the prices of the Hamburg and Munich market can not be confirmed unambiguously, indicating that a long-run equilibrium relationship does not exist. In this regime price differences between Hamburg and Munich are lowest, reducing the margin for profits resulting from interregional trade. This might explain why cointegration can not be identified clearly. In addition, since apples have been stored for quite some time at this point of time, once they are taken out of the warehouse, the apples perish very fast which is a further factor restricting interregional trade in this regime.

VI. CONCLUSIONS AND DISCUSSION

In this paper we propose a three-step procedure to estimate a regime-dependent VECM accounting for threshold effects not only in the short-run adjustment towards the long-run equilibrium but in the long-run equilibrium relationship as well. This type of non-linearity allows the long-run equilibrium relationship to move back and forth between regimes as a function of the size of an exogenous threshold variable with respect to a threshold value, rather than hypothesizing a one-off break in this relationship. This model seems to be particularly suitable in settings of irregular seasonal price transmission, typical for fresh fruits and vegetables, in which the use of seasonal dummy variables to account for seasonal variation in the equilibrium relationship might not be sufficiently flexible. The proposed price transmission model estimation strategy is unique in utilizing the test by GONZALO AND PITARAKIS (2006) on the hypothesis of linear versus threshold cointegration.

In our application to the German wholesale market for apples we find clear evidence of threshold cointegration with the share of German apples traded in total wholesale market trade serving as the threshold variable. We identify two thresholds in the cointegration regression and distinguish four price transmission regimes which are characterized by different equilibrium relationships as well as short-run adjustment processes towards this equilibrium. Our econometric results fit well with the actions on the German apple market.

This research will be extended to further enlighten the connection between the size of the price difference between the markets in Hamburg and Munich and the degree of market integration. Actual costs of transporting apples from one market to the other will be gathered and be compared to the size of price differences in each regime.

In addition, the factors inducing differences in price transmission elasticity implying non-linear cointegration will be further investigated. In this study the risk of successfully selling apples in another market probably plays a major role whereas variation in transport costs is of minor importance.

Further, the price responses in one market to a price shock in the other market for each regime could be analyzed by impulse-response functions.

The model could be extended and regard for threshold effects induced by the ECT term, as in the threshold VECMs, also. Yet, this faces the problem that with increasing number of regimes in a model, the number of observations entering into the estimation of the parameters of a regime decreases.

ACKNOWLEDGEMENTS

We are grateful to Heribert Tintinger of the Bundesanstalt für Landwirtschaft und Ernaehrung (BLE) in Bonn for making data on apple trade on German wholesale markets available. Linde Goetz gratefully acknowledges financial support by the Schaumann-Stiftung. Work leading to this paper took place while Stephan von Cramon was a Fulbright Scholar visiting the Department of Applied Economics and Management of Cornell University, where Christopher Barrett provided many important ideas and suggestions. The usual disclaimer applies.

REFERENCES

ANDREWS, D. (1993): Tests for parameter instability and structural change with unknown change point, Econometrica, Vol. 61: 821-856.

BAFFES, J. AND B. GARDNER (2003): The Transmission of World Commodity Prices to Domestic Markets Under Policy Reforms in Developing Countries, Policy Reform, Vol. 6(3): 159-180.

BAKUCS, L.Z. AND I. FERTÖ (2005): Marketing Margins and Price Transmission on the Hungarian Pork Market, in: BROSIG, S. AND H. HOCKMANN (eds.): How effective is the Invisible Hand? Agricultural and Food Markets in Central and Eastern Europe. Studies on the Agricultural and Food Sector in Central and Eastern Europe, IAMO, Vol. 31: 134-149.

BALCOMBE, K., BAILEY, A. AND J. BROOKS (2007): Threshold Effects in Price Transmission: The Case of Brazilian Wheat, Maize and Soya Prices, American Journal of Agricultural Economics 89(2): 308-323.

BRUEMMER, B., S. VON CRAMON-TAUBADEL, S. AND S. ZORYA (forthcoming): A Markov-switching vector error correction model of vertical price transmission between wheat and flour in Ukraine, European Review of Agricultural Economics.

BALKE, N.S. AND T.S. FOMBY (1997): Threshold Cointegration, International Economic Review 38: 627-645.

BARASSI, M.R. AND A.M.R. TAYLOR (2004): A test for Change in the Cointegrating Rank (mimeo, Birmingham: University of Birmigham, 2004).

BARASSI, M.R. AND A. GHOSHRAY (2007): Structural Change and Long-run Relationships between US and EU Wheat Export Prices, Journal of Agricultural Economics, Vol. 58(1): 76-90.

CANDELON, B. AND H. LÜTKEPOHL (2001): On the Reliability of Chow Type Tests for Parameter

Constancy in Multivariate Dynamic Models, Economic Letters, Vol. (73): 155-160.

CHAVAS, J.-P. AND A. MEHTA (2004): Price Dynamics in a Vertical Sector: The Case of Butter, American Journal f Agricultural Economics, Vol. 86(4): 1078-1093.

BUNDESANSTALT FÜR LANDWIRTSCHAFT UND ERNÄHRUNG (BLE): Observation wholesale markets 2003-2005, unpublished.

DERCON, S. (1995): On Market Integration and Liberalisation: Methods and Application to Ethiopia, Journal of Development Studies, Vol. 32(1): 112-143.

DICKEY, D.A. AND W.A. FULLER (1984): Likelihood ration statistics for autoregressive time series with a unit root, Econometrica, Vol. 49: 1057-1072.

GOETZ, L. AND VON S. CRAMON-TAUBADEL (2006): Asymmetric Price Transmission in the Israeli Citrus Export Sector in the Aftermath of Liberalization, Contributed Paper presented at the "1. Mediterranean Conference of Agro-Food Social Scientists", April 23-25, Barcelona, Spain.

GONZALO, J. AND T. LEE (1998): Pitfalls in testing for long run relationships, Journal of Econometrics, Vol. 86: 129-154.

GONZALO, J. AND J. PITARAKIS (2006): Threshold Effects in Cointegrating Relationships, Oxford Bulletin of Economics and Statistics, 68 (Supplement): 813-833.

GOODWIN, B.K. AND N. PIGGOTT (2001): Spatial market integration in the presence of threshold effects, American Journal of Agricultural Economics 83: 302–317.

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

GUILLOTREAU, P., LE GRE, L. AND M. SIMIONI (2005): Price-Cost Margins and Structural Change: Sub-Contracting within the Salmon Marekting Chain, Review of Development Economics, Vol. 9(4): 581-597.

HANSEN, B.E. AND B. SEO (2002): Testing for two-regime threshold cointegration in vector error correction models, Journal of Econometrics, Vol. 110: 293-318.

HANSEN, G. (1996): Residual-based tests for cointegration in models with regime shifts, Journal of Econometrics, Vol. 70: 99-126.

JOHANSEN, S. (1988): Statistical Analysis of Cointegration Vectors, Journal of Economic Dynamics and Control 12 (June-Sept. 1988): 215-238.

KENNEDY, P. (2003): A Guide to Econometrics, 5th Edition, Blackwell Publishers.

KWIATOWSKI, D., PHILLIPS, P., SCHMIDT, P. AND Y. SHIN (1992): Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root, Journal of Econometrics, Vol. 54: 159-178.

JOHANSEN, S., MOSCONI, R. AND B. NIELSEN (2000): Cointegration analysis in the presence of structural breaks in the deterministic trend, Econometrics Journal, Vol. 3: 216-249.

LO, M.C. AND E. ZIVOT (2001): Threshold cointegration and non-linear adjustment to the law of one price, Macroeconomic Dynamics, Vol. 5: 533–576.

MEYER, J. (2004): Measuring market integration in the presence of transaction costs - a threshold vector error correction approach, Agricultural Economics, Vol. 31: 327-334.

RODRÍGUEZ, G.M. AND J.J.C. HERNÀNDEZ (2005): Modelling weekly Canary tomato exports, Agricultural Economics, Vol. 33: 255-267.

SERRA, T., GIL, J.M. AND B.K. GOODWIN (2006): Local polynomial fitting and spatial price relationships: price transmission in EU pork markets, European Review of Agricultural Economics 33(3): 415–436.

SANJUAN, A. AND P.J. DAWSON (2003): Price transmission, BSE and structural breaks in the UK meat sector, European Review of Agricultural Economics, Vol. 30(2): 155-172.

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

WANG, D. AND W. TOMEK (2007): Commodity Prices and Unit Root Tests, American Journal of Agricultural Economics, Vol. 89(4): 873-889.

Author:	Linde Goetz
Institute:	Institute for Agricultural Development in Middle and Eastern Europe (IAMO)
Street:	Theodor-Lieser-Strasse 2
City: ^	06120 Halle/Saale
Country:	Germany
Email:	goetz@iamo.de