Cross-commodity Price Transmission and Integration of the EU Livestock Market of Pork and Beef: Panel Time-series Approach

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Abstract: At this study we analyze horizontal cross-commodity price transmission and integration of the EU livestock market of pork and beef. The study seeks to investigate whether or not there are long-term and short-term relationships between pork and beef prices in the EU. Our focus is on cross-commodity price transmission which provides valuable insights into market integration and efficiency. We utilize recently developed panel time-series techniques. Our data consists of monthly data on pork and beef prices in the EU member states during the period from February 1995 to June 2014. The estimation results reveal that there exists bi-directional relationship between pork and beef prices in the EU in long run. Cross-commodity price transmission between pork and beef has increased remarkably during the past ten years in the EU15 member states. Also the convergence to the equilibrium has sped up. In short run, we found evidence only for price transmission from pork prices to beef prices, not vice versa. Overall, short-run dynamics is significantly different in the EU livestock market compared to long run dynamics.

Key words: price transmission, market cointegration, panel time-series, pork and beef prices

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1. Introduction

In recent years, the empirical research on agricultural price transmission has gathered considerable attention. It is frequently utilized concept in market integration analysis. Price transmission can be separated into two types: horizontal and vertical. Vertical price transmission refers to price linkages along a supply chain, whereas horizontal price transmission means the price linkage occurring among different markets at the same position in the supply chain.

The literature analyzing vertical price linkages has concentrated on evaluations of the links between farm, wholesale and retail prices. The asymmetric vertical price transmission, i.e., increasing and decreasing prices at one level of supply chain transmit at different rates to another level, has aroused considerable discussion in agricultural economics (Vavra and Goodwin, 2005). The price relationships along the supply chain provide insights into marketing efficiency and consumers' and farmers' welfare (Aguiar and Santana, 2002). Meyer and von Cramon-Taubadel (2004) and Frey and Manera (2005) provide reviews of the literature on asymmetry price transmission.

The notion of horizontal price transmission usually refers to price linkages across market places (*spatial price transmission*). However, it can also concern the transmission across different agricultural commodities (*cross-commodity price transmission*)[1], from non-agricultural to agricultural commodities (for example from energy prices to agricultural prices; Serra et al., 2010 and Hassouneh et al., 2012), and across different purchase contracts for the same commodity (for example from futures to spot markets and vice versa; Baldi et al., 2013).

The key underlying theoretical explanation of spatial price transmission is the spatial arbitrage and the consequent Law of One Price (LOP)[2]. On the contrary, for cross-commodity price transmission, the co-movement of prices is mostly driven by the substitutability and complementarity relations among the products (Saadi, 2011), which, in turn, depends on the respective demand functions and on the underlying preferences. Substitutability and complementarity essentially implies that the prices of two commodities have a long-run relationship and shocks to one of the markets will get transmitted to another if markets are integrated. Understanding of inter-commodity price relationships and shock transmissions becomes important when *price volatility and* market failures formed the basis for public interventions. Even though, the background theory differs between spatial and cross-commodity price transmission, the empirical framework and the econometric implications of these different cases of horizontal price transmission are the same.

At this study we analyze horizontal cross-commodity price transmission and integration of the EU livestock market of pork and beef. The study seeks to investigate whether or not there are short-term and long-term relationships between pork and beef prices in EU. The contribution of this study to the existing literature on the price transmission in agricultural market is twofold. The first contribution is the focus on cross-commodity price transmission. Studies on this field are not as frequent as studies on vertical or spatial price transmission although they provide valuable insights into market integration.

The second contribution of our study refers to econometric methods: we utilize recently developed panel time-series techniques. In other words, to study the causality relationship between pork and beef prices, we catch both time-series and cross-country variation in our data set. Our data consists of monthly data on pork and beef prices in the EU member countries during the period from February 1995 to June 2014.

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[1] Detailed discussion: Esposti and Listorti (2012, 2013).
[2] The Law of One Price (LOP) is an economic concept which posits that "a good must sell for the same price in all locations". The intuition behind LOP is based on the assumption that differences between prices are eliminated by market participants taking advantage of arbitrage opportunities.

Panel models have many advantages compare to time series or cross-section models: panel models make more information available, hence more degrees of freedom and more efficiency. They allow controlling for individual heterogeneity. They also allow identifying effects that cannot be detected in simple time series or cross-section data. One benefit is better properties of the testing procedures when compared to more standard time series methods. In the panel framework, we can analyze long-run relationship across the panel while allowing the associated short-run dynamics and fixed effects to be heterogeneous across different members of the panel (Banerjee, 1999; Maddala and Wu, 1999).

To examine the linkages between pork and beef prices, we follow the standard three-step approach consisting of (i) assessing the stationarity of the time series, (ii) in case the variables are not stationary, checking whether they are characterized by a cointegration relationship, (iii) in case cointegration holds, testing for long-run and short-run relationships between pork and beef prices by estimating error correction mechanism (ECM), which permits to analyse the long-run relationship between the variables jointly with the short-term adjustment towards the long-run equilibrium.

In analyses of the price linkages in the meat sector, panel causality approach has been utilized mainly in vertical price transmission studies, for example, Boetel and Liu (2008), Goodwin and Holt (1999), Goodwin and Harper (2000) and Carraro and Stefani (2010).

In our study, testing for unit root is performed using the panel unit root test of Levin, Lin and Chu (2002), Im, Pesaran and Shin (2003), Breitung (2000), Maddala and Wu (1999), Choi (2001), Pesaran (2003) and Im, Lee and Tislau (2005, 2010). The panel cointegration test based on Pedroni (1999, 2004), Kao (1999), Westerlund (2005, 2007), Westerlund and Edgerton (2008) are considered. In addition, we carry out Johansen's Fisher panel cointegration test suggested by Maddala and Wu (1999). Moreover, Mean-Group estimator (MG) by Pesaran and Smith (1995), Pooled Mean Group estimator (PMG) by Pesaran, Shin and Smith (1999), group-mean Dynamic OLS (DOLS) and group-mean Fully Modified OLS (FMOLS) by Pedroni (2000) are implemented to test for Granger-causality in panel models. We use the software Stata for empirical analyses.

The main objective of the study is to analyze the dynamics of meat price transmission and market integration between pork and beef in the EU. The specific objectives are:

- To examine the trend in prices of pork and beef in the EU,
- To analyze the price integration between pork and beef markets in the EU,
- To identify the long and short run price transmission of pork and beef in the EU.

The rest of the paper is organized as follows: In the next section we briefly discuss on the concept of market integration and price transmission and how it can be modelled in the panel framework. We also take a quick snapshot on studies on the EU meat market integration. Section 3 defines the data and its limitations. Section 4 gives the results and discussion and the last section concludes.

2. Background

2.1. Market integration and horizontal price transmission

Market integration is an indicator that explains how much different markets are related to each other. Price transmission reflects the extent of market integration and the extent of market efficiency. Consequently, analysis of relationships between prices is a common tool in market integration analysis. Even though theory in general argues that other variables (product attributes) are equally important in describing and explaining market integration, this interest in prices can be justified because data on prices are easier to obtain, and often the only data available.

Price transmission can be understood from two aspects: vertical and horizontal. The literature analyzing vertical price linkages has concentrated on evaluations of the links between farm, wholesale and retail prices.

In general, the primary attention in studies that analyze vertical price transmission is the magnitude, speed and nature of the price adjustments through the supply chain and especially the extent to which adjustments are asymmetric.[3]

Although studies into vertical price transmission are the most frequent, horizontal price transmission is the focus of this study. The literature analyzing horizontal price linkages dates back more than one-hundred years and is typically concerned with links between prices at different locations (spatial price arbitrage). Spatial price arbitrage is an equilibrium concept, where in a well-functioning market, transactions between spatially dispersed agents will ensure that price shocks occurring in one market evoke responses in other market places (Serra et al., 2006). In theory, prices of homogeneous goods in two separate locations will differ by, at most, the cost of moving the commodity form the cheapest market place to the most expensive one. This arbitrage condition is equivalent to the weak version of the Law of One Price (LOP) (Fackler and Goodwin, 2001). Numerous concepts such as market integration and market efficiency have been used to describe spatial price arbitrage (Serra et al., 2006 and Fousekis, 2007). The majority of the studies on spatial price relationships have been focused on food and raw material markets.

Markets are not related only between regions but also between commodities. Horizontal price transmission also concerns the transmission across different agricultural commodities (*cross-commodity price transmission*) (Esposti and Listorti, 2013). In this case of horizontal price transmission the co-movement of prices is mostly driven by the substitutability and complementarity relations among the commodities (Saadi 2011), which, in turn, depends on the respective demand functions and on the underlying preferences. However, the question is not whether the two commodities are complements or substitutes, but what degree they are first and what degree they are second.

When goods and information flow freely, shocks occurring in one market will evoke responses in other markets. In cross-commodity price transmission studies are often examined the question of how shocks in one of the commodity markets will get transmitted to the rest if markets are integrated. As the price of one commodity increases following an international price shock, demand may shift towards another commodity resulting in higher prices on these as well. Thus, stabilisation of prices of essential agricultural commodities continues to remain an area of major concern for policy makers. An understanding of inter-commodity price relationships and shock transmissions forms the basis for public interventions (Alderman, 1993; Rashid, 2011; Sharma and Kumur, 2001). Price instability affects both producers and consumers and has macroeconomic implications as well. World widely, this was an important aspect during the global food crisis from the early 2007 to the middle 2008, during which the prices of a number of food commodities increased sharply. It was followed by a period of collapsing prices in the second half of 2008.

There are only few studies on cross-commodity prices transmission. For example, Asche et al. (2005) examined cross-commodity market integration between wild and farmed salmon on the Japanese market and found that the species were close substitutes on the market, and that the expansion of farmed salmon had resulted in price decreases for all salmon species. Nielsen et al. (2007) found that markets for farmed trout are related toothed fish markets in Germany, and that markets for these trout are more closely linked to markets for captured fish than to farmed salmon. Timmer (2009) addresses the long-run relationships among the prices of the three basic cereal staples, rice, wheat and corn (maize) in Asian market. He found that wheat and corn are more closely connected with each other than with rice markets. Rashid (2011) examined inter-commodity price relationships to assess the relative importance of each of the three major cereals (maize, wheat, teff) in Ethiopian cereal markets. He concluded that maize is the most significant in exacerbating price variability with respect to the persistence of shocks to itself and the two other cereals. He noticed that focusing on maize helps to stabilize prices and also to reduce costs of stabilization. Sharma and Kumur (2001) found complex long run relationships among rice, wheat, groundnut seed, mustard seed, cottonseed, groundnut oil, vanaspati oil, mustard oil and other edible oils in Indian market. They concluded that simply analyzing the price behavior of one commodity while ignoring the behavior of the prices of substitutes will not be meaningful for price stabilization purposes. The earlier researches includes for example Alderman

[3] Studies on vertical price transmission: Peltzman (2000), Meyer and von Cramon-Taubadel (2004), McCorriston and Sheldon (1996), Vavra and Goodwin (2005).

(1993) who investigated how information is transmitted across commodities (maize, millet, sorghum) in Ghana. He noted imperfections in the way markets process information: the lagged price of maize conveys information that is not contained in the past price of sorghum or millet.

Although the studies specifically related to cross-commodity price movements are only a few, after the biofuels issue has being raised by the recent food crisis, a considerable number of papers have been dealing with the impact of crude oil on major agricultural commodities.[4]

2.2. Meat market integration in the EU

The integration of agricultural markets has always been very important for the member states of the EU. Therefore, they merged the organization of the sector in the (Common Agricultural Policy CAP). Under the EU's Common Agricultural Policy, the agricultural markets are required to become spatially integrated within and between all member states. In an integrated market, price information related to the production costs should be efficiently transmitted between the member states. Specially, the pork market is said to be homogenous because of its concentration on some dominating countries. Several authors found a stable long-run equilibrium using cointegration tests on small samples of countries, for instance Sanjuan and Gil (2001) and Serra et al. (2006). Nevertheless, investigating a larger sample of 14 EU countries, Fousekis (2007) found that the prices are not homogenous. The bovine market of the EU is more heterogeneous due to the disturbances caused by policy measures (CAP) and a large variety of production strategies (e.g. suckler cow husbandry, bull-mast or as a side product of milk production). Because of a large number of exogenous shocks, such as policy changes, the standard cointegration measures are problematic. For that reason Ihle et al. (2012) used a methodology that is robust for break points and found cointegration between the calf prices of four EU countries.

The developmental steps towards market efficiency in the EU pork and beef markets are analyzed by Liu (2011). She evaluated the extent to which the Finnish domestic meat market responds to changes in the European price (German and Danish prices). According to Liu, 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. 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.

Meyer (2012) applied panel convergence tests with individual adjustment processes in order to analyze the convergence of the EU livestock markets of pork and beef. The estimation results reveal that the overall price heterogeneity is reducing within the EU for both beef and pork. Meyer confirms that exogenous changes, such as the enlargement and policy measures (European Food Monitoring Tool and Mid-Term Review) improved the functioning of the internal markets. According to the study, larger heterogeneity of the beef prices is the result of the still remaining differences of policy measures within the member states of the EU. Focusing on the analysis of the EU enlargement in 2004, the study found a stronger convergence of the prices of the new member states compared to the old member states, which confirms that the accession countries are catching-up. Nevertheless, the prices in the more segmented beef markets of the accession countries were less strongly progressing towards homogeneity. A detailed study of the EMU and non-EMU countries indicated that the dropping of the currency risk has indeed had an influence on agricultural markets. The countries within the Eurozone converge faster than the other countries, which significantly reduced the welfare losses of consumers and producers. The study also reveals that at the beginning of the ongoing eurocrisis the prices within the EU only temporarily dispersed, but then converged again.

[4] For example Arshad and Hameed (2009), Saghaian (2010), Baffes (2007), Muhammad and Kebede (2009), Bakhat and Wurzburg (2013), Ciaian and Kancs (2011) and Kristoufek, Janda and Zilberman (2012).

2.2. Modelling price transmission in the panel framework

Let us consider agricultural prices observed over three different dimensions: space, commodity and time. The generic price is $p_{i,k,t}$ where: $i=1,...,j,...,N$ is the market place (spatial dimension); $k=1,...,h,...,K$ is the commodity; $t=1,...,s,...,T$ is the period of observation (time dimension). By more conventionally distinguishing between a cross-sectional dimension, given by the combination of the dimensions ik, and a time dimension t, we can identify any generic price observation as $p_{ik,t}$ (scalar) and any generic price series (vector) as $\mathbf{p}_{ik}$. Notice that here we consider the logarithms of prices. This monotonic transformation facilitates the economic interpretation of results, in particular considering that regression coefficients may be interpreted as elasticities. Consequently, henceforth $\mathbf{p}_{ik}$ identifies the time series of the price logarithm of the k^{th} commodity in the i^{th} market place.

The behavior of $p_{i,k,t}$ over its three dimensions can be represented within suitable structural models as the combination of market fundamentals such as supply, demand and stock formation. However, such models are complex and hardly tractable in the empirical analysis, whereas the investigation of price evolution and linkage is more frequently afforded within reduced-form models. When all the three dimensions are explicitly considered, reduced-form models are actually and by far more feasible and of immediate use to generate price predictions, i.e. the estimation of $E(p_{i,k,t} \mid p_{j,k,t-s}, p_{i,h,t-s}, p_{i,k,t-s})$, given the available observations.

By distinguishing a cross-sectional dimension, ik, and a time dimension, t, a generic reduced-form model of price formation and transmission over these two dimensions is the following:

$$p_{ik,t} = \alpha_{ik} + \sum_{s=1}^{s=S<T} \rho_s p_{ik,t-s} + \sum_{s=0}^{s=S<T} \sum_{jh \neq ik} \omega_{ik,jh}^{s} p_{jh,t-s} + \varepsilon_{ik,t} \quad (2.1)$$

where S is the maximum time lag and $\varepsilon_{ik,t} \sim N(0, \sigma_{ik,t}^2)$. In a more compact matrix form equation can be written as:

$$\mathbf{P} = \boldsymbol{\alpha} + \sum_{s=0}^{s=S<T} \mathbf{P}_s \mathbf{W}_s + \boldsymbol{\varepsilon}_t \quad (2.2)$$

where $\mathbf{P}$, $\mathbf{P}_s$ and $\boldsymbol{\varepsilon}_t$ are $(T\times(N\times K))$ matrices, s expresses the time lag, $\boldsymbol{\alpha}$ is a $(T\times(N\times K))$ matrix of time invariant parameters, that is $\alpha_{ik,t} = \alpha_{ik,t-s} = \alpha_{ik}, \forall i, j, s$ (any column of $\boldsymbol{\alpha}$ contains T elements with constant value α_{ik}) and $\boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \boldsymbol{\Omega}_t)$. $\mathbf{W}_s$ is a $((N\times K) \times (N\times K))$ matrix of unknown parameters incorporating the correlation across prices within both the time and cross-sectional (space-commodity) dimensions. The diagonal elements, $\omega_{ik,ik}^{s}$, indicate the auto-correlation over time, with the exclusion of the matrix $\mathbf{W}_0$, where diagonal elements are evidently $\omega_{ik,ik}^{0} = 0 \ \forall \ ik$. The off-diagonal elements, $\omega_{ik,jh}^{s}$, represent the cross-sectional dependence of prices; in other words, they express the interdependence among the different prices and, therefore, the degree and the direction of transmission of the price shocks.

In particular:

- if $h=k$ but $i\neq j$, we are considering the price transmission for the same commodity across space, that is, different market places. In this case, under perfect spatial arbitrage, the validity of the Law of One Price (LOP) implies that $\omega_{ik,jk} = 1$;
- if $i=j$ but $h\neq k$, we are considering the price transmission between two different commodities in the same market. In this case, elements $\omega_{ik,ih}$ indicate the degree of substitutability between the different goods. $\omega_{ik,ih}$ will be close to 1 (-1) under perfect substitutability (complementarity) between h and k, while it will be close to 0 under low substitutability (complementarity).

As $\mathbf{p}_{ik}$ indicates the logarithms of prices, the elements of $\mathbf{W}_s$ express the price transmission elasticities. Within the logarithmic form, the implicit assumption is that all factors possibly contributing to price differentials but not explicitly taken into account in the model (for example, transportation and transaction costs) are a constant proportion of prices.

These constant multiplicative terms (that can be naturally intended as percentages) apply to price $p_{ik,t-s}$, to obtain $p_{jh,t}$ and are captured by the elements of $\boldsymbol{\alpha}$.

If the matrix of unknown parameters, $\mathbf{W}_s$, contains all the information about price linkages over the three dimensions, we can expect that the transmission equations (2.1-2.2) get rid of possible autocorrelation and heteroskedasticity across both the time and cross-sectional dimensions; we can assume that spherical error terms are restored: $\boldsymbol{\varepsilon}_t \sim N(\mathbf{0}, \sigma\mathbf{I})$ and $E(\varepsilon_t, \varepsilon_{t-s}) = 0$. The proper specification of (2.1-2.2) aims indeed at restoring such conditions.[5]

3. Data description

In order to analyze the price integration of pork and beef in the EU, we use the data on monthly pork and beef prices in the EU member states, which are obtained from Tike (Agricultural Statistics Finland)[6]. Tike is responsible for agricultural price monitoring in Finland and price monitoring reports include data on both Finland and other EU countries.

In our dataset, the EU member countries are categorized into three groups of EU15[7], EU25[8] and EU27[9]. Data on EU15 covers the period from February 1995 to June 2014. Correspondingly, the data on EU25 covers the period from January 2005 to June 2014 and the data on EU27 covers the period from January 2007 to June 2014. In other words this means that the single price time series from EU15 includes 221 observations; the single price time series from EU25 consists of 102 observations and the single price time-series from EU27 includes 78 observations. In empirical analysis we carry out our investigations by utilizing six different data specifications: we have one group of EU member states (EU15 countries) when investigating period 2/1995-6/2014, we have two groups of EU member states (EU15 and EU25) when investigating period 1/2005-6/2014 and we have three groups of EU member states (EU15, EU25 and EU27) when investigating period 1/2007-6/2014. By repeating the estimations with different group of member states we can make conclusions regarding stability of the results over time and also regarding cross-sectional stability.

We have constituted our monthly price data by taking averages from weekly data. Weekly data includes many missing values which are problematic in our investigation. We have estimated and filled some missing values with the help of data authors. As a result, we have succeed to create balanced monthly level panel data. However, one should notice that Malta is missing in the panels because its time-series includes serious shortages[10]. Regarding other countries, quality of data is good and time-series are completed.

Carcass classification plays an important role in the European Union meat market. The aim of the carcass classification scheme is to ensure a common classification standard throughout the European Union which enables the EU to operate a standardized price reporting system. The EU classification of the analyzed pork is labeled “E”, which indicates that 55% or more of the carcass has to be lean meat. The quality of the beef is correspondingly “R3”. According to the EU grading scheme, it is qualitatively good meat, which means that the overall profiles are straight, the muscle development is good and the content of fat is medium. These two are also the most typical carcass classes.

The descriptive statistics of pork and beef prices are reported for each member state in Tables 3.1 and 3.2. Notice that in empirical analysis we use the natural logarithms of these variables in order to obtain the elasticities.

[5] More detailed discussion: Esposti and Listorti (2013)

[6] I am grateful to Pia Outa and Lauri Juntti for sending me the excellent price data.

[7] EU15 member states are Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and United Kingdom.

[8] EU25 member states are EU15 plus Cyprus, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Slovak Republic and Slovenia.

[9] EU27 member states are EU25 plus Bulgaria and Romania.

[10] Despite the fact that Malta is missing in our data sets, we use terms “EU25” and “EU27” in our empirical investigations.

Table 3.1. Descriptive statistics of pork prices in the EU member states

PORK	2/1995-6/2014 (EU15)				1/2005-6/2014 (EU25)				1/2007-6/2014 (EU27)			
	Mean	Std.	Min	Max	Mean	Std.	Min	Max	Mean	Std.	Min	Max
Austria	148.5	21.8	86.2	227.1	151.6	15.7	122.7	195.3	152.6	16.9	122.7	195.3
Belgium	140.7	21.3	82.3	228.9	141.9	13.8	117.4	183.5	142.8	14.8	117.4	183.5
Bulgaria									179.0	14.8	153.9	230.7
Cyprus					173.7	24.1	126.9	234.0	172.2	24.9	126.9	234.0
Czech Republic					155.0	16.9	122.2	196.4	157.5	17.7	122.2	196.4
Denmark	131.8	19.4	88.2	195.9	134.4	16.5	107.0	176.0	136.9	17.4	107.0	176.0
Estonia					152.5	13.5	133.4	184.4	156.1	12.7	133.6	184.4
Finland	143.0	15.3	109.2	184.8	149.7	14.4	130.7	184.8	153.4	13.9	135.1	184.8
France	140.3	19.5	93.9	206.4	142.3	15.4	116.0	190.0	143.4	16.3	116.0	190.0
Germany	150.5	22.1	84.5	232.1	155.4	15.4	118.3	197.9	156.6	16.3	118.3	197.9
Greece	172.8	24.9	110.4	241.4	179.0	18.1	130.6	213.9	177.8	18.9	130.6	213.9
Hungary					153.6	15.9	123.3	193.1	156.1	16.6	126.3	193.1
Ireland	136.8	16.6	88.7	179.7	142.8	14.1	117.5	174.2	144.7	14.9	117.5	174.2
Italy	159.4	22.4	99.7	226.7	162.4	21.6	120.7	226.7	165.9	22.4	125.4	226.7
Latvia					160.9	17.3	129.4	203.4	163.8	18.1	129.4	203.4
Lithuania					156.7	16.9	126.4	206.9	160.1	17.0	126.5	206.9
Luxembourg	157.7	22.3	104.1	257.2	155.7	14.3	126.2	196.4	156.9	15.1	126.2	196.4
Netherlands	132.3	21.3	64.3	226.5	139.0	14.6	110.9	179.0	140.3	15.4	110.9	179.0
Poland					148.1	20.6	112.2	197.5	152.6	20.3	112.2	197.5
Portugal	154.6	22.9	85.3	223.5	160.0	16.4	129.0	201.0	161.6	16.4	129.0	201.0
Romania									164.7	17.1	136.7	205.6
Slovakia					157.5	17.7	124.4	201.8	160.7	18.1	124.4	201.8
Slovenia					150.4	14.8	118.6	191.4	151.0	16.2	118.6	191.4
Spain	150.0	24.8	74.2	217.1	157.4	20.9	118.1	217.1	159.7	21.7	118.1	217.1
Sweden	144.7	19.3	104.4	198.2	152.7	19.8	118.1	198.2	157.1	19.8	118.1	198.2
UK	155.1	19.6	89.9	201.5	163.2	16.8	127.8	201.5	167.1	16.7	127.8	201.5

Table 3.2. Descriptive statistics of beef prices in the EU member states

BEEF	2/1995-6/2014 (EU15)				1/2005-6/2014 (EU25)				1/2007-6/2014 (EU27)			
	Mean	Std.	Min	Max	Mean	Std.	Min	Max	Mean	Std.	Min	Max
Austria	303.4	41.4	215.0	410.0	335.8	34.9	243.3	410.0	343.6	35.1	243.3	410.0
Belgium	259.1	29.9	201.8	332.3	274.0	26.2	209.4	332.3	280.4	24.5	244.9	332.3
Bulgaria									242.2	44.8	163.6	360.5
Cyprus					286.5	22.9	235.5	345.0	293.3	20.5	251.7	345.0
Czech Republic					294.7	30.5	235.5	362.4	303.6	27.8	251.7	362.4
Denmark	301.2	49.3	219.2	413.8	340.7	39.6	259.3	413.8	350.7	37.5	277.4	413.8
Estonia					262.5	41.3	179.3	357.0	275.6	35.4	179.3	357.0
Finland	315.0	42.4	242.5	424.9	346.5	35.4	286.8	424.9	358.3	30.4	305.7	424.9
France	300.8	41.9	215.0	401.4	332.5	35.1	262.0	401.4	337.7	37.0	262.0	401.4
Germany	296.2	49.5	168.2	423.3	334.9	39.7	261.6	423.3	343.6	39.7	261.6	423.3
Greece	392.2	27.8	313.1	455.5	413.2	21.5	362.6	455.5	422.1	12.7	371.4	455.5
Hungary					242.9	19.0	185.1	296.0	239.4	19.5	185.1	296.0
Ireland	281.5	52.6	188.4	437.7	322.1	46.3	238.5	437.7	334.1	44.5	272.0	437.7
Italy	328.6	39.9	197.3	422.3	359.7	29.3	277.0	422.3	366.8	27.2	277.0	422.3
Latvia					207.1	39.4	127.9	321.0	217.9	36.8	147.0	321.0
Lithuania					250.8	46.1	164.3	335.2	261.3	45.5	178.5	335.2
Luxembourg	301.7	38.9	182.3	399.7	327.6	33.8	247.8	399.7	335.2	33.2	247.8	399.7
Netherlands	277.0	40.4	171.1	380.3	303.7	29.1	244.0	380.3	309.6	29.0	256.8	380.3
Poland					271.5	38.0	210.4	354.7	281.7	36.2	217.7	354.7
Portugal	324.0	31.4	247.2	390.7	347.0	21.5	283.9	390.7	351.6	17.4	308.4	390.7
Romania									249.4	40.4	159.9	346.1
Slovakia					291.5	41.4	182.4	390.7	302.5	39.5	216.5	390.7
Slovenia					314.4	32.8	254.1	392.0	323.2	31.0	258.7	392.0
Spain	305.4	40.6	208.1	400.7	336.7	32.4	279.0	400.7	343.7	31.7	283.9	400.7
Sweden	287.8	49.0	211.3	460.3	316.7	54.3	231.3	460.3	329.6	54.0	231.3	460.3
UK	292.4	55.9	195.3	450.9	330.5	56.6	238.7	450.9	345.2	54.4	276.8	450.9

The descriptive statistics reveal, among other things, that in the new member states pork prices are higher than in the old member states. However, in the case of beef, the situation is opposite. In figures 3.1, 3.2 and 3.3 are showed how the pork and beef prices has evolved during the period 2/1995(EU15) or 1/2004(EU25) or 1/2007(EU27)-6/2014 in the EU member states. In order to make the figures more comparable, we have indexed the data. From the user's point of view, the price indexes are the most practical, because index revisions do not interrupt the series. Additionally, visual inspection of the pork and beef price time series, and in advance of the subsequent econometric analysis, may help to assess sustainability issues in individual cases.

All in all, the figures confirm that there is some relationship between the pork and beef prices. The relationship between the price series has strengthened *since* the beginning of the 2000s in the EU15 member states. Before that, the developments in the pork and beef price series were divergent. Notable, fluctuations in pork price series are much larger than fluctuations in beef price series in the EU15. However, when examining the price series of new member states the situation is opposite: beef prices have fluctuated more than pork prices.

The figures also reveal that there exist seasonal cycles particularly in pork prices and mainly in the Southern European countries. In order to control for the potential seasonal differences in our econometric investigations in the next chapter, we add dummy variables for the months of January and March through December. For example, Osborn (1993) argued that seasonal non-stationarity can often be adequately represented by seasonal dummies.

Figure 3.1. The development of pork and beef prices in the EU15 member states during the period 2/1995-6/2014, weekly data, index: week 5/1995=100.

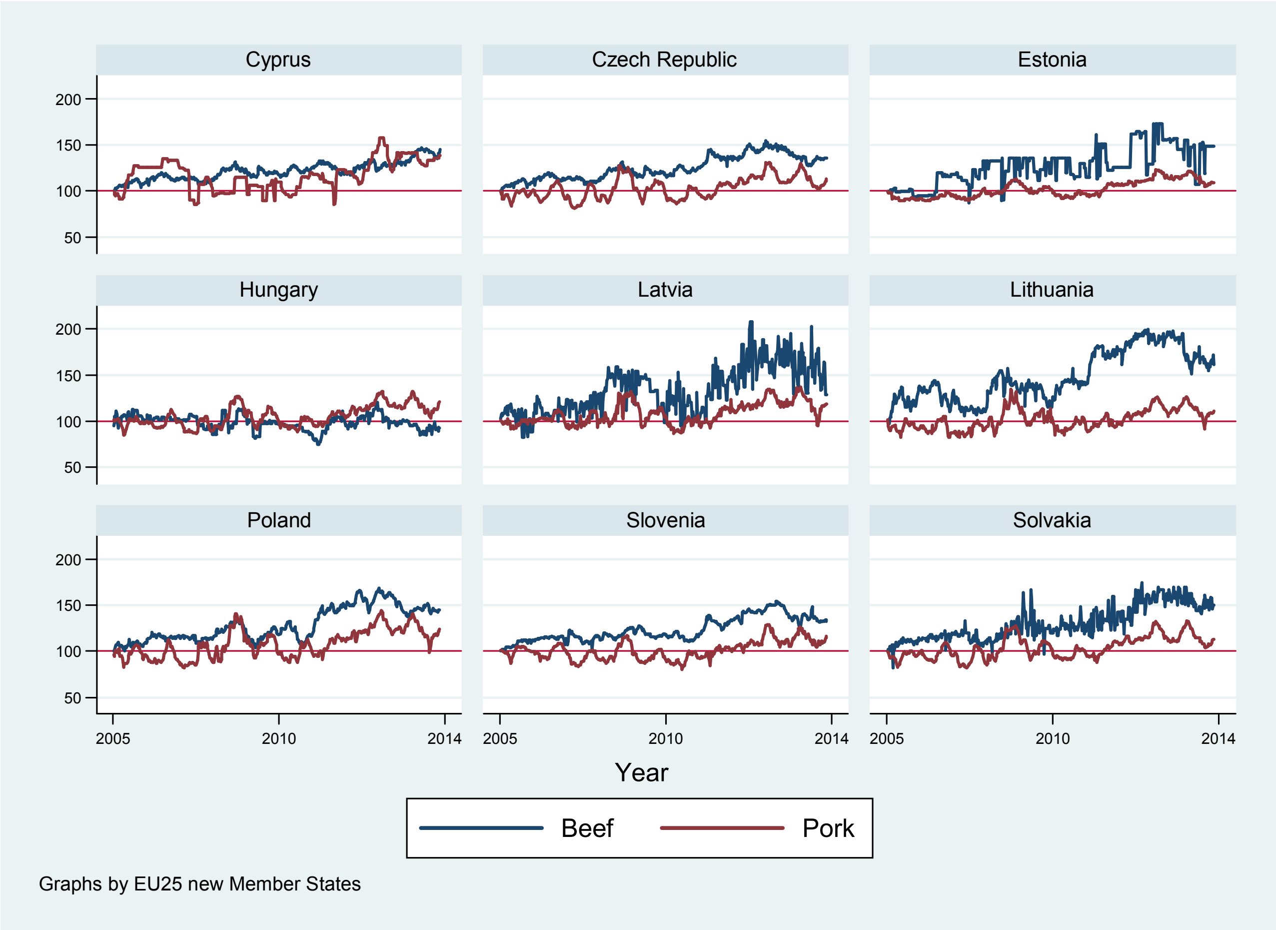

Figure 3.2. The development of pork and beef prices in the EU25 member states (excl. EU15) during the period 1/2005-6/2014, weekly data, index: week 1/2005=100.

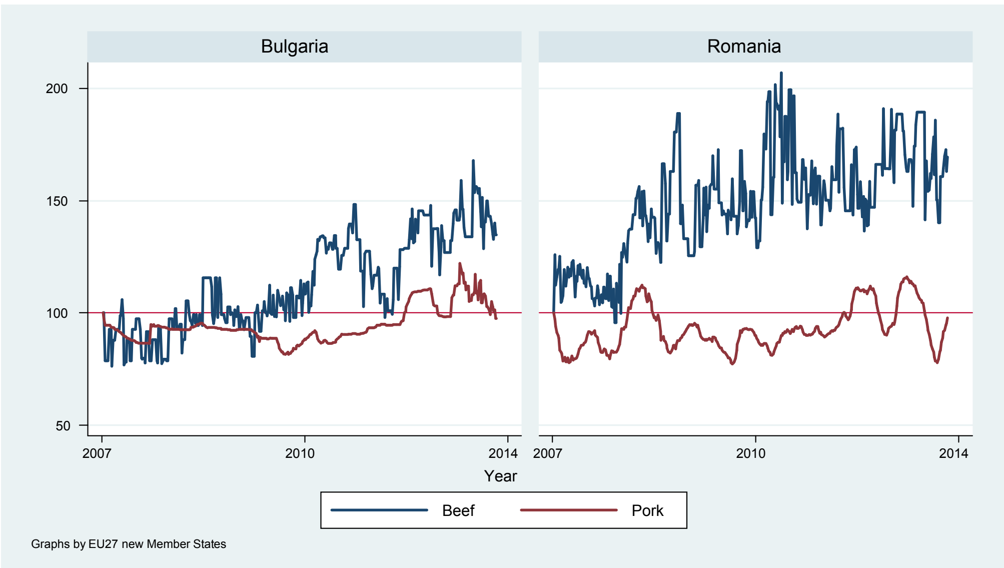

Figure 3.3. The development of pork and beef prices in the EU27 member states (excl. EU25) during the period 1/2007-6/2014, weekly data, index: week 1/2007=100.

4. Empirical results

4.1. Panel unit root results

A variety of procedures for the analysis of unit roots in a panel context have been developed. The logic behind the use of a panel unit root test is to combine the information from time series with the information from cross-sectional units. There may be considerable potential for tests that can be employed in an environment where the time series may be of limited length, but very similar data may be available across a cross–section of countries, regions, firms, or industries. Since the power of unit root tests depend on the total variation in the data used (both in the number of observations and their variation), panel unit root tests are more powerful than standard time-series unit root tests. The variation across cross-section units improves estimation efficiency, leading to smaller standard errors and, consequently, to higher t-ratios and potentially more precise parameter estimates. With the increasing availability of quite rich panel data sets in a number of contexts, these kinds of tests would seem very attractive. One of the advantages of panel unit root tests is that their asymptotic distribution is standard normal. This is in contrast to individual time series unit roots which have non-standard asymptotic distribution.

However, a variety of issues arise when panel data are employed in testing for unit roots. Early panel unit root tests generally ignore cross-sectional dependence that is common for most of the macroeconomic series. Despite the shortcoming of early unit root tests, they help researchers to develop new tests to deal with dependence across cross-section units. These recent tests try to find out the same question of non-stationarity with early ones but have different approaches to resolution and different implications for empirical research. Therefore, in order to compare results predicated by these tests and to discover what the impact of cross-sectional dependence assumption is on unit root testing, we use more than one unit root tests. The first-generation tests we conduct are Levin, Lin and Chu (2002), Im, Pesaran and Shin (2003), Breitung (2000), Maddala and Wu (1999) and Choi (2001). Table 4.1 show the statistics and the p-values from the panel unit root tests concerning beef price series and Table 4.2 concerning pork price series. After the first-generation unit root tests, we conduct Pesaran's (2003) second-generation unit root test (Table 4.3), which takes the cross-sectional dependence in account. We also employ unit root test based on the Lagrangian multiplier (LM) principle developed by Im and Lee (2001) which applies when a structural break occurs at different time period in each time series as well as when the structural break occurs in only some of the time series (Table 4.4)[11]. The proposed test is not only robust to the presence of structural breaks, but is also powerful in the basic case where no structural breaks are involved.

In all of the tests we use the natural logarithms of the prices.

The first sections in Table 4.1 and 4.2 show the empirical result with price series that include individual effects only, while the latter sections include individual effects and individual trends. The trends amount to fixed effects in the first difference specification. Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution, χ^2 with $2*N$ degrees of freedom. IPS, LLC and Breitung tests assume asymptotic normality. Selection of lags is based on the Schwarz criterion (SIC)[12] and the Newey-West bandwidth[13] selection using the Bartlett-Kernel[14] method.

[11] To our pork and beef price series structural breaks may have been caused by the enlargement of the EU, policy measures (European Food Monitoring Tool, Mid-Term Review), CAP reforms, EMU and epidemic and pandemic diseases, for example BSE.

[12] Estimating the lag length of autoregressive process for a time series is a crucial econometric. When fitting models, it is possible to increase the likelihood by adding parameters, but doing so may result in overfitting. Schwarz criterion (SIC) (or Bayesian information criterion (BIC)) resolves this problem by introducing a penalty term for the number of parameters in the model. Thus, SIC is a criterion for model selection and it is based on maximizing or minimizing the likelihood function.

[13] Because the PP test differs from the ADF test in the treatment of serial correlation, the information-based method (such as SIC) does not apply to the PP test. In the PP test, an issue concerns the choice of a lag truncation parameter is a choice of the bandwidth (autocovariance lag). Newey and West (1994) proposed a method for estimating the optimal bandwidth from truncated sample autocovariances.

[14] A weighting function used in non-parametric estimation techniques.

The alternative hypothesis of a stationary process differs between panel unit root tests. Namely, the alternative hypothesis of LLC and Breitung tests assume that all panels are stationary. In other tests the alternative hypotheses assume that some or at least one cross section is stationary. The null hypothesis is the same in all the tests: all panels contain unit roots.

According to the results, more or less, a unit root is detected for the level variables, while the first differences appear to be stationary. For the beef and pork price series, only the LLC and Breitung tests reject the null in some cases, while other three tests can not reject the unit root hypothesis. After first-differencing the variables and repeating the tests, nonstationarity is eliminated as all panel unit root tests reject the null of nonstationarity for the first-differenced variables at the 1% level of significance. We can thereby conclude that according to first generation panel unit root tests first-differenced variables are stationary so that panel variables are integrated of order one, *I(1)*. These results imply that the long-run panel variables are not stationary, and that these variables could be cointegrated.

Table 4.1. Results of panel unit root tests: Beef

ln beef	Individual effects						Individual effects + individual trend					
	EU15	EU15	EU25	EU15	EU25	EU27	EU15	EU15	EU25	EU15	EU25	EU27
	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014
Series in levels												
Levin, Lin and Chu	-3.452	-4.083	-7.121	-4.914	-2.660	-2.497	-6.336	-4.485	-9.282	-5.642	-2.144	-3.951
	(0.009)***	(0.006)***	(0.005)***	(0.010)**	(0.015)**	(0.019)**	(0.001)***	(0.034)**	(0.000)***	(0.004)***	(0.057)*	(0.018)**
Im, Pesaran and Shin	0.665	0.496	0.551	0.728	0.617	0.382	0.854	0.636	0.379	0.938	0.477	0.513
	(0.551)	(0.438)	(0.476)	(0.581)	(0.527)	(0.359)	(0.743)	(0.522)	(0.041)**	(0.862)	(0.024)**	(0.39)**
Breitung	0.363	1.193	-0.870	1.851	-0.927	-1.225	-0.344	0.942	-0.323	0.252	-0.288	-0.982
	(0.247)	(0.646)	(0.019)**	(0.976)	(0.009)***	(0.003)***	(0.007)***	(0.084)*	(0.038)**	(0.020)**	(0.034)**	(0.007)***
Fisher ADF	-1.428	-2.046	-1.872	-2.781	-2.547	-1.892	-1.676	-2.546	-2.047	-2.924	-2.492	-1.973
	(0.669)	(0.772)	(0.727)	(0.831)	(0.808)	(0.764)	(0.721)	(0.799)	(0.723)	(0.854)	(0.802)	(0.793)
Fisher PP	-2.038	-1.514	-1.113	-1.871	-1.090	-1.555	-1.865	-1.232	-0.881	-1.660	-1.059	-1.436
	(0.657)	(0.726)	(0.909)	(0.831)	(0.966)	(0.838)	(0.743)	(0.924)	(0.975)	(0.816)	(0.957)	(0.875)
Series in first differences												
Levin, Lin and Chu	-127.452	-107.363	-76.481	-169.717	-150.626	-90.990						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Im, Pesaran and Shin	-60.384	-45.279	-78.919	-89.258	-90.316	-82.623						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Breitung	-30.782	-19.677	-16.511	-38.462	-24.953	-12.874						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Fisher ADF	-10.950	-15.323	-12.122	-18.645	16.548	-13.578						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Fisher PP	-39.570	-32.451	-25.811	-40.624	-24.975	-29.557						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. p-values are reported in parentheses. Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Selection of lags based on SIC. Newey-West bandwidth selection using Bartlett kernel. Null hypothesis: unit root.

Table 4.2. Results of panel unit root tests: Pork

ln pork	Individual effects						Individual effects + individual trend					
	EU15	EU15	EU25	EU15	EU25	EU27	EU15	EU15	EU25	EU15	EU25	EU27
	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014
Series in levels												
Levin, Lin and Chu	-4.434	-2.895	-2.553	2.076	2.792	4.617	-6.271	-4.390	-1.723	2.472	4.766	7.588
	(0.004)***	(0.005)***	(0.018)**	(0.257)	(0.333)	(0.179)	(0.002)***	(0.006)***	(0.014)**	(0.351)	(0.468)	(0.763)
Im, Pesaran and Shin	-0.622	-0.752	-0.434	-0.445	-0.279	-0.390	-0.277	-0.0896	-0.812	-0.543	-0.614	-1.154
	(0.356)	(0.292)	(0.458)	(0.457)	(0.572)	(0.469)	(0.014)**	(0.006)***	(0.299)	(0.041)**	(0.343)	(0.097)
Breitung	-1.466	-2.441	1.390	0.313	0.253	0.218	-3.897	-5.126	0.364	-2.352	0.221	0.410
	(0.019)**	(0.005)***	(0.051)	(0.365)	(0.286)	(0.224)	(0.001)***	(0.001)***	(0.418)	(0.089)	(0.263)	(0.475)
Fisher ADF	-0.804	-0.736	-0.378	-0.649	-0.589	-0.425	-0.765	-0.704	-0.315	-0.686	-0.625	-0.552
	(0.964)	(0.957)	(0.922)	(0.945)	(0.940)	(0.931)	(0.950)	(0.941)	(0.925)	(0.946)	(0.948)	(0.937)
Fisher PP	-0.632	-0.532	-0.311	-0.619	-0.628	-0.455	-0.661	-0.474	-0.393	-0.556	-0.512	-0.420
	(0.943)	(1.000)	(1.000)	(0.967)	(0.960)	(1.000)	(0.940)	(1.000)	(1.000)	(0.974)	(1.000)	(1.000)
Series in first differences												
Levin, Lin and Chu	-148.661	-112.414	-96.526	-224.133	-197.067	-190.471						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Im, Pesaran and Shin	-62.335	-71.456	-78.980	-85.438	-46.177	-50.512						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Breitung	-74.589	-51.126	-43.229	-80.165	-62.834	-37.244						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Fisher ADF	-24.878	-20.247	-12.332	-17.711	-15.948	-13.286						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						
Fisher PP	-47.254	-52.815	-39.729	-58.610	-32.882	-25.569						
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***						

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. p-values are reported in parentheses. Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality. Selection of lags based on SIC. Newey-West bandwidth selection using Bartlett kernel. Null hypothesis: unit root.

Unit root tests that assume cross-sectional independence can have low power if estimated on data that have cross-sectional dependence. As a result, Pesaran's (2007) CIPS test for unit roots is calculated. This is a second generation unit root test that allows for cross-sectional dependence. Pesaran included cross-sectional averages of the lagged levels as the common factor to filter out the cross-section dependence. The statistic is constructed from the average of the cross-sectionally ADF (CADF) *t*-statistics. The averaging of the group-specific results follows the procedure in the Im, Pesaran and Shin (2003) test and brings out the CIPS *t*-statistic. Under the null of nonstationarity the test statistic has a non-standard distribution. The critical values are tabulated by the author for different combinations of N and T.

The panel unit root test results using CIPS test are reported in Table 4.3. These tests were estimated with a constant term or with a constant term and trend. We choose to include maximum 5 lags in each regression. We get qualitatively similar findings when the number of lags in the regression is increased. Values of statistics for different number of lags are reported for comparison purposes in order to assess the robustness of the conclusion regarding non-stationarity of the data.

In this case the series were not stationary at the levels. So we made the regression analysis with first differences of the variables. The results of the CIPS test indicate that each series contains a unit root.

Table 4.3. Results of Pesaran's CIPS test

ln beef	Individual effects						Individual effects + individual trend					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
Series in levels												
LAG 1	-1.753	-1.556	-1.485	-1.521	-1.319	-1.210	-2.020	-1.813	-1.747	-1.778	-1.596	-1.502
LAG 2	-1.347	-1.598	-1.091	-1.319	-1.372	-1.063	-1.872	-1.833	-1.638	-2.556**	-1.617	-1.29
LAG 3	-1.424	-1.105	-0.891	-1.328	-0.953	-0.814	-1.951	-1.720	-1.420	-1.853	-2.181**	-1.362
LAG 4	-1.440	-1.219	-1.024	-2.045*	-1.425	-0.758	-2.752**	-1.971	-1.944*	-1.578	-1.855	-1.496
LAG 5	-1.395	-1.346	-0.971	-1.757	-1.152	-1.543						
Series in first differences												
LAG 1	-3.256***	-3.071***	-2.827***	-2.790***	-2.548***	-2.501***						
LAG 2	-3.21*5**	-2.877***	-2.716***	-2.714***	-2.684***	-2.383***						
LAG 3	-3.033***	-2.861***	-2.690***	-2.665***	-2.481***	-2.306**						
LAG 4	-4.013***	-3.057***	-2.842***	-2.848***	-2.752***	-2.535***						
LAG 5	-3.873***	-2.846***	-2.894***	-2.806***	-2.757***	2.664***						
ln pork	Individual effects						Individual effects + individual trend					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
Series in levels												
LAG 1	-2.455	-2.673	-2.690	-2.818	-2.881	-2.891	-2.874	-2.980**	-2.818	-3.042**	-2.955	-3.076*
LAG 2	-2.171	-2.253	-2.545	-2.736	-2.808	-2.782	-2.864	-2.881	-2.870	-2.938	-2.844	-3.215**
LAG 3	2.067	-2.329	-2.381	-2.619	-2.792	-2.670	-2.806	2.654	-2.482	-2.846	-2.691	-2.982*
LAG 4	-2.114	-2.472	-2.413	-2.557	-2.755	-2.706	-2.947**	-2.756	-2.715	-2.929	-2.756	-3.044**
LAG 5	-1.981	-2.290	-2.333	-2.784	-2.692	-2.916	-3.055**	-2.845	-2.524	-2.992*	-2.721	-2.890
Series in first differences												
LAG 1	-5.017***	-5.637***	-5.096***	-5.454***	-5.673***	-5.565***						
LAG 2	-4.868***	-4.937***	-4.673***	-5.024***	-4.881***	-5.290***						
LAG 3	-4.608***	-4.871***	-4.923***	-4.895***	-4.866***	-5.014***						
LAG 4	-4.875***	-5.336***	-4.944***	-5.231***	-5.690***	-5.280***						
LAG 5	-4.717***	-5.476***	-4.833***	-5.084***	-5.322***	-5.417***						

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Null hypothesis: unit root.

The first and second generation panel unit root tests may suffer from significant loss of power if the data contains structural breaks. It is complicated to distinguish between unit root and stationary processes with structural breaks. Today, the panel data unit root tests which allow for structural breaks have received considerable attention among econometricians. Therefore, we extend our analysis by performing Im et al. (2005) panel LM test (ILT), that is an extension of the Lee and Strazicich's (2004) minimum LM test with one structural break, as well as the Im et al. (2010) test that allow for the presence of heterogeneous structural breaks (ILT*) and cross-sectional dependence in the data (ILT_{CA}*). The objective of this analysis is to test unit root hypotheses in the presence of structural change at the unknown time of the break.

In Table 4.4 we present the results of two versions of the panel LM test that is with and without cross-sectional dependence, and perform the tests considering a heterogeneous break only in level and both in level and trend. We employ the one-break version of the panel LM tests of Im et al. (2005) (assuming cross-sectional independence) and Im et al. (2010) (assuming cross-sectional dependence). The lag order selection are chosen using the recursive *t*-statistic procedure with an upper bound of $k_{max}=2$. The null hypothesis of the

LM panel test is that all series contain unit roots, with the alternative that some of the series in the panel are stationary.

The results show that in almost all cases, after taking into account the fact that the two meat price series are subject to a structural break in mean and both in mean and the slope of the series, the null hypothesis of a unit root is not rejected, indicating that these series are not stationary with the presence of a structural break. The previous results from unit root tests without breaks generally indicate that pork and beef price series follow unit root process. Allowing for a structural break do not change the results from non-stationary to stationary. However, it should be noted that if we allow more than one structural break point, the results could be altered.

To make the analysis robust, the results of panel data unit root tests are compared with those obtained with individual unit root tests. Appendix 1 includes LM unit root test results for each individual country assuming one structural breakpoint. We do not consider structural break-dates as being exogenously determined, but we test endogenously for them, i.e. assuming the break-date to be unknown. According to the results in 55 of the 67 pork series and in 49 of the 67 beef series, the unit root null cannot be rejected at the 10% significance level. According to the break locations, there seems to be many important events that have caused significant breaks in the time paths of price series in the EU member states.

Table 4.4. Results of panel LM tests with one structural break

In pork	Level						Level + trend					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
C-S Independence												
ILT	-0.621	-0.375	-0.098	-0.846	-1.310	-1.210						
ILT*	1.574	2.885	0.753	2.267	0.555	1.068	1.075	1.852	0.712	2.086	0.471	0.656
C-S Dependence												
ILT_{CA}*	5.217	4.267	2.982	6.334	1.587	3.014	-3.576**	-3.184**	-2.095	-2.894	-1.648	-1.286
In beef	Level						Level + trend					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
C-S Independence												
ILT	-1.523	-1.222	-1.825	-1.085	-2.368	-1.454						
ILT*	5.852	4.561	6.113	4.219	2.696	3.224	2.005	2.854	3.045	1.328	1.789	1.151
C-S Independence												
ILT_{CA}*	9.652	10.854	7.562	8.881	5.754	6.229	-2.343	-1.975	-2.858	-1.567	-2.641	-2.753

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Null hypothesis: unit root.

We employed a series of panel unit root tests that assume cross sectional independence (the so called first generation panel data unit root tests) and cross-sectional dependence (second generation panel data unit root tests). These tests do not allow for structural breaks. For this reason we completed our analysis by employing the Lagrange Multiplier (LM) panel unit root test with a break. In summary, the results of panel unit root test we have employed indicate quite unanimously that both beef and pork series contain a unit root and are integrated of order one, *I(1)*. Next, we can carry out the second part of the empirical analysis (panel cointegration tests) and check for the existence of a long-run relationship between the price variables.

4.2. Cointegration test results

Using unit root tests we can verify the stationarity of the series. However, empirical questions often concern multivariate relationships; it becomes essential to find out if a particular set of variables is cointegrated. In the time series framework, cointegration refers to the idea that if a set of variables is individually integrated of order one, it is possible that some linear combinations of these variables are stationary. In this case, the vector of slope coefficients is referred to as the cointegrating vector.

Compared to panel unit root tests, the analysis of cointegration in panels is still at an early stage of development although during the recent years there have seen quite big steps forward. So far, the focus of the panel cointegration literature has been on residual-based approaches, although there have been a number of attempts to develop system approaches as well. As is the case for panel unit root tests, panel cointegration tests are based on homogeneous and heterogeneous alternatives. The residual-based tests were developed to ward against the spurious regression problem that can arise in panels when dealing with *I(1)* variables. Kao (1999) and Pedroni (1999, 2004) were among the first to propose residual-based tests for the null hypothesis of no cointegration in cross-sectionally independent panels. However, it has since then become clear that these tests do not work in general, as cross-section dependence is likely to be the rule rather than the exception. In fact, as Gengenbach, Palm, and Urbain (2006) showed the presence of unattended cross-section dependence in the form of non-stationary common factors can actually cause the test statics to diverge as N and T grows. As a response to this, they propose to estimate separately the common and idiosyncratic components of X_{it} and Y_{it} using the principal components method of Bai and Ng (2004), and then to test for cointegration in the resulting component estimates. Banerjee and Carrion-i Silvestre (2006), Westerlund (2007) and Westerlund and Edgerton (2008) propose a similar test but instead of applying the principal components method to X_{it} and Y_{it} directly, they apply it to the residuals of a first-stage regression of Y_{it} onto X_{it}. Cointegration requires that both the common and idiosyncratic components of the residuals are stationary.

Residual based panel cointegration tests result in low power for small samples. In the presence of cross section dependencies, the tests are subject to large size distortions. The situation gets worse if the number of cross sections is increased. To overcome these deficits, panel cointegration tests have been developed that control for the dependencies via a common factor structure (Banerjee and Carrion-i-Silvestre, 2006). The existence of panel cointegration can be taken into account by the second generation panel cointegration tests among of which we employ Westerlund (2007) and Westerlund and Edgerton (2007). They are based on the structural dynamic and they are more powerful when compared with the residual-based panel cointegration tests. The main idea is to test the null hypothesis of no cointegration by inferring whether the error-correction term in the conditional panel error-correction model is equal to zero.

The last cointegration test we perform is Johansen-Fisher based panel cointegration test for vector autoregressive models to panel data. It proceeds by applying the Johansen test to each cross-section unit individually and then combining the marginal significance values for each of these to arrive at a single p-value.

In Table 4.5, we consider residual-based Pedroni test with deterministic intercept and trend. Deterministic time trend has been included to account for cross sectional dependence.

Among the seven Pedroni's panel cointegration tests, four are based on the within dimension (panel cointegration statistics) and the three others on the between dimension (group mean panel cointegration statistics). All tests are based on the null hypothesis of no cointegration for all countries. Under the alternative hypothesis, for the panel statistics, there is cointegration for all countries i. However, the group statistics allow for heterogeneity across countries under the alternative hypothesis.

The estimations are performed in both directions with the dependent and independent variables interchanged to find evidence for the existence of a bi-directional relationship between the pork and beef prices. The majority of Pedroni's test statistics for heterogeneous panels indicate the possibility of a bi-directional cointegrating (or long-run equilibrium) relationship between pork and beef prices. Overall, we can claim that only one among the seven statistics used by the Pedroni test (ie. Panel ν-statistics) does not reject the null hypothesis of no cointegration. The cointegrating relationship seems to be stronger when we use the beef price series as a dependent variable.

Table 4.5. Results of Pedroni residual cointegration tests

	Dep. var. of coint. reg. ln Beef						Dep. var. of coint. reg. ln Pork					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
Panel Cointegration Statistics (Within-Dimension)												
Panel v-statistics	2.262	2.698	1.638	3.747	2.545	2.290	1.954	2.478	1.025	0.885	1.425	1.653
	(0.031)**	(0.011)**	(0.093)*	(0.000)***	(0.016)**	(0.029)**	(0.057)*	(0.012)**	(0.116)	(0.305)	(0.088)*	(0.071)*
Panel PP ρ-statistics	-5.823	-3.875	-3.112	-5.706	-2.435	-2.809	-2.435	-2.603	-2.058	-2.914	-2.311	-2.221
	(0.000)***	(0.000)***	(0.013)**	(0.000)***	(0.021)**	(0.015)**	(0.012)**	(0.008)***	(0.055)*	(0.005)***	(0.010)**	(0.021)**
Panel PP t-statistics	-5.064	-5.888	-4.446	-4.773	-3.283	-2.849	-2.838	-2.031	-3.047	-1.878	-3.002	-2.622
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.002)***	(0.007)***	(0.015)**	(0.083)*	(0.023)**	(0.004)***	(0.020)**	(0.017)**
Panel ADF t-statistics	-10.577	-12.544	-14.886	-11.429	-9.005	-7.029	-6.132	-7.117	-5.027	-8.165	-5.749	-3.819
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.007)***	(0.000)***	(0.071)*	(0.000)***	(0.054)*	(0.125)
Group Mean Panel Cointegration Statistics (Between-Dimension)												
Group PP ρ-statistics	-3.959	-5.711	-2.262	-3.344	-2.748	-1.855	-2.845	-3.559	-2.007	-3.985	-2.360	-1.762
	(0.000)***	(0.000)***	(0.025)**	(0.002)***	(0.011)**	(0.071)*	(0.000)***	(0.000)***	(0.028)**	(0.000)***	(0.025)**	(0.131)
Group PP t-statistics	-4.564	-5.633	-7.947	-6.385	-5.706	-6.91	-2.811	-3.384	-3.875	-6.135	-4.564	-5.828
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.005)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
Group ADF t-statistics	-12.078	-15.125	-14.281	-13.112	-9.179	-10.733	-7.316	-8.304	-6.926	-9.001	-8.078	-6.404
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.009)***
Number of test statistics that reject H0 at 5% level	7/7	7/7	6/7	7/7	7/7	6/7	6/7	6/7	4/7	6/7	5/7	4/7

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. p-values are reported in parentheses. Selection of lags based on SIC. Newey-West bandwidth selection using Bartlett kernel. Null hypothesis: no cointegration.

The Kao (1999) test follows the same basic approach as the Pedroni's tests. Totally, Kao proposed four DF-type panel cointegration tests based on the OLS residuals from the homogeneous panel regression. Kao also uses ADF-type panel cointegration test for cointegration in heterogeneous panel. In this test version, the parameters are allowed to differ across the cross-sections. In our study, we only report Kao's ADF- type test.

In the null hypothesis, the residuals are nonstationary (there is no cointegration) and in the alternative hypothesis, the residuals are stationary (there is a cointegrating relationship among the variables). Table 4.6 reports the results of Kao's residual panel cointegration tests, which reject the null of no cointegration at the 5% significance level in most of the specifications. Findings of Kao's residual cointegration test corroborates the findings of Pedroni's residual cointegration test in Table 4.5.

Both the Pedroni and the Kao tests use the Schwartz Information Criterion (SIC) to automatically select the appropriate lag length. Further, spectral estimation is undertaken by the Bartlett kernel with the bandwidth selected by the Newey-West algorithm. Deterministic time trends are included in all specifications.

Table 4.6. Results of Kao residual cointegrtion test

	Dep. var. of coint. reg. ln Beef						Dep. var. of coint. reg. ln Pork					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
ADF-statistics	-1.642	-2.552	-0.950	-3.251	-1.238	-1.444	-1.322	-3.247	-0.789	-2.661	-1.521	-1.728
	(0.002)***	(0.000)***	(0.041)**	(0.000)***	(0.071)*	(0.025)**	(0.005)***	(0.000)***	(0.066)*	(0.000)***	(0.035)**	(0.002)***

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. p-values are reported in parentheses. Selection of lags based on SIC. Newey-West bandwidth selection using Bartlett kernel. Null hypothesis: no cointegration.

Since we might have cross-sectional dependence in our series, cross-sectional dependence in cointegration vectors is likely. Therefore, we perform the Westerlund (2007) cointegration test with bootstrap under the assumption of cross-sectional dependence.

Westerlund (2007) developed four new panel cointegration tests that are based on structural dynamics and do not impose any common-factor restriction. The idea is to test whether the error-correction term in the panel error-correction model is equal to zero. The four kinds of tests can be divided into two groups by the difference of the alternative hypotheses. The two tests called group-mean tests are designed to test the alternative hypothesis that at least one unit is cointegrated, while the panel tests are designed to test the alternative hypothesis that the panel is cointegrated as a whole. These two tests can be used in both cases of cross-sectional dependency and independence cases. These tests allow also heterogeneity among the units forming the panel.

We adopt the Westerlund (2007) panel cointegration tests selecting the lead and lag orders based on the minimum AIC (Akaike's Information Criterion) and the Bartlett kernel window width. We perform cointegration tests with both a constant and a trend. To take consideration of the cross-sectional dependence we introduce bootstrap into the test to get the robust critical values for the test statistics. The number of replication for the bootstrap is 2000.[15]

The Westerlund's test takes no cointegration as the null hypothesis. Table 4.7 below report the panel cointegration results. All the group mean panel cointegratin statistics and most of the panel cointegration statistics reject the null of no cointegration at the 5% level of significance. Overall, the null hypothesis of no-cointegration is rejected in both asymptotic standard distribution and in bootstrap method. The results suggest that cointegration relationship exists between the series and they are expected to move together in the long run.

Table 4.7. Results of Westerlund's second-generation panel cointegration tests

	Dep. var. of coint. reg. **ln Beef**						Dep. var. of coint. reg. **ln Pork**					
	EU15	EU15	EU25	EU15	EU25	EU27	EU15	EU15	EU25	EU15	EU25	EU27
	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014
Group Mean Panel Cointegration Statistics												
G_τ	-9.211	-12.755	-7.542	-8.221	-13.635	-11.79	-10.461	-9.357	-7.159	-8.446	-10.954	-12.349
asymptotic p-value	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
bootstap p-value	(0.002)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
G_α	-8.928	-10.212	-8.317	-11.426	-15.662	-9.343	-9.523	-9.876	-10.225	-12.313	-12.845	-10.612
asymptotic p-value	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
bootstap p-value	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
Panel Cointegration Statistics												
P_τ	-5.715	-8.145	-4.021	-3.646	-7.722	-7.148	-5.213	-7.858	-4.118	-3.429	-7.058	-6.924
asymptotic p-value	(0.001)***	(0.000)***	(0.002)***	(0.004)***	(0.000)***	(0.002)***	(0.001)***	(0.000)***	(0.002)***	(0.005)***	(0.000)***	(0.003)***
bootstap p-value	(0.007)***	(0.002)***	(0.005)***	(0.010)**	(0.003)***	(0.003)***	(0.008)***	(0.002)***	(0.006)***	(0.010)**	(0.003)***	(0.004)***
P_α	-5.224	-8.457	-3.228	-6.113	-3.832	-4.557	-4.888	-8.371	-3.347	-6.089	-3.665	-4.426
asymptotic p-value	(0.000)***	(0.000)***	(0.000)**	(0.000)***	(0.000)***	(0.000)***	(0.000)**	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
bootstap p-value	(0.045)**	(0.016)**	(0.218)	(0.036)**	(0.184)	(0.056)*	(0.050)*	(0.020)**	(0.192)	(0.038)**	(0.185)	(0.058)*

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Test regression is fitted with a constant and one lag and lead with the kernel bandwidth being set according to the rule $4(T/100)^{2/9}$. p-values are reported in parentheses. The p-values are for a one-sided test based on 2000 bootstrap replications. Null hypothesis: no cointegration.

In order to assess the robustness of our findings, we also implemented the panel cointegration test proposed by Westerlund and Edgerton (2007). Unlike the other panel data cointegration tests, here the null hypothesis is now cointegration. This test relies on the popular Lagrange multiplier test of McCoskey and Kao (1998), and permits correlation to be accommodated both within and between the individual cross-sectional units. The test is also robust to unknown heterogeneous breaks in the intercept and/or the slope of the cointegrating regression. In addition, the bootstrap suggested by Westerlund and Edgerton (2007) is based on the sieve-sampling scheme[16], and has the appealing advantage of significantly reducing the distortions of the asymptotic test.

The results reported in Table 4.8 for a model including a constant and a trend clearly indicate the cointegrating relationship between pork and beef prices in the EU since the null hypothesis of cointegration is always accepted. This result is not modified if one refers to the bootstrap critical values compared to asymptotic ones. Results also confirm that the cointegrating relationship is robust to potential level and regime shifts.

Table 4.8. Results of second-generation panel cointegration test proposed by Westerlund and Edgerton

	Dep. var. of coint. reg. **ln Beef**						Dep. var. of coint. reg. **ln Pork**					
	EU15	EU15	EU25	EU15	EU25	EU27	EU15	EU15	EU25	EU15	EU25	EU27
	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014	2/1995-6/2014	1/2005-6/2014	1/2005-6/2014	1/2007-6/2014	1/2007-6/2014	1/2007-6/2014
LM-stat	1.054	1.347	1.548	2.012	1.758	1.964	2.042	2.187	1.838	2.573	2.225	2.349
asymptotic p-value	(0.515)	(0.386)	(0.211)	(0.251)	(0.187)	(0.198)	(0.335)	(0.309)	(0.227)	(0.264)	(0.349)	(0.376)
bootstap p-value	(0.319)	(0.204)	(0.251)	(0.128)	(0.157)	(0.141)	(0.102)	(0.082)	(0.126)	(0.066)	(0.116)	(0.084)

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Test regression is fitted with a constant and one lag and lead with the kernel bandwidth being set according to the rule $4(T/100)^{2/9}$. p-values are reported in parentheses. The p-values are for a one-sided test based on 2000 bootstrap replications. The Westerlund and Edgerton (2007) tests are performed using the Stata "xtwest" command. Null hypothesis: cointegration.

[15] We then used Stata command -xtwest- to test for cointegration.
[16] Sampling method proposed by Rietveld (1978), a simple and practical strategy for selecting line items with PPS (probability to proportional to line item size).

The last panel cointegration test applied is the Johansen-Fisher type panel cointegration test developed by Maddala and Wu (1999). They use Fisher's result to propose an alternative approach to test for cointegration in panel data by combining tests from individual cross-sections to obtain a test statistic for the full panel. The Johansen-Fisher panel cointegration test is a panel version of the individual Johansen cointegration test. The Johansen-Fisher panel cointegration test is based on the aggregates of the p-values of the individual Johansen maximum eigenvalues and trace statistic.

The results from Johansen-Fisher panel cointegration test are presented in Table 4.9. We used the Akaike Information Criterion (AIC) and the Schwarz Information Criterion (SIC) to determine the optimal lag length. We perform the cointegration test with both a constant without trend and a constant with trend.

Based on the maximal eigenvalue statistics, we reject the null hypothesis of no cointegrating relationship. Results from the trace statistics also support the finding of maximal eigenvalue statistics where there is rejection towards null hypothesis of $r=0$ and $r\leq 1$, respectively. In a nutshell, based on the statistical results of the Johansen-Fisher panel cointegration test, there is sufficient evidence to conclude the existence of the long run relationship between the pork and beef prices.

Table 4.9. Results from the Johansen-Fisher panel cointegration test

	Individual effects						Individual effects + individual trend					
	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014	EU15 2/1995-6/2014	EU15 1/2005-6/2014	EU25 1/2005-6/2014	EU15 1/2007-6/2014	EU25 1/2007-6/2014	EU27 1/2007-6/2014
Maximum eigenvalue												
r=0	772.135	683.428	597.194	802.645	753.461	628.107	882.336	752.671	651.544	885.658	795.190	664.506
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
r≤1	456.274	351.927	405.618	527.946	315.655	379.509	511.683	426.377	332.961	495.842	436.813	398.770
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
Trace test												
r=0	746.353	633.586	546.217	692.590	705.644	342.818	852.643	729.526	597.457	835.281	369.509	375.477
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***
r≤1	456.235	351.918	405.672	527.946	315.618	379.508	852.659	729.547	597.419	835.293	369.528	375.465
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. p-values are reported in parentheses. Selection of lags based on AIC and SIC. Null hypothesis: no cointegration.

In summary, the results of panel cointegration tests are generally highly significant which gives strong evidence that the variables have a long run relationship.

Based on the maximal eigenvalue statistics, we reject the null hypothesis of no cointegrating relationship. Results from the trace statistics also support the finding of maximal eigenvalue statistics where there is rejection towards null hypothesis of $r=0$ and $r\leq 1$, respectively. In a nutshell, based on the statistical results of the Johansen-Fisher panel cointegration test, there is sufficient evidence to conclude the existence of the long run relationship between the pork and beef prices.

In summary, the results of panel cointegration tests are generally highly significant which gives strong evidence that the variables have a long run relationship.

4.3. Estimation results

The cointegration tests are only able to indicate whether or not the variables are cointegrated and if a long-run relationship exists between them. Since they do not indicate the direction of causality, the long-run equilibrium coefficients should be estimated by using appropriate estimators.

When panels of data are available, there exist a number of alternative estimation methods that vary on the extent to which they account for parameter heterogeneity. The dynamics of traditional estimators is simply pooled and treated as homogeneous. Only the intercepts are allowed to differ cross countries. Early and prominent examples include fixed effects (E), random effects (RE), and generalized methods of moments (GMM). These methods are typically focused on solving the problem of fixed effect heterogeneity in the case of large N and small T panels; whereas they are not designed to correct for the endogeneity induced by the latent heterogeneity. Pesaran and Smith (1995) show that the traditional procedures for the estimation of pooled models can produce inconsistent and potentially misleading estimates of the lagged dependent variable's parameter in dynamic panel data models if latent heterogeneity is present.

In this study, we utilize panel-based vector error correction models (VECM) including MG (Mean Group) and PMG (Pooled Mean Group) estimation methods. The MG estimator allows for complete (short-run and long-run) parameter heterogeneity across panel cross-sections. The PMG estimator is an intermediate case between the averaging and pooling methods of estimations.[17]

The first technique, MG, introduced by Pesaran and Smith, (1995) calls for estimating separate regressions for each country and calculating the coefficients as unweighted means of the estimated coefficients for the individual countries. It allows for all coefficients to vary and be heterogeneous in the long-run and short-run. Pesaran and Smith showed that the MG method produces consistent estimates of the average of the parameters when the time-series dimension of the data is sufficiently large. However, for small N the average estimators (MG) in this approach are quite sensitive to outliers and small model permutations (Favara, 2003).

The MG estimator does not take into account that some economic conditions tend to be common across countries in the long run. The PMG estimator does because it combines both pooling and averaging. The long run coefficients are constrained to be the same across sections, while the intercepts, short run coefficients, and error variances are allowed to differ. The PMG estimator is believed to offer the best compromise between consistency and efficiency, because one would expect the long-term growth path to be driven by a similar process across EU member states while the short-term dynamics around the long-term equilibrium path differ because of idiosyncratic member state features and shocks to fundamentals. The PMG estimator is appropriate when data have complex country-specific short-term dynamics which cannot be captured imposing the same lag structure on all countries. This estimator combines the properties of efficiency of the pooled dynamic estimators while avoiding the inconsistency problem deriving from slope heterogeneity. Moreover, since the PMG estimator does not impose any restriction on short–term coefficients, it provides important information on country specific values of the speed of convergence towards the long-run relationship linking pork and beef price series.

The variables are *I(1)* and cointegrated, which means that the error term is *I(0)* process for all i. A principal feature of cointegrated variables is their responsiveness to any deviation from long-run equilibrium. This feature implies an error correction model in which the short-run dynamics of the variables in the system are influenced by the deviation from equilibrium. Since the model is nonlinear in the parameters, Pesaran, Shin, and Smith (1999) develop a maximum likelihood method to estimate the parameters.

The error-correction model to be estimated is given by the following equations:

$$\Delta lnbeef_{it} = \varphi\left(lnbeef_{i,t-1} - \theta_i lnpork_{it}\right) + \sum_{j=1}^{p-1} \gamma_{ij}\, \Delta lnbeef_{i,t-1} + \sum_{j=0}^{q-1} \delta_{ij}\, \Delta lnpork_{i,t-j} + \mu_i + \epsilon_{it} \quad (4.1)$$

$$\Delta lnpork_{it} = \varphi'\left(lnpork_{i,t-1} - \theta'_i lnbeef_{it}\right) + \sum_{j=1}^{p-1} \gamma'_{ij}\, \Delta lnpork_{i,t-1} + \sum_{j=0}^{q-1} \delta'_{ij}\, \Delta lnbeef_{i,t-j} + \mu'_i + \epsilon'_{it} \quad (4.2)$$

where the number of groups $i=1,2,\dots,N$ and the number of periods $t=1,2,\dots,T$. δ_{ij} are the coefficient vectors, γ_{ij} are scalars and μ_i is the group specific effect. θ_i is the vector which contains the long-run relationships between the variables and the parameter φ_i is the error-correcting speed of adjustment term. The long run coefficients are estimated using the Maximum Likelihood (ML) estimation

[17] OLS estimators are super-consistent in the case of co-integrated variables, but they are based on strong homogeneity assumptions among countries by imposing single slope coefficient in pooled estimation, which is inappropriate for this study regarding potential country heterogeneity. This is the reason for using MG and PMG estimators instead of traditional panel techniques.

The long-run relationship imposes under the null hypothesis the condition*:* $\varphi_i = 0$ for all i. This hypothesis means that there is no long-run stable relationship between the independent variable and the dependent variable in the model. The decision rule says that when the error correction term (ECT) is negative and significant, the null hypothesis of no causality would be rejected.

ECT is a measure of the extent by which the observed values in time *t-1* deviate from the long-run equilibrium relationship. Since the variables are cointegrated, any such deviation at time *t-1* should induce changes in the values of the variables in the next time point, in an attempt to force the variables back to the long-run equilibrium relationship.

In practice, the MG and PMG procedure involves first estimating autoregressive distributed lag (ARDL) models separately for each country *i*. In a series of papers, Pesaran and Smith (1995), Pesaran (1997), and Pesaran and Shin (1999) show that one can use the ARDL approach to produce consistent and efficient estimates of the parameters in a long-run relationship between both integrated and stationary variables, and to conduct inference on these parameters using standard tests. The main requirements for the validity of this methodology are that, first, there exists a long-run relationship among the variables of interest and, second, the dynamic specification of the model is sufficiently augmented so that the regressors become weakly exogenous and the resulting residual is serially uncorrelated[18].

Tables 4.10 and 4.11 report MG and PMG estimates of the ECM. We have already revealed that there is bi-directional relationship between pork and beef prices in the EU. Thus, the dependent variable in the estimation results of Table 4.10 is *lnbeef* and correspondingly the dependent variable is *lnpork* in Table 4.11 results. Lags are chosen on the basis of AIC and are allowed to vary across countries. We report results both with and without trend. A constant country-specific term is included. Natural logarithms are used in order to obtain directly the elasticities.

We analyze the stability of our results over time by re-estimating the models for EU15 countries in two sub periods (1/2005-6/2014 and 1/2007-6/2014) and for EU25 countries in one subperiod (1/2007-6/2014). This is also a check for cross-sectional stability: we can compare if new member state distort the estimation results.

Table 4.10. MG and PMG regression results, dependent variable: lnbeef

MG	EU15 2/1995-6/2014		EU15 1/2005-6/2014		EU25 1/2005-6/2014		EU15 1/2007-6/2014		EU25 1/2007-6/2014		EU27 1/2007-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run												
ln pork	0.251	0.253	0.282	0.288	0.213	0.217	0.423	0.426	0.409	0.413	0.392	0.394
	(10.385)***	(12.127)***	(16.291)***	(18.754)***	(6.372)***	(6.479)***	(29.490)***	(30.572)***	(27.477)***	(28.582)***	(25.115)***	(25.626)***
time trend	-0.001		0.002		-0.002		0.001		-0.002		-0.001	
	(-2.714)**		(2.217)**		(-2.456)**		(1.694)		(-1.880)		(-1.755)	
Short Run												
ECT *(adj. speed)*	-0.040	-0.042	-0.061	-0.065	-0.055	-0.057	-0.098	-0.099	-0.085	-0.085	-0.101	-0.105
	(-11.118)***	(-13.546)***	(-15.643)***	(-17.119)***	(-14.554)***	(-15.725)***	(-20.439)***	(-21.477)***	(-16.986)***	(-18.442)***	(-22.971)***	(-24.811)***
Δ ln pork	-0.003	-0.003	-0.010	-0.011	0.006	0.007	0.050	0.051	0.061	0.062	0.036	0.038
	(-1.828)*	(-1.859)*	(-1.231)	(-1.261)	(2.290)**	(2.429)**	(3.154)**	(3.157)**	(3.793)***	(3.846)***	(2.208)**	(2.211)**
Δ ln $pork_{t-1}$	-0.001	-0.001	-0.002	-0.002	-0.001	-0.001	0.015	0.017	0.023	0.023	0.009	0.009
	(-0.944)	(-0.954)	(-1.107)	(-1.136)	(-0.753)	(-0.761)	(2.685)**	(2.682)**	(2.508)**	(2.518)**	(1.957)*	(1.974)*
Δ ln $pork_{t-2}$	0.001	0.001	0.002	0.003	0.001	0.001	0.001	0.001	0.002	0.003	0.001	0.001
	(0.409)	(0.401)	(0.807)	(0.873)	(0.545)	(0.555)	(1.224)	(1.239)	(1.587)	(1.626)	(0.371)	(0.390)
Δ ln $beef_{t-1}$	0.617	0.621	0.692	0.695	0.448	0.451	0.523	0.524	0.438	0.438	0.644	0.655
	(3.056)**	(3.172)**	(3.408)**	(3.422)**	(2.637)**	(2.676)**	(3.665)***	(3.712)***	(3.355)***	(3.363)***	(5.654)***	(5.690)***
Δ ln $beef_{t-2}$	0.005	0.005	0.006	0.007	0.004	0.005	0.002	0.002	0.007	0.008	0.005	0.005
	(1.023)	(1.089)	(1.189)	(1.229)	(1.061)	(1.130)	(1.003)	(1.018)	(0.726)	(0.764)	(0.808)	(0.810)
Observations	3315	3315	1530	1530	2550	2550	1170	1170	1950	1950	2106	2106
Log likelihood	-514.654	-555.265	-328.647	-344.380	-275.581	-302.088	-189.618	-171.282	-148.723	-159.335	196.736	-219.481
Hausman test	0.221	0.195	0.089	0.102	0.242	0.255	0.075	0.071	0.152	0.153	0.179	0.185
	(0.640)	(0.628)	(0.778)	(0.725)	(0.562)	(0.528)	(0.898)	(0.910)	(0.679)	(0.676)	(0.642)	(0.659)

PMG	EU15 2/1995-6/2014		EU15 1/2005-6/2014		EU25 1/2005-6/2014		EU15 1/2007-6/2014		EU25 1/2007-6/2014		EU27 1/2007-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run												
ln pork	0.193	0.197	0.215	0.219	0.147	0.155	0.315	0.331	0.281	0.308	0.301	0.310
	(16.476)***	(17.554)***	(19.889)***	(21.216)***	(10.124)***	(12.397)***	(33.262)***	(36.661)***	(32.910)***	(35.090)***	(29.431)***	(33.995)***
time trend	-0.002		-0.001		-0.001		-0.002		-0.001		-0.003	
	(-2.113)*		(-1.686)		(-1.958)*		(-1.224)		(-1.517)		(-2.437)**	
Short Run												
ECT *(adj. speed)*	-0.035	-0.038	-0.057	-0.06	-0.047	-0.054	-0.086	-0.093	-0.071	-0.073	-0.082	-0.088
	(-8.530)***	(-9.158)***	(-12.227)***	(-12.972)***	(-13.114)***	(-13.785)***	(-17.218)***	(-19.076)***	(-14.962)***	(-15.921)***	(-22.124)***	(-21.850)***
Δ ln pork	-0.002	-0.003	-0.017	-0.019	-0.005	-0.005	-0.037	-0.047	-0.072	-0.074	-0.055	-0.068
	(-1.485)	(-1.556)	(-1.237)	(-1.281)	(-2.028)*	(-2.123)*	(-2.173)*	(-2.445)**	(-3.906)***	(-3.949)***	(-2.682)**	(-3.019)***
Δ ln $pork_{t-1}$	-0.001	-0.001	-0.001	-0.002	0.004	0.005	0.008	0.010	0.007	0.007	0.003	0.005
	(-0.662)	(-0.733)	(-0.940)	(-1.019)	(1.750)	(1.828)	(1.938)*	(2.181)*	(-1.874)	(1.955)*	(1.478)	(1.610)
Δ ln $pork_{t-2}$	-0.001	-0.002	0.002	0.0032	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001
	(-0.274)	(0.359)	(0.652)	(0.695)	(-0.286)	(-0.334)	(-0.655)	(-0.680)	(-0.717)	(-0.771)	(-0.423)	(-0.467)
Δ ln $beef_{t-1}$	0.549	0.555	0.576	0.606	0.367	0.372	0.432	0.444	0.608	0.615	0.584	0.588
	(4.263)***	(4.407)***	(2.556)**	(2.752)**	(2.437)**	(2.481)**	(2.981)**	(3.117)**	(4.258)***	(4.336)***	(3.715)***	(3.743)***
Δ ln $beef_{t-2}$	0.002	0.002	0.004	0.005	0.002	0.002	0.001	0.003	0.003	0.003	0.002	0.004
	(0.686)	(0.837)	(0.954)	(0.981)	(0.930)	(0.975)	(1.566)	(1.618)	(1.228)	(1.262)	(0.715)	(0.734)
Observations	3315	3315	1530	1530	2550	2550	11700	1170	1950	1950	2106	2106
Log likelihood	-589.244	-525.467	-369.413	-391.582	-337.911	-315.390	-173.930	-178.510	-156.977	-157.65	-239.292	-246.629
Hausman test												

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Automatic selection of lags based on Akaike information criterion. t-statistic are reported in parentheses. The Hausman test is a test of poolability of the long-run coefficient (i.e. of the restriction that all countries have the same long-run elasticity). All equations include a constant country-specific term.

[18] Augmenting the model with lags addresses the potential endogeneity of remittances. In this respect, Pesaran (1997) and Pesaran and Shin (1999) show that for inference on the long-run parameters, sufficient augmentation of the order of the autoregressive distributive lag model can simultaneously correct for the problem of residual serial correlation and endogenous regressors.

Table 4.11. MG and PMG regression results, dependent variable: lnpork

	MG											
	EU15		EU15		EU25		EU15		EU25		EU27	
	2/1995-6/2014		1/2005-6/2014		1/2005-6/2014		1/2007-6/2014		1/2007-6/2014		1/2007-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run												
ln beef	0.091	0.092	0.108	0.114	0.067	0.071	0.194	0.196	0.136	0.142	0.217	0.222
	(1.790)	(1.848)	(1.736)	(1.812)	(1.304)	(1.516)	(3.775)***	(3.888)***	(2.746)**	(2.851)**	(3.882)***	(4.219)***
time trend	-0.001		-0.002		-0.002		-0.001		-0.001		-0.003	
	(-1.462)		(-1.276)		(-1.637)		(-1.446)		(-1.285)		(-2.483)**	
Short Run												
ECT *(adj. speed)*	-0.024	-0.026	-0.045	-0.048	-0.033	-0.037	-0.053	-0.059	-0.072	-0.075	-0.081	-0.082
	(-1.480)	(-1.507)	(-2.726)**	(-2.812)**	(-1.619)	(-1.719)	(-4.120)***	(-4.275)***	(-6.733)***	(-6.846)***	(-8.191)***	(-8.219)***
Δ ln beef	0.001	0.001	0.002	0.002	0.003	0.004	0.023	0.025	0.027	0.027	0.028	0.028
	(0.692)	(0.744)	(1.012)	(1.078)	(1.599)	(1.623)	(2.880)**	(2.919)**	(3.026)**	(3.047)**	(3.219)**	(3.248)**
Δ ln $beef_{t-1}$	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.008	-0.009	-0.01	-0.011	-0.004	-0.005
	(-0.755)	(-0.785)	(-0.567)	(-0.582)	(-0.753)	(-0.760)	(-1.636)	(-1.648)	(-2.031)*	(-2.114)*	(-1.617)	(-1.646)
Δ ln $beef_{t-2}$	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.002	0.002	0.002	0.002
	(0.367)	(0.380)	(0.513)	(0.552)	(0.629)	(0.655)	(0.838)	(0.889)	(1.412)	(1.420)	(1.488)	(1.495)
Δ ln $pork_{t-1}$	0.559	0.567	0.537	0.542	0.396	0.406	0.471	0.481	0.522	0.529	0.492	0.497
	(5.539)***	(5.563)***	(6.076)***	(6.137)***	(4.164)**	(4.242)**	(4.418)***	(4.555)***	(5.418)***	(5.476)***	(4.714)***	(4.661)***
Δ ln $pork_{t-2}$	0.003	0.003	0.004	0.003	0.001	0.003	0.002	0.002	0.005	0.006	0.006	0.006
	(0.664)	(0.734)	(1.340)	(1.253)	(0.417)	(0.936)	(0.705)	(0.818)	(1.418)	(1.567)	(1.645)	(1.681)
Observations	3315	3315	1530	1530	2550	2550	1170	1170	1950	1950	2106	2106
Log likelihood	-396.751	-411.259	-252.674	-270.125	-198.647	-213.346	-145.782	-161.838	-140.238	-158.639	-201.961	-222.377
Hausman test	0.071	0.075	0.042	0.043	0.035	0.038	0.055	0.059	0.092	0.095	0.101	0.105
	(0.910)	(0.902)	(0.958)	(0.955)	(0.972)	(0.968)	(0.929)	(0.922)	(0.795)	(0.783)	(0.730)	(0.722)

	PMG											
	EU15		EU15		EU25		EU15		EU25		EU27	
	2/1995-6/2014		1/2005-6/2014		1/2005-6/2014		1/2007-6/2014		1/2007-6/2014		1/2007-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run												
ln beef	0.066	0.067	0.08	0.083	0.042	0.046	0.119	0.122	0.099	0.103	0.105	0.108
	(1.576)	(1.617)	(1.655)	(1.690)	(1.077)	(1.145)	(5.361)***	(5.552)***	(3.183)**	(3.652)***	(5.166)***	(5.648)***
time trend	-0.003		-0.003		-0.002		-0.002		-0.004		-0.006	
	(-1.919)*		(-2.027)*		(-1.954)*		(-1.228)		(-2.734)**		(-4.203)**	
Short Run												
ECT *(adj. speed)*	-0.02	-0.021	-0.038	-0.041	-0.027	-0.030	-0.051	-0.054	-0.066	-0.067	-0.074	-0.076
	(-1.223)	(-1.254)	(-1.648)	(-1.706)	(-1.565)	(-1.587)	(-2.380)**	(-2.582)**	(-2.828)**	(-2.901)**	(-5.317)***	(-5.356)***
Δ ln beef	0.001	0.001	0.003	0.004	0.005	0.006	0.005	0.005	0.006	0.007	0.007	0.008
	(0.490)	(0.534)	(0.873)	(0.919)	(1.727)	(1.774)	(1.828)	(1.830)	(1.285)	(1.366)	(1.469)	(1.590)
Δ ln $beef_{t-1}$	-0.001	-0.001	-0.001	-0.002	-0.002	-0.003	-0.005	-0.007	-0.008	-0.009	-0.007	-0.007
	(-0.667)	(-0.738)	(-0.944)	(-1.010)	(-1.759)	(-1.828)	(-1.953)*	(-2.007)*	(-2.186)*	(-2.224)*	(-1.955)*	(-1.939)*
Δ ln $beef_{t-2}$	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.002	0.003	0.003
	(0.276)	(0.276)	(0.337)	(0.354)	(0.591)	(0.608)	(0.963)	(0.947)	(1.174)	(1.227)	(1.662)	(1.690)
Δ ln $pork_{t-1}$	0.518	0.520	0.555	0.589	0.351	0.354	0.454	0.462	0.479	0.483	0.467	0.476
	(6.030)***	(6.105)***	(5.929)***	(5.959)***	(4.818)***	(4.880)***	(5.243)***	(5.316)***	(5.697)***	(5.731)***	(4.619)***	(4.641***)
Δ ln $pork_{t-2}$	0.002	0.002	0.003	0.003	0.002	0.002	0.001	0.002	0.003	0.004	0.003	0.004
	(0.473)	(0.528)	(0.716)	(0.780)	(0.631)	(0.742)	(0.238)	(0.553)	(0.717)	(0.914)	(1.115)	(1.165)
Observations	3315	3315	1530	1530	2550	2550	11700	1170	1950	1950	2106	2106
Log likelihood	-410.693	-433.584	-295.642	-316.198	-337.525	-354.612	-168.767	-173.973	-155.951	-171.360	-207.691	-219.634
Hausman test												

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Automatic selection of lags based on Akaike information criterion. t-statistic are reported in parentheses. The Hausman test is a test of poolability of the long-run coefficient (i.e. of the restriction that all countries have the same long-run elasticity). All equations include a constant country-specific term.

Overall, our baseline results in Table 4.10 seem to be relatively robust: there is long run equilibrium between pork and beef prices. Pork price is a significant factor affecting beef price in the EU in long run. Since variables are defined in logarithmic terms, the estimated coefficients are directly the price elasticities of beef price with respect to pork price. All MG and PMG regressions results in Table 4.10 show a significant and positive long run coefficient. On the contrary, the magnitude of the long-term coefficient is affected by the estimation method. By choosing the MG rather than the PMG method slightly higher values are obtained. We employ the Hausman test of long-run homogeneity restriction to choose the appropriate estimator. Namely, homogeneity of long run coefficients implied by PMG estimating procedure cannot be assumed a priori but needs to be tested. When long-run homogeneity restriction can be accepted, PMG estimates would be more efficient compared to MG. Respectively, when the slope coefficients are heterogeneous, then PMG estimates would be inconsistent and the MG estimator provides consistent estimates of the mean of long-run coefficients. According to test results shown at the bottom of Table 4.10, we cannot reject the null of long-run homogeneity restriction, implying that the PMG estimator is efficient under the null hypothesis and is preferred over the MG estimator.

We can now concentrate on the estimation results of PMG estimation method. The PMG estimates find strong evidence of positive relationship between the EU pork and beef prices. According to the results, there exists cross-commodity price transmission from pork prices to beef prices in the EU. All PMG regressions results in Table 4.11 show a significant and positive coefficient of 0.15 to 0.33 in the long run, implying that a 1% increase in pork prices, ceteris paribus, would raise beef prices 0.15-0.33%. According to the results of PMG estimation the relationship between the two price series is the strongest and, thus, the price transmission is the most significant in the old member states, EU15 countries. Also, in the case of EU15, we notice that the price transmission is more apparent and the coefficients are higher when we investigate estimation results from subperiod 1/2007-6/2014 compared to estimation results from subperiod 1/2005-6/2014 or from the full period 2/1995-6/2014. This indicates that integration of the EU livestock market of pork and beef has increased during the investigation period. This is also valid when analyzing price transmission and market integration between pork and beef in the EU25 countries: the significance of long run coefficient is higher and the long run price elasticity measured by the estimated long run coefficient is larger in estimation results from subperiod 1/2007-6/2014 than in estimation results from the full period 1/2005-6/2014.

The estimated error-correcting speed adjustment term ECT is, as expected, significantly negative and less than 1 in absolute value, implying convergence towards long run equilibrium relationship. The lowest value of the error correction coefficient in the PMG estimation results is 0.035 and the highest is 0.093, implying a speed of adjustment of about 1-2 years. The speed of adjustment has also improved during the investigation period; In the case of EU15 countries, the adjustment term is 0.056 (trend)/0.060 (no trend) in the estimation results from the period 1/2005-6/2014 and 0.086 (trend) 0.093 (no trend) in the results from the period

1/2007-6/2014. This means that after the year 2005 the convergence to the equilibrium has sped up from 1.5 years to 1 year. Price transmission between the EU meat markets has fastened significantly but is still quite slow. This indicates that there exist some rigidities that decelerates the price transmission process.

Furthermore, we can note that in short run the variation in pork prices has affected beef prices only after 2007. This means that short run relationship between pork and beef prices has only existed since 2007. Moreover, short term price transmission from pork price to beef price in the EU is interestingly negative according to our PMG estimation results. This means that increase in pork price lower the beef price in short term. However, this short term effect is minor although statistically significant.

By repeating the estimation of the model for EU15 in period 1/2005-6/2014 and for EU15 and EU25 in period 1/2007-6/2014 we also check the sensitivity of the results at the type of countries included in the panel; i.e. we check what kind of influence the new member states have had on the results. We notice that the panel estimates indicate cross-sectional stability over results. The inclusion of the new member states to the data has not changed the interpretation of the estimation results.

In table 4.11 we repeat the estimation but now we specify pork as a dependent variable and beef as an independent variable. Consequently, we are interested in the price elasticities of pork with respect to beef price. The long-run slope homogeneity hypothesis of PMG is tested via the Hausman test. On the basis of the Hausman test it is not possible to reject the hypothesis of poolability of the long-run elasticity of the beef price and we can therefore focus on the results of PMG estimation method. All in all, MG estimation method provides larger coefficients than PMG method. According to the results of PMG estimation, there exists cross-commodity price transmission in long run from beef prices to pork prices in the EU only in the estimation results from subperiod 1/2007-6/2014. This can be interpreted as overall increased integration and efficiency in livestock market in the EU during the past ten years; not only the variation of pork price will transmit to the prices of other meat types but also vice versa. There is bi-directional relationship between pork and beef prices in the EU in long run although price transmission coming from the fluctuations in pork price is larger and more significant. Notably, this bi-directional relationship has not been valid until 2007 in our investigation. In table 4.11 the estimation results from subperiod 1/2007-6/2014 show a significant and positive coefficient of 0.099 to 0.122 in the long run, implying that a 1% increase in beef price ceteris paribus, would raise pork price 0.1-0.12% in the EU.

The error correction terms are negative and significant; hence the null hypothesis of no long-run relationship is rejected. According to the results of PMG estimation, the speed of adjustment from the deviation in the long run relationship between the beef and pork prices is between -0.051 and -0.076. The model implies moderate adjustment inertia; it converges to the equilibrium, with 5-7.5% percent of discrepancy corrected in each period. No significant price transmission from beef price to pork price is found in short run in the EU livestock market. Overall, this suggests that short-run dynamics is significantly different in livestock market compared to long run dynamics.

4.4. Robustness check against estimation method

Next, we use alternative heterogeneous panel cointegration techniques that correct for endogeneity, namely Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS). We use FMOLS and DOLS between-dimension estimators (Group Mean Estimator) proposed by Pedroni (2001). Unlike the PMG, which uses a maximum likelihood method, FMOLS and DOLS are based on a modified OLS[19]. FMOLS and DOLS estimators allow us to relax the assumption of long-run homogeneity. Both methods allow for regressors' complete endogeneity, and treat all parameters, i.e. dynamics and cointegrating vectors, as heterogeneous across panel members. FMOLS runs a static OLS with fixed effects for each panel member individually and uses the

[19] It is worth noting that under endogeneity and no cointegration, using OLS produces first-order bias, and external instruments are needed. However, under endogeneity and cointegration**,** OLS produces superconsistent estimates and second-order endogeneity bias (i.e. inconsistent estimates of standard errors). Internal instruments are then used (such as in FMOLS and DOLS).

estimated residuals to (non-parametrically) build member-specific adjustment terms, which are then used to correct for each member's endogeneity. Differences in regressors are used as internal instruments. Instead, DOLS individually corrects for endogeneity parametrically by running OLS with fixed effects for each panel member including leads and lags of differenced regressors.

It is worth to note that FMOLS and DOLS are static models and thus, we can only estimate long term relationship between pork price and beef price in the EU member states.

The DOLS estimates employ two lags and two leads. Overall, the results in Table 4.12 are robust with respect to the choice of the lag structure. A country-specific constant has been incorporated while a time trend was not included.

Table 4.12. FMOLS and DOLS regression results

	EU15 2/1995-6/2014		EU15 1/2005-6/2014		EU25 1/2005-6/2014		EU15 1/2007-6/2014		EU25 1/2007-6/2014		EU27 1/2007-6/2014	
	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS	FMOLS	DOLS
Independent variable: ln pork												
ln beef	0.164	0.203	0.149	0.165	0.132	0.172	0.333	0.361	0.311	0.357	0.375	0.390
	(7.453)***	(9.048)***	(6.761)***	(8.819)***	(6.520)***	(7.476)***	(13.062)***	(14.281)***	(12.373)***	(12.982)***	(15.207)***	(15.916)***
Observations	3315	3315	1530	1530	2550	2550	1170	1170	1950	1950	2106	2106
Independent variable: ln beef												
ln pork	0.109	0.131	0.127	0.156	0.119	0.123	0.288	0.312	0.189	0.216	0.231	0.268
	(2.459)**	(3.109)***	(3.061)***	(3.795)***	(2.227)***	(2.258)***	(7.136)***	(7.788)***	(5.161)***	(5.907)***	(6.344)***	(6.765)***
Observations	3315	3315	1530	1530	2550	2550	1170	1170	1950	1950	2106	2106

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. t-statistic are reported in parentheses.

These results of DOLS and FMOLS estimates confirm the existence of a long-run relationship between pork and beef prices in the EU countries. All the long run coefficients obtained by applying FMOLS and DOLS have the same signs although higher magnitudes as the previous PMG results, suggesting equilibrium derived from the PMG method are roughly equal to those from FMOLS and DOLS. The results from FMOLS and DOLS estimation also confirm the observation that price transmission and market integration has increased between pork market and beef market in the EU during the observation period. In addition these estimation results confirm the cross-sectional stability: results are not alert the type of countries included in the panel (EU15, EU25 or EU27).

4.5. Robustness check against definitions of the data sample

In this section, the robustness of the relation between pork and beef price series is checked against alternative definitions of the data sample. We address the following questions. How is the relation changed over time? Are there countries with a significantly different behavior?

We first check the stability of our results over time via PMG estimation. The entire estimation period of 2/1995-6/2014 is divided into 4 subperiods (2/1995-12/1999, 1/2000-12/2004, 1/2005-12/2009, 1/2010-6/214) and estimation is repeated for every subperiod including EU15 member states. These estimation results are presented in Appendix 2.

The estimation results in Tables 4.10 and 4.11 already revealed that integration in the EU livestock market of pork and beef has increased during the investigation period. The purpose of this robustness check is to get a better picture of this development and also examine the meat market integration and cross-commodity price transmission between pork and beef during the early stage of the EU.

The estimation results show that the long-term elasticity and the speed adjustment term have changed substantially between the subperiods considered. Overall, estimations suggest a significantly different and higher relation between price series for the latter subperiods. We can observe that EU meat market was diverged before year 2005. It cannot be found statistically significant relation between pork and beef prices during the periods 2/1995-12/1999 and 1/2000-12/2004. Since 2005, however, EU meat market integration has increased sharply and price transmission between meat markets has become significant. This is true in both cases: when we have *lnpork* as an independent variable and when we have *lnbeef* as an independent variable. This is clear evidence that meat market in Europe is well functioned and efficient and the direction of the EU livestock market is towards stronger integration.

Next, we proceed by studying robustness of our results across countries. The findings from our estimation might be affected also by the relative small number of countries in the data sample. As a further robustness check we have re-estimated the model via PMG estimation method excluding from the data one country at a time. This permits to understand whether the results are strongly driven the behavior of a single country. Figures in Appendix 3 plot the value of the long-run elasticity and speed adjustment term on the country excluded from the sample.

According to the data availability, there are figures for EU15 member states, EU25 member states (excl. EU15) and EU27 member states (excl. EU25). In the first set of figures independent variable is *lnbeef*. In the latter set *lnpork* is treated as an independent variable.

In the first figures, EU15 countries that appear to influence significantly on the estimation of the long-run elasticity are Denmark, Germany, Greece, Netherlands and Portugal. When Denmark, Germany, and Netherlands are excluded from the sample, the estimated elasticity is significantly lower: values of the long-run elasticity falling outside the 95% confidence band estimated from the whole sample. This indicates that the presence of these countries contributes to keep high the value of the long-run elasticity estimated on the whole sample. Furthermore, these results can also be interpreted that Denmark, Germany and Netherlands have the most significant influence on the meat market integration and meat price transmission in the EU. Without these countries EU meat market would not be as integrated and as efficient. They are also the leaders in the EU pork market. According to the results, Germany is the leader when we consider effectiveness of the meat market measured by the speed of price transmission (responses are the fastest)[20] but Denmark is the price leader which has the greatest impact on the cross-commodity price transmission in the EU meat market (responses are the largest)[21].

Respectively, Greece and Portugal appears to decrease significantly the value estimated across the whole panel of EU15 member states: the exclusion leads to an estimated elasticity of about 0.22, while the estimated elasticity of the whole panel data sample is 0.197. This also means that Greece and Portugal have negative effect on meat market integration in the EU and their inclusion reduces effectiveness of the EU meat market. When we focus on new member states, inclusion of Cyprus, Slovakia, Slovenia and Bulgaria has also statistically significant negative influence on transmission between pork and beef prices and on the meat market integration in the EU.

When considering the speed of adjustment terms – the values of variables come back to the long-run equilibrium level – Denmark and Germany have the most significant influence on the meat market integration. When they are excluded from the data sample, the estimated speed of adjustment term is statistically significantly lower. Greece has opposite influence. In the figures on the speed of adjustment terms covering EU25 member states (excl. EU15), Latvia and Lithuania have statistically significant negative effect on the meat market integration. When they are excluded from the sample, the estimated speed of adjustment term is significantly higher.

In the latter set of figures, *lnpork* is treated as an independent variable. According to the results, EU15 countries influencing statistically significantly on the estimation results of the long-run elasticity are France,

[20] The estimated speed of adjustment term is the lowest when Germany is excluded from the data.
[21] The estimated long-run elasticity is the lowest when Denmark is excluded from the data.

Germany, Greece, Ireland, and the UK. Presence of France, Germany, Ireland and the UK contributes to keep high the value of the long-run elasticity estimated on the whole sample. When they are one at time excluded from the data, the estimated coefficients of long-run elasticity fall outside the 95% confidence band. Without these countries price transmission from beef to pork in the EU would not be as large and as efficient. They appear to be the leaders in the EU beef market. Presence of Greece in the data sample affects another direction: Greece has negative effect on meat market integration in the EU and its inclusion reduces effectiveness of the EU meat market. When we focus on the new member states, inclusion of Slovakia and Slovenia has statistically significant negative influence on price transmission from beef to pork and on the meat market integration among the EU25 member states.

When *lnpork* is an independent variable and *lnbeef* is an explanatory variable, the estimated speed of adjustment term is significantly higher when Greece, Cyprus, Solavakia, Slovenia and Bulgaria are one by one excluded from the data set. France, Germany, Ireland and the UK have instead opposite influence and they have statistically significant positive effect on the meat market integration when we are considering price transmission from beef prices to pork prices.

All in all, the impacts, though significant, appear to be quite moderate and they are not strong enough to alter qualitative results. Moreover, they do not alter considerably quantitative results. We may conclude that our results are robust against estimation method and also against data sample definition.

5. Concluding remarks

The main objective of the study is to analyze the dynamics of price transmission and market integration in the EU livestock market of pork and beef. Our focus is on horizontal cross-commodity price transmission which is not as common approach as vertical or spatial price transmission. However, it provides us interesting information on the connection between commodity prices and integration of commodity markets.

We utilize recently developed panel time-series techniques to analyze the price linkages in the EU meat sector. Our data consists of monthly data on pork and beef prices in the EU member countries during the period from February 1995 to June 2014.

The results indicate that there exists bi-directional relationship between pork and beef prices in the EU in long run. According to the estimation results price transmission is the most significant in the old member states, EU15 countries. Price transmission coming from the fluctuations in pork price is larger and more significant than price transmission coming from the fluctuations in beef price. Depending on investigation period, 1% increase in pork prices, ceteris paribus, would raise beef prices 0.15-0.33%. Long run price transmission from beef price to pork price can be observed only in the results from subperiod 1/2007-6/2014. These results imply that a 1% increase in beef price ceteris paribus, would raise pork price 0.1-0.12% in the EU.

Cross-commodity price transmission between pork and beef has increased remarkably during the past ten years in the EU15 member states. It cannot be found statistically significant relation between pork and beef prices before year 2005. However, ever since EU meat market integration has increased sharply and price transmission between meat markets has become significant. This is also valid when analyzing price transmission and market integration between pork and beef in the EU25 and EU27 countries. This indicates that integration of the EU livestock market of pork and beef has increased during the investigation period.

Also the convergence to the equilibrium has sped up. In the case of EU15 countries after the year 2005 the convergence to the equilibrium has sped up from roughly 1.5 year to 1 year. Although the equilibrium is achieved faster than before, the convergence process is still quite slow. Price transmission from pork to beef in the EU livestock market is strong in magnitude but there still exist some rigidities that decelerates the process on the whole.

In short run, we found evidence only for price transmission from pork prices to beef prices in the EU, not vice versa. Short run relationship between pork and beef prices has only existed since 2007. Moreover, according to the estimation results, short term price transmission from pork price to beef price is interestingly negative. This means that increase in pork price lower the beef price in short term. Overall, short-run dynamics is significantly different in the EU livestock market compared to long run dynamics. This can partly be explained by different time lags in the adaptation of markets. The inclusion of the new EU member states to the data has not changed the interpretation of the estimation results.

We also checked whether the results are strongly driven the performance of a single country. The results indicate that Danish, German and Dutch pork markets and Irish, French, German and British beef markets have the most significant influence on the meat market integration and meat price transmission in the EU. Without these countries EU meat market would not be as integrated and as efficient. Denmark, Germany and Netherlands are also the leaders in the EU pork market. Similarly, Ireland, France, Germany and the UK appear to be the leaders in the EU beef market. Countries that have negative effect on meat market integration in the EU are most clearly Greece, Cyprus, Slovakia, Slovenia and Bulgaria. However, the impacts caused by the significance of meat market of a single country, appear to be moderate – though statistically significant – and they do not alter our main results.

Based on our empirical results we can conclude that the EU livestock market is interactional and integrated. The meat market integration in the EU has proceeded rapidly in recent years. Reduction of barriers to the trade of agricultural products has made the integration possible. Furthermore, it is still important to ensure that regulation measures of the Union treat all member state equally: prevent the barriers to trade between the member countries, harmonize the conditions of competition and monitor that common rules are widely abided by as they stand.

Today, operational environment of the EU meat market is characterized by rapid and significant price fluctuations. The major price fluctuations are affecting increasingly operation of the meat market. The integration of meat market has intensified at the same time as the price fluctuations have generalized. Consequently, this means that the market integration has promoted the price fluctuations to spread between product markets. As the market and price risks grow, managing them is crucial in the EU agriculture policy.

From the point of view of the market integration, the EU meat market is well-functioned and effective. The impacts of the policy measures go faster and faster through the whole meat market. Due to the price integration, targeted policy measures for the one branch of the meat market has impact on the other meat market branches too. Policy decisions concerning the EU meat market must be designed by the perspective of the performance of the whole meat market.

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Appendix 1. LM unit root test results for individual EU member states assuming one structural breakpoint.

PORK	EU15 (2/1995-6/2014)			EU25 (1/2005-6/2014)			EU27 (1/2007-6/2014)		
	Univariate LM Stat	Optimal Lag	Break Location	Univariate LM Stat	Optimal Lag	Break Location	Univariate LM Stat	Optimal Lag	Break Location
Austria	-4.987**	6	10/2004	-6.451***	4	3/2011	-6.623***	3	3/2011
Belgium	-6.212***	5	12/2003	-7.397***	3	9/2012	-8.055***	3	9/2012
Denmark	-8.758***	3	2/2001	-10.113***	3	10/2010	-12.438***	3	10/2010
Finland	-7.055***	2	7/2001	-8.656***	2	1/2011	-9.174***	2	1/2011
France	-6.826***	8	4/2003	-7.904***	6	7/2010	-7.843***	5	7/2010
Germany	-8.492***	4	3/2001	-9.649***	2	4/2011	-10.963***	2	4/2011
Greece	-3.367	7	5/2006	-4.007*	5	2/2013	-3.898*	4	2/2013
Ireland	-5.813***	5	8/2003	-4.937***	5	5/2012	-5.674***	5	5/2019
Italy	-4.064*	5	6/2005	-4.711**	5	7/2012	-6.058***	4	7/2012
Luxembourg	-5.824***	7	9/2003	-5.868***	6	11/2012	-7.187***	4	11/2012
Netherlands	-9.142***	3	6/2001	-10.674***	4	3/2011	-10.008***	4	3/2011
Portugal	-3.548	6	5/1998	-4.216*	3	1/2013	-5.286***	3	1/2013
Spain	-5.371**	5	10/2000	-6.542***	4	6/2012	-6.131***	4	6/2012
Sweden	-6.809***	5	3/2002	-7.396***	2	8/2011	-7.770***	2	8/2011
United Kingdom	-7.471***	7	11/2001	-9.060***	3	1/2012	-8.898***	3	1/2012
Cyprus				-5.112***	6	12/2010	-5.973***	5	12/2010
Czech Republic				-6.087***	7	12/2009	-6.244***	5	12/2009
Estonia				-6.596***	5	4/2011	-6.111***	4	4/2011
Hungary				-3.859*	5	2/2011	-3.563	1	2/2011
Latvia				-3.635*	4	9/2009	-4.502**	4	9/2009
Lithuania				-4.738**	3	1/2010	-5.562***	3	1/2010
Poland				-6.953***	4	10/2010	-7.324***	4	10/2010
Slovakia				-5.044**	6	7/2010	-5.129**	5	7/2010
Slovenia				-5.729***	1	3/2011	-6.376***	3	3/2011
Bulgaria							-6.275***	2	3/2012
Romania							-4.996***	3	5/2013

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively.

BEEF	EU15 (2/1995-6/2014)			EU25 (1/2005-6/2014)			EU27 (1/2007-6/2014)		
	Univariate LM Stat	Optimal Lag	Break Location	Univariate LM Stat	Optimal Lag	Break Location	Univariate LM Stat	Optimal Lag	Break Location
Austria	-5.893***	4	2/2000	-7.045***	3	10/2010	-7.545***	4	10/2010
Belgium	-7.114***	3	3/2003	-8.205***	3	5/2011	-8.803***	4	5/2011
Denmark	-6.252***	3	9/2001	-6.989***	3	3/2013	-11.806***	2	3/2013
Finland	-5.387***	5	2/2002	-6.364***	3	1/2012	-10.532***	3	1/2012
France	-3.071	4	9/2009	-3.552	4	8/2012	-3.683	1	8/2012
Germany	-7.646***	3	6/2001	-7.898***	4	5/2012	-8.199***	4	5/2012
Greece	-2.640	5	5/2005	-2.877	4	1/2007	-2.898	4	1/2007
Ireland	-5.257***	5	2/2004	-4.763**	5	6/2011	-5.094***	5	6/2011
Italy	-3.165	3	9/2003	-3.332	5	11/2011	-3.541	4	11/2011
Luxembourg	-5.764***	6	11/2003	-6.286***	3	1/2013	-6.687***	4	1/2013
Netherlands	-7.259***	2	2/2002	-7.648***	2	9/2011	-8.190***	3	9/2011
Portugal	-5.863***	3	3/2004	-5.905***	2	2/2008	-5.673***	5	2/2008
Spain	-3.617	6	8/2004	-3.070	4	4/2013	-3.486	5	4/2013
Sweden	-5.836***	3	7/2007	-6.450***	1	6/2011	-7.117***	2	6/2011
United Kingdom	-8.451***	5	1/2008	-7.937***	4	2/2012	-7.493***	4	2/2012
Cyprus				-6.491***	5	9/2007	-6.832***	6	9/2007
Czech Republic				-7.514***	5	10/2009	-8.091***	4	10/2009
Estonia				-5.284***	5	2/2006	-5.333***	2	11/2012
Hungary				-2.525	5	1/2011	-2.661	3	1/2011
Latvia				-3.061	6	11/2011	-2.864	4	11/2011
Lithuania				-3.357	4	9/2010	-3.070	6	9/2010
Poland				-5.153***	4	11/2010	-5.486***	5	11/2010
Slovakia				-4.266**	6	1/2011	-4.962***	5	1/2011
Slovenia				-4.189**	5	12/2009	-4.221**	1	12/2009
Bulgaria							-7.920***	4	1/2010
Romania							-5.873***	3	9/2007

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively.

Appendix 2. Pooled Mean Group ECM estimates, EU15 over different sub-periods

Independent variable: *lnbeef*

	2/1995-12/1999		1/2000-12/2004		1/2005-12/2009		1/2010-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run								
ln pork	0.121	0.128	0.145	0.153	0.288	0.296	0.315	0.324
	(1.173)	(1.228)	(1.692)	(1.809)	(24.582)***	(25.074)***	(38.616)***	(39.445)***
time trend	0.001		0.002			0.003		0.005
	(1.556)		(2.724)**			(3.883)***		(4.690)**
Short Run								
ECT *(adj. speed)*	0.042	0.044	-0.068	-0.069	-0.183	-0.190	-0.204	-0.209
	(0.791)	(0.836)	(-1.567)	(-1.615)	(-20.770)***	(-22.181)***	(-21.338)***	(-21.919)***
Δ ln pork	0.002	0.002	0.001	0.002	0.022	0.022	0.025	0.027
	(0.660)	(0.691)	(1.072	(1.109)	(3.136)**	(3.123)**	(3.395)***	(3.417)***
Δ ln $pork_{t-1}$	-0.001	-0.001	-0.001	-0.001	0.009	-0.008	0.010	0.012
	(-0.553)	(-0.573)	(-0.817)	(-0.829)	(2.850)**	(2.866)**	(3.125)**	(3.157)**
Δ ln $pork_{t-2}$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
	(0.517)	(0.506)	(0.584)	(0.582)	(0.749)	(0.720)	(1.019)	(1.033)
Δ ln $beef_{t-1}$	0.456	0.462	0.604	0.611	0.531	0.535	0.489	0.492
	(4.125)***	(4.201)***	(5.228)***	(5.321)***	(4.919)***	(5.074)***	(5.461)***	(5.547)***
Δ ln $beef_{t-2}$	0.009	0.009	0.010	0.009	0.004	0.004	0.008	0.006
	(2.081)*	(2.063)*	(2.314)**	(2.276)**	(1.871)	(1.850)	(1.929)*	(1.818)
Observations	885	900	900	900	900	900	810	810
Log likelihood	-102.541	-89.578	-125.673	-100.284	-222.548	-208.649	-267.590	-238.181

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Automatic selection of lags based on Akaike information criterion. t-statistic are reported in parentheses. All equations include a constant country-specific term.

Independent variable: *lnpork*

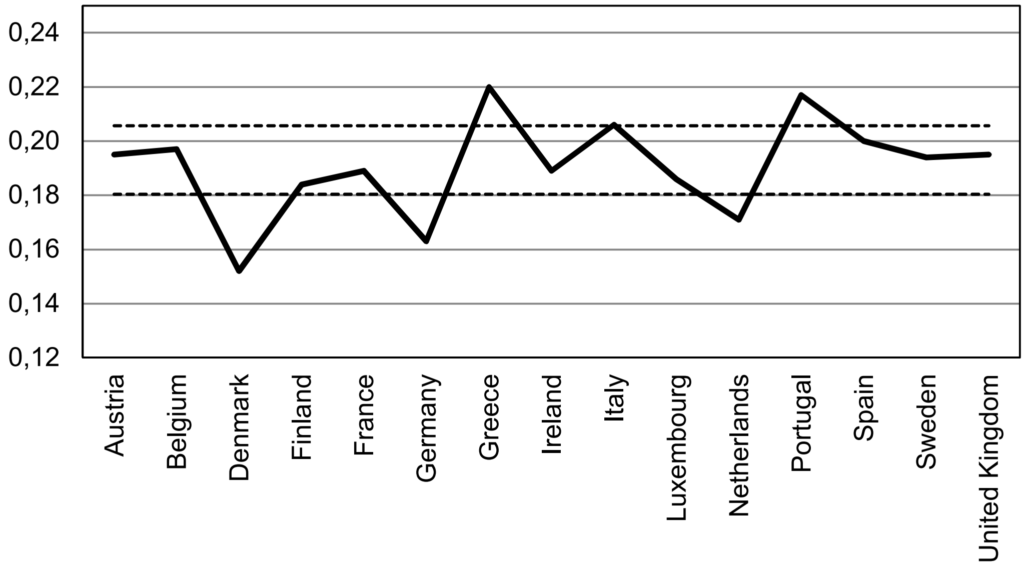
	2/1995-12/1999		1/2000-12/2004		1/2005-12/2009		1/2010-6/2014	
	trend	no trend	trend	no trend	trend	no trend	trend	no trend
Long Run								
ln beef	0.051	0.053	0.072	0.076	0.109	0.112	0.137	0.142
	(0.890)	(0.932)	(1.257)	(1.295)	(1.945)*	(2.163)**	(5.368)***	(5.559)***
time trend	-0.005		-0.005		-0.002		-0.002	
	(-3.118)**		(-2.926)**		(-1.647)		(-1.779)	
Short Run								
ECT *(adj. speed)*	0.014	0.016	-0.004	-0.004	-0.073	-0.077	-0.110	-0.112
	(0.135)	(0.191)	(0.813)	(0.866)	(-4.614)***	(-4.658)***	(-9.229)***	(-9.289)***
Δ ln beef	0.001	0.001	0.001	0.001	0.010	0.009	0.017	0.020
	(0.410)	(0.448)	(0.382)	(0.402)	(2.375)**	(2.327)**	(3.456)**	(3.713)***
Δ ln $beef_{t-1}$	-0.001	-0.001	-0.001	-0.001	-0.003	-0.003	-0.006	-0.008
	(-0.103)	(-0.125)	(-0.187)	(-0.218)	(-0.591)	(-0.644)	(-1.226)	(-1.232)
Δ ln $beef_{t-2}$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
	(0.151)	(0.108)	(0.219)	(0.252)	(0.723)	(0.764)	(1.513)	(1.567)
Δ ln $pork_{t-1}$	0.701	0.708	0.722	0.731	0.687	0.689	0.743	0.752
	(9.147)***	(9.306)***	(10.148)***	(10.135)***	(8.990)***	(8.981)***	(9.357)***	(9.456)***
Δ ln $pork_{t-2}$	0.005	0.005	0.004	0.004	0.002	0.003	0.004	0.004
	(0.747)	(0.793)	(1.156)	(1.188)	(0.917)	(0.944)	(1.362)	(1.381)
Observations	885	900	900	900	900	900	810	810
Log likelihood	-77.890	-90.261	-82.696	-98.987	-131.254	-140.383	-144.229	-153.627

Note: ***, **, * denote significance at 1%, 5% and 10%, respectively. Automatic selection of lags based on Akaike information criterion. t-statistic are reported in parentheses. All equations include a constant country-specific term.

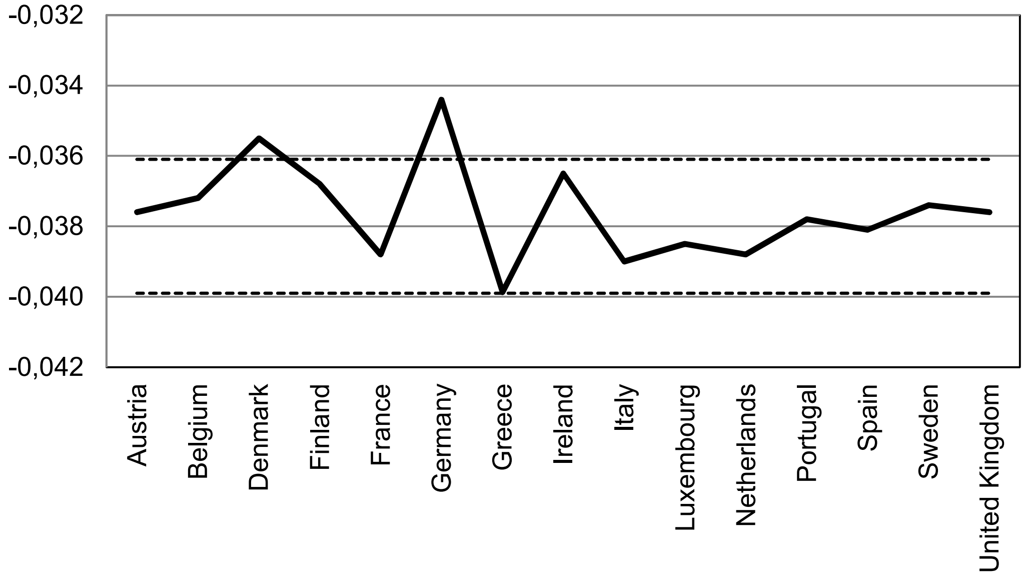
Appendix 3. Cross-sectional stabilities of the long-run elasticity and the speed of adjustment EU15 (period: 2/1995-6/2014)

lnbeef as an independent variable

Long-run elasticity



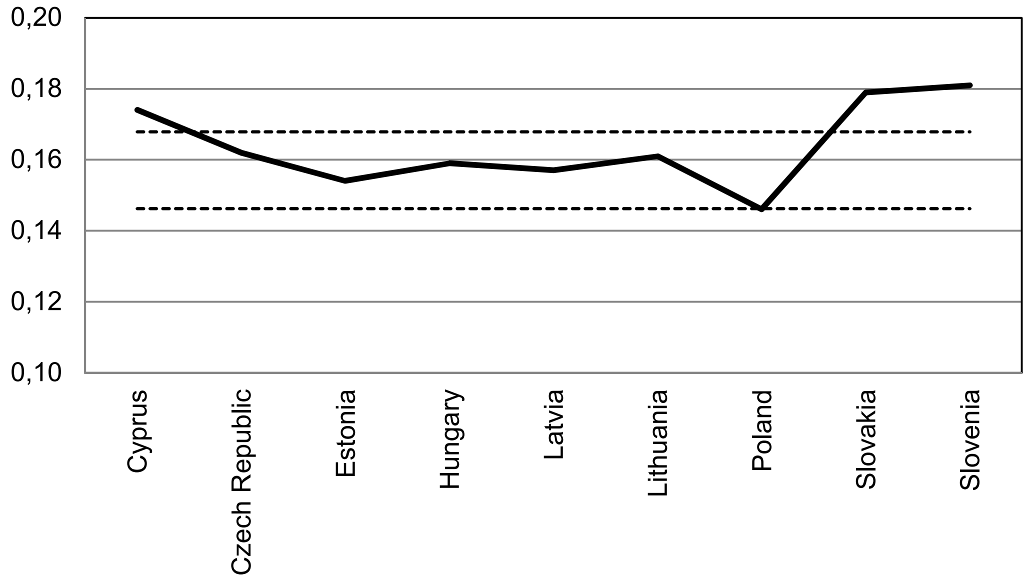
Speed of adjustment



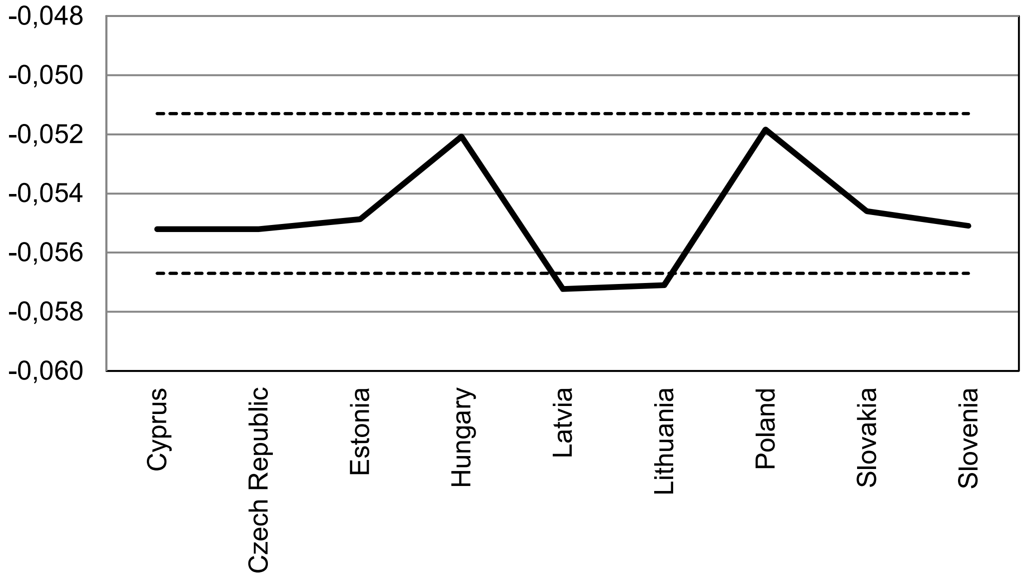
EU25 *(excl. EU15;* period: 1/2005-6/2014)

lnbeef as an independent variable

Long-run elasticity



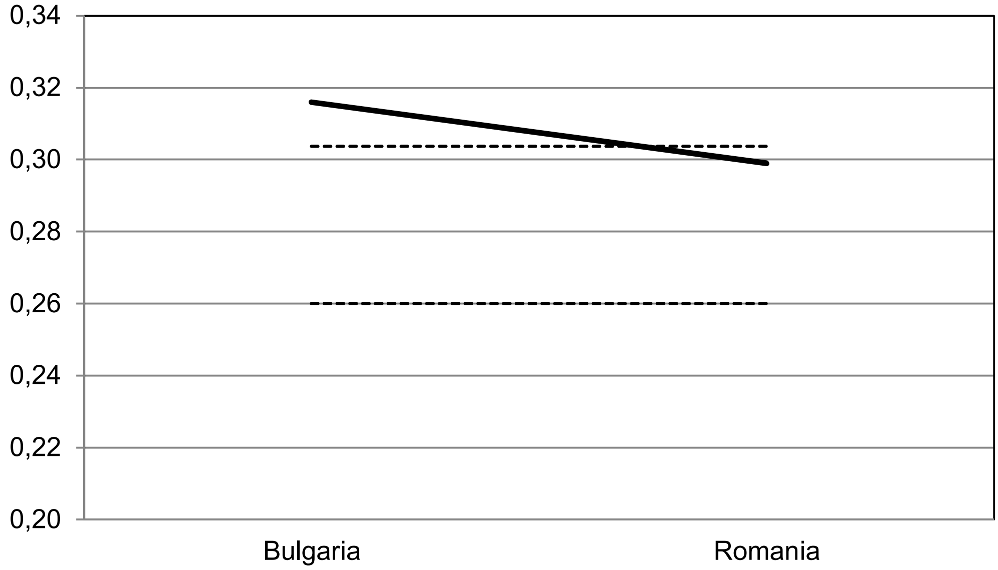
Speed of adjustment



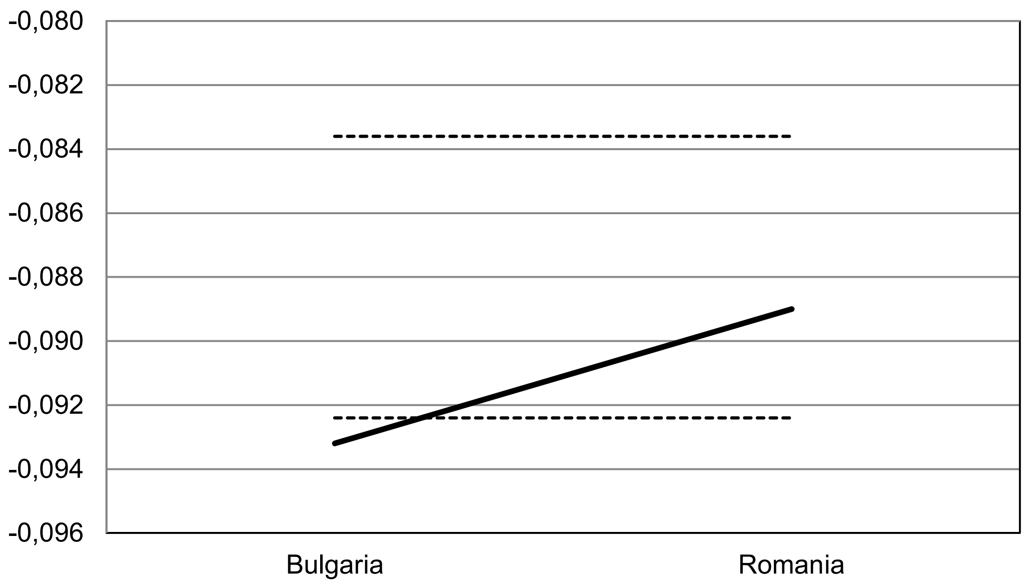
EU27 *(excl. EU25;* period: 1/2007-6/2014)

lnbeef as an independent variable

Long-run elasticity



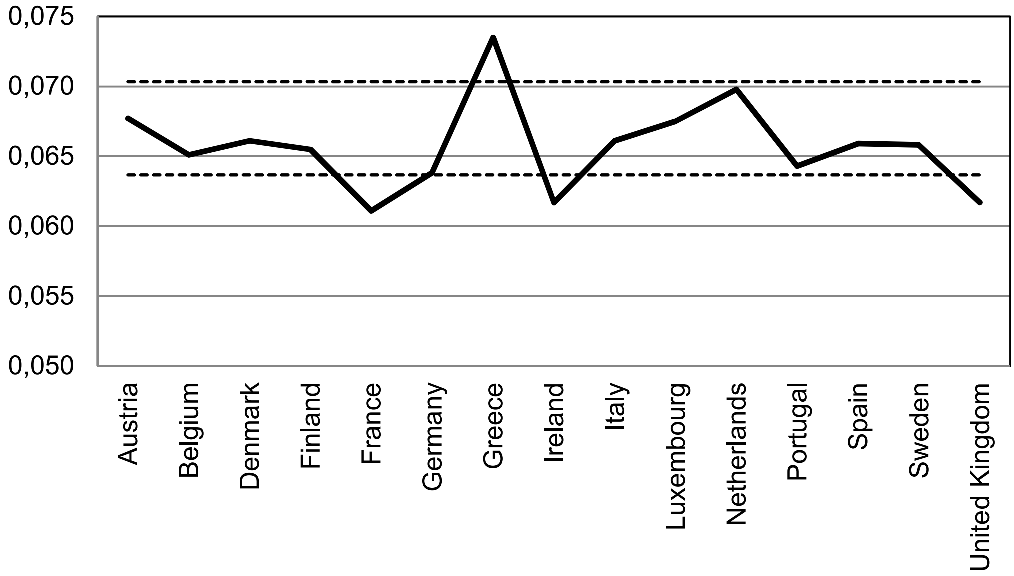
Speed of adjustment



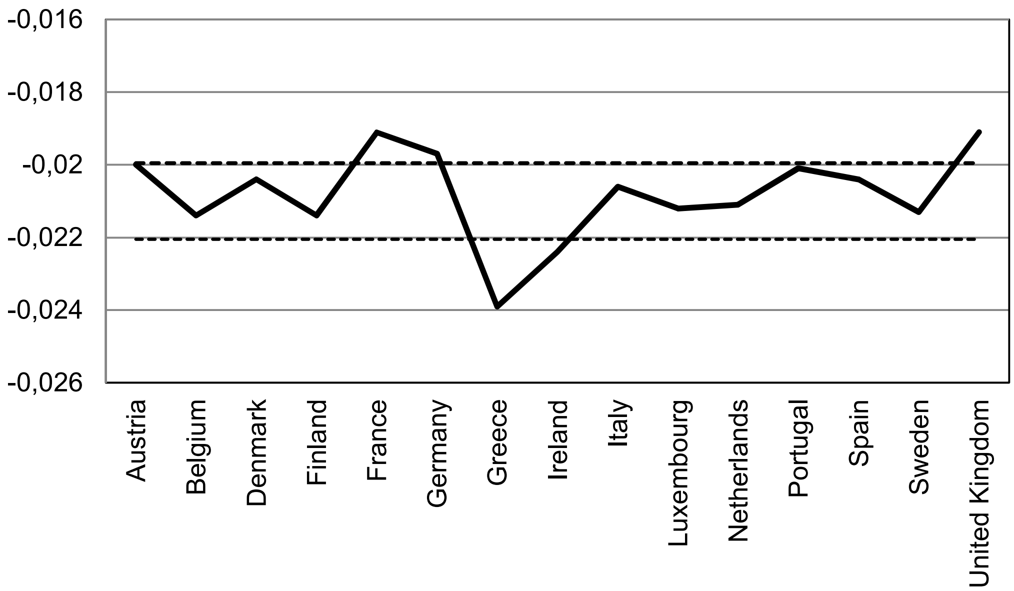
EU15 (period: 2/1995-6/2014)

lnpork as an independent variable

Long-run elasticity



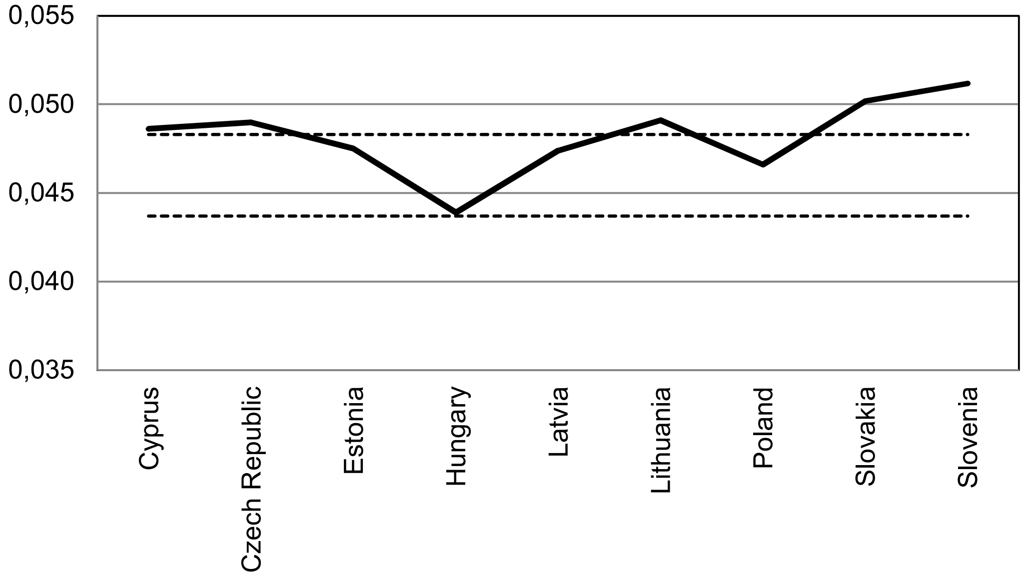
Speed of adjustment



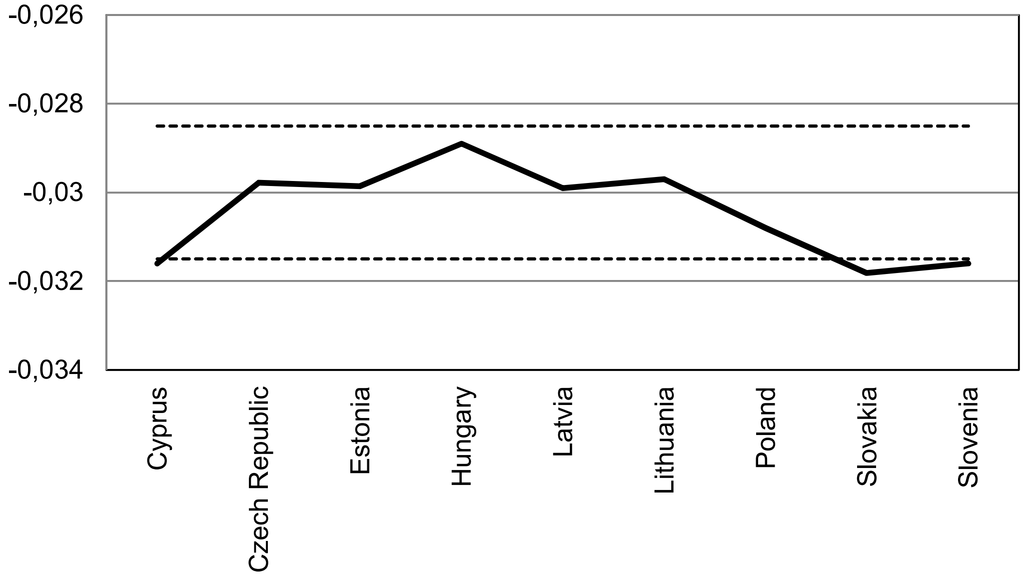
EU25 *(excl. EU15;* period: 1/2005-6/2014)

lnpork as an independent variable

Long-run elasticity



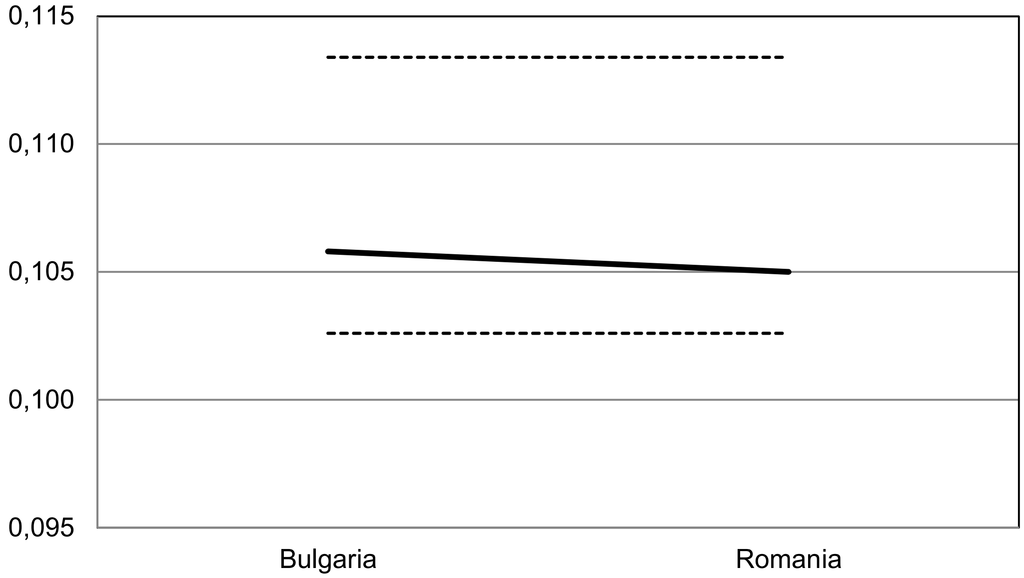
Speed of adjustment



EU27 *(excl. EU25;* period: 1/2007-6/2014)

lnpork as an independent variable

Long-run elasticity



Speed of adjustment