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THRESHOLD COINTEGRATION AND ASYMMETRIC PRICE TRANSMISSION IN FINNISH BEEF AND PORK MARKETS

Arto Luoma
Jani Luoto
Marko Taipale

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ABSTRACT: This study examined the transmission of producer price changes to consumer prices in Finnish beef and pork markets. Both meat varieties were studied based on monthly observations from 1981 through May 2003. According to the earlier studies price transmission can be asymmetric e.g., producer price rises move faster and/or more completely to consumer prices than corresponding price reductions. This may happen, for example, because of adjustment costs or market power. Adjustment was studied with co-integration threshold models and price change threshold models. In addition nonlinear impulse response functions were calculated based on 0,1 € change in producer price. It was not possible to detect possible asymmetric price transmission, since the time period under study was characterized by strong structural changes in the formation of producer prices, However, the results suggest that it is the consumer price that responds to the long-term disequilibrium of the consumer and producer prices. The consumer price also reacts to changes in the producer price in the short term.

Key words: Price transmission, beef, pork, non-linearity

ARTO LUOMA – JANI LUOTO – MARKO TAIPALE. 2004. KYNNYSYHTEIS-INTEGROITUVUUS JA EPÄSYMMETRINEN HINTOJEN VÄLITTYMINEN SUOMEN NAUDAN- JA SIANLIHAN MARKKINOILLA. Pellervon taloudellisen tutkimuslaitoksen työpapereita n:o. 70. 30 s. ISBN 952-5299-79-1 (PAP), ISBN 952-5299-85-6 (PDF), ISSN 1455-4623.

TIIVISTELMÄ: Tutkimuksessa tarkastellaan tuottajahintojen muutosten välittymistä kuluttajahintoihin Suomen naudan- ja sianlihan markkinoilla. Molempien lihamarkkinoiden aineistona oli kuukausittaisia havaintoja vuodesta 1981 vuoden 2003 toukokuuhun. Aiempien tutkimusten mukaan hintojen välittyminen voi olla epäsymmetristä, jolloin esimerkiksi tuottajahinnan nousut siirtyvät eri tavalla kuluttajahintoihin kuin tuottajahintojen laskut. Syynä tähän voivat olla esimerkiksi sopeutumiskustannukset tai markkinavoima. Sopeutumista tutkittiin yhteisintegroituvuuteen perustuvilla kynnysmalleilla ja hintamuutosten suuruuteen perustuvilla kynnysmalleilla. Lisäksi laskettiin epälineaarisia impulssivastefunktioita, jotka perustuivat 10 sentin tuottajahinnan nousuun tai laskuun. Mahdollista hintojen epäsymmetristä välittymistä ei voitu luotettavasti tutkia, koska tutkittavana oleva ajanjakso käsitti voimakkaan rakennemuutoksen tuottajahinnan muodostumisessa. Tulosten perusteella kuitenkin havaitaan kuluttajahinnan reagoivan kuluttaja- ja tuottajahintojen pitkän aikavälin epätasapainoon. Kuluttajahinta reagoi myös lyhyellä aikavälillä tuottajahinnan muutoksiin.

Avainsanat: Hintatransmissio, naudanliha, sianliha, epälineaarisuus

CONTENTS

TIIVISTELMÄ (NON-TECHNICAL SUMMARY IN FINNISH)

Viimeisen kymmenen vuoden aikana on maataloustuotteiden tuottajahinnoissa tapahtunut merkittäviä muutoksia EU:ssa. Erityisesti naudan- ja sianlihan hinnat ovat heilahdelleet rajusti niin eläintautiongelmien kuin maatalouspolitiikan muutostenkin takia. Suomessa tuottajahintoja pudotti lisäksi 1990-luvun puolessa välisissä maan EU-jäsenyys. Tässä tutkimuksessa tarkastellaan naudan- ja sianlihan tuottajahinnoissa tapahtuneiden muutosten välittymistä näiden tuotteiden kuluttajahintoihin Suomessa. Raportissa on analysoitu hintamuutosten siirtymisten suuruutta, ajoitusta ja symmetrisyyttä.

Yleisen käsityksen mukaan raaka-aineiden hintojen korotukset siirtyvät ainakin nopeammin ja kenties myös täydellisemmin lopputuotteiden hintoihin kuin vastaavat laskut. Hintojen välittyminen on tällöin epäsymmetristä. Tämä voi johtua siitä, että markkinoilla on vähän toimijoita, jolloin yritysten on mahdollista käyttää markkina-voimaansa tuotteiden hinnoittelussa. Epäsymmetrisyys voi tarkoittaa myös sitä, että hintamuutosten välittyminen vaihtelee sen mukaan, miten suuria alkuperäiset hinta-muutokset ovat.

Tässä tutkimuksessa käytetään lineaarisia sekä kahden ja kolmen sopeutumisalueen epälineaarisia virheenkorjausmalleja. Erilaiset sopeutumisalueet määritetään tuottaja- ja kuluttajahintojen välisen laskennallisen pitkän aikavälin tasapainosta poikkeamisen suuruuden perusteella tai pelkästään tuottajahinnan muutoksien suuruuden perusteella. Lisäksi tuottajahintojen nousujen ja laskujen vaikutuksia kuluttajahintoihin arvioitiin epälineaarisella impulssivasteanalyysillä.

Tutkimusaineistona on kuukausittaiset naudan- ja sianlihan tuottajahinnat vuodesta 1981 vuoden 2003 toukokuuhun. Kuluttajahinnaksi samalle ajanjaksolle on muodostettu kori, joka on saatu Tilastokeskuksen keräämistä kuluttajahintatilaston hinnoista eri ruhon osille painottaen näiden keskimääräistä osuutta kokonaisessa eläimessä.

Tutkimusajanjaksolla tapahtui lihamarkkinoilla suuria rakenteellisia muutoksia, jotka vaikeuttivat myös hintavälittymisen tutkimista. 1980-luvulla suunnilleen vuosikym-menen loppuun saakka elettiin hintasääntelyn jälkeistä aikaa. Tuottajahinnat nousivat maltillisesti ja kuluttajahinnat selvästi tuottajahintoja nopeammin. Koska hintojen liikkeiden välillä ei näyttäisi tällä ajanjaksolla olevan tasapainotilannetta, vaan hinnat jatkuvasti eriytyivät toisistaan, hintojen sopeutumisen symmetrisyydestä ei voida saada selkeitä tuloksia. Kuluttajahintojen jatkuvasti nopeampi nousu saattoi aiheutua monista tekijöistä: kaupan ja teollisuuden kustannukset ja marginaalit saattoivat nousta. Kulu-tuksessakin saattoivat painottua vähitellen kalliimmat ruhon osat. Eriytymisen taustalla olleisiin tekijöihin ei kuitenkaan voitu tässä tutkimuksessa mennä.

1990- ja 2000-luvuilla samanlaista jatkuvaa eriytymistä ei enää ollut varsinkaan sianlihassa. EU-jäsenyys sai aikaan muutoksia hintarakenteissa siihen suuntaan, että raaka-aineen ts. tuottajahinnan osuus kuluttajahinnasta pieneni, mutta nämä eivät näy aineistossa jatkuvana tuottaja- ja kuluttajahintojen eriytymisenä. Naudanlihan osalta rakenteelliseksi muutokseksi voidaan kylläkin tulkita periodin loppu, jolloin kuluttajahinnat nousivat naudanlihassa selvästi ilman merkittäviä muutoksia tuottajahinnassa ja syntyi saman tyyppinen eriytymisen tilanne kuin 1980-luvulla. Tähän ajankohtaan ajoittui myös euron käyttöönotto ja kuluttajahintaindeksin muutos, minkä vuoksi on vaikea erottaa eri tekijöiden vaikutusta toisistaan ja tulkita oliko kysymyksessä kertahyppäys.

Toisella periodilla hintasopeutumisen nopeus on hyvin samanlaista sekä tuottajahinnan laskulle että nousulle. Tilastollisesti merkitsevää epäsymmetrisyyttä ei ollut havaittavissa, etenkään kun määriteltiin sopeutumisalueet laskennallisesta tasapainosta poikkeamisen avulla. Sopeutumiskertoimet ovat kuitenkin yleensä melko pieniä tai eivät merkittävästi poikkea nollasta. Virheenkorjausmallien sopeutumisparametrien perusteella ei voitu luotettavasti tutkia hintojen välittymisen epäsymmetrisyyttä, koska tarkasteltavalla aikavälillä sian- ja naudanlihan hintarakenteessa tapahtui merkittäviä muutoksia. Koska tuottaja- ja kuluttajahintojen suhde pieneni sianlihan osalta vuoteen 1995 asti ja naudanlihan osalta lähes koko tutkittavan ajanjakson, ei voida varmuudella puhua pitkän aikavälin tasapainotilasta tai poikkeamasta siitä.

Impulssivasteanalyysi kertoo konkreettisemmin, kuinka tuottajahinnan nousut ja laskut välittyvät kuluttajahintoihin. Epälineaarisissa malleissa on kuitenkin se haittapuoli, että pitkään ennustejaksoon liittyy suuri epävarmuus ja tulokset riippuvat valitusta ajankohdasta. Nyt käytettiin viimeistä toukokuun 2003 havaintoa ja siten sen hetkistä markkinatilannetta lähtökohtana. Sianlihalla kuluttajahinta muuttuu symmetrisesti suhteessa tuottajahinnan laskuihin ja nousuihin. Kun mallina oli pitkän aikavälin tasapainosta poikkeamisen suuruus, niin 10 sentin nousu tai lasku kiloa kohden tuottajahinnassa siirtyi noin 1,75-kertaisesti kuluttajahintaan 1,5 vuoden aikana.

Naudanlihamarkkinoiden monien muutosten vuoksi malli ei toiminut yhtä hyvin kuin sianlihalla ja vastaavat tulokset olivat erikoisia siten, että tuottajahinnan noususta näytti seuraavan kuluttajahinnan lasku pitkällä aikavälillä. Tämä tulos on seurausta hintojen kehityksestä tarkasteltavan aikajakson loppupuolella. Tämän vuoksi kiinnitettiin mallin sopeutumiskertoimet vakioiksi, jotka eivät riipu poikkeamasta pitkän aikavälin tasapainosta. Näin saatiin impulssivastefunktio, jossa 10 sentin muutos kuluttajahinnassa siirtyy 1,5 vuoden aikana 0,6-kertaisena.

1. INTRODUCTION

Prices are the cornerstones of modern market economies. With different products, marketing channels can geographically be extremely long or the marketing channel may consist of many vertically integrated levels. Moreover, it is usually the prices that connect various stages. A common belief is that price transmission between different stages in the marketing chain is not symmetric. Usually, it is claimed that input price increases are transferred more rapidly to consumer prices than corresponding price reductions. This view is strongly supported by the extensive study of Pelzman (2000). He examined not only many consumer product markets but also numerous producer goods markets, and in more than two thirds of the cases found asymmetry in the price transmission. This finding is highlighted by noting that this is not a prediction of economic theory, and hence the prevailing theory is claimed to be wrong.

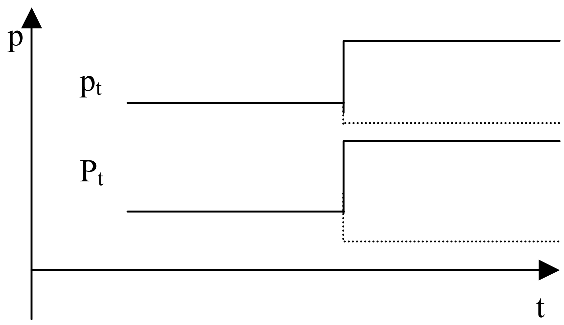
Agricultural markets have been one of the central targets for the analysis of price transmission. The purpose of this study is to analyse Finnish pork and beef markets in order to determine how producer price changes are passed on to consumer prices and how and which prices do adjust after a possible disequilibrium in meat markets. Using conventional linear models in a nonlinear situation leads to the wrong conclusions being drawn. However, the cost of using non-linear models comes with their added complexity. Nevertheless, if nonlinearity is rejected, one is free to carry out the estimation with linear models.

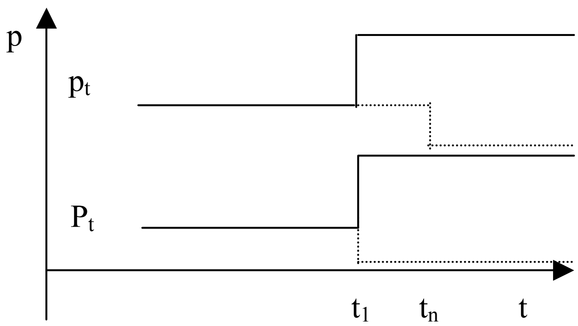
This paper is organized as follows. Chapter two takes asymmetry as a working hypothesis and explains it in terms of both economic theory and applicable econometric models. Reviewing the development of econometric methods to up-to-date models is important. The next subsection deals with relevant previous studies and explains the exact focus of this paper. Chapter three introduces the data used and the preliminary analyses of the time series properties. The models to be estimated are also naturally introduced. Next, chapter four first discusses the model selection procedure and then reports the estimation results. The final chapter provides a short summary of the paper and concludes with a discussion of the main results.

2. ASYMMETRIC PRICE TRANSMISSION AND ECONOMETRIC TESTING

2.1 Asymmetric price transmission and its causes

Von Cramon-Taubadel (1998) provides a concise discussion of the definition of asymmetric price transmission and its possible causes. Asymmetry of price adjustment can exist either with respect to magnitude or speed. A combination of these two is also possible. In the case of magnitude (Figure 1a), long-term elasticities of price transmission differ depending on the direction of the initial price change. This happens because input price P_t rise is moved more completely to output price p_t than the corresponding input price reduction. Accordingly, in the case of speed (Figure 1b), short-term elasticities are different. At the time of the input price rise, t_1, the output price responds immediately whereas the reaction to an input price drop takes n periods of time. Price transmission can of course also be asymmetric in the other direction, i.e. input price reductions are transferred more completely or faster to output prices.





Source: Meyer & von Cramon-Taubadel (2002)

Figure 1a. *Asymmetric price transmission (magnitude)*

Figure 1b. *Asymmetric price transmission (speed)*

One major cause of asymmetric price transmission are adjustment costs such as the cost of making new labels and informing market partners about price changes. Adjustment takes place only after the input price change has been sizable enough. In addition, it is notable that different firms may have different adjustment costs. For example, meat packers who face high fixed costs and excess capacity may reduce their margins because of competition and therefore producer prices may also rise faster in the case of increased demand than they fall in the case of weakened demand. General inflation also may affect the type of asymmetry. When input prices rise, firms usually correct not only for this but also for a general and possibly anticipated rise in operating costs. If the input price lowers, inflation moderates the possibilities for lowering output prices. Perishable products are one special case. Retailers may be reluctant to raise their prices immediately, because they fear reduced sales and accumulated spoiled storage.

Another important cause of asymmetric price transmission is imperfect competition among middlemen between farms and consumers and the resulting market power. Oligopolistic actors may then use their market power and react more quickly to reduced margins than to stretched ones. Market power is also the most likely explanation for asymmetric price transmission in the long run. However, the opposite is also possible if firms care about their market share and raise prices less aggressively than they lower them. Hence, it is not clear a priori what effect market power has on price transmission. In addition, especially for agricultural markets, a point worth mentioning is government intervention. Government policies often support producer prices. Producer price decreases are then easily interpreted to be only temporary, but price increases have a more permanent nature. In addition, Meyer and von Cramon-Taubadel (2002) list other factors, for example local monopolies, as a cause of asymmetric adjustment.

2.2 Econometric testing of price transmission

A large amount of empirical literature has examined price linkages between different markets. Von Cramon-Taubadel (1998) discussed the development of testing price transmission in general and specifically asymmetric tests. Early studies before co-integration methods used variants of the following model:

(1) $$\Delta p_t = c + \beta^{+}\Delta P_t + \beta^{-}\Delta P_t + \varepsilon_t.$$

In this model, Δ is a difference operator so that $\Delta p_t = p_t - p_{t-1}$. Again, p_t is the output price and P_t is the input price. The response of the former to the latter is decomposed into positive and negative changes. Term c is a constant, while β^{+} and β^{-} are adjustment coefficients of the positive and negative changes in P. If the test result is that $\beta^{+} = \beta^{-}$, then price adjustment is symmetric. It is possible to distinguish between short-term and long-term adjustment by adding lags to equation (1). Long-term symmetry is detected by testing whether the sums of the coefficients in these polynomials are equal and a testable condition for short-term symmetry is that these polynomials are identical.

Next, von Cramon-Taubadel (1998) made a fundamental clarification. The estimation of equations like (1) does not fully consider the time series properties of the data used, which are typically non-stationary leading to problems in the testing. Making the data stationary by differencing is part of the solution to the problem, but equation (1) is still incompatible with co-integration and long-term information between time series. A proper way to proceed is to use the error correction models introduced by Engle and

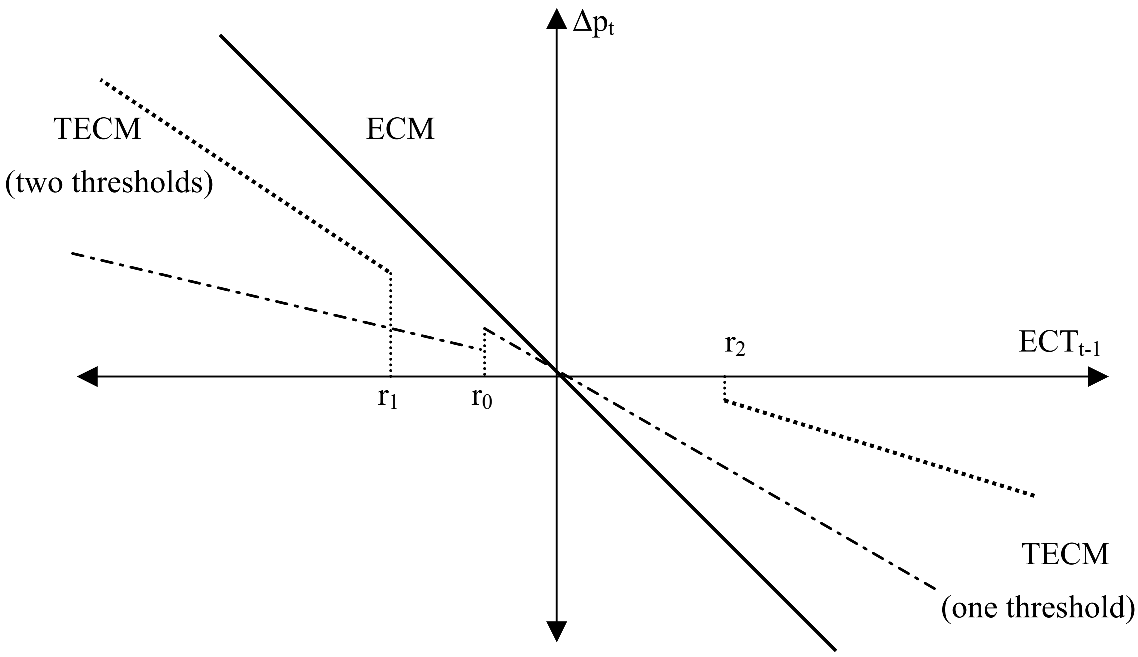
Granger (1987). Their model utilizes both short-term dynamics and long-term information. A typical formulation for an error correction model is:

(2) $$\Delta p_t = c + \beta_1 \Delta P_t + \alpha ECM_{t-1} + \beta_2(L)\, \Delta\, p_{t-1} + \beta_3(L) \Delta\, P_{t-1} + \varepsilon_t,$$

where αECM is the error correction term and $\beta_2(L)$ and $\beta_3(L)$ are lag polynomials. Long-term information is given by the cointegrating relation $p_t = c_0 + \beta P_t + u_t$, which in its basic form is just a static regression model in levels. The lagged value of the error term in the cointegrating regression is given by $ECM_{t-1} = u_{t-1} = p_{t-1} - c_0 - \beta P_{t-1}$. In order for the system to be stable in single equation models the error correction coefficient α has to be negative, as explained next. When α is negative, a positive residual in the cointegrating regression makes the error correction term negative in equation (2). Just the opposite occurs when a negative residual is multiplied by another negative number and the total effect is positive, which also means a converging situation.

The standard error correction model has two drawbacks. The first disadvantage is that it assumes symmetric adjustment to long-run equilibrium irrespective of the sign of the shock. This feature has been modelled more realistically by Granger and Lee (1989), who introduced a non-symmetric error-correction model. The idea is simply to segment the ECM term to positive and negative ones. This model nests earlier model (2) and the equality of the two different ECM coefficients can be tested by an F-test. The non-symmetric error correction model was further developed by Enders and Granger (1998). They argued that standard unit root tests are misspecified if adjustment is asymmetric. They therefore calculated new critical values for unit root tests.

The second disadvantage with the conventional error correction model is that after a shock it assumes instant adjustment towards the equilibrium and hence does not consider, for example, possible transaction costs. A new family of models that tackles the instant adjustment problem was started from the article by Balke and Fomby (1997) with the illuminating title “Threshold cointegration”. Cointegration as a whole is still maintained but between estimable thresholds r_1 and r_2 there may be a range of unit root adjustment. Deviation from equilibrium will result in a price change only if the deviation is larger than the threshold value. Figure 2 clarifies the intuition behind nonsymmetric error correction and threshold cointegration models.



Source: Meyer (2003)

Figure 2. *Impact of the error correction term on the price adjustment*

Δp_t is the price adjustment of the output price and it is described as a function of deviations from the long-term equilibrium, ECT. Figure 2 nests three different error correction models. The simplest model is the usual one (thick continuous line) with instant and symmetric adjustment. The next step is to add one threshold r_0 and the model becomes a non-symmetric one. The threshold r_0 usually lies close to zero. Whether the deviation from the long-term equilibrium is larger or smaller than r_0 the adjustment takes place at a different speed (different angle of the adjustment line). Therefore, one can speak of a two-regime model. The third, threshold cointegration model, has two thresholds, namely r_1 and r_2, and hence three different regimes. The difference compared to the two-regime model lies in the region between r_1 and r_2. Now the figure has been drawn in a way that there is no adjustment if deviations from the long-term equilibrium are small enough. In practise there would be small, random price movements and hence the adjustment is called random walk in the middle regime. Although justified by economic theory, the adjustment in the middle regime is not random walk by definition, but it has to be different from outer regimes. By the same token the absolute value of r_1 and r_2 need not to be equal, even though in empirical applications it is often assumed to be so. Formally, a three-regime error correction model can be described as in Goodwin and Harper (2000):

(3) $$p_t = \begin{cases} c^1 + \sum_{j=1}^{K}(\beta_j^1 \Delta P_{t-j+1}) + \alpha^1 ECM_{t-1} + \varepsilon_t^1, & \text{if } ECM_{t-1} \le r_1, \\ c^2 + \sum_{j=1}^{K}(\beta_j^2 \Delta P_{t-j+1}) + \alpha^2 ECM_{t-1} + \varepsilon_t^2, & \text{if } \mathrm{r}_1 < ECM_{t-1} \le r_2, \\ c^3 + \sum_{j=1}^{K}(\beta_j^3 \Delta P_{t-j+1}) + \alpha^3 ECM_{t-1} + \varepsilon_t^3, & \text{if } ECM_{t-1} > r_2. \end{cases}$$

Recall that α^1, α^2 and α^3 are the long-term adjustment parameters, which show the speed at which the system moves back to equilibrium if there is a deviation from the long-term equilibrium. In addition there are short-term effects described by the difference terms. Altogether, the price adjustment process is unique at any point in time because it depends on both the ECM term and lagged price differences.

2.3 Previous studies on asymmetric price transmission

Von Cramon-Taubadel (1998) examined the German pork market with an earlier non-symmetric error correction model and found evidence of asymmetric price transmission in the form that wholesale prices react more strongly to squeezed than stretched margins. Abdulai (2002) presented an application of the Enders and Granger method to the Swiss pork market. His result was that producer price increases move faster to consumer prices than reductions in producer prices. In addition, as producer prices are rigid, it is the retail prices that respond and only to negative disequilibria. These results parallel Abdulai's impulse response analysis, which describes dynamic interrelationships among different variables. Impulse response functions show that the margin between producer and consumer prices returns more rapidly to the normal level after a producer price hike than a producer price reduction. In non-linear cases impulse response analysis is more complicated than in the linear world. Technical details behind non-linear impulses are explained shortly in the appendix of our study. In this study impulse response analysis is also carried out, but the approach of the calculations is similar to Lloyd et al. (1999), where analysis focuses on the response of the consumer price to a 0.1 € monetary shock in the producer price rather than the response of the margin between different prices. Besides ignoring the possible existence of two thresholds, the studies of von Cramon-Taubadel and Abdulai employ single equation models. In this case it might be a problem if both price series are endogenous. Endogeneity causes estimation to be biased and inconsistent. However, von Cramon-Taubadel considered this possibility and tested for exogeneity conditions. In addition, he also reported impulse response functions, which are in accordance with error correction models.

From a methodological point of view, an interesting approach is that of Obstfeld and Taylor (1997), who used the idea of two thresholds, which accordingly defines three different regimes. They applied maximum likelihood estimation to study purchasing power parity (PPP) and the law of one price with separate commodity categories. They used a symmetric threshold and the speed of adjustment in outer regimes was the same. The non-adjustment region was based on transaction costs and uncertainty and together they enable less than perfect arbitrage.

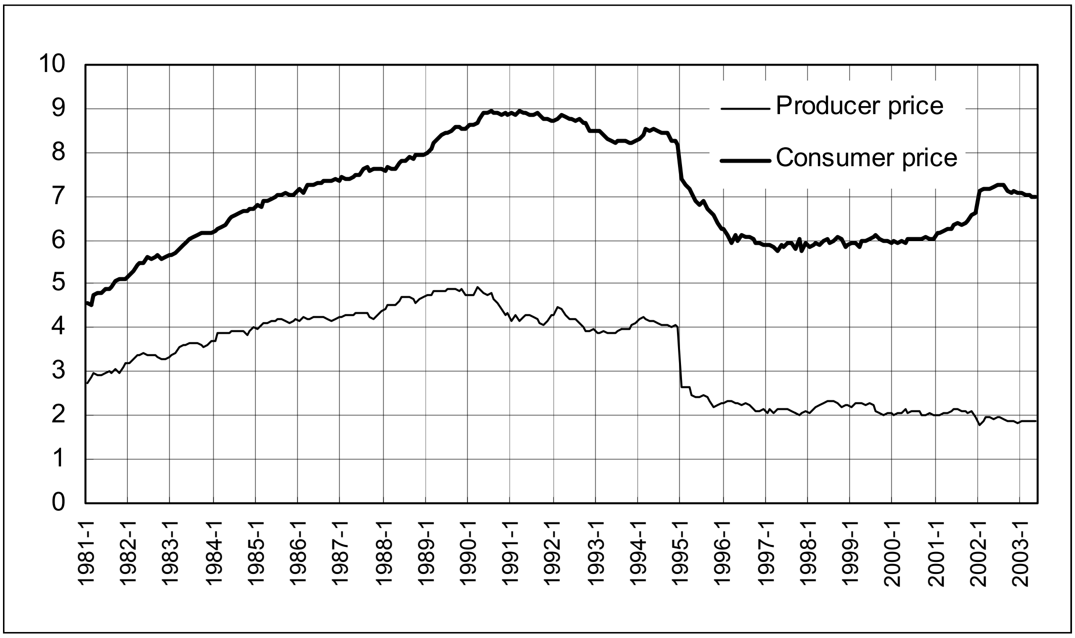
Goodwin & Holt (1999) and Goodwin & Harper (2000) used grid search to determine the thresholds and find important asymmetries in U.S. beef and pork farm, wholesale and retail prices. Asymmetry appeared so that large negative deviations from the equilibrium were accompanied by especially significant error correction terms. In addition, price interrelationships existed between wholesale and retail prices rather than between farm and wholesale prices. They also reported impulse response functions. Based on Hansen and Seo (2002), Meyer (2003) also estimated a threshold vector error correction model in order to analyze price transmission between pork markets in Germany and the Netherlands. Because the analysis was between markets at different locations it is called spatial transmission, which is naturally different from vertical price transmission. Meyer's report has also two threshold values, r, which are symmetric around zero. Vector models take care of the possible endogeneity problem and make the analysis of adjustments more complete by allowing feedback from both variables. The study of Forbes et al. (1999) of index futures markets using a Bayesian estimation method is parallel to our paper.

In all previous published studies the cointegrating vector has either been assumed to be given, when appropriate, or has been estimated separately. In our paper, the cointegration model is estimated simultaneously with other parameters of the model. To our knowledge, the other contribution of this paper is to also define different regimes by the magnitude of producer price changes. Previous studies have used discrepancies from the cointegration relation as a decisive factor in defining different regimes, but recalling the definition one can see that an equilibrium error may result from either variable. If one is especially interested in the effect of producer price changes, it is illustrating to also use that change as decisive factor in determining regimes.

3. DATA AND ESTIMATION METHOD

3.1 The data and their time series properties

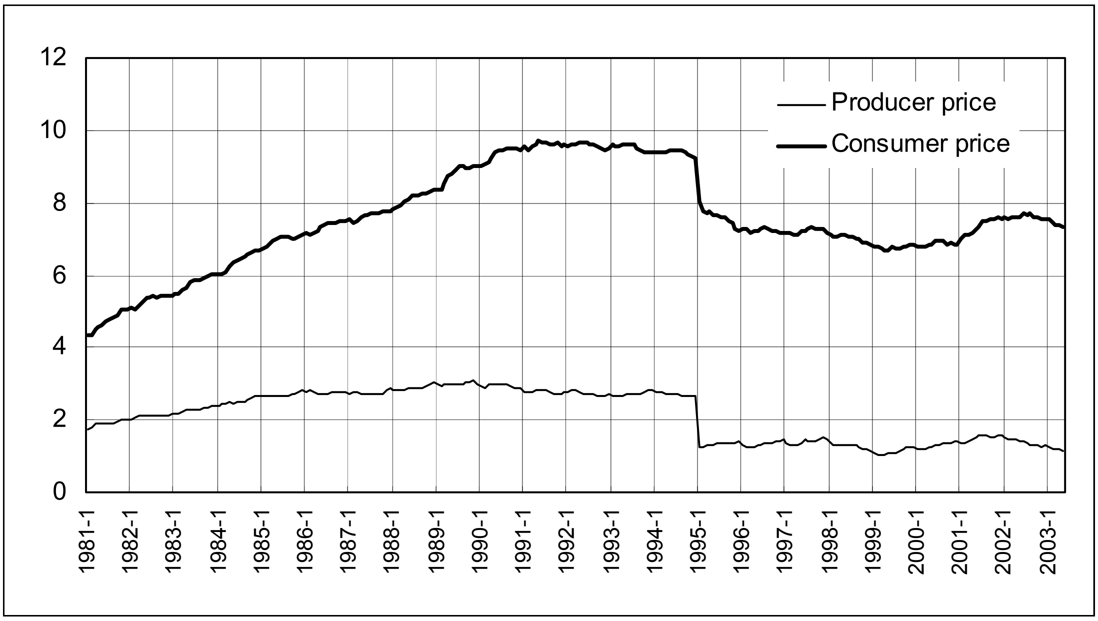
This study uses monthly data for retail and producer prices from January 1981 through May 2003, which makes a total of 269 observations. Producer prices for beef and pork were obtained from the Monthly Review of Agricultural Statistics provided by the Board of Agriculture Statistical Office.[1] The producer price data used are the average monthly prices for all beef and pork meat. The figures are totals for beef, including cows and heifers, and totals for pork, including sows. As the consumer price for beef and pork we use various meat cuts that Statistics Finland reports in the monthly publication Consumer Price Statistics. Details on the calculation of the average retail carcass are provided in Appendix I. Prices are measures in euros per kilogram. The time series of the data used are provided in the next two figures.



Source: Statistics Finland, Ministry of Agriculture and Forestry

Figure 3. *Beef prices in Finland €/kg*

[1] Nowadays referred to as the Information Centre of the Ministry of Agriculture and Forestry



Source: Statistics Finland, Ministry of Agriculture and Forestry

Figure 4. *Pork prices in Finland €/kg*

An overall look at the figures already provides some insights. On both markets, producer and consumer prices seem to move in parallel. However, a closer look reveals some interesting differences. During the economic boom period of the Finnish economy at the end of the 1980s, consumer prices seemed to rise faster than producer prices. After that, the Finnish meat industry started to prepare for possible EU membership and producer prices began to decline. Beef consumer prices did not at first adjust at all, and in the case of pork the adjustment of consumer prices was very moderate. After EU membership all prices dropped, but the adjustment was faster for producer prices. The BSE crisis seems to have had a small effect on Finnish beef markets and perhaps caused a small rise in pork consumer prices.

In order to ensure appropriate statistical inference, standard augmented Dickey-Fuller (ADF) unit root-tests were performed for the time series in order to detect possible non-stationarity. The results are presented in Appendix II. The lag length for DF tests was defined so that residual autocorrelation disappeared. No lags were needed for beef producer prices, while for beef consumer prices three lags was adequate. Both price series were found to be integrated of the order one, i.e. I(1). Pork producer price test results are identical to those for beef, but pork retail prices are somewhat different. No lags are needed in the DF test, but then the series seems to be stationary. Adding one lag or dropping the three first observations makes the series I(1). Our conclusion is that pork retail prices should be handled as I(1). This borderline property is also seen in the next step, which is to test co-integration. These results are based on the method of

Johansen (1991) and also appear in Appendix II. The beef series is co-integrated but the pork series seems to be stationary. However, the test statistic, 4.1, only slightly exceeds the 5% critical value of 3.8. In addition, the critical value is based on a greater amount of data. Theoretically, it would be strange if producer and consumer prices drift apart, so we assume them to be cointegrated. In addition, our estimation method allows the error correction coefficient to be non-significant in some regimes if needed and does not force cointegration to work.

3.2 A Bayesian multiple-regime vector error-correction model

In our analysis we use the Bayesian estimation technique. Especially with non-linear models, Bayesian methods have some advantages over classical ones. For example, Bayesian methods take into account possible multiple peaks in parameter likelihoods (Koop & Potter, 1999). Specifically, we adopt a vector error correction approach. In the following, P_t is again the producer price and p_t is the retail price. Coefficients c^j are constants in different regimes. Matrix A^j_i includes the short-term coefficients for ΔP_t and Δp_t. The error correction terms are $\alpha^j \beta' y_{t-1}$. Vectors α^j include the long-term adjustment parameters related to error correction terms. Let us define an R+1 regime multivariate threshold autoregressive model with p lags:

(4) $$\Delta y_t = \sum_{j=1}^{R+1} I(r_{j-1} < z_{k,t-1} \le r_j)\left(c^j + \sum_{i=1}^{p} A_i^j \Delta y_{t-i} + \alpha^j \beta' y_{t-1} + \varepsilon_t^j \right)$$

where I(.) is an indicator function, $r_0 = -\infty$, $r_{R+1} = \infty$, $c^j = (c^j_1\ c^j_2)'$, $\beta = (\beta_1\ 1)$, $\alpha^j = (\alpha^j_1\ \alpha^j_2)'$, $y_t = (P_t\ p_t)'$ and ε^j_t is an error term with zero mean and Σ^j covariance matrix. In this paper, in contrast to previous studies, we use two different specifications to trigger the regime switches. Specifically, for a given but unknown threshold r the change between regimes is determined either by $z_{1,t} = \beta' y_t$ or $z_{2,t} = \Delta P_t$.

This concise definition nests the error-correction model used by Goodwin and Harper (2000) described in the previous section. In this formulation by Goodwin and Harper and many others, the data is organized into different regimes according to the size of the long-term equilibrium error. Therefore, we will call it a cointegration threshold model. In a three-regime model the first regime is the situation where $z_{1,t} = \beta_1 P_t + p_t$ has a value smaller than r_1. Based on the long-term cointegration relationship the retail price is then too low with respect to the producer price (or equivalently the producer price is too high with respect to the retail price). In other words, the margin is narrower than usual. A point of interest is naturally how disequilibrium vanishes. This is described mainly by α-coefficients in equation (4). The second or middle regime prevails when both producer and retail prices are in accordance with the long-term equilibrium and the discrepancy is located between r_1 and r_2. Adjustment costs and heterogeneous economic actors may cause these thresholds and it is very likely that within the thresholds there is no response of retail price to small movements in producer price. The margin can then be described as a random walk process. The third regime is opposite to the first one. Then, $z_{1,t}$ is large and the margin is stretched (consumer price too high or producer price too low). It is also possible that there is only one threshold, which means two regimes but still asymmetric price adjustment. The absence of thresholds reduces the model to a symmetric error-correction model.

In our paper the number of regimes is determined using the Bayesian Information Criterion (BIC), formally:

$$BIC = -2 \log likelihood + b \log(N),$$

where N is the number of observations and b the number of parameters in the model. A motivation to use various information criteria is to find a model with a large likelihood but at the same time to penalize for choosing a complex model with many parameters.

An alternative mechanism that divides the system into different regimes is the magnitude of producer price changes, ΔP_{t-1}. Later we will term this a price change threshold model. We are interested in the question of whether large price increases and decreases cause different reactions. However, we still want to utilize the long-term information implied by co-integration between producer and retail prices.

Model estimation and parameter distribution simulations are performed with R software. Bayesian analysis usually exploits subjective prior information. This can be subject to criticism and therefore we use noninformative priors. Details of the estimation routine, simulation methods and prior information used are provided in Appendix III.

4. ESTIMATION RESULTS

4.1 Model selection

Before we can proceed to the estimation results we must choose the most appropriate model regarding the number of regimes and lags. The BIC values were computed for a different number of regimes and lags in order to choose an appropriate model. These results are presented in Tables 1 and 2 for beef and pork. The more negative the BIC, the better the model is with respect to the information it provides. Recall from equation (4) that in both tables case z_1 is the regime based solely on cointegration and case z_2 is the regime based on price changes.

All the models we have estimated are special cases of model (4). However, we have inserted two indicator variables to eliminate the effect of the outlying observation (the 168th observation of the differenced series[2]). The first indicator obtains the value 1 at t=168 and is 0 otherwise. The second indicator obtains 1 at t=169. In the beef series and all price change models we have used only the first indicator. In this way we obtained the smallest BIC values.

Table 1. *The values of BIC in the pork price models*

No. of lags (p)	Linear	2-regime	3-regime	2-regime	3-regime
0	-3326	-3339	-3318	-3305	-3219
1	-3344	-3319	-3278	-3307	-3273
2	-3309	-3281	-3223	-3262	-3218
3	-3290	-3239	-3172	-3230	-3169
		case z_1	case z_1	case z_2	case z_2

Table 2. *The values of BIC in the beef price models*

No. of lags (p)	Linear	2-regime	3-regime	2-regime	3-regime
0	-2853	-2856	-2836	-2843	-2811
1	-2834	-2835	-2799	-2819	-2775
2	-2815	-2806	-2759	-2797	-2733
3	-2809	-2773	-2722	-2771	-2702
		case z_1	case z_1	case z_2	case z_2

On the basis of the BIC tables above the best model for pork is the linear model with one lag and for beef the 2-regime model with no lags. All price change models had worse BIC values than the corresponding cointegration threshold models.

[2] The level of producer and consumer prices changed because of Finland's membership in the European Union.

We also performed some likelihood ratio tests to test equality of the long-term adjustment coefficients α^j and the short-term dynamic matrices A^j across regimes. The test results are shown in Appendix IV. Since the adjustment coefficients are not equal in the pork cointegration threshold model we decided to use a 2-regime model with one lag in our final analysis. If all parameters but the adjustment coefficients are fixed to be the same in both regimes the BIC obtains the value -3346 which is better than that of the corresponding linear model.

On the basis of the test results the adjustment coefficients are different for beef in the 3-regime cointegration threshold models. However, if we estimate the model we obtain parameters such that the consumer price increases and the producer price decreases in the case of stretched margin. This counter-intuitive result reflects the price development during 2001 and 2002 when the consumer price was rising and the producer price was constant or slightly decreasing. Therefore, we decided to use a 2-regime model with no lags assuming that all parameters but the covariance matrices are equal in both regimes. The BIC for this model is -2866.

When price change regime models are used there are two significant (p=0.02) test results indicating that there might be some differences in the short term dynamics between the regimes. However, we did not include these models in the final analysis, since the BIC values for these models were worse than for the corresponding linear models.

4.2 Pork co-integration threshold model

In Table 3 we present the parameter estimates for the first model, in which the thresholds are based on the value of the linear combination $z_{1,t-1} = \beta_1 P_{t-1} + p_{t-1}$. In Appendix V we present histograms for parameter distributions.

Table 3. *Pork co-integration threshold model parameter estimates*

Regime	1		2	
Variable	Median	Credible interval	Median	Credible interval
c_1	0.028	(-0.001, 0.104)	-0.007	(-0.097, 0.071)
c_2	0.054	(-0.108, 0.126)	0.292	(0.025, 0.471)
a_{11}	0.164	(-0.072, 0.515)	0.278	(0.132, 0.437)
a_{12}	-0.024	(-0.193, 0.105)	0.044	(-0.045, 0.129)
a_{21}	0.453	(0.082, 1.034)	0.363	(0.101, 0.596)
a_{22}	0.155	(-0.096, 0.360)	0.083	(-0.070, 0.227)
α_1	-0.005	(-0.027, 0.145)	0.001	(-0.014, 0.019)
α_2	-0.013	(-0.257, 0.005)	-0.059	(-0.093, -0.008)
β_1	-1.648	(-2.842, -1.441)		
r	3.520	(-0.529, 4.484)		

Regime 1 consists of observations from 1981 to 1988 (see Figure 5 a). That time there was a structural change in pork prices and the consumer prices were increasing rapidly relative to the producer prices. The proportion of the producer price to the consumer price was about 40 % in 1981 and from 1995 to 2003 it was 15 - 20%. In regime 1, the adjustment parameters were not significantly different from zero and there was no cointegration.

Regime 2 includes the observations from 1989 to 2003. This time the structural change is probably over and one can start speaking about a long-term equilibrium and deviations from it. It can be seen from the estimates of the adjustment coefficients that producer prices do not react but retail prices respond to correct for discrepancies from the long-term equilibrium. A practical understanding of the meaning of adjustment parameters can be obtained by calculating the time it takes for the equilibrium error to halve. The formula for this is $\ln(0.5)/\ln(1+\alpha_2)$. With –0.059 the result is about 11 months.

However, this is not the whole story. Besides error correction there is also adjustment from the short-term dynamics. The interpretation of lagged price differences is performed in the same way as in a typical regression model. The coefficient for the retail price change as a response to the producer price change, a_{21}, is significant and positive in both regimes.

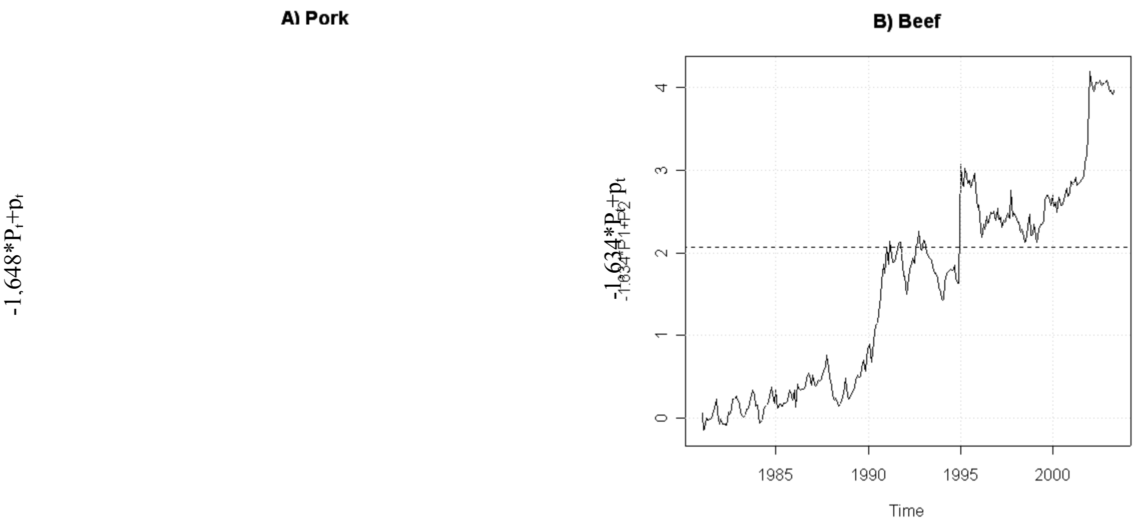


Figure 5. *Graphs of the assumed stationary linear combination of producer and consumer prices of a) pork and b) beef. The dotted line indicates the threshold between the regimes.*

Since it is impossible to say anything about the possible asymmetry of price transmission on the basis of the previous model, we estimated the following three-regime model:

$$(5) \qquad \Delta y_t = \begin{cases} \alpha^1(\beta' y_{t-1} - r_2) + \varepsilon^1, & \text{when } \beta' y_{t-1} \le r_1, \\ \alpha^2(\beta' y_{t-1} - r_2) + \varepsilon^2, & \text{when } r_1 < \beta' y_{t-1} \le r_2, \\ \alpha^3(\beta' y_{t-1} - r_2) + \varepsilon^3, & \text{when } \beta' y_{t-1} > r_2, \end{cases}$$

where $y_t = (P_t,p_t)'$, $\beta = (\beta_1,1)'$ and P_t and p_t are the producer and consumer prices, respectively. As in the previous model, regime 1 represents the first part of the series by the end of 1988 when the prices were changing structurally. Regime 2 is a squeezed margin case in which the producer price is too high relative to the long run price equilibrium or the consumer price is too low, or both. In regime 3, the margin is stretched, because the producer price is too low and/or the consumer price is too high.

The estimation results for this model are given in Table 4. The estimates are maximum likelihood estimates and the confidence intervals for the α^i are calculated assuming that β_1, r_1 and r_2 are estimated without error. The BIC for this model is 3352. There seems to be no significant difference in adjustment speed between regimes 2 and 3. However, the result confirms our earlier result that the consumer price alone adjusts towards the equilibrium.

Table 4. *Pork co-integration threshold 3-regime model parameter estimates.*

Regime	1		2		3	
Variable	Estimate	Confidence interval	Estimate	Confidence interval	Estimate	Confidence interval
α_1	-0.005	(-0.003, -0.007)	0.003	(-0.016,0.022)	-0.014	(-0.035, 0.008)
α_2	-0.016	(-0.013, -0.020)	-0.071	(-0.097,-0.045)	-0.084	(-0.129,-0.039)
β_1	-1.66					
r_1	3.49					
r_2	4.98					

Model (5) is different from model (4) in that it does not include the constants c^j. If these constants are included and they are estimated without restrictions, each regime has its own drift. Since in model (5) the drift is assumed to be 0 in all regimes, long term adjustment is needed to explain the increase of the prices in regime 1. Therefore, when model (5) is used, the adjustment coefficients α_i are statistically significant in this regime, while they were not significant when model (4) was used.

4.3 Beef co-integration threshold model

The results from the beef co-integration threshold model are reported in Table 5.

Table 5. *Beef co-integration threshold model parameter estimates*

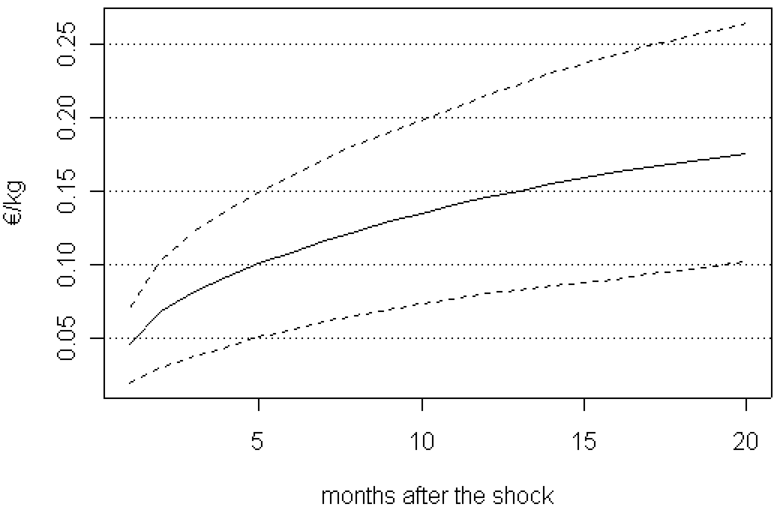
Regime Variable	Median	Credible interval
c_1	0.017	(0.001, 0.072)
c_2	0.044	(0.016, 0.123)
α_1	-0.009	(-0.021, -0.002)
α_2	-0.022	(-0.033, -0.014)
β_1	-1.634	(-1.928,-0.988)
r	2.065	(1.079,3.650)

Both adjustment coefficients α_1 and α_2 are negative. This implies that in the case of the squeezed margin both prices increase. Since the consumer price increases more rapidly, the margin increases. In the case of the stretched margin both prices decrease but the consumer price more rapidly. The half life of the disequilibrium error is $\log(0.5)/\log(1+\alpha_1\beta_1 + \alpha_2) \approx 95$ months or 8 years. Thus, the adjustment speed is very slow.

In this model we do not have short term dynamics. The price series is a two dimensional random walk process except for the cointegrating relation which binds the two component series together. The model is nonlinear only because the covariance matrices are assumed to be unequal in different regimes. This accounts for the fact that the error variance is large for the consumer price in the second regime. The first regime roughly consists of the first half of the series and the second regime of the second half (see Figure 5b). There is a clear increasing trend in the margin of the prices. The proportion of the producer price to the consumer price was about 60 % in 1981 and about 27 % in the beginning of 2003.

4.4 Impulse response analysis

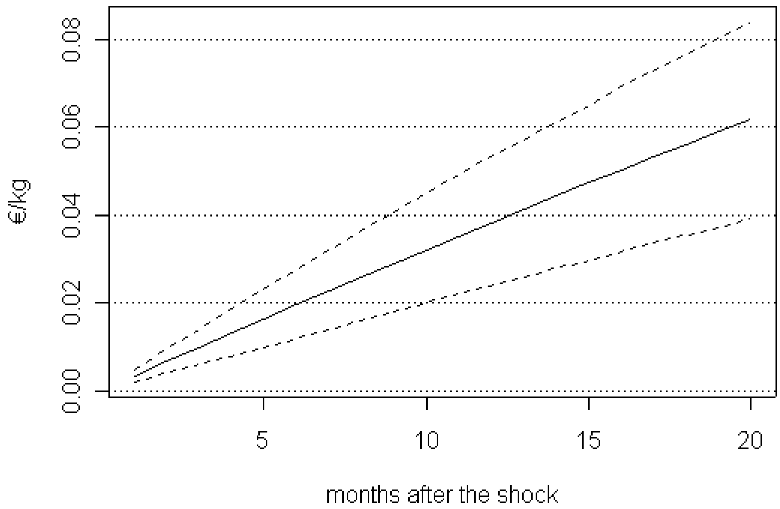
In Section 2.3 it was briefly mentioned that impulse response analysis describes dynamic interrelations between the variables in the system after a specific shock has occurred. Usually, the size of the shock is assumed to be either unity or the standard deviation of the error term. In the latter case it is easy to interpret the shock as an average one. In linear models the sign and the timing of the shock are irrelevant, but in nonlinear models both of these factors have to be taken into account. For example, the reaction of the retail price to a producer price shock depends on the sign of the shock, i.e. whether the producer price increases or decreases. In this paper we use 0.1 €/kg as the shock and the timing is the end of the data. Impulse responses were calculated to both negative and positive shocks. Since they turned out to be symmetric for our data sets, only responses to a positive shock are shown. Exact numbers are also provided for some months. In the figures we report impulse responses up to 20 months.



Time months	0	2	9	18
Producer price	10	14	14	14
Consumer price	0	7	13	17

Figure 6. *The response of consumer prices to a 0.1 € rise in producer prices in the pork cointegration threshold model*

In Figure 6 the response of the consumer price is given for pork. We have used the 2-regime model represented in Section 4.2. The immediate adjustment in the following month equals about 0.05 euros. However, adjustment does not stop after one month, but continues and after 20 months the permanent reaction is about 1.75 times greater than the original shock. The reactions to positive and negative shocks are symmetric, since the latest data point belongs to the second region and small shocks cannot move the system to the first region. Figure 7 presents the corresponding results for beef markets.



Time months	0	2	9	18
Producer price	10	10	11	12
Consumer price	0	1	3	6

Figure 7. *The response of consumer prices to a 0.1 € rise in producer prices in the beef cointegration threshold model*

In the beef market models, the reaction of the consumer price to a change in the producer price is very slow. Although the price changes are statistically significant, their economic meaning is marginal. This is due to the adjustment coefficients, which have small absolute values. They are so small, since in the last part of the time series there is no evidence of price adjustment. In fact, it is possible that producer price changes are followed by a consumer price reaction of opposite sign. As noted earlier, there was a time period at the end of the time series, when the consumer price was rising and the producer price slightly declining.

5. SUMMARY AND CONCLUSIONS

The purpose of this study was to determine how producer price changes are transmitted to retail prices. The emphasis was on testing asymmetric price transmission with threshold models, because asymmetry is highly likely and ignoring it leads to an econometric misspecification. Besides empirical studies, the economic theory relating to asymmetric price transmission was also briefly reviewed. The time series models used in this paper were formulated in two different ways. The first method was the usual convention presented in the literature where the error correction model is divided into different regimes based on a stationary linear combination of the producer and retail prices. This specification is founded solely on the cointegration property of the price series. The second model still utilizes long-term information provided by cointegration, but the splitting of the data into different regimes is determined by the producer price changes. Parameter estimates and impulse response functions were only computed for the first model, since the second model appeared to be inferior to the corresponding linear model.

On the basis of our analysis, the first part of the pork price series was not cointegrated and it only indicated a structural change in the formation of the consumer price. In the second part the results did not statistically support the hypothesis of asymmetric price transmission. However, it should be noted that it is not obvious when the prices have converged on a condition such that the possible asymmetry can be reliably studied. Besides, if asymmetry is small, a longer time interval would be needed to obtain statistically significant results. As far as the beef markets are concerned, they were characterized by a continuous structural change. The proportion of the producer price to the consumer price was decreasing all the time. In such a situation it is impossible to say anything about the asymmetry of price transmission.

It is usually the retail price that responds, which is understandable given the fact that producing meat is a biological process with production lags and inelastic supply in the short term. Our result concerning the price adjustment of the pork prices is similar to those reported by von Cramon-Taubalel (1998) and Abdulai (2002) in the stretched margin case. In the squeezed margin case our estimate for adjustment speed is much smaller. The reports of Goodwin and Holt (1999) and Goodwin and Harper (2000) were applications of three regime models, so their results are not directly comparable with ours from two regimes. Moreover, their parameter estimates are closer to the results of Cramor-Taubadel and Abdulai than to our results.

REFERENCES

Abdulai, A. 2002. Using threshold cointegration to estimate asymmetric price transmission in the Swiss pork market. Applied Economics, 34.

Balke, N.S. & Fomby, T.B. 1997. Threshold Cointegration. International Economic Review. Vol. 38, No 3.

Enders, W. & Granger, C.W.J. 1998. Unit-Root tests and Asymmetric Adjustment With an Example Using the Term Structure of Interest Rates. Journal of Business & Economic Statistics. Vol. 16, No 3.

Engle, R.F. & Granger, C.W.J. 1987. Co-integration and Error Correction: Representation, Estimation, and Testing. Econometrica, 55.

Forbes, C.S., Kalb, G, R, J. & Kofman, P. 1999. Bayesian Arbitrage Threshold Analysis. Journal of Business & Economic Statistics. Vol. 17, No 3.

Goodwin, B.K. & Harper, D.C. 2000. Price Transmission, Threshold Behaviour, and Asymmetric Adjustment in the U.S. Pork sector. Journal of Agricultural and Applied Economics. 32,3.

Goodwin, B.K. & Holt, M.T. 1999. Price Transmission and Asymmetric Adjustment in the U.S.Beef sector. American Journal of Agricultural Economics. Vol. 81, No 3.

Granger, C.W.J. & Lee, T.H. 1989. Investigation of production, Sales and Non-symmetric Error Correction models. Journal of Applied Econometrics. Vol. 4.

Hansen, B.E. & Seo, B. 2002. Testing for two-regime threshold cointegration in vector error-correction models. Journal of econometrics, 110.

Johansen, S. 1991. Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. Econometrica, 59.

Koop, G. & Potter, S.M. 1999. Bayes factors and non-linearity: Evidence from economic time series. Journal of Econometrics, 88

McCorriston, S., Lloyd T., Morgan W. & Rayner, A. 1999. Ministry of Agriculture, Fisheries and Food. "Report of an Investigation into the Relationship between Producer, Wholesale and Retail Prices of Beef, Lamb and Pork. MAFF, London.

Meyer, J. 2003. Measuring market integration in the presence of transaction costs: A threshold vector error correction approach. 25th IAAE Conference in Durban/South Africa – Contributed paper.

Meyer, J. & Von Cramon-Taubadel, S. 2002. Asymmetric price transmission: a survey. Paper at the Xth EAAE Conference in Zaragoza.

Obstfeld, M. & Taylor, A.M. 1997. Nonlinear Aspects of Goods-Market Arbitrage and Adjustment: Heckscher's Commodity Points Revisited. Journal of the Japanese and International Economies 11.

Pelzman, S. 2000. Prices Rise Faster than They Fall. Journal of Political Economy. Vol. 108, No 3.

Potter, S.M. 1998. Nonlinear Impulse Response Functions. Federal Reserve Bank of New York.

Von Cramon-Taubadel, S. 1998. Estimating asymmetric price transmission with the error correction representation: An application to the German pork market. European Review of Agricultural Economics. 25.

Zellner, A. 1971. An Introduction to Bayesian Inference in Econometrics. New York. Wiley.

APPENDIX I

Details of consumer retail carcass/basket

Beef	share in consumer basket, %
Topside	11.9 (9.5 after 1993)
Shoulder	7.6
minced meat	80.5
sliced topside (beginning of 1993)	2.4

From the beginning of 1993 the share of topside is reduced because it is the source for sliced topside. As the exact share of sliced topside is not so important, the share is determined so that the value of the consumer basket continues smoothly at the change point.

Pork	share in consumer basket, %
pork chop	30.6 (12.5 after 1988 and 20.6 after 2002)
side of bacon (not after beginning of 2002)	36.1
smoked ham	33.3 (58.8 after 2002)
pork sirloin (beginning of 1988)	12.5 (after 1988 and 20.6 after 2002)

The same principle of smoothing of the consumer basket value applies here. Pork sirloin is almost same as pork chop without bones.

APPENDIX II

Unit-root tests 1981 (2) to 2003 (5)
Critical values: 5%=-2.872 1%=-3.456; Constant included

	t-adf	beta Y_1	\sigma	lag	t-DY_lag	t-prob	F-prob
Beef P_t	-0.2581	0.9984	0.1021	0			

	t-adf	beta Y_1	\sigma	lag	t-DY_lag	t-prob	F-prob
Beef p_t	-1.8274	0.99154	0.086004	3	3.8698		0.0001
Beef p_t	-1.8686	0.99113	0.088277	2	4.7119	0.0000	0.0001
Beef p_t	-2.0881	0.98971	0.091779	1	0.64002	0.5227	0.0000
Beef p_t	-2.1264	0.98955	0.091676				0.0000

	t-adf	beta Y_1	\sigma	lag	t-DY_lag	t-prob	F-prob
Pork P_t	-0.71825	0.99395	0.093983	0			

	t-adf	beta Y_1	\sigma	lag	t-DY_lag	t-prob	F-prob
Pork p_t	-2.6763	0.98869	0.091632	1	4.6475		0.0000
Pork p_t	-3.1012*	0.98648	0.095127				0.0000

Beef:

Ho:rank=p	-Tlog(1-\mu)	using T-nm	95%	-T\Sum log(.)	using T-nm	95%
p = 0	22.07**	21.91**	14.1	23.28**	23.1**	15.4
p ≤ 1	1.205	1.196	3.8	1.205	1.196	3.8

Pork:

Ho:rank=p	-Tlog(1-\mu)	using T-nm	95%	-T\Sum log(.)	using T-nm	95%
p = 0	34.28**	34.03**	14.1	38.46**	38.17**	15.4
p ≤ 1	4.179*	4.148*	3.8	4.179*	4.148*	3.8

APPENDIX III

The *t*th rows of Y (N×m), X (N×(m+1)), Z (N×m), and E (N×m) are $\Delta y'_t$, $(1,\Delta y'_{t-1})$, y'_{t-1} and ε'_t, respectively, and let W = (X Zβ). The matrices Y^j, W^j and E^j include observations from Y, W and E that belong to the j*th* regime. We can then write model (4) in a matrix form:

$$Y^j - W^j B^j = E^j$$

where B^j is obtained by stacking $c^{j\prime}$, $A_i^{j\prime}$ and $\alpha^{j\prime}$. The likelihood function is a product of R+1 likelihood functions (see, e.g., Forbes, Kalb and Kofman (1999) and Zellner (1971))

$$L(r_1,..,r_R,\beta,\alpha^j,A^j,\Sigma^j) \propto \prod_{j=1}^{R+1} (2\pi)^{-mN^j/2} \left|\Sigma^j\right|^{-N^j/2} \exp\left\{-\frac{1}{2} trace \Sigma^{j^{-1}} E^{j\prime} E^j\right\}$$

We use a non-informative prior for the parameters:

$$q(r_1,.., r_R,\beta,\alpha^j,A^j,\Sigma^j) \propto \prod_{j=1}^{R+1} \left|\Sigma^j\right|^{-(m+1)/2} .$$

Then, the posterior distribution can be written in the form

$$p(r_1,..,r_R,\beta,\alpha^j,A^j,\Sigma^j \left| d\right.) \propto \prod_{j=1}^{R+1} (2\pi)^{-mN^j/2} \left|\Sigma^j\right|^{-(N^j+m+1)/2} \exp\left\{-\frac{1}{2} trace \Sigma^{j^{-1}} E^{j\prime} E^j\right\}$$

where d is data. The conditionals for the parameters in $vec(B^j)=b^j$ and covariance matrix Σ^j are

$$b^j \mid \Sigma^j, r_1,\dots,r_R, \beta, d \sim N(\hat{b}^j, \Sigma^j \cdot (W^j W^j)^{-1})$$

$$\Sigma^j \mid r_1,\dots,r_R, \beta, d \sim IW(N^j - k, S^j)$$

where IW(v, S) is inverted-Wishart with v degrees of freedom and with matrix parameter S, $N(b,\Sigma^j \cdot (W^j W^j)^{-1})$ is a multivariate normal with covariance matrix $\Sigma^j \cdot (W^j W^j)^{-1}$, $\hat{b}^j$ $vec(\hat{B}^j)$, $\hat{B}^j = (W^{j\prime} W^j)^{-1} W^{j\prime} Y^j$ and $S^j = (Y^j - W^j \hat{B}^j)'(Y^j - W^j \hat{B}^j)$. Finally, integrating B^j and Σ^j out we get the joint marginal posterior for $r_1,\dots,r_R$ and β

$$p(r_1,..,r_R,\beta|d) \propto \prod_{j=1}^{R} \left|W^j{}'W^j\right|^{-m/2} \left|S^j\right|^{-(N^j-k)/2} \prod_{i=1}^{m} \Gamma[(N^j-k+1-i)/2]$$

The joint marginal posterior for the parameters $r_1,\ldots,r_R$ and β can be simulated, for example, using the MCMC methods. We tested Metropolis-Hasting algorithm and acceptance rejection sampling and found these techniques to be computationally difficult. However, in our two-dimensional case (r and β_1), we were successful in using a simple following algorithm. We used a grid search to find the 20 000 most probable values of the distribution of parameters r and β_1 and used these values to simulate the posterior distribution.

In the case of nonlinear impulse response we refer to Potter (1998) and denote the conditional expectation of Y_{t+n} given information up to time t by $E[Y_{t+n}|y^t(\omega), \theta]$ for $n > 0$, where $y^t(\omega)$ identifies the individual realization of the time series up to time t at ω ($y^t(\omega) = y_t,\ldots,y_1$) and θ includes all parameters. A nonlinear impulse response function can then be defined by the difference between an expectation conditioned on the sample path $y^t(\omega')$ and sample path $y^t(\omega)$, where $y^t(\omega')$ is equal to $y^t(\omega)$, except for the element y_t, which is perturbed by some constant δ:

$$\text{nlinf}(y^t(\omega), \delta, \theta) = E[Y_{t+n}|y^t(\omega'), \theta] - E[Y_{t+n}|y^t(\omega), \theta]$$

To estimate $\text{nlinf}(y^t(\omega), \delta, \theta)$ we generated 1 000 random observations from the distribution $f(y_{t+n}|y^t(\varpi),\theta)$ and calculated the corresponding perturbed values. Then we calculated the mean of the differences of the perturbed and unperturbed observations. We also repeated this procedure 1000 times allowing for randomness in the parameter vector θ to obtain 95 % probability limits for the impulse response functions.

APPENDIX IV

We tested the equality of the adjustment vectors α^j and the short term dynamic matrices A^j in the different regimes using likelihood ratio tests. If $\hat{\theta}$ and $\hat{\theta}_0$ are the maximum likelihood estimates of the full and restricted models, respectively, and L(.) denotes the likelihood function, then $-2(\log(L(\hat{\theta}_0))-\log(L(\hat{\theta})))$ is approximately $\chi2$(d)-distributed where d is the number of restrictions.

***Table 1**.* *Tests for the pork price series*

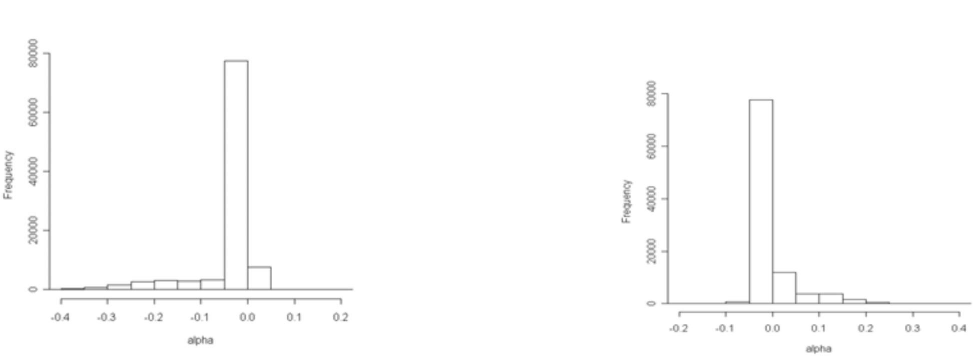
H_0:	TEST	df	p-value	TEST	df	p-value
0 degrees						
a) 2 regimes						
$\alpha^1=\alpha^2$	25.9	2	2.4 e-6	2.42	2	0.30
b) 3 regimes						
$\alpha^1=\alpha^2=\alpha^3$	24.5	4	6.4 e-5	8.84	4	0.065
1 degree						
a) 2 regimes						
$\alpha^1=\alpha^2$	12.4	2	0.002	2.63	2	0.27
$A^1=A^2$	2.48	4	0.65	11.4	4	0.022
b) 3 regimes						
$\alpha^1=\alpha^2=\alpha^3$	16.0	4	0.0030	8.77	4	0.067
$A^1=A^2=A^3$	5.25	8	0.73	10.7	8	0.22
	case z_1			case z_2		

Table 2. *Tests for the beef price series*

H_0:	TEST	df	p-value	TEST	df	p-value
0 degrees						
a) 2 regimes						
$\alpha^1=\alpha^2$	1.92	2	0.38	0.44	2	0.80
b) 3 regimes						
$\alpha^1=\alpha^2=\alpha^3$	20.68	4	0.00037	1.35	4	0.85
1 degree						
a) 2 regimes						
$\alpha^1=\alpha^2$	5.01	2	0.081	1.307	2	0.52
$A^1=A^2$	8.82	4	0.066	7.52	4	0.11
b) 3 regimes						
$\alpha^1=\alpha^2=\alpha^3$	12.7	4	0.013	1.92	4	0.75
$A^1=A^2=A^3$	26.3	8	0.00095	17.84	8	0.022
	case z_1			case z_2		

APPENDIX V

Pork co-integration threshold model parameter histograms



regime 1 regime 2

Figure 1. *The histograms of α_1 in regimes 1-2*

regime 1 regime 2

Figure 2. *The histograms of α_2 in regimes 1-2*

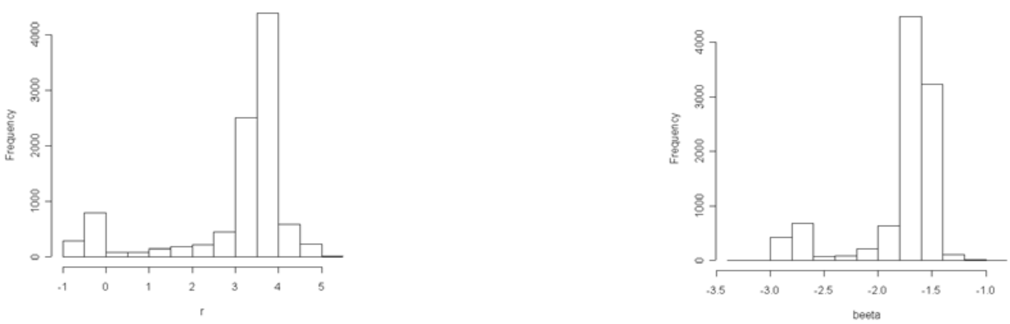


Figure 3. *The histograms of r and β_1*

Beef co-integration threshold model parameter histograms

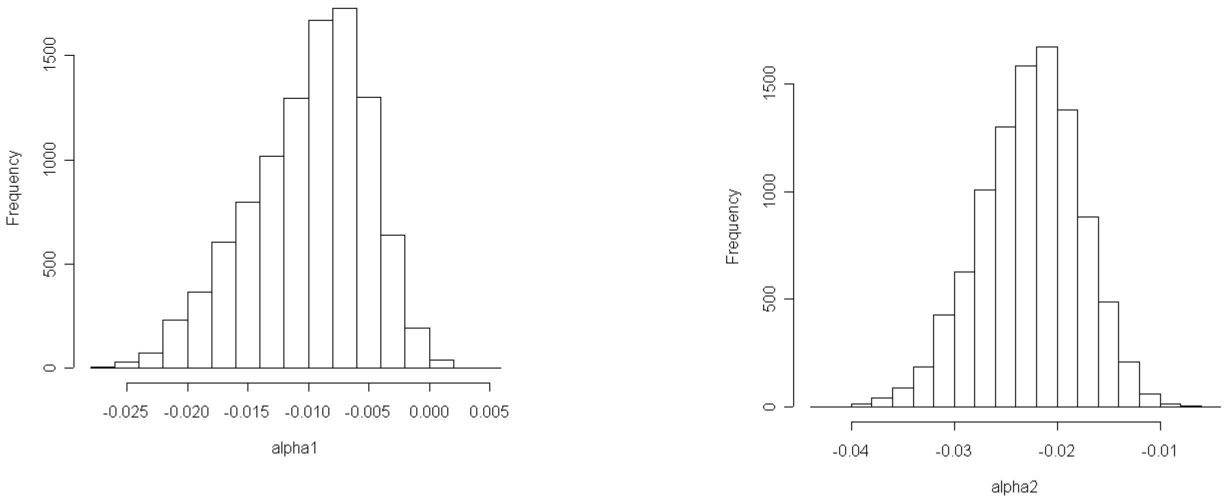


Figure 4. *The histograms of α_1 and α_2*

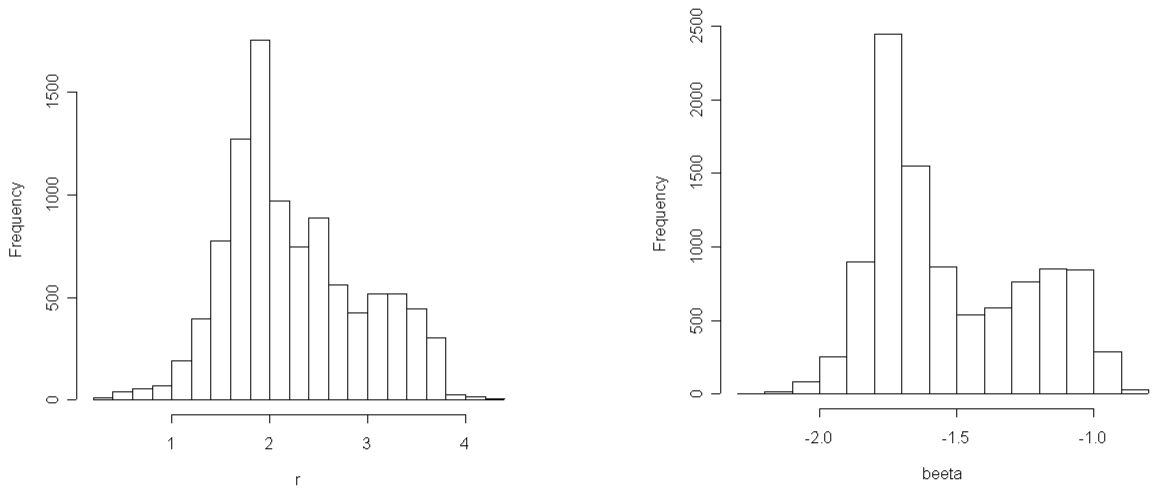


Figure 5. *The histograms of r and β*

PELLERVON TALOUDELLINEN TUTKIMUSLAITOS PTT
Pellervo Ekonomiska Forskningsinstitut
Pellervo Economic Research Institute
Eerikinkatu 28 A, 00180 Helsinki, Finland
puh. (09) 3488 844, telefax (09) 3488 8500, s-posti: econ.res@ptt.fi

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