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Horizontal price transmission of the Finnish meat sector with major EU players

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Abstract

The integration of the Finnish meat market in the EU has important implications for domestic agricultural policy. Our aim is to estimate the characteristics of the Finnish pork and beef markets in relation to those of Germany and Denmark. Our analysis uses symmetric and asymmetric threshold error correction models. Both pork and beef prices in Finland are found to have slowly cointegrated with German prices, but the cointegration relationship of the two counties is only found to be symmetric for pork prices, while it is asymmetric for beef prices. The producer price for pork in Finland is symmetrically cointegrated with the Danish price, but the Finnish and Danish beef prices show a random walk. This implies that the price transmission to the Finnish pork producer market from the EU market is smoother and more efficient than for the beef market. However, the speed of transmission is still slow compared to that between the Danish and German markets.

Key Words: cointegration, asymmetric, error correction, thresholds, pork and beef prices

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1. Introduction

The economic structure of Finnish agriculture and agrifood radically changed across all sectors in 1995 when Finland joined the EU. The commitment to the Common Agricultural Policy (CAP) of the EU, in particular, directly or indirectly affected the prices for Finnish agricultural goods in both the short- and long-term. The immediate impact of the change was clearly seen in the meat sector in 1995. Meat prices fell to about half of the level of 1994 before Finnish accession the EU, even though the producers received subsidies for the transitional period. The producer price for meat in Finland has become much more volatile, and the price level has followed the average price in the EU quite closely ever since. This indicates that the prices for Finnish agricultural products, including meat products, have become more subject to the changes in other EU countries. Meanwhile, the trading volume of meat between Finland and other EU countries has fluctuated since 1995 (see Figure 1). Pork and beef have been the main meat products for both the Finnish domestic market and Finland's trading partners. Each year, the trading volume of pork and beef,[1] excluding process meat products, has accounted for over 38% of total trade in the meat and offal sector. As illustrated in Figure 2, pork has always been the main traded meat in the meat sector.

In the EU, Germany is the largest producer and consumer of meat, while Denmark is one of the major meat producers, and in particlar a leading pork exporter. Thus, the trade between Finland and these two countries, and particularly the imports from them, dominates in comparison to the trade between Finland and the other EU-27 countries. Figure 3 presents the import of pork and beef from Germany and Denmark to Finland between 1995 and 2009. Clearly, the import of pork from Germany to Finland has gradually increased, and as a result it reached 7.4 thousand tons in 2009 from 3.8 tons in 1995, a 1930-fold increase within 14 years. In comparison, the import of pork from Denmark to Finland has steadily declined from a peak volume in 1999. The import of beef from the two countries to Finland has shown a similar pattern, with Denmark first leading and Germany later catching up, especially after 2001.

Under the EU's Common Agricultural Policy, the same agricultural products are required to become spatially integrated within and between all member states. In an integrated market, price information related to the production costs should also be efficiently transmitted between the member states. National governments of EU member states and their regulations should help to attain the goal of an integrated and efficient market. The basis for price transmission is traditionally

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founded on the concept of the law of one price (LOP) (Samuelson, 1952; Takayama and Judge, 1972) or, at an aggregate level, on purchasing power parity (PPP) (Goodwin and Piggot, 2001; Jamaleh, 2002; Seo, 2003; Brooks and Melyukina 2003). The LOP or PPP postulates that under conditions of perfect competition, price transmission is complete, with the equilibrium prices of a commodity sold on foreign and domestic markets differing only by the transfer costs if the markets share a common currency.

In practice, perfect price transmission is not realistic, especially for a small open economy such as that in Finland, and a high degree of integration with other member of states in production costs is also known to be unrealistic, as the production costs in Finland are substantially higher than in the main agricultural areas of Europe. The Finnish costs cannot, therefore, be integrated with the competitive production costs and product prices in the EU. As a result, price transmission models are mostly used to provide important implications as to how changes in either supply or demand in one market will transmit to another. Consequently, price transmission can reflect the extent of market integration and the extent of market efficiency. Thus, in analysing the efficiency of the meat market, a fundamental issue is the extent to which the Finnish domestic meat market responds to changes in the European price. Price transmission from the major European markets to domestic markets is central in understanding the extent of integration of economic agents into the market process (FAO, 2003)[2]. In other words, to understand the form of prices in the domestic market of Finland, the relationship between the Finnish market and such important European markets as Germany and Denmark has become an interesting topic of study. The issues of market efficiency and the extent of price transmission of market information have attracted considerable attention during the last couple of years, as the price of food in Finland has dramatically increased. Many questions have been asked about price transmission between the Finnish market and the EU market, and questions such as whether farmers in Finland have benefitted from the price increases need to be carefully addressed.

Another issue that is important for spatial market integration is its importance in regulating the structure of food processing through antitrust legislation, because domestic retail businesses in Finland are highly concentrated. Since the most common reason for the prevalence of vertical restraints in the food sector is the increasing market power of food retailers (McCorrison 2002), market concentration can be expected to have important implications, particularly in the Finnish

[2] FAO, 2003, http://www.fao.org/docrep/006/y5117e/y5117e06.htm

market. A concentrated market structure is known to be a key condition providing firms with incentives for oligopolistic behaviour, such as non-cooperative tacit collusion, strategic price signalling and strategic investment (Tirole 1992). From the perspective of the Finnish meat market, the problem is that domestic retailing is more concentrated than domestic processing. Even though the processing industry is also quite concentrated, Finnish processors are too small to cope in the overly competitive European-wide and global export markets. An important question then is what types of public policies would efficiently regulate the domestic meat processing industries, and what means could be adopted or promoted to improve their competitiveness. The crucial research question linked to the policies regulating structural development in domestic meat processing is the size of the Finnish market. In economic concepts, the issue is how well the Finnish meat market is integrated into the European-wide meat market, and what are the characteristics of spatial transmission of price information between the Finnish and other European markets.

The objective of this paper is to estimate the characteristics of the Finnish pork and beef markets in relation to their markets elsewhere in Europe. More specifically, the study aims to quantify the elasticity of price transmission between the Finnish meat market and other European meat markets, focusing on producer prices for pork and beef in Finland, Germany and Denmark.

Some attempts have been made in the literature to investigate price transmission between Finland and other European countries, but the results have remained mixed with respect to different products. For example, the broiler price was not found to be cointegrated at all (Xing, 2008), while the producer price for pork meat was cointegrated with that in Germany, but the degree of market integration was very low (Jalonoja et al, 2005). To some extent, the previous studies have become outdated, since they do not include the most recent price pick for the last two years. Given the discrepancy in the literature as to whether the meat market in Finland during the last two years has altered [and?] the extent to which the integration with other EU markets may have changed over time, this article seeks to further explore this issue. A re-examination of this question is especially necessary in the light the possible structural change in the meat market within the last two years.

2. Theoretical model

Based on the law of one price (Krugman and Obstfel, 1997; Mundlack and Larson, 1992), the domestic price for meat can be written as a function of the international meat price, the nominal exchange rate and the transaction costs. In market integration studies, econometric analysis is mostly carried out on the logarithms of the prices in question. Thus, the Finnish domestic producer price for pork or beef can be written as a bivariate function of the logarithm formed from the German or Danish price for pork or beef, as shown in model (1):

$$\ln p_{it}^{Fin} = \beta_0 + \beta \ln p_{it}^{j} + \mu_{it} \tag{1}$$

where β_0 is a constant term that captures transactions costs and β is a coefficient representing the elasticity of price transmission, which is assumed to have the value of one for perfectly integrated markets and when a strong form of the LOP holds. Meanwhile, β also gives the relationship between the market prices. If $\beta = 0$ there is no relationship between the prices, while if $0 < \beta < 1$ there is a relationship between the prices, but the relative price is not constant, and the goods are imperfect substitutes. $\ln p_{it}^{Fin}$ represents the price on the Finnish market for the products *i*; $\ln p_{it}^{Fin}$ represents the price on the German or Danish market; the subscript index *i* represents pork or beef, while the subscript *j* signifies the country (Germany or Denmark); β is assumed to be unity, as if there is perfect integration between $\ln p_{it}^{fin}$ and $\ln p_{it}^{j}$; and $\mu_{it} \sim IID(\mu, \sigma^2)$ is the error term, uncorrelated with other explicative variables of the model.

However, the model presented in equation (1) was found to have shortcomings associated with the nature of the unit-root nonstationarity of most commodity price data and nonstationary price data. The presence of nonstationarity in the price series commonly used to test spatial market integration invalidates conventional approaches to inference such as model (1). (Engle and Granger, 1987). Recognition of this issue has stimulated an extensive body of literature applying unit root and cointegration tests to evaluations of spatial integration (Bessler and Fuller, 1994; Conforti, 2004). A frequently used technique to identify cointegrated behaviour and meanwhile separate out the short-term adjustment component and the long-term equilibrium component is the error correction model (ECM). Using cointegration theory, the ECM can rewrite equation (1) as a bivariate equation as follows:

$$\Delta \ln P_t^j = \Phi_0 D_t + \sum_{k=1}^{p-1} \Gamma_k \Delta P_{t-k}^j + \alpha(\beta' P_{t-1}^j + \beta_0) + \varepsilon_t \tag{2}$$

where Δ is the difference operator, $\ln P_t^j$ is a 2×1 vector of dependent variables (pairwise combinations of prices for Finnish meat with German and Danish meat), and Φ_0 is a 2×1 vector of coefficients for a deterministic term consisting of a vector of D_t possible trend dummies and intercept terms. Each Γ_k represents a 2×2 matrix of coefficients for corresponding meat prices. Γ_k also demonstrate the short-term dynamics of the system, given that a long-term cointegration relationship exists between Finnish meat and German or Danish meat, represented by the error correction term (*ECT*) $(\beta' P_{t-1}^{j} + \beta_0)$. In the *ECT*, β contains the cointegrating vectors or long-term equilibrium of the prices, and the loading factor α shows the speed of adjustment towards the long-term equilibrium following a short-term deviation. Within this context, short-term adjustments are directed by, and consistent with, the long-term equilibrium relationship, allowing the researcher to assess the speed of adjustment that shapes the relationship between the two prices. Usually, $0 < |\alpha| < 1$, and in the context of market integration and price transmission studies, the value of α can be seen as a proxy for the extent to which policies, transaction costs and other distortions delay full adjustment to the long-term equilibrium (Sharma, 2002). Finally, the error term vector ε_t denotes a 2×1 vector of mutually orthogonal random price disturbances, assumed to be serially uncorrelated with a zero mean and constant variance.

Given that the cash and futures prices are both nonstationary and I(1), the error-correction specification is estimated using the method of Johansen (1990). This is based on reduced rank restrictions on the vector autoregressive representation $\alpha\beta'$, i.e., if the two series are cointegrated, then rank $\alpha\beta' = r < 2$. Johansen's (1992) sequential likelihood ratio test is used to determine the cointegration relationship, which is a trace statistic denoted by LR_{tr}. The trace statistic is shown in function (3):

$$LR_{tr}(r \mid k) = -T \sum_{i=r+1}^{k} \ln(1 - \hat{\lambda}_i) \qquad (3),$$

where λ_i denotes the i-th largest eigenvalue of the matrix Π in function (2). The maximum eigenvalue statistic tests the null hypothesis of r cointegrating relations against the alternative of r+1 cointegrating relations. This test statistic is computed as function (3):

$$LR_{\max}(r \mid r+1) = -T \ln(1 - \lambda_{r+1}) = LR_{tr}(r \mid k) - LR_{tr}(r+1 \mid k) \qquad (4),$$

for r = 0, 1, 2….., k-1.¨

In more detail, Johanson tests according to equations (3) and (4) could test both the unrestricted model (with a trend) and restricted model (without a trend). Thus, the test for the cointegrating relationship between the Finnish meat price and German/Danish meat price, where n = 2, becomes the test for the null hypothesis: $r = 0$ and $r = 1$ with and without a trend, starting without trend. If $r = 0$ is rejected and $r = 1$ cannot be rejected, a cointegrating relation is found between the Finnish meat price and German/Danish meat price. Otherwise, if $r = 0$ cannot be rejected, there is no cointegration relationship between the prices.

There are two cases in which cointegration analysis cannot make inferences. One is that it cannot be used to make inferences about the direction of causation between variables. The other is the case when a non-cointegration relationship between prices is found. Therefore, further Granger causality tests are performed. In the first case, Granger's causality test (1988) can be used to determine the direction of causality. At least one-way causality is a necessary condition, and the test provides additional evidence as to whether a cointegration relationship holds between two series (Granger, 1986). More specifically, a weak exogeneity test is imposed on the vector α, in which at least one parameter has to be non-zero. In the latter case, when the two series are found not to be cointegrated, Granger-type tests require transformations to induce stationarity without the *ECT*.

The linear VECM, as show in equation (2), has been noted to be highly sensitive to structural breaks (Shepherd, 2004; Brummer *et al.*, 2009), which might cause the instability of parameters and price asymmetry (Von Gramon-taubadel, 2003; Ben-Kaabia, 2005). Therefore, instability tests, namely the Chow breakpoint test and cointegrating vector stability test (Hansen and Johansen, 1999), are used to examine the long-term stability of the linear VECM and potential structural breaks and changes. More specifically, the Chow breakpoint test is applied to examine the hypothesis of a stable long-term relationship and the cointegrating vector stability test is used to evaluate the stability of the recursively estimated eigenvalues of the cointegrating vector, which in this study is the estimated eigenvalue of ECT $\alpha(\beta' P_{t-1}^{j} + \beta_0)$ from equation (2).

Furthermore, an important issue in the empirical application of price transmission to the Finnish meat market from other European markets explored here is to test the linearity of the VECM against non-linear models. By doing this, the linear VECM could be tested to determine whether the producer prices for both pork and beef have been symmetrically transferred to Finnish producers from other major European meat markets. The presence of asymmetries in the price transmission mechanism has been investigated for a wide variety of countries and commodities (Frey and Manera, 2008). This has mostly been done by releasing the adjustment mechanism in order to allow the regime shifts to depend on whether prices are increasing or decreasing (Von Cramon-Taubadel, 2004; Keele, 2005). Hansen (1999) and Hansen and Seo (2002) developed a sup-LM test for the linear VECM against a bivariate threshold vector error correction model (TVECM) with a maximum of two thresholds and three regimes. If the linearity of the VECM shown in equation (2) is rejected, the TVECM can be applied as follows:

$$\Delta \ln P_t^j = \begin{cases} c^L + \alpha^L ECT_{t-1} + \sum_{i=1}^{k-1} \Gamma_i^L \Delta \ln P_{t-1} + \varepsilon_t^L, & if \quad ECT_{t-1} < \hat{\gamma}^L \\ c^M + \alpha^M ECT_{t-1} + \sum_{i=1}^{k-1} \Gamma_i^M \Delta \ln P_{t-1} + \varepsilon_t^M, & if \quad \hat{\gamma}^L < ECT_{t-1} < \hat{\gamma}^H \\ c^H + \alpha^H ECT_{t-1} + \sum_{i-1}^{k-1} \Gamma_i^H \Delta \ln P_{t-1} + \varepsilon_t^H, & if \quad ECT_{t-1} > \hat{\gamma}^H \end{cases} \quad (5),$$

where $\hat{\gamma} = (\hat{\gamma}^L, \hat{\gamma}^H)$ are the estimated thresholds that segment the different regimes. ECT_{t-1} = $(\Delta \ln p_{t-1}^{Fin} - \beta_1 \Delta \ln p_t^{Ger})$ for the bivariate TVECM of Finnish meat prices and the corresponding German meat prices and ECT_{t-1} = $(\Delta \ln p_{t-1}^{Fin} - \beta_2 \Delta \ln p_t^{Den})$ for the bivariate TVECM of Finnish meat prices and the corresponding Danish meat prices. Setting β_i = 1, ECT_{t-1} is the price spread between Finnish meat prices and German/Danish meat prices in logarithmic form. Since both thresholds are unknown, they need to be estimated along with the remaining parameters of the model. Combining the strategy proposed by Lo and Zivot (2001) and Hansen and Seo (2002), the thresholds could be estimated through a likelihood ratio (LR) programme, in which thresholds $\hat{\gamma} = (\hat{\gamma}^L, \hat{\gamma}^H)$ are first set up as a grid search to minimize the log determinant of the residual covariance matrix of the TVECM, which is analogous to maximizing the standard LR tests (Ben-Kaabia et al. 2005). Secondly, the covariance matrices of the VECM and TVECM with one threshold and with two thresholds are computed and compared as follows:

$$LR_{ij} = T(\ln(\det \hat{\Sigma}_i) - \ln(\det \hat{\Sigma}_j)) \quad (6),$$

where $\hat{\Sigma}_i$ and $\hat{\Sigma}_j$ are the residual covariance matrices of the VECM and TVECM with the *i*th regime numbers varying from 0 to 3. Thus, the first test would be a test of the linearity of the VECM against non-linearity. If the test is rejected we choose threshold vector error correction with either 1 or 2 thresholds. Consequently, analysis of the TVECM will follow the steps listed below:

1) Grid searching for the threshold based on research by Hansen and Seo (2002);
2) The identification of one or two thresholds and definition of two or three regimes, correspondingly;
3) Estimation of the TVECM.

3. Data and preliminary tests

3.1 Data

The data consist of two groups of price series: one comprises producer prices for pork and the other producer prices for beef extracted from Finland, Germany and Denmark. The Finnish and Danish data are from the Information Centre of the Ministry of Agriculture and Forestry (TIKE). The German data are from the German Centre for Documentation and Information in Agriculture (ZADI). This is the scientific information institute of the Federal Ministry of Consumer Protection, Food and Agriculture. The prices used in the study are the prices that are paid to the producer for one slaughtered kilogram of meat at the gate of the slaughterhouse, and transportation costs to the slaughterhouse are not therefore included. The prices are the average prices of the EUROP quality classes, which have been weighted according to the slaughter weight.

The data used in both groups are weekly and the periods covered in two groups are slightly different due to missing data. The pork price group is dated from the 10th week of 1995 to the 22nd week of 2009, and the period for the beef price group extends from the 5th week of 1995 to the 23rd week of 2009 (see Figures 4a and 4b). Both groups of data are stabilized by converting them to logs and are displayed in Table 1. They are labelled as lgpork_F, lgpork_G, lgpork_D, lgbeef_F, lgbeef_G and lgbeef_D, representing the logged producer prices for pork and beef in Finland (F), Germany (G) and Denmark (D). Table 1 also presents summary statistics for the logarithms of the two groups of price series. Table 1 indicates that the mean values of beef prices within a group vary little in comparison with the mean values of pork prices among the three selected countries. The standard deviations of the Finnish meat prices are the lowest among the countries, implying that the Finnish producer prices for meat are the most stable. All the series exhibit slight kurtosis.

3.2 Unit root test

To decide whether cointegration analysis is needed for the test, the time series properties of the price series are analysed using three unit root tests, including the traditional ADF and the alternative KPSS test (Kwiatkowski *et at.*, 1992). While the ADF states the null hypothesis of non-stationarity or the presence of a unit root, the KPSS test defines stationarity as the null. The Monte Carlo simulations by Schwert (1989) demonstrated that ADF tests have low power and are sensitive to the choice of lag length. The unit root tests are known to have low power problems in small samples, particularly if the series include structural breaks. However, the KPSS test is more robust in dealing with the problem of structural breaks. Thus, both unit root tests are applied to statistically determine the order of integration of the time series used in cointegration analysis. (Murthy and Nath, 2003).

The results of all the tests are presented in Table 2. ADF tests with or without a trend indicate the existence of a unit root for most of the price series of Finnish and Danish meat. Two exceptional series are apparent for German meat market, which was known to experience a structural break in 2000. The KPSS test yielded contradictory results, and proved to be a more robust test if structural breaks are encountered (Leybourne & Newbold 2000). Thus, it was concluded that there is strong evidence for nonstationarity in all the price series. For the first difference series, the results of all the unit root tests indicated that they are stationary, and the results are not reported here. Thus, all the selected meat series are integrated with order 1, designated as I(1).
(Place Table 2 here)

4. Empirical results

4.1 Cointegration analysis

Given that the data are nonstationary, the next step is to analyze the price transmission from the German/Danish producers to the Finnish meat producers with a cointegration framework. More specifically, we test for the presence of cointegration with and without a linear trend between meat prices, which includes tests between Finnish meat prices and German/Danish prices, in comparison to a test between German and Danish meat prices. The results of Johanson tests are listed in Table 3a-3b. Akaike’s information criterion was used to determine the optimal order of lags (3 lags for each series). The trace statistics indicate that for both the with-trend and without-trend specification, we can reject the null hypothesis of no cointegrating vector (r = 0) in favour of one cointegrating vector (r = 1) for all bivariate cointegrating tests on the group of pork prices. This result has two implications: one is that Finnish producer prices for pork are integrated to the European market process, which is represented here by the variables lgpork_G and lgpork_D. The other is that for the group of beef prices, the test between Finnish producer prices (lgbeef_F) and Danish producer prices (lgbeef_D) failed to reject the null hypothesis of no cointegrating relationship. This suggests that the Finnish and Danish producer prices for beef do not co-move. In comparison, lgbeef_F and lgbeef_G as well as lgbeef_D and lgbeef_G were found cointegrated, at least without trend. These results were in line with expectations, as Germany is the main beef exporter and importer of the EU, and of both Finland and Denmark. The import of beef from Germany to Finland has steadily grown during the last decades, except for the downturn during the BSE crisis between 2001 and 2002, while in comparison the import from Denmark to Finland has declined (see Figure 2).

Under a cointegration relationship, with equation (2) it is also possible to check whether the signs of coefficients are in line with the predictions of economic theory. This is carried out by analysis of the coefficients of the variables of the first cointegration equation normalized. In this study, for the bivariate series between the Finnish and German/Danish meat prices, the normalization is imposed on the Finnish meat price; for the bivariate series between German and Danish meat prices, the normalization is imposed on the Danish meat price.

Tables 4a and 4b present the coefficient estimates of long-term *ECT* for the tested bivariate VECM. Naturally, $\beta = 1$ for the variables on which normalization is imposed, while α represents the adjustment coefficients in the corresponding bivariate VECM.

1) Table 4a – pork prices. Firstly, all estimated values for the elasticity of producer prices from one market with respect to the other market, β, are correctly signed and statistically significant. For example, in the pairwise combination of lgpork_F and lgpork_G, the estimated value of the elasticity of price transmission into Finnish pork prices with respect to German pork, β_{Ger}, equals 0.69. This suggests that variations in the German market are not fully transmitted to the Finnish market, which is expected to be caused by high transaction costs in the Finnish market. By comparison, the law of one price holds very well in the pairwise combination of lgpork_D and lgpork_G, where β_{Ger} equals 0.98, having an elasticity of transmission of unity, in line with the prediction of economic theory. Secondly, all the signs for the adjustment coefficients, α, are correctly signed given that the deviations from the long-term equilibrium are obtained from the co-integrating vector normalized with respect to lgpork_Fin and lgpork_Den. However, the signs of α_{Ger} and α_{Fin} in the pairwise combination of lgpork_F and lgpork_G, and of lgpork_F and lgpork_D, respectively, are statistically nonsignificant. Hence, adjustment towards a long-term equilibrium only takes place through changes in the Finnish pork price (lgpork_F). For the pairwise combination of lgpork_D and lgpork_G, it seems that adjustment toward a long-term equilibrium is two directional. Bearing in mind the mixed results from the unit root tests, especially for the series lgpork_G, these results further support the validity of the co-integrating relationship in the equation, as at least one-way causality is found in the lagged ECT term (Granger, 1986). Finally, all the significant values of α are less than 6%, which suggests that the adjustment process is relatively slow. More specifically, between lgpork_F and lgpork_G, for example, α_{Fin} equals approximately 3%, which implies that after a shock, 3% of the departure from the long-term equilibrium will disappear each week. Notably, the Finnish producer price for pork adjusts at almost the same speed (3%) to the long-term equilibrium that is produced together with either the German or Danish producer price. By comparison, the Danish price eliminates the deviation at an approximately two-

fold higher speed of 6% in each period from the equilibrium that is produced together with the German price. There are various possible reasons for a slow adjustment in price transmission, including policies, the number of stages in marketing and the corresponding contractual arrangements between economic agents, storage and inventory holding. Unlike Denmark, Finland has a significant domestic market for pork meat, as most pork produced in Finland is domestically consumed, and self sufficiency in the pork sector in 2009 was reported to be 112% (Statistics Finland, 2010). As the domestic market is of a significant size, one should expect that any shock deviation from equilibrium that may come from the European market would take more time to fad away in the Finnish market when compared to the Danish market, which is one of the major exporters for pork meat in Europe.

2) Table 4b – beef prices. This table contains only two pairwise combinations, because the Finnish and Danish producer prices for beef were found to have no cointegration relationship, and this result is therefore not presented in Table 4b. Firstly, both values of β have negative signs and are statistically significant, but their magnitudes are different. In particular, when the price in the German producer's market increases by 1%, the Danish market grows by 0.86%, which implies that its elasticity is quite close to 1. In comparison, the value of β between lgbeef_F and lgbeef_G amounts to 1.63, indicating that information is transmitted with significant distortions between lgbeef_F and lgbeef_G. This might be caused by structural changes or breaks during the estimated period, and the linearity of the VECM might not serve the data very well, which needs further testing. Table 4b demonstrates that when restrictions on the long-term β parameters are imposed, short-term deviation from the equilibrium presented by α is eliminated at a speed of less that 1% in each period between lgbeef_F and lgbeef_G, as compared to 8% for the producer prices between lgbeef_D and lgbeef_G. Apparently, such results were able to detect the characteristics of each market. Compared to the Danish market, the Finnish producer's market in the beef sector is more segmented and geographically more distant from Germany, and shocks occurring in Germany take much longer to reach to Finland compared to Denmark. Another reason for the very different speed of adjustment is that Germany is a more important trading partner for Denmark in the beef sector in comparison to Finland. For example, in 2009, Denmark imported 7 million tons of beef products from Germany, almost 3 times more than Finnish imports from this country, and Germany accounted for one third of Danish imports from European countries[3]. Meanwhile, Germany

[3] Referring to the EU15 countries.

imported approximately 8 million tons of beef products from Denmark, which is twice as much as the imports from Finland.

4.2 Weak exogeneity and Granger causality test

4.2.1 Weak exogeneity test for cointegrated bivariate VECMs.

A series is regarded as weakly exogenous if it leads other series in the long term without being influenced by other series (Carter and Mohapatra, 2008). A weakly exogenous series can therefore be useful in explaining variations in the ‘nonexogenous’ series (Leatham 2001). Tables 5a and 5b present the results of the weak exogeneity test for the bivariate VECMs. First, for group of pork prices, uniformly, the null hypothesis that lgpork_G is weakly exogenous for the long-term equilibrium relationship with both lgpork_F and lgpork_D is not rejected at the 5% significance level. This indicates that the German producer price for pork is the leader for the pork group, i.e. it is not affected by short-term interruptions in the equilibrium. It is also worthwhile noting that price variations originating in the Danish producer’s market have a much stronger impact on the German than the Finnish producer’s market. Not surprisingly, the Finnish producer price was found to be the price taker in both the German and Danish markets, and it adjusted itself to restore market equilibrium once shocks had taken place. Second, for the group of beef prices, the German price is still the leader of the equilibrium relationship, regardless of which partner the equilibrium is built up with. The hypothesis that lgbeef_F is weakly exogenous with respect to lgbeef_G is rejected at the 5% significance level. In comparison, weak exogeneity of lgbeef_G with respect to lgbeef_F cannot be rejected at the 5% significance level, but interestingly it can be rejected if the significance level is extended to 10%. However, the P-value of the test for the weak exogeneity of lgbeef_G with respect to lgbeef_D is much higher (0.34) compared to the one for lgbeef_G with respect to lgbeef_F (0.07). This indicates that the price variation originating from German producers affects Danish producers more than those in Finland.

4.2.2 Granger causality test with non-cointegrated bivariate series

As no cointegrating relationship could be found between lgbeef_F and lgbeef_D, the relationship between lgbeef_F and lgbeef_D is displayed by causality testing. Table 6 reports the results of the bivariate causality test and a summary of the causality result based upon the noncointegrated data. Given the lack of cointegration, the tests must be undertaken on I(0), i.e. first-differenced data only. The results presented in Table 6 suggest that Granger causality between Δlgbeef_F and Δlgbeef_D is not statistically significant at the 5% significance level in either direction. However, if the significance level is extended to 10%, the Δlgbeef_D is found to causally lead Δlgbeef_D. This result, together with the non-cointegration relationship between lgbeef_F and lgbeef_D, suggests that the producer price for beef in Finland and that in Denmark behave like driftless random walks.

4.3 Linearity test of the VECM and estimated coefficients of the VECM and TVECM

Hansen and Johanson (1999) provide recursive statistics for the stability of the eigenvalue produced from the linear VECM. The purpose of the cointegration vector stability test is to examine whether significant structural changes occur during the test period. Rejection of the hypothesis of stability of the recursively estimated eigenvalues would indicate that the error correction vector is not stable over time. If this is the case, the validity of the linear VECM may be questionable. Figures 5a to 5e illustrate the recursive τ statistics of the linear VECM model presented in equation (2). Clearly, all the τ test statistics are less than the critical value for the 1% significance level, which is 1.6 over the entire sample period, implying that the parameters estimated from equation (2) are stable. However, if the significance level is extended to 5%, with a corresponding critical value of 1.4, there is a peak in the τ statistics of Figure 5d that exceeds the critical value. This implies that a linear VECM might not be the most stable model for the bivariate function of lgbeef_F and lgbeef_G.

As non-linearities possibly exist in the adjustment process, a test of the null of linearity against the alternative of a TVECM is next performed. Meanwhile, the number of regimes for systems can be determined by applying equation (6), the likelihood ratio (LR) test provided by Hansen (1999) and Lo and Zivoc (2001).

The test results are presented in Tables 7a and 7b. The asymptotic distributions of LR_{23} are non-standard and bootstrap P values and critical values are calculated by a method used by Hansen and Seo (2002) and Lo and Zivot (2001). Clearly, for the group of pork prices, all the tests suggest that the linear VECM is preferred and thus no further TVECM analysis is necessary. However, the hypothesis of linearity for the bivariate lgbeef_F and lgbeef_G suggests one significant threshold, which is consistent with the previous results of the stability test illustrated in Figure 5d. Since only one threshold is found significant, the TVECM could be simply done in the following function:

$$\Delta \ln beef\_F = \begin{cases} c^L + \alpha^L ECT_{t-1} + \sum_{i=1}^{k-1} \delta_i^L \Delta \ln beef\_G_{t-k} + \sum_{i-1}^{m-1} \varphi_i^L \Delta \ln beef\_F_{t-m} + \varepsilon_t^L, & if \quad ECT_{t-1} < \hat{\gamma}^L \\ c^H + \alpha^H ECT_{t-1} + \sum_{i-1}^{k-1} \delta_i^H \Delta \ln beef\_G_{t-k} + \sum_{i=1}^{m} \varphi_i^H \Delta \ln beef\_F_{t-m} + \varepsilon_t^H, & if \quad ECT_{t-1} > \hat{\gamma}^H \end{cases} \quad (7),$$

The estimated parameters of the VECM and TVECM for the two groups of meat prices are presented in Tables 7a and 7b, respectively.

1) For pork prices, all the estimated ECT terms are significant and consistent with the results presented in Table 8a. This confirms that the Finnish producer price is slowly cointegrated towards a long-term equilibrium with both German and Danish producer prices for pork. In the short term, however, the situation is different: the Finnish producer price for pork reacts more spontaneously to shocks coming from the domestic market. A shock to the German producer price does not generate any spontaneous response in the Finnish producer price, or does not share a common reaction time, while conversely, the Danish producer price reacts immediately to variation in the German producer price, indicating that the Danish producer market is more sensitive to changes taking place in central European, as represented by Germany. Interestingly, shocks originating from Denmark were found to positively and significantly affect the Finnish producer's market. However, the magnitude of the effect was smaller than that originating from the domestic market. Taken together, this suggests that in the short term the Finnish producer price reacts quickly and spontaneously to shocks coming from the domestic market. In comparison, the Danish producer price reacts more rapidly and significantly to shocks coming from central Europe, such as Germany
2) For beef prices, Table 8b first reports a summary of the estimated parameter of the bivariate TVECM of lgbeef_F and lgbeef_G with one detected threshold, 0.176. Thus, only two regimes are included in the test. Apparently, the parameter estimate for β in the TVECM appears to be quite close to a unit coefficient, compared to -1.63 in the VECM, which indicates that the law of one price holds relatively well when the asymmetry of the price transmission is accounted for.

More specifically, the first regime occurs when $ECT_{t-1} < 0.176$, namely the normal regime, and the second regime occurs when $ECT_{t-1} > 0.176$, namely the extreme regime, i.e. when the Finnish price for beef is at least 19% higher than the German price[4]. In the extreme regime the series consisted of 56 observations, which covered the whole of 2001, accounting for 7.4% of the total observations. In November 2000, Germany reported the discovery of domestic cases of bovine spongiform encephalopathy (BSE). In the following year, 2001, there were estimated to be 500 cases of BSE in Germany, and sales of beef products dropped by 50% because of public fears of mad cow disease or BSE. Correspondingly, the producer price for beef dropped to a historically low level. Both exports and imports of beef products suffered from large losses. The result suggests that in the extreme regime, the Finnish producer price for beef has minimal error-correction effects but quite a large effect resulting from short-term German dynamics. This indicates that the Finnish producer price for beef did not adjust itself with respect to the German producer price into the long term equilibrium, but the dramatic drop in the German producer price strongly and negatively affected the Finnish producer price in the short term. By comparison, in the normal regime, the Finnish producer price for beef cointegrates slowly towards a long-term equilibrium with the German producer price. Meanwhile, the Finnish producer price is minimally affected by the short-term dynamics of the producer price in Germany. However, the Finnish domestic dynamics are dominant in the short term under the normal regime.

Finally, the adjustment speed of the Danish producer price for beef towards a long-term equilibrium with respect to the German producer price is about 8%, which is 5 times faster than the speed of adjustment between Finnish and German prices. Together with the lack of cointegration between Finnish and Danish producer prices, all the results reflect the fact that, besides being a remote and small trader in the EU, Finland has a dominant domestic market for producers in beef sector.

5. Conclusions

We examined the price cointegration relationship between the Finnish pork and beef markets and those in Germany and Denmark using both a bivariate symmetric error correction model and bivariate asymmetric threshold error correction model, which recognizes the non-stationary nature of the price data and allows for asymmetric price responses. Symmetric models were able to fit

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4

$$\ln beef\_F - 1.06 \ln beef\_G > 0.176 \Rightarrow beef\_F > beef\_G^{1.06} \times e^{0.176} \Rightarrow beef\_F > 1.19 \times beef\_G^{1.06}$$

most bivariate price series, except for the bivariate series between Finnish and German beef prices, for which one threshold was identified, and thus a two-regime threshold TVECM was applied.

A cointegrating relationship was found for most of bivariate price series, except for the Finnish and Danish producer prices for beef, and further causality testing confirmed that the Finnish and Danish beef prices move as a driftless random walk. In the both symmetric and asymmetric vector error correction models, we found that the LOP held relatively well in the Finnish producer's meat market compared with those in Germany and Denmark. However, the speed of adjustment towards long-term equilibrium was found to be slower compared to the speed of the bivariate price series of Germany and Denmark. This seems to be consistent with the different trading activities among the countries, i.e. trade between Finland and German is not as active as that between Denmark and Germany in the pork and beef markets. Another possible reason is that the meat sector in Finland is still very much dominated by domestic consumption, and the high self-sufficiency indicates that domestic price shocks are still the dominant price changes in Finnish meat price dynamics, at least at the producer's level.

However, there is a very interesting and important phenomenon in the asymmetric price case, i.e. the bivariate price series between Finnish and German beef prices. The estimated model identified one threshold and two regimes. Error correction appears to only occur in the typical regime, but not in the extreme regime, which covered the BSE period. In the short term, the dominating effect in the typical regime came from the domestic market, but in the extreme regime, the dominating effect came from the German market. This suggests that the Finnish domestic market has a dominant influence on the beef producer price most of the time, but is still highly vulnerable to the short-term effects of a large negative shock in the German market.

This study has very important economic implications at three levels. First, better and statistically tested knowledge on the transmission of price information can be used to justify domestic agricultural policies and infer whether the law of one price holds at the domestic producer's level compared to the European market. Secondly, the result concerning asymmetric prices is beneficial in assessing the efficiency and competitiveness of the Finnish meat market. Thirdly, better knowledge of the regime structure and volatility processes for pork and beef prices and the sources of this volatility will be of interest to farmers and extension agents needing to make and advise on investment decisions during the ongoing and very rapid structural adjustment in Finnish agriculture.

The relatively slow and sluggish response of Finnish domestic prices to price shocks in foreign markets supports the view that the Finnish meat chain, which is a combination of co-operative processors and publicly quoted companies, can smooth out some of the short-term price fluctuations and high price volatility observed abroad. Another reason for the sluggish price movements may lie in the structure of the delivery and pricing contracts between the meat processors and meat purchasing groups at the wholesale level. The economic performance and efficiency of these contracts cannot explicitly be examined using reduced form price models, and this topic is therefore left for future research.

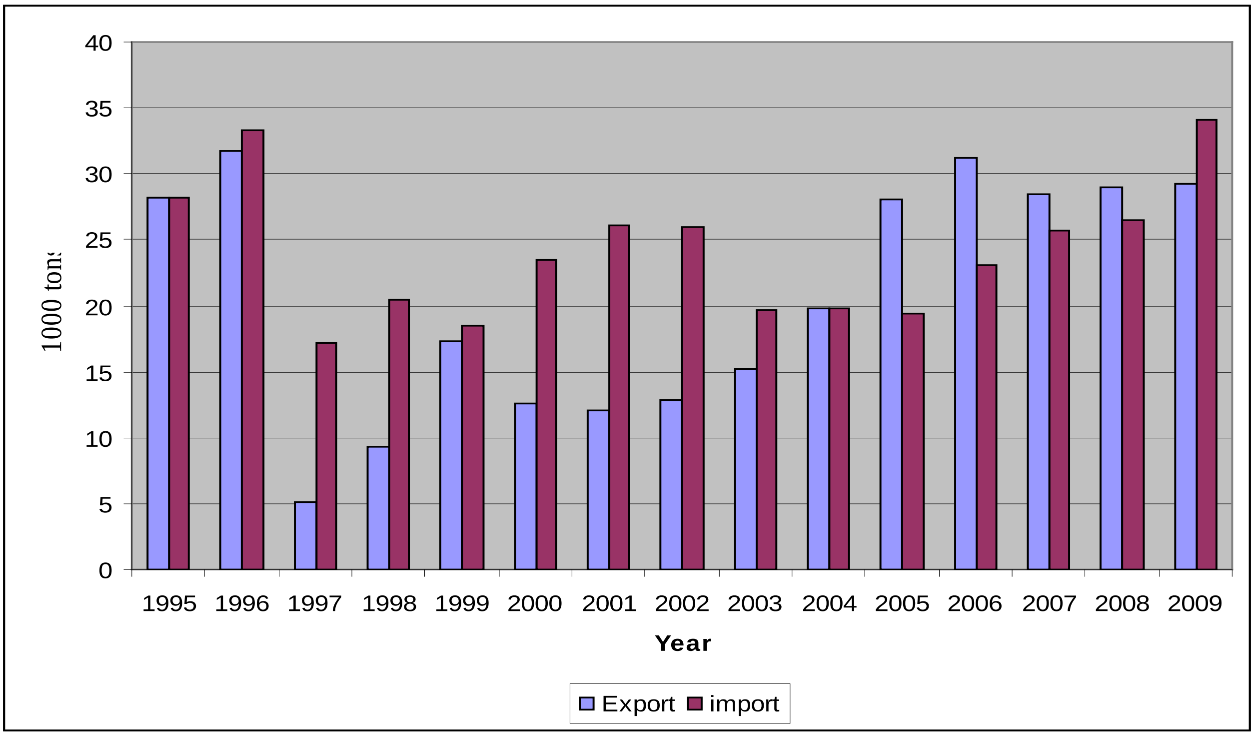

Figure 1. Trading volume of Finnish meat and offal (by HS6 G_02) from 1995 to 2009.

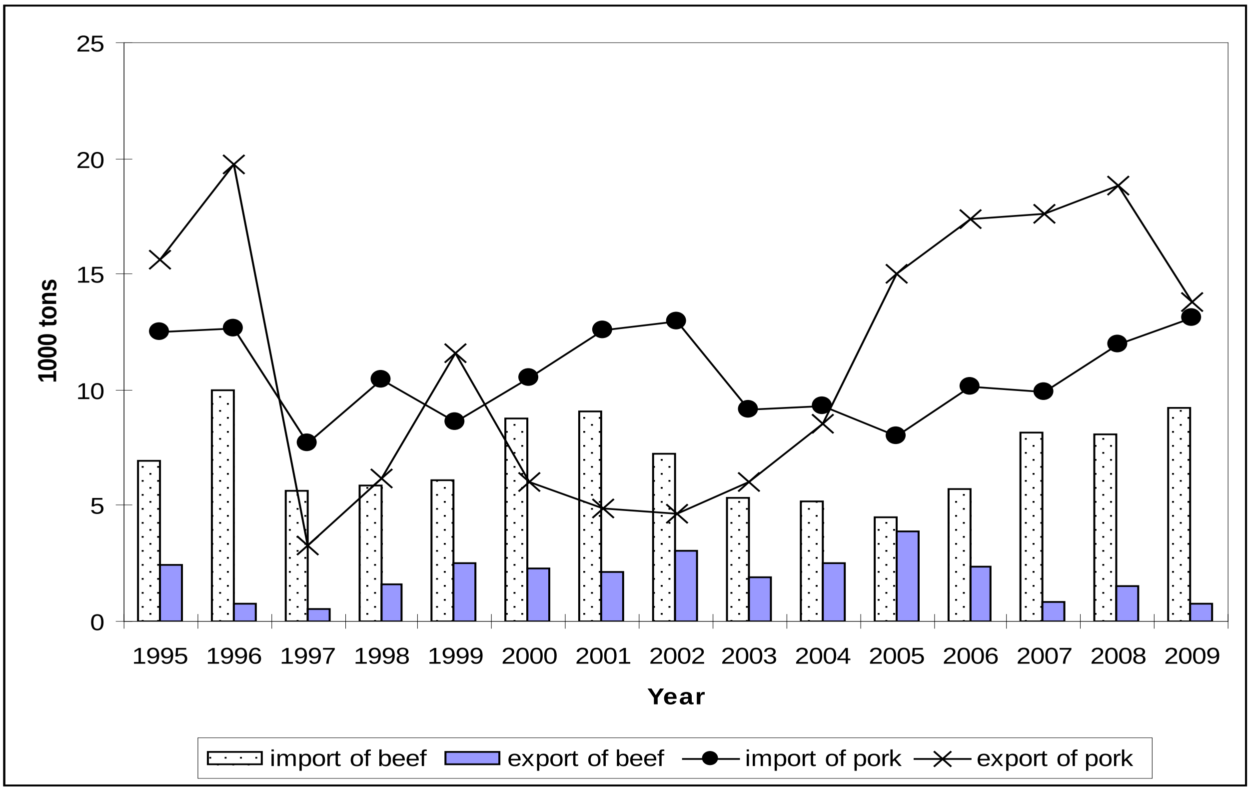

Figure 2. Trading volume of Finnish beef (by HS6 G_020110-020230) and pork (by HS6 G_020311-020329) from 1995 to 2009.

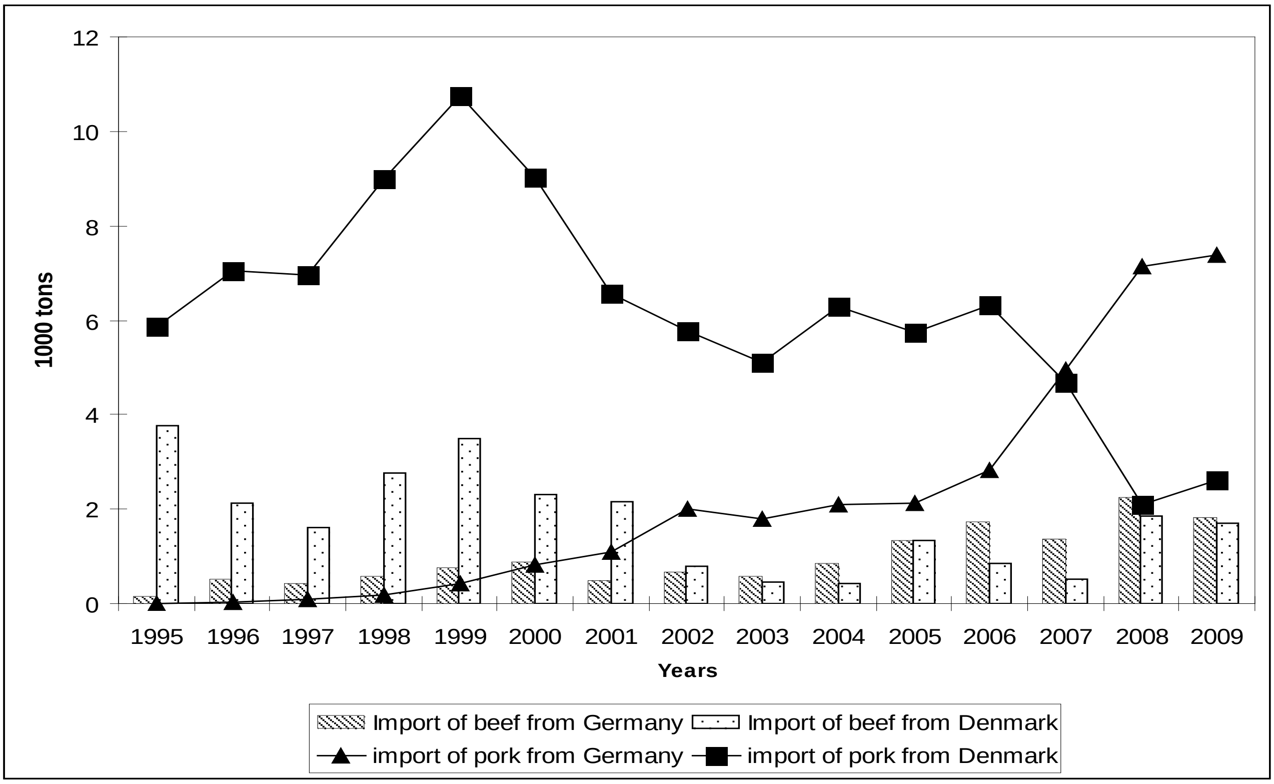

Figure 3. Import of beef (by HS6 G_020110-020230) and pork (by HS6 G_020311-020329) from Germany/Denmark to Finland between 1995 and 2009.

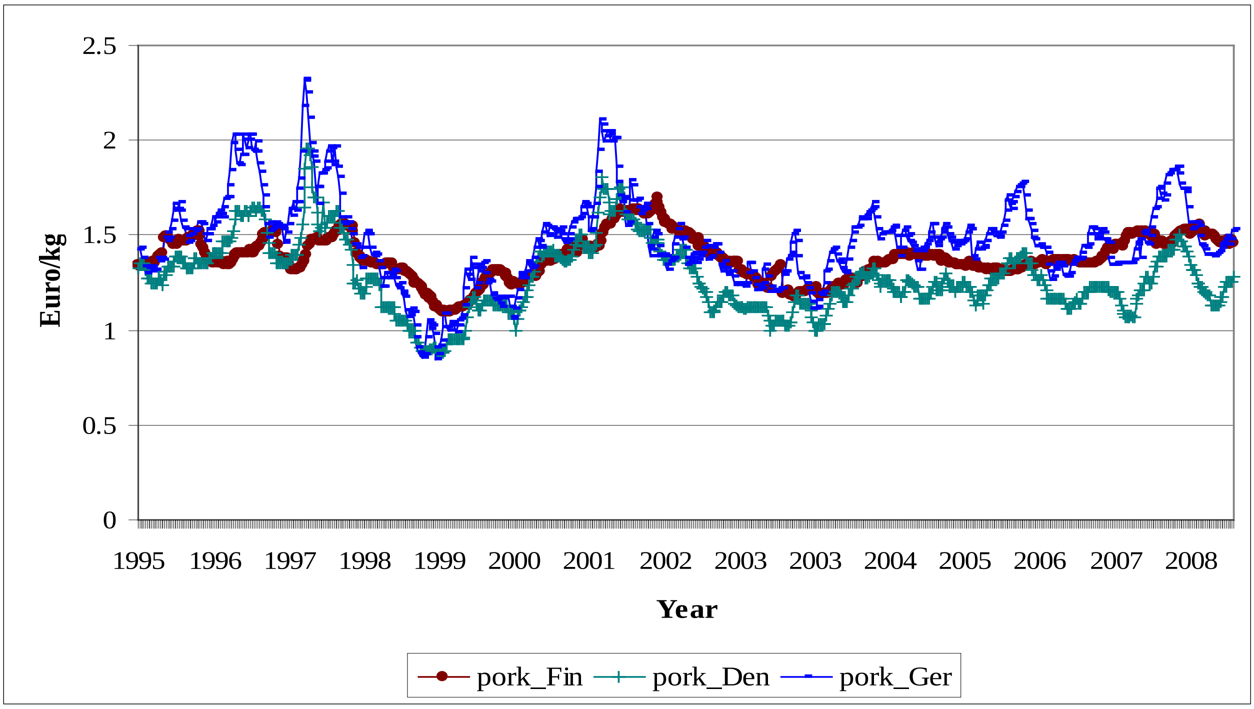

Figure 4a. Producer prices for Finnish, German and Danish pork in 1995-2009 (Euros/kg).

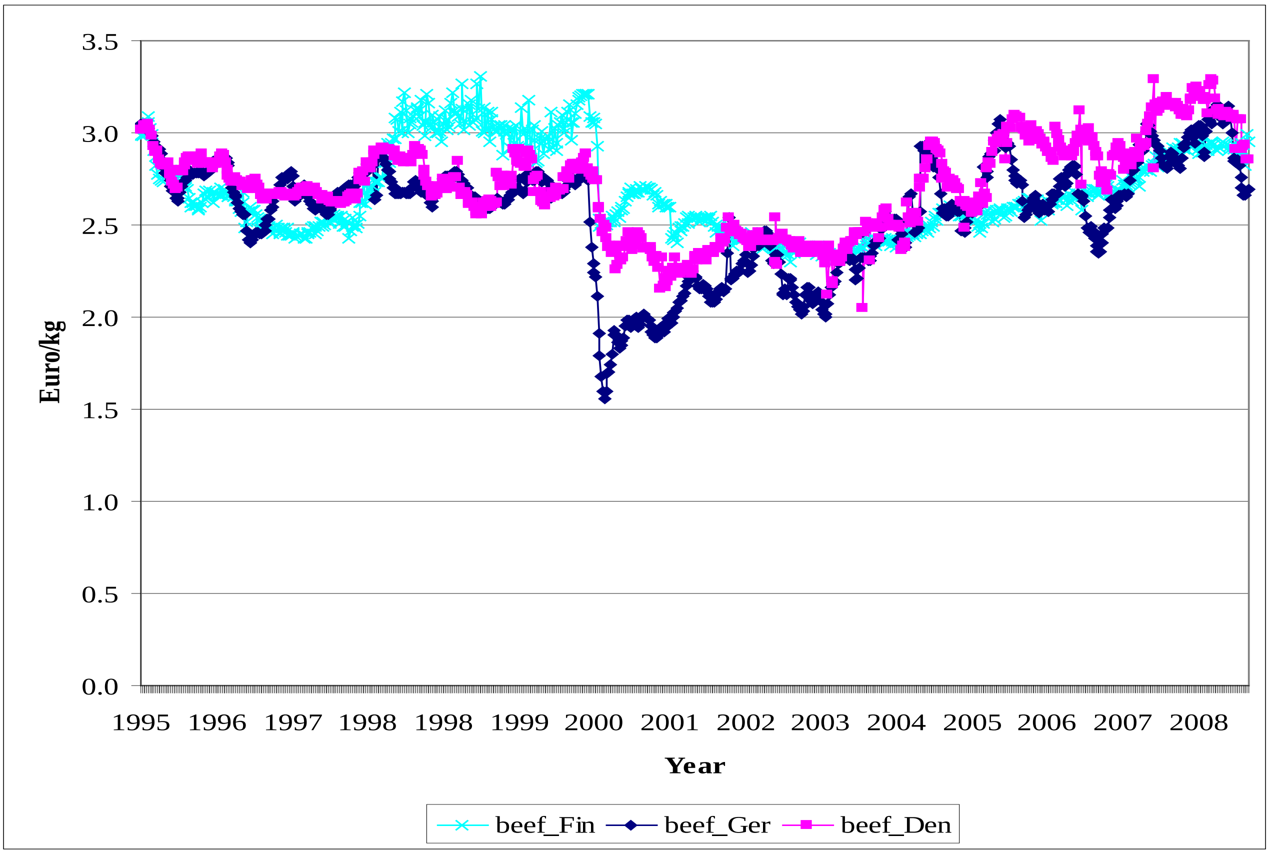

Figure 4b. Producer prices for Finnish, German and Danish beef in 1995-2009 (Euros/kg). [NOTE: on the vertical axis the decimal symbols need to be converted from commas (,) to points(.)]

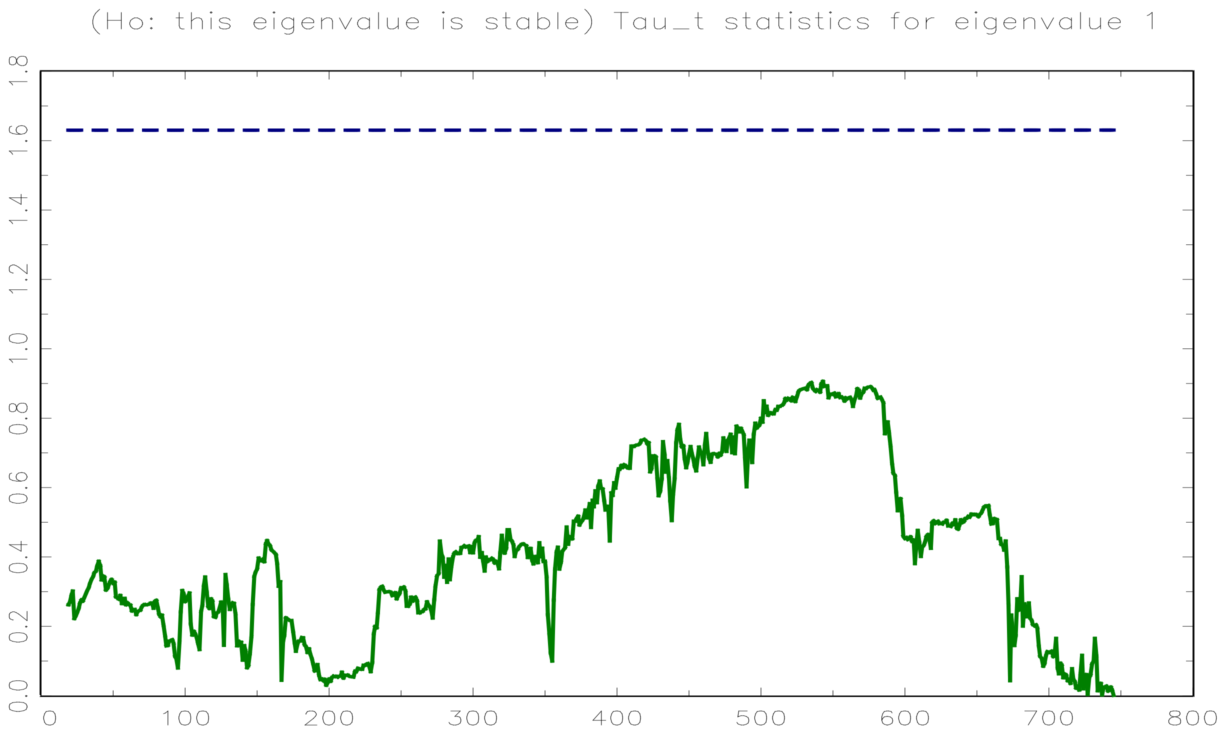

Figure 5a. Recursive eigenvalue τ statistics for lgpork_F and lgpork_G.

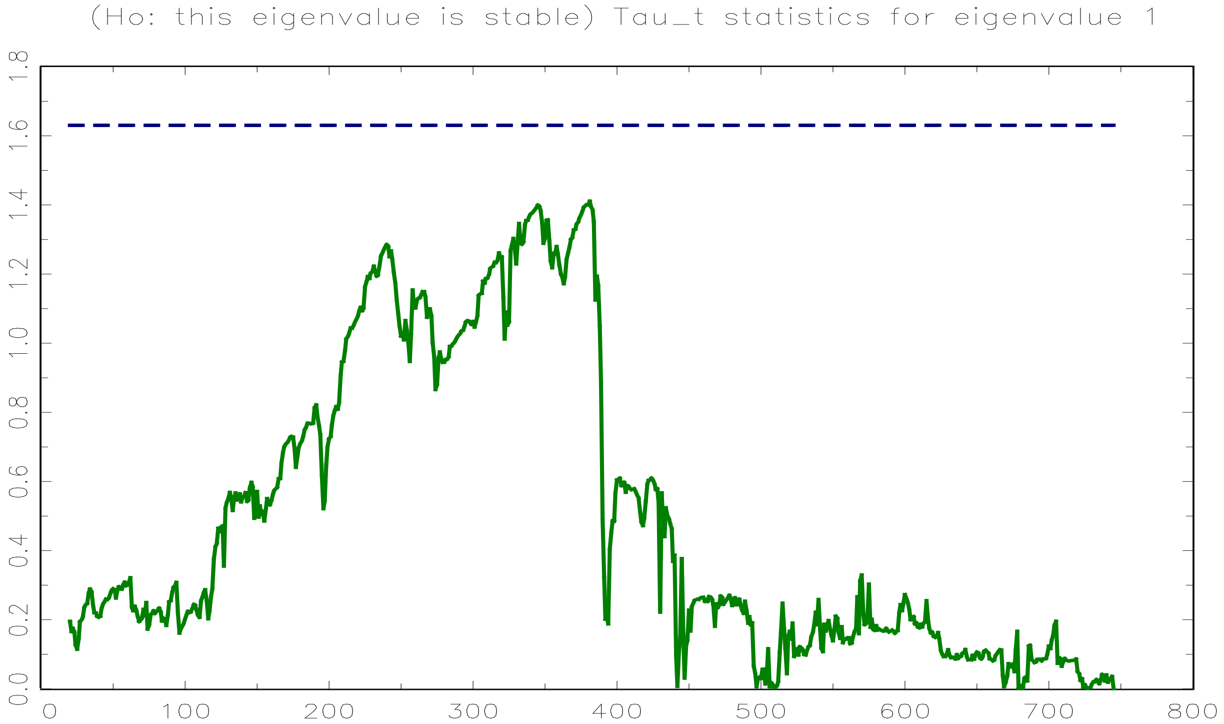

Figure 5b. Recursive eigenvalue τ statistics for lgpork_F and lgpork_D.

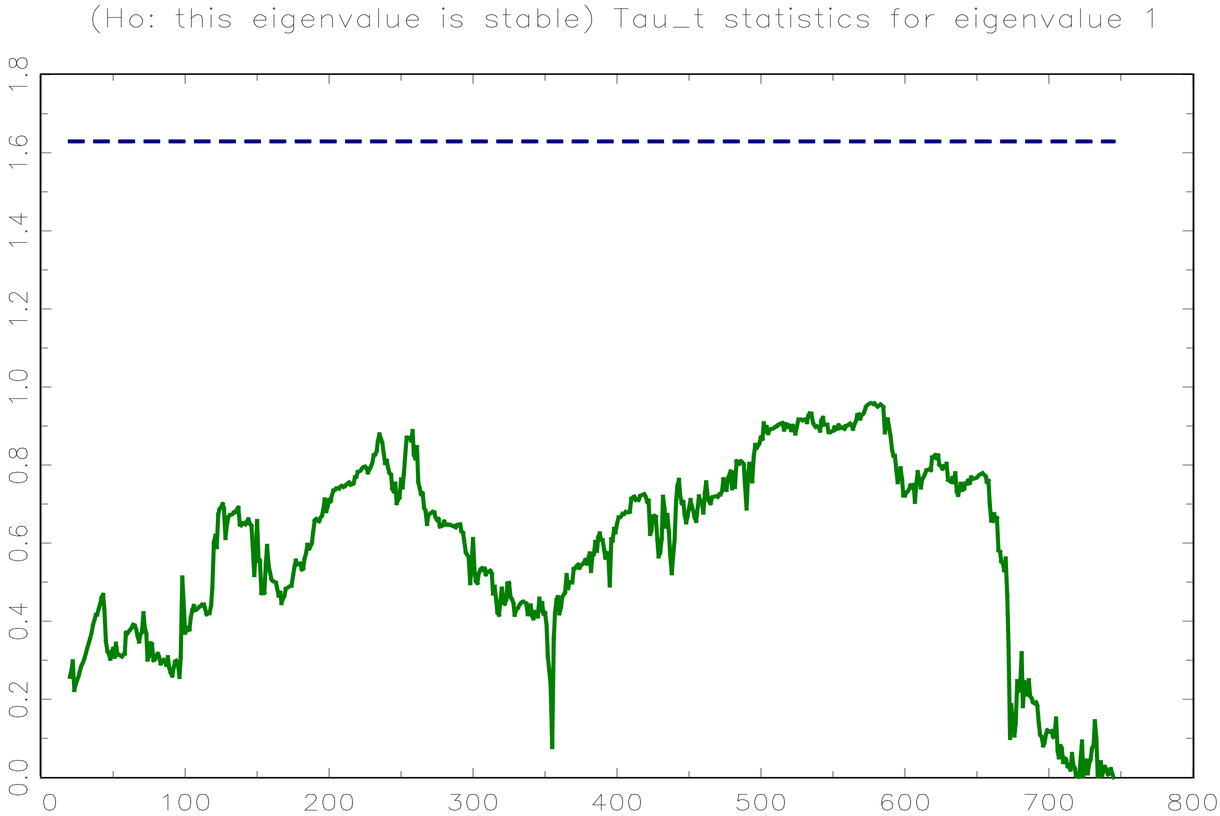

Figure 5c. Recursive eigenvalue τ statistics for lgpork_D and lgpork_G.

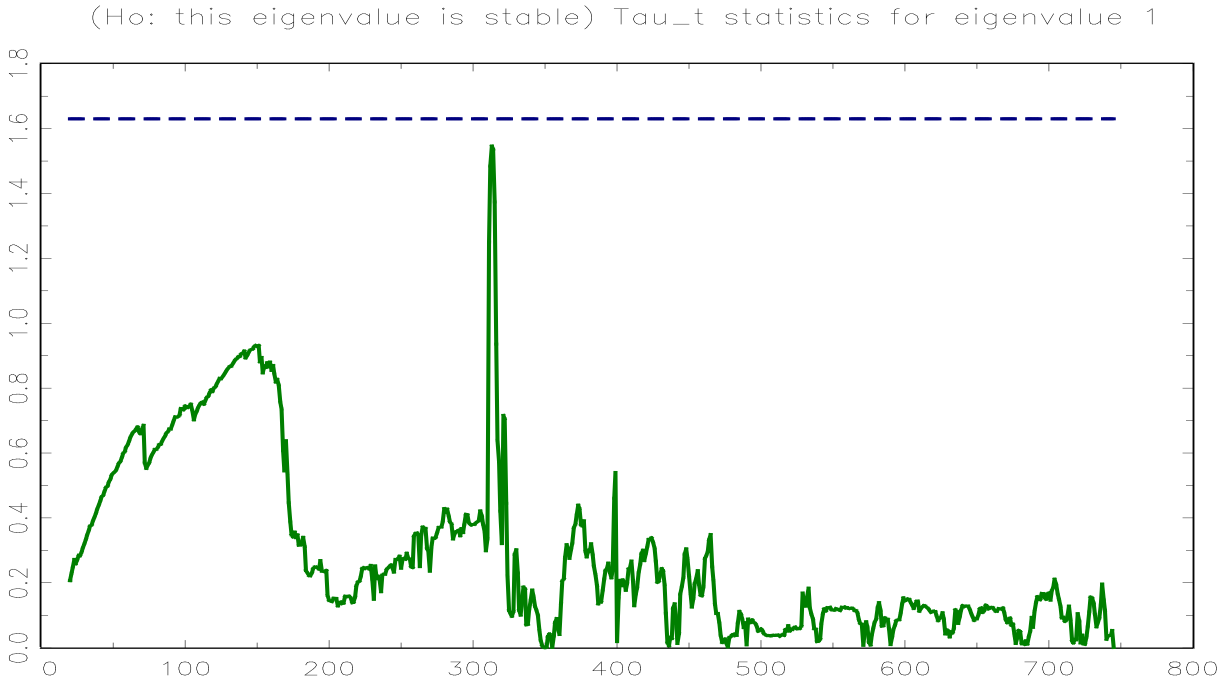

Figure 5d. Recursive eigenvalue τ statistics for lgbeef_F and lgbeef_G.

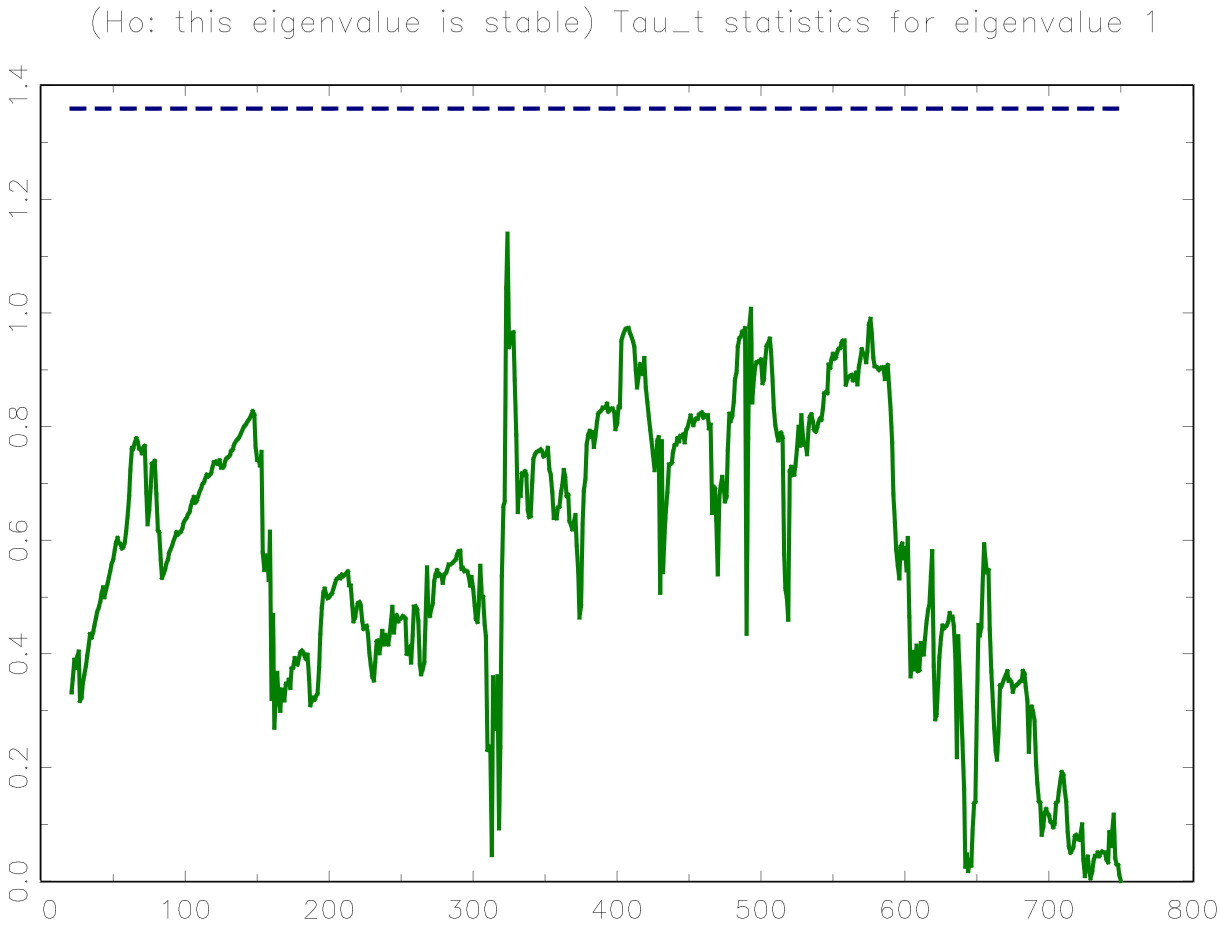

Figure 5e. Recursive eigenvalue τ statistics for lgbeef_D and lgbeef_G.

Table 1. Descriptive statistics for the price series

	Group of pork prices [a)]			Group of beef prices [b)]		
	lgpork_F	lgpork_G	lgpork_D	lgbeef_F	lgbeef_G	lgbeef_D
Mean	0.32	0.37	0.23	0.98	0.94	0.99
Median	0.31	0.38	0.22	0.96	0.97	1.00
Maximum	0.53	0.84	0.67	1.20	1.15	1.19
Minimum	0.09	-0.17	-0.13	0.83	0.44	0.72
Standard Deviation	0.09	0.16	0.15	0.09	0.13	0.094
Skewness	-0.29	-0.30	0.07	0.42	-1.05	-0.24
Kurtosis	2.99	4.04	3.06	2.05	3.70	2.33
J-B Normality	10.1	44.5	0.71	5.77	12.01	6.56
Observations	744	744	744	749	749	749

Note: [a)] The sample period for the series of pork prices extends from the 10th week of 1995 to the 22nd week of 2009.

[b)] The sample period for the series of beef prices extends from the 5th week of 1995 to the 23rd week of 2009.

Table 2. Unit root test results

	Intercept included	Intercept and linear time trend included
Price Series	Test statistics for ADF *t* test[1)]	Test statistics for ADF *t* test
lgpork_F	-2.03 (lag 2)	-2.67 (lag 8)
lgpork_G	-3.07* (lag 2)	-3.20 (lag 3)
lgpork_D	-2.63 (lag 1)	-2.84 (lag 2)
lgbeef_F	-1.62 (lag 15)	-1.53 (lag 15)
lgbeef_G	-2.92* (lag 6)	-2.95(lag 6)
lgbeef_D	-2.15 (lag5)	-2.38(lag5)
Critical value at 5%	-2.86	-3.42
	Test statistics for KPSS LM test[2)]	Test statistics for KPSS LM test
lgpork_F	1.10*	0.25*
lgpork_G	0.94*	0.71*
lgpork_D	2.85*	0.72*
lgbeef_F	0.59*	0.46*
lgbeef_G	1.55*	1.55*
lgbeef_D	2.51*	2.51*
Critical value at 5%	0.46	0.15

Note: Optimal lag lengths were determined by Akaike's Information Criterion. [1)] ADF test hypothesis H_0: The series has a unit root and the critical value follows Davidson and Mackinnon (1993). [2)] KPSS hypothesis H_0: The series is stationary and the critical value follows Kwiatkowski et al. (1992). [3])

Table 3a. Bivariate cointegration test for prices of pork groups

Tested groups		H_0 : rank ($\alpha\beta'$) = r	Trace test statistics	5% Critical value
lgpork_F and lgpork_G	Model without trend	r=0	51.50*	20.26
		r=1	7.00	9.16
	Model with trend	r=0	52.07*	25.87
		r=1	6.97	12.51
lgpork_F and lgpork_D	Model without trend	r=0	43.70*	20.26
		r=1	8.41	9.16
	Model with trend	r=0	51.63*	25.87
		r=1	8.48	12.52
lgpork_G and lgpork_D	Model without trend	r=0	46.13*	20.26
		r=1	7.95	9.16
	Model with trend	r=0	56.67*	25.87
		r=1	8.18	12.52

Table 3b. Bivariate cointegration test for prices of beef groups

Tested groups		H_0 : rank ($\alpha\beta'$) = r	Trace test statistics	5% Critical value
lgbeef_F and lgbeef_G	Model without trend	r=0	20.64*	20.26
		r=1	6.44	9.16
	Model with trend	r=0	21.36	25.87
		r=1	7.11	12.51
lgbeef_F and lgbeef_D	Model without trend	r=0	12.37	20.26
		r=1	4.09	9.16
	Model with trend	r=0	13.99	25.87
		r=1	5.72	12.52
lgbeef_G and lgbeef_D	Model without trend	r=0	41.96*	20.26
		r=1	7.13	9.16
	Model with trend	r=0	51.18*	25.87
		r=1	8.43	12.52

Note: Critical values are from MacKinnon-Haug-Michelis (1999). (*) indicates a rejected null hypothesis.

Table 4a. Estimates of ECM coefficients of the linear VECM for prices of pork groups corresponding to Equation (2)

Tested bivariate series	Estimates of loading factor α	Restrictions on cointegrating vector β
lgpork_F and lgpork_G (lag2)	α_{Fin} =-0.028 (0.004)*	$\beta_{Fin} = 1$
	α_{Ger} =0.013 (0.01)	β_{Ger} = -0.69 (0.08)*
lgpork_F and lgpork_D (lag2)	α_{Fin} =-0.029 (0.004)*	$\beta_{Fin} = 1$
	α_{Den} =0.008 (0.009)	β_{Den} = -0.73 (0.09)*
lgpork_D and lgpork_G (lag3)	α_{Den} =-0.052 (0.009)*	$\beta_{Den} = 1$
	α_{Ger} =0.023 (0.014)*	β_{Ger} = -0.98 (0.065)*

Table 4b. Estimates of ECM coefficients of the linear VECM for prices of beef groups corresponding to Equation (2)

Tested bivariate series	Estimates of loading factor α	Restrictions on cointegrating vector β
lgbeef_F and lgbeef_G (lag 2)	$\alpha_{Fin} =$ -0.0175 (0.0054)**	$\beta_{Fin} = 1$
	$\alpha_{Ger} =$ 0.0083 (0.003)*	β_{Ger} = -1.63 (0.38)*
lgbeef_D and lgbeef_G (lag 3)	$\alpha_{Den} =$ -0.077 (0.013)*	$\beta_{Den} = 1$
	$\alpha_{Ger} =$ 0.012 (0.011)	β_{Ger} = -0.82 (0.07)*

Note: Standard errors for parameters are shown in parentheses in Tables 4a and 4b. An asterisk (*) denotes variables significant at 5%.

Table 5a. Test for long-term Granger causality for the group of pork prices

Tested bivariate series	Hypotheses	LR test statistics	P-value
lgpork_F and lgpork_G	$\alpha_{Fin} = 0$	36.5	0.00
	$\alpha_{Ger} = 0$	1.39	0.24
lgpork_F and lgpork_D	$\alpha_{Fin} = 0$	26.11	0.00
	$\alpha_{Ger} = 0$	0.51	0.47
lgpork_D and lgpork_G	$\alpha_{Den} = 0$	22.25	0.00
	$\alpha_{Ger} = 0$	3.51	0.06

Table 5b. Test for long-term Granger causality for the group of beef prices

Tested bivariate series	Hypotheses	LR test statistics	P-value
lgbeef_F and lgbeef_G	$\alpha_{Fin} = 0$	3.97	0.05
	$\alpha_{Ger} = 0$	3.35	0.07
lgbeef_D and lgbeef_G	$\alpha_{Den} = 0$	25.61	0.00
	$\alpha_{Ger} = 0$	0.90	0.34

Table 6. Test of bivariate causality for non-cointegrated lgbeef_F and lgbeef_D

Hypotheses	F-statistic	P-value
H_0: Δlgbeef_F does not Granger-cause Δlgbeef_D	1.62	0.19
H_0: Δlgbeef_D does not Granger-cause Δlgbeef_F	2.37	0.09

Table 7a. Tests for non-linearities in price adjustment in the bivariate VECM for the group of pork prices

Tested bivariate series	lgpork_F and lgpork_G	lgpork_F and lgpork_D	lgpork_D and lgpork_G
LM test statistics	LR_{13} = 16.31	LR_{13} = 16.07	LR_{13} = 24.03
P-value	0.84	0.76	0.30

Note: The LR_{13} tests the null of linear cointegration against the alternative of threshold cointegration following Hansen and Seo (2002)

Table 7b. Tests for non-linearities in price adjustment in the bivariate VECM for the group of beef prices

Tested bivariate series	lgbeef_F and lgbeef_G	lgbeef_D and lgpork_G
Test statistics	LR_{13}=31.05 LR_{23}=20.94	LR_{13}=25.94
P-value	P_{13}=0.05 P_{23}=0.15	P_{13}=0.34

Note: The tests are implemented in R statistics. The LR_{13} tests the null hypothesis of linear cointegration against the alternative of threshold cointegration following Hansen and Seo (2002), and LR_{23} tests the null hypothesis of a two-regime TVECM against the alternative of a three-regime TVECM (Lo and Zivot, 2001)

Table 8a. Estimated parameters of the linear VECM normalized on one endogenous variable for the group of pork prices

Bivariate VECM of lgpork_F and lgpork_G normalized on lgpork_F		
	Coefficient	t-statistic [p-value]
ECT_{t-1}	-0.028	-6.22[0.00]
Δlgpork_F(-1)	0.091	1.99[0.05]
Δlgpork_F(-2)	0.087	2.19[0.03]
Δlgpork_G(-1)	-0.013	-1.11[0.26]
Δlgpork_G(-2)	-0.006	-0.53[0.59]
R-square	0.09	
Durbin-Watson stat	2.01	
B-G Serial correlation LM Test	3.49 [0.47]	
ARCH(1)	0.87[0.49]	
Bivariate VECM of lgpork_F and lgpork_D normalized on lgpork_F		
ECT_{t-1}	-0.029	-5.70[0.00]
Δlgpork_F(-1)	0.083	2.30[0.02]
Δlgpork_F(-2)	0.084	2.34[0.02]
Δlgpork_D(-1)	0.037	1.96[0.05]
Δlgpork_D(-2)	0.021	1.08[0.38]
R-square	0.10	
Durbin-Watson stat	2.01	
B-G Serial correlation LM Test	1.15[0.57]	
ARCH(1)	1.07[0.37]	
Bivariate lgpork_D and lgpork_G normalized on lgpork_D		
ECT_{t-1}	-0.052	-5.46[0.00]
Δlgpork_D(-1)	0.115	1.98[0.05]
Δlgpork_D(-2)	0.023	0.65[0.51]
Δlgpork_D(-3)	0.072	2.19[0.02]
Δlgpork_G(-1)	0.152	5.77[0.00]
Δlgpork_G(-2)	0.165	4.89[0.00]
Δlgpork_G(-3)	0.029	0.99[]0.32]
R-square	0.35	
Durbin-Watson stat	2.00	
B-G Serial correlation LM Test	0.04[0.95]	
ARCH(1)	0.75[0.63]	

Table 8b. Estimated parameters of the TVECM and VECM normalized on one endogenous variable for the group of beef prices

Bivariate TVEM of lgbeef_F and lgbeef_G normalized on lgbeef_F (refer to equation 7)		
cointegrating vector (lgbeef_F, lgbeef_G)	(1, -1.062)	
Threshold	0.176	
	Coefficient	standard error
Typical Regime when ECT_{t-1}<0.176		
ECT_{t-1}	-0.0142	0.004*
Δlgbeef_F(-1)	-0.525	0.000***
Δlgbeef_F(-2)	0.289	0.000***
Δlgbeef_F(-3)	-0.126	0.001**
Δlgbeef_G(-1)	0.0073	0.845
Δlgbeef_G(-2)	0.032	0.402
Δlgbeef_G(-3)	-0.01	0.804
Extreme regime when ECT_{t-1}> 0.176		
ECT_{t-1}	0.0738	0.15
Δlgbeef_F(-1)	-0.164	0.138
Δlgbeef_F(-2)	0.0631	0.591
Δlgbeef_F(-3)	-0.471	0.000***
Δlgbeef_G(-1)	0.713	0.000***
Δlgbeef_G(-2)	-0.629	0.000***
Δlgbeef_G(-3)	0.491	0.000***
R-square	0.17	
Durbin-Watson	2.00	
B-G Serial correlation LM Test	0.70 [0.71]	
ARCH(1)	1.08[0.35]	
Observations in regime 1	693 accounting for 92.6% of total observations	
Observations in regime 2	56 accounting for 7.4% of total observations	

Bivariate VECM of lgbeef_D and lgpork_G normalized on lgbeef_D		
ECT_{t-1}	-0.077	-5.71[0.00]
Δlgbeef_F(-1)	-0.497	-13.71[0.00]
Δlgbeef_F(-2)	-0.246	-6.18[0.00]
Δlgbeef_F(-3)	-0.116	-3.23[0.00]
Δlgbeef_G(-1)	0.079	1.81[0.07]
Δlgbeef_G(-2)	0.052	1.15[0.24]
Δlgbeef_G(-3)	0.041	0.93[0.35]
R-square	0.26	
Durbin-Watson	1.98	
B-G Serial correlation LM Test	2.73[0.25]	
ARCH(1)	0.15[0.99]	

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