

# Structural change in European calf markets: decoupling and the blue tongue disease

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## Abstract

European cattle markets have recently undergone significant change. We explore the simultaneous impacts of agricultural policy reform and the occurrence of an animal health crisis on spatial interdependencies of calf prices of four major European Union markets. The markets are found to be integrated. Price shocks are rapidly absorbed. We find that the member state specific implementations of the 2003 Common Agricultural Policy reforms significantly affected prices of both the national market and of other member states. The blue tongue disease further induced structural change. Using counterfactual scenarios, we show that the decoupling of payments from production led to reduced calf prices.

**Keywords:** animal health crisis, cattle market, decoupling, policy reform, price transmission

**JEL classification:** Q18, Q13, C32

## 1. Introduction

Reforming the European Union's (EU) Common Agricultural Policy (CAP) has been one of the constants of European agricultural markets. The first major effort in transforming the interventionist market and price support policies was the MacSharry Reform of 1992. This was followed by the Agenda 2000 and its mid-term review, which in 2003 was repackaged into arguably the most fundamental reforms, often referred to as the 'Fischler Reforms'. A key element of these was the decoupling of the domestic support to farmers, which aimed to sever the link between direct payments and

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production decisions. Unlike previous reforms, member states were allowed considerable discretion over the timing and degree of implementation granted in most major European agricultural subsectors. Discretion was most apparent in the cattle sector. The detachment of direct payments from the quantity of animals slaughtered impacted beef production profitability as different production incentives were provided at the national level. These incentives were transmitted to calf markets in the form of a reduced willingness to pay for calves in beef and veal production, thus affecting the quantities and prices of traded animals. The calf trade within the EU is not subject to direct policy interventions. Instead, trade is indirectly influenced by the restrictions laid out in the regulation on the protection of animals during transport<sup>1</sup> (European Council, 2005). Temporary trade restrictions may also be imposed due to sanitary requirements resulting from animal epidemics.

We seek to address the research gap identified by the OECD (2006: 7): ‘Disentangling the effect of recent important policy changes from other factors that also affect production and trade is extremely difficult... There is a need for research in this area.’ To our knowledge, we are the first to investigate the spatial price relationships among EU calf markets subjected to two sources of structural change. We explore how spatial price relationships among EU calf markets are impacted by heterogeneity in implementation of the transition to a more decoupled agricultural policy. In addition to the 2003 reforms, we take into account the potential effects of the extensive animal movement restrictions resulting from the infection of Central Europe with the blue tongue (BT) disease.

Weekly price data are used within a multivariate cointegration framework with exogenous variables to assess price interdependencies among four major national EU calf markets. By considering all price series simultaneously we overcome the omitted variable problem typical of pair-wise cointegration studies which exclude potentially relevant price series and error correction terms (Gonzalez-Rivera and Helfand, 2001). As we simultaneously consider the effects of policy change and trade restrictions, we suspect the data to contain structural breaks. Since standard unit root and cointegration tests are sensitive to this problem, we use a recently developed non-parametric unit root test and a robust cointegration test.

While a large number of price transmission studies have been conducted on USA and international agricultural markets, few have examined intra-EU price relations (among the exceptions are Gordon, Hobbs and Kerr, 1993; Zanas, 1993; Serra, Gil and Goodwin, 2006). A number of investigations, however, have focused on national beef markets within the EU. For example, Lianos and Katranidis (1993) study the beef market in Greece and Bakucs and Fertó (2005) in Hungary. Reztis and Stavropoulos (2010) assess the impact of the CAP framework and its reforms on beef prices and

1 Regulation No 1/2005 specifies a number of requirements for animal transport, depending on the duration and distance of transport, e.g. transport of calves younger than 10 days (14 days) is prohibited (must not exceed 8h).

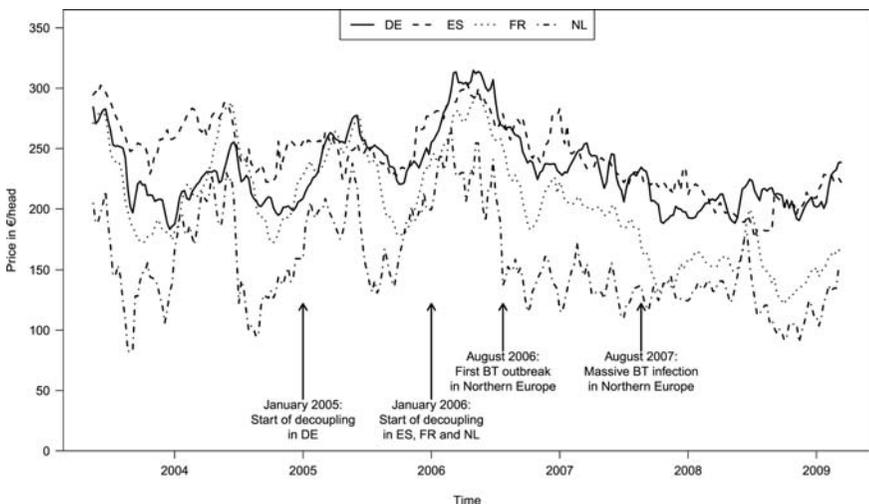
consumer behaviour in Greece. And Junker, Komorowska and van Tongeren (2009) and Hassouneh, Serra and Gil (2010) investigate the impact of the animal disease BSE on trade and prices. Concerning the US beef market, Livanis and Moss (2005) study the impact of food safety on market relationships while Lloyd *et al.* (2001) and Schlenker and Villas-Boas (2009) examine how BSE impacts consumer behaviour and prices in the UK and USA, respectively.

In the next section, we discuss the data and the policy environment. We follow with the methodological framework, estimation results and a comprehensive interpretation. Finally concluding remarks are provided.

## 2. Data and policy environment

Weekly post 2003 CAP reform data are used to investigate the dynamics and interrelationships of four major EU live calf markets: Germany (DE), France (FR), the Netherlands (NL) and Spain (ES) (Figure 1). The choice of countries is motivated by their importance in the EU calf trade. Annually, the number of new born calves in the EU 27 is some 30–33 million animals of which approximately 10 per cent are traded within the Union. The Netherlands and Spain are the largest importers, with 800,000 and 400,000 head, respectively. While Germany is a large net exporter with net exports near 400,000, France is the largest exporter (more than 900,000 head) and the fourth largest importer (about 100,000 head) (ZMP, 2009a, 2009b).

The data include prices of young male calves aged 8 days to 4 weeks from Week 20 (May 15) of 2003 to Week 17 (April 30) of 2009; there are 310



**Fig. 1.** Weekly calf prices for Germany, France, the Netherlands and Spain. *Source:* Authors' depiction based on European Commission (2009a). All graphs are prepared using R (R Development Core Team, 2010).

observations. The data are collected by each member state and transmitted to the European Commission (European Commission, 2002). The prices are representative national averages weighted by the relative importance of each breed and quality.<sup>2</sup>

We now turn to the construction of the variables which quantify the decoupling policies and the emergence of BT.

## 2.1. Policy variables

The 2003 CAP reforms eliminated the link between headage and payments. Under the previous system, farmers could apply for various slaughter premia: steers EUR 150 (up to two payments), bulls EUR 210, adult animals EUR 80 and calves EUR 50 per animal. Nearly all of these coupled direct payments were replaced by the single farm payment (SFP) based typically on historical payments between 2000 and 2002. While the aim of the reforms was full decoupling of payments, individual member states had the option to either fully or partially decouple.<sup>3</sup>

The SFP was implemented in Germany in 2005, while France, Spain and the Netherlands began 1 year later. Germany chose to fully decouple payments in 2005. France and Spain partially decoupled in 2006. The Netherlands also started decoupling payments in 2006 but to a much lesser extent. Slaughter premia for calves and adult animals partially remained in France, Spain and the Netherlands, whereas in Germany they were included in the SFP. We expect these heterogeneous implementations would lead to differing national production incentives, since beef production would be linked to payments differently.

The number of animals of each category receiving slaughter premia is reported annually by the European Commission (2009b). Using these data, the total annual payments were computed for each country. Based on these numbers, we constructed three policy indices  $pol_{DE}$ ,  $pol_{FR}$  and  $pol_{NL}$  which reflect the degree of decoupling in Germany, France and the Netherlands, respectively<sup>4</sup> (Table 1). The variables are computed for each year between 2005 and 2009 relative to the average coupled payments in the base period of 2002–2004. An index of decoupling in country  $Z$  and year  $t$  can be

2 While we assume commodity homogeneity, we recognise that different breeds and animal types exist among countries. However, animal number data suggest that the mixture of animals in each country has remained constant over the sample period. Hence, we assume that the price differentials reflecting these quality differences also remain constant.

3 For detailed accounts on decoupling see, e.g. OECD (2000) or OECD (2006).

4 We use the variable  $pol_{FR}$  for both French and Spanish policies since both countries adopted virtually the same policy. For 2008 and 2009, no expenditure figures were available. Thus, we used a pragmatic approach for predicting the missing values for  $pol_{FR}$  and  $pol_{NL}$  (Germany completely liberalised) because some variability in the policy variables is needed in order to avoid perfect multicollinearity. Animal numbers receiving premia are extrapolated by drawing from a normal distribution with the mean and standard deviation of the animal numbers of 2006 and 2007 for each country. We are aware that the chosen approach is only approximate. However, it suffices to meet the target of giving meaningful estimates by resembling the level and variability in the animal numbers in 2006 and 2007.

**Table 1.** Policy variables quantifying the degree of decoupling

Year	2003	2004	2005	2006	2007	2008	2009
pol <sub>DE</sub>	0	0	100	100	100	100	100
pol <sub>FR</sub>	0	0	7	77	78	78	77
pol <sub>NL</sub>	0	0	2	24	24	23	25

Source: Authors' calculations based on European Commission (2009b).

Note: The values of the policy variables in 2005 of 7 per cent in France and Spain and 2 per cent in the Netherlands were due to natural variations in the national animal numbers which received payments. Decoupling started only 1 year later.

defined as:

$$\text{pol}'_Z = 1 - \frac{\text{premia paid by } Z \text{ in } t}{\text{average premia paid by } Z \text{ in base period}} \quad (1)$$

The closer the index is to 100, the higher the degree of decoupling, that is, 0 and 100 mean fully coupled and fully decoupled, respectively.<sup>5</sup>

Theory suggests an inverse relationship between decoupling and calf prices. Beef production is a function of a number of inputs, including young calves. The demand for calves is given by the marginal value product of calves used in beef production. The slaughter premia are paid to the company delivering the cattle to the slaughterhouse, i.e. in most cases the cattle farmer or fattener. Premia coupled to production shift the factor demand curve for calves outward. However, if the premia are reduced or eliminated, the derived demand curve for calves shifts downward due to a reduction in the marginal value product. With constant marginal costs of calf production, the price of calves will fall; even for a modest upward shift in the supply curve, the sign of the price change would remain unchanged. Thus, we expect a negative effect of decoupling in a country on its equilibrium price for calves.

## 2.2. Blue tongue outbreak in central Europe

Non-policy shocks can also impact market relationships; the BSE crisis was a prominent example. BT is the most recent shock to EU beef markets. BT is a seasonal non-contagious viral disease of ruminants. The disease is mainly transmitted by a midge species that can cause mouth ulcers and in some cases a 'blue tongue' in the animal (Conraths *et al.*, 2009). Although fatality rates from BT in cattle are low, it has important consequences for the dairy and slaughter cattle sector because it can reduce milk yields up to 50 per cent as well as cow fertility. While prevalent in Sub-Saharan Africa over the past couple decades, the disease spread northward to become widespread in the Mediterranean region. It was first detected in Central Europe in August

<sup>5</sup> Note that this measure differs from other decoupling measures, e.g. as suggested in Cahill (1997).

2006 and rapidly spread from the south western part of the Netherlands into the neighbouring countries and in 2007 to the UK.

The BT virus appears in a number of variations which are called serotypes.<sup>6</sup> The outbreak of BT in Central Europe in August 2006 became an important media topic although the number of animals infected remained low in this year. In August 2007, a massive outbreak was recorded in Germany, France and the Netherlands (Conraths *et al.*, 2009). However, the number of cases subsequently declined in Germany and the Netherlands due to vaccination programs introduced in 2008. Before January 2008, cattle in Spain were infected with serotype 1 only but serotype 8 later entered the country coming from Central Europe via southern France.

In an effort to stem the spread of BT, the European Commission adopted strict control measures which included vaccinations and restrictions on the movement of cattle, sheep and goats (European Commission, 2007). When a case of the disease was confirmed, restriction and surveillance zones with radii of 100 and 150 km, respectively, were established (European Commission, 2000). Movement of animals out of the restricted zones was prohibited. Additionally, national import restrictions were occasionally issued by several member states as by France and Spain for German exports. Germany, however, was allowed to continue exports to the Netherlands, the most important destination for its calves.<sup>7</sup> Due to the severe disruption of trade flows, substantial implications for spatial calf price relationships are expected. Calf prices before and after the peak outbreak of BT in August 2007 are shown in Table 2. Both the means and standard deviations of prices were considerably lower in the period after August 2007 than before. The decrease in the price levels and the disappearance of the seasonalities (Figure 1) occurred due to two reasons. The supply structure of calves in the member states was altered by the policy changes and beef prices, exacerbated by the implemented movement restrictions. On the other hand, demand for calves from fatteners had decreased as production costs rose and consumers increasingly purchased chicken meat and pork (BROFA, 2007).

### 3. Methodology and economic background

Market integration and price transmission studies provide valuable insights into the understanding of spatial market networks. Non-integrated markets or the weak transmission of price signals may be due to trade or domestic policies, exchange rate rigidities or transactions costs. If these impediments can be identified, actions can be taken to improve market efficiency. With this knowledge policy makers can be informed of the effectiveness and consequences of implemented measures.

6 These differ by the chemical substances on the surface of the virus. Vaccination is specific to each serotype. This analysis focuses on BT serotype 8 since it is the version of the disease which first occurred and had the largest economic impact in Central Europe.

7 As both countries were subject to restricted zones of the same serotype no movement restrictions were imposed between them.

**Table 2.** Prices (€/head) before and after the massive blue tongue outbreak

	Before August 2007		After August 2007	
	Mean	Standard deviation	Mean	Standard deviation
DE	241	30	207	12
ES	223	21	152	14
FR	255	34	208	16
NL	167	40	127	18

Source: Authors' calculations based on [European Commission \(2009a\)](#).

We begin by discussing the relations between the notions of market integration and price transmission. [Fackler and Goodwin \(2001: 978\)](#) refer to market integration as 'a measure of the expectation of the price transmission ratio'. However, [Barrett and Li \(2002\)](#) define the concept as tradability of a commodity as either established by trade flows or the indifference of agents to trade. Our understanding comes closest to the definition of [Gonzalez-Rivera and Helfand \(2001: 576\)](#) who define it as 'the set of locations that share both the same commodity and the same long run information'. We submit that the mere tradability condition does not suffice to ensure that markets are integrated. For instance, the situation in which a state trading agency uses prohibitive border protection measures to disconnect domestic from international markets, while still exporting domestic products, can hardly be viewed as integrated markets.

Price transmission analysis is based upon the economic concept of the Law of One Price (LOP) in its weak (strong) form, which states that the prices of a homogenous commodity in one market differ at most (exactly) from the costs of transporting the commodity between them. However, in the short-run prices can deviate from this long-run equilibrium condition due to various sources of shocks. When such disequilibrium occurs, price signals will elicit the movement of products between surplus and deficit markets, thus restoring the long-run equilibrium. This economic concept can be empirically investigated using cointegration analysis, where the cointegrating relationship is interpreted as the long-run equilibrium. The existence of such a relationship implies a stationary residual term which is interpreted as a temporary and stochastic deviation from the equilibrium.

If prices are found to be cointegrated, the system can be written as a vector error correction model (VECM) as:

$$\begin{aligned}
 \Delta p_t &= \alpha \beta' p_{t-1} + \sum_{i=1}^k \Gamma_i \Delta p_{t-i} + \varepsilon_t = \alpha e q_{t-1} + \sum_{i=1}^k \Gamma_i \Delta p_{t-i} + \varepsilon_t \\
 &= \Pi p_{t-1} + \sum_{i=1}^k \Gamma_i \Delta p_{t-i} + \varepsilon_t,
 \end{aligned} \tag{2}$$

where  $p_t$  is an  $n$ -dimensional vector containing the prices of a homogenous product in  $n$  spatially separated markets, and  $\Delta p_t = p_t - p_{t-1}$ . The  $n \times r$  matrix  $\beta$  contains the coefficients of  $r$  linear combinations of prices  $p$ , which quantify the structural relationships of the long-run equilibria, so that  $eq_{t-1}$  is the  $r$ -dimensional vector of deviations from the equilibria. The term  $\alpha$  denotes the  $n \times r$  loading matrix. It contains the rates at which the price differences  $\Delta p_t$  react on the past equilibrium deviations  $eq_{t-1}$ . The  $n \times n$  matrices  $\Gamma_i$  contain the short-run reactions of current price differences on past ones and  $\varepsilon_t$  denotes a Gaussian white noise error term. Since calf trade among the four countries is likely to exhibit complex interdependencies, we adopt a multivariate approach.

We follow Gonzalez-Rivera and Helfand (2001) and propose to regard a set of  $n$  markets as integrated if all markets are connected by either direct or indirect trade flows *and* if they are driven by one and only one common factor implying the existence of  $r = n - 1$  long-run price relationships. Thus, market integration implies a long-run view on market interdependencies and can be evaluated by assessing trade flows and by cointegration testing.

In contrast, the transmission of price shocks between integrated markets has a long-run and a short-run dimension. In the long run, it is quantified by the parameters of the prices in the  $j$ th cointegration relationship, i.e. by the  $j$ th column,  $j = 1, \dots, r$ , of the cointegration matrix  $\beta$ .<sup>8</sup> Hence it is a gradual measure since the respective coefficients of  $\beta$  can take continuous values. The closer the coefficients are to zero, the weaker the price transmission is in the long run. If these coefficients can be restricted to unity, the long-run price transmission is said to be complete and the strong form of the LOP holds. In this case, a price change in one market is completely transmitted to the other market.

The short-run dimension of price transmission refers to the parameters in the  $j$ th row of the loading matrix  $\alpha$ . They quantify the magnitudes to which each of the  $n$  prices reacts on the  $j$ th element of the disequilibrium vector  $eq_{t-1}$  from period to period, i.e. the speeds of adjustment of a price shock. The sign of the respective element of  $\alpha$  signals the direction of the adjustment while its absolute magnitude usually lies in the unit interval. Thus, price transmission in the short run is also a gradual measure.

In this interpretation, market integration and price transmission describe different aspects of spatial market interdependencies. Each of them has a one-to-one correspondence to parameters estimated in the cointegration methodology framework, facilitating the interpretation. In combination with a comprehensive dynamic analysis, they are capable of offering detailed insights into spatial price interdependencies.

<sup>8</sup> When using logged data these parameters can be interpreted as long-run price transmission elasticities.

### 3.1. Model design

The final specification of the VECM in equation (2) is based on the context of the EU calf market. First, we augment the cointegration space with a constant, a time trend and the three exogenous policy variables,  $pol_{DE}$ ,  $pol_{FR}$  and  $pol_{NL}$ . Second, we include  $k = 2$  lags of the price differences, selected by the Akaike Information Criterion, and the dummy variable  $d_{2003}$  for the year 2003 outside the cointegration space. With respect to the latter, there was a dramatic fall in calf prices in all countries during the first year of the sample period as a result of a number of exogenous events in the year 2003. Among these events was the EU enlargement in early 2004. Another notable event was the response of calf prices to the peak in milk prices in 2002, which encouraged milk production and thus increased calf numbers. Additionally, the implementation of the 2003 reforms in each member state was not fully determined in early 2003. Seasonality was also included outside the cointegration space as significant seasonal patterns are suggested by Figure 1. Upon exploring this possibility, the results of various likelihood-ratio tests suggested the inclusion of 51 weekly dummies  $d_w$ ,  $w = 1, \dots, 51$ .

## 4. Empirical results

### 4.1. Unit root tests

A major challenge is how to deal with potential structural breaks in both the univariate series and the cointegration relationships due to the BT outbreak. In the presence of structural breaks, standard unit root and cointegration tests are misleading because test statistics are inflated and suffer a considerable loss in power (Gregory and Hansen, 1996; Aparicio, Escribano and Sipols, 2006). We adopt the recently developed forward backward range unit root test (FB-RUR) to obtain valid inference. Proposed by Aparicio, Escribano and Sipols (2006), this non-parametric test counts the number of cumulative minima and maxima. This number is small (large) for a stationary (unit-root) series because its variance is constant (increasing). The statistic is robust to structural breaks and outliers. Whenever the test statistic is smaller than the critical value, the null hypothesis of a unit root is rejected. As shown in Table 3, all series, except that for the Netherlands, exhibit a unit root. Nevertheless, we treat it as non-stationary as recommended by Juselius (2008: Chapter 2).

**Table 3.** Results of the FB-RUR test

Series	DE	ES	FR	NL
FB-RUR statistic	1.947	2.433	2.839	1.379***

*Note:* The critical values for the 5 and 1 per cent significance level are 1.866 and 1.582, respectively. Three asterisks denote significance at the 1 per cent level.

## 4.2. Cointegration tests

Cointegration relationships may also be subject to structural breaks. In this case, performing an adequate cointegration test poses a challenge since the few theoretical results that do exist consider breaks only in the constant of the cointegration space. To our knowledge, no tests are available for multiple cointegration relationships when there are potential breaks in the slope coefficients in the presence of exogenous variables in the cointegration space. We are precisely interested in this issue: the effects of a changing policy environment and an animal disease on the long-run price equilibria.

Gregory and Hansen (1996) developed several tests to identify break points in the cointegration parameters. These tests, however, are only suitable for multivariate models with a single cointegration relationship. The only cointegration tests applicable to a multivariate VECM with multiple cointegration vectors are a modified version of the Johansen-trace test (Johansen, 1995) and the Saikkonen–Lütkepohl test (Saikkonen and Lütkepohl, 2000) both of which are not capable to identify break points. The limiting distribution of the Johansen trace test depends not only on deterministic terms in the cointegration relationship, but also on the number and the locations of structural breaks. Hence, we draw upon the Saikkonen–Lütkepohl test which is robust to breaks in the constants of the cointegration space. In Table 4, strong evidence is provided for  $n - 1 = 3$  cointegration relationships in the four-variate system. Therefore, we conclude that the four markets are integrated; as a result, all cointegration relationships are bivariate.

Further, we test whether the constants of the cointegration relationships are subject to BT-induced structural breaks. Due to the above-mentioned problems, we adopt a pragmatic testing approach and use the Gregory–Hansen test for the four-variate system with one cointegration relationship to determine the potential break point (Table 5). We find a significant structural break in week 35/2007 which closely corresponds to the peak outbreak of BT. Indeed, this is strong evidence that the massive outbreak of BT, that

**Table 4.** Results of the Saikkonen–Lütkepohl cointegration test

$H_0$	$\text{Rank}(\Pi) \leq 0$	$\text{Rank}(\Pi) \leq 1$	$\text{Rank}(\Pi) \leq 2$	$\text{Rank}(\Pi) \leq 3$
Test statistic	18.55	35.97	14.26	0.71
$p$ -value	<0.001	0.001	0.022	0.455

**Table 5.** Results of the structural break Gregory–Hansen test

ADF* statistic	95% Critical value	Observation	Year	Week
-5.29	-5.28	224	2007	35

is, the trade measures implemented in response, significantly impacted the long-run calf price relationships. We add a shift dummy  $d_{\text{AUG}07}$  into the cointegration space, which equals unity from Week 35 (August) 2007 on. The fully specified VECM<sup>9</sup> becomes:

$$\Delta p_t = \alpha\beta'(p'_{t-1} \text{ const trend pol}_{\text{DE}} \text{ pol}_{\text{FR}} \text{ pol}_{\text{NL}} d_{\text{AUG}07})' + \sum_{i=1}^2 \Gamma_i \Delta p_{t-i} + \sum_{w=1}^{51} d_w + d_{2003} + \varepsilon_t. \quad (3)$$

### 4.3. VECM results

The unrestricted multivariate VECM equation (3) is estimated with the Johansen procedure (Johansen, 1995) on which we impose several over-identifying restrictions based on the theoretical expectations. We normalise the bivariate cointegration relationships on DE, ES and FR, respectively, because the Netherlands was by far the largest importer of young calves among the four markets. Hence, all long-run price equilibria are expressed relative to the Dutch price.

Next, we test for the strong form of the LOP. The coefficients of the Dutch price can only be restricted in the relationships with the Spanish and French prices, respectively. Furthermore, German decoupling policy should not impact the ES–NL or the FR–NL relationships. However, excluding the German price from the ES–NL relationship jointly with the other hypotheses was strongly rejected. The expectation that the common French and Spanish policy should not play a role in the DE–NL relationship is confirmed by the joint Wald test. We do not expect the BT outbreak in 2007 to impact the DE–NL relationship since both countries were subject to restricted zones of the same disease serotype and no bilateral trade restrictions were issued. Hence, the movement of animals should not have been affected. The joint Wald test for all these hypotheses yields a  $p$ -value of 0.153.

Restrictions on the adjustment and the short-run parameters are also imposed. Since economic theory cannot provide hypotheses about each of the 252 parameters, we use sequential elimination to identify valid restrictions. According to the largest reduction of the Hannan–Quinn criterion, we identify a set of 28 exclusion restrictions. A likelihood ratio test of a  $p$ -value of 0.246 indicates that these restrictions cannot be rejected. Following Hendry and Juselius (2001: 104) we account for outliers in the residuals by

9 We thank an anonymous referee for suggesting an interesting alternative multivariate model specification with nonlinear adjustments using the BT dummy as an exogenously determined threshold. This would allow assessing whether the impact of the policy variables on long-run price transmission differs before and after the massive BT outbreak. However, due to the limited number of observations after the structural break, which is only slightly larger than the number of model parameters to be estimated, this strategy is not feasible.

**Table 6.** *p*-Values of misspecification tests

Series	LM-test	Jarque-Bera test	ARCH-LM test	Multivariate ARCH-LM test
DE		0.4500	0.9440	
ES		0.0008	0.2403	
FR		0.1896	0.4319	
NL		0.1005	0.9868	
Multivariate test	0.1676			0.3935

using the identification criterion  $|\hat{\varepsilon}_t| > 3.3 \hat{\sigma}_{\varepsilon}$  and include 11 identified outliers as impulse or transitory dummies in the autoregressive component. The model is re-estimated with these restrictions by a two stage procedure (Lütkepohl and Krätzig, 2004). Misspecification tests applied to the residuals suggest that the chosen specification adequately describes the data generating process (Table 6).

## 5. Discussion

The estimated cointegration relationships for the restricted VECM are shown in Table 7. The coefficients of NL (second column) represent the long-run price transmission elasticities. The LOP in its strong form is found to hold between ES–NL and FR–NL. We conclude that price transmission in the long run is complete for these pairs. The price transmission elasticity between DE–NL cannot be restricted to 1, although it is reasonably close to 1.

The magnitudes of the remaining coefficients are plausible. The coefficients of the policy variables denote the average change of the price in the first column in response to increased decoupling. For example, an increase in decoupling in Germany by 10 percentage points is expected to result in a 0.7 per cent decrease in the German calf price. Decoupling in France and Spain also led to decreases in domestic calf prices.<sup>10</sup>

The estimated coefficients of the BT dummy are also of plausible magnitude. The trade restrictions issued as a result of the massive outbreak of the disease in August 2007 indeed impacted price relationships. These trade measures led to a near 14 per cent drop in the Spanish price. The BT dummy, however, is not statistically significant in the FR–NL relationship. Since both France and the Netherlands experienced the BT outbreak and thus belonged (partially) to the same restricted zone, they were not subject to trade restrictions (European Commission, 2007).

The estimated adjustment coefficients are shown in Table 8. These estimates provide information on how the national prices responded to deviations

<sup>10</sup> The effects of this variable for the common French and Spanish policy on both equilibrium prices are negative.

**Table 7.** Cointegration relationships of the restricted VECM

	NL	Constant	Trend	pol <sub>DE</sub>	pol <sub>FR</sub>	pol <sub>NL</sub>	$d_{AUG\ 07}$
DE	1.173 (0.055)	-0.932 (0.307)	0.002 (<0.001)	-0.0007 (<0.001)	—	-0.006 (0.003)	—
ES	1.000 (-)	0.349 (0.082)	0.002 (<0.001)	-0.0022 (<0.001)	-0.009 (0.004)	0.023 (0.013)	-0.135 (0.037)
FR	1.000 (-)	0.116 (0.124)	> -0.001 (<0.001)	—	-0.013 (0.004)	0.039 (0.013)	0.014 (0.036)

*Note:* The prices in the first column are a function of the variables in the remaining columns with the reported coefficients. Standard errors are given below in parentheses.

**Table 8.** Adjustment coefficients of the restricted VECM

Cointegration relationship	DE–NL	ES–NL	FR–NL
DE	–0.077 (0.018)	—	0.062 (0.019)
ES	0.062 (0.017)	–0.101 (0.020)	—
FR	0.102 (0.021)	—	–0.128 (0.021)
NL	—	0.134 (0.027)	—

Note: Standard errors in parentheses.

from the long-run price equilibria. They are of expected magnitude and sign. The Dutch price appears weakly exogenous in the DE–NL and FR–NL relationships. Interestingly, several prices which are not part of the respective cointegration relationship show significant adjustment, e.g. the French price significantly responds to deviations from the DE–NL equilibrium. This underscores the adequacy of the multivariate approach chosen; important variables would have been omitted if only bivariate relationships were specified.

French and Dutch calf prices respond the fastest to equilibrium deviations. Spanish and particularly German prices react much slower. The French price is not only sensitive in the long run to policy change, but also shows a similar sensitivity regarding its reactions to deviations from price equilibria in the short run. The general picture is that market prices respond quickly to disequilibria. Adjustment speeds vary between 6 per cent to more than 13 per cent of equilibrium errors. Thus, price transmission among the four markets is not only high in the long run but also in the short run.

### 5.1. Counterfactual simulations

Two counterfactual simulations illustrate the effect of decoupling (see also OECD, 2000) on the equilibrium prices of each cointegration relationship. We compare estimated equilibrium prices based on the values of the observed policy variables at certain points in time to the hypothetical levels of these policy variables.<sup>11</sup> The two scenarios presented are based on the 12-week average Dutch price before the respective date. Although the equilibrium prices are calculated for the pairwise cointegration relationships of the restricted model, we note that the model coefficients were estimated in a multivariate system. They therefore encompass both the effects of a country's own policy choices on its domestic price and the effects of the policy choices of all other countries in the system. Therefore, a change in the equilibrium price is not the sole consequence of the country's own decoupling choice, but also the choices of the other countries.

Scenario I evaluates the situation for 1 January 2005. It compares the actual policy decisions with the more conservative assumption that each of the four

<sup>11</sup> These equilibrium prices may differ from the observed market prices; they describe the long-run market equilibrium for a given policy constellation.

**Table 9.** Scenario I—fully coupled policies on 1 January 2005

Country		DE	FR	ES	NL
Actual policy	Observed price (€/head)	200	246	196	145
	Policy variable	100	7.5	7.5	1.6
	Equilibrium price (A)	151	156	198	—
Scenario I	Policy variable	0	0	0	0
	Equilibrium price (B)	163	162	254	—
	Ratio (B) to (A)	1.08	1.04	1.28	—

**Table 10.** Scenario II—most protective policy choice on 1 January 2007

Country		DE	FR	ES	NL
Actual policy	Observed price (€/head)	235	254	206	143
	Policy variable	100	78	78	24
	Equilibrium price (C)	164	144	227	—
Scenario II	Policy variable	25	25	25	25
	Equilibrium price (D)	171	304	434	—
	Ratio (D) to (C)	1.04	2.12	1.91	—

countries would have decided for zero decoupling. Germany took the most liberal policy decision to completely decouple on this date even though the start of the implementation of the reforms could have been delayed until January 2007. Table 9 illustrates the expected depressing effect of decoupling on the equilibrium price in each country. The actually implemented decoupling policy led to lower equilibrium prices (A) in comparison to the hypothetical case of zero liberalisation (B). The French equilibrium price appears to be the least impacted by the chosen decoupling policy. The German equilibrium price would have been 8 per cent higher without decoupling. In contrast, the Spanish equilibrium price could have been expected to be nearly 30 per cent higher if none of the countries had decoupled. These findings are driven by the coefficient estimates of the German policy variable in the cointegration relationships (Table 7). The Spanish equilibrium price reacts significantly to German policy changes while the French price is not affected.

Scenario II assesses the hypothetical situation of a more protective choice of decoupling policies on 1 January 2007, the date of the mandatory movement towards decoupling for all countries. With the exception of the Netherlands, chosen national policies were quite liberalised at this time (see Table 1). The scenario assumes hypothetical values of the policy variables at 25 per cent, which approximates the observed situation in the Netherlands. In Table 10 it is apparent that the effects in Scenario II are much stronger than 2 years earlier in Scenario I. Equilibrium prices would have been much higher if Germany, France and Spain had opted for considerably more

**Table 11.** Descriptive statistics of the estimated error correction terms

	DE–NL	ES–NL	FR–NL
Minimum	−0.1419	−0.4334	−0.1317
Median	0.2490	−0.0029	0.2190
Maximum	1.1004	0.7324	0.7570

*Note:* For calculating the observed magnitudes of relative price deviations from equilibrium, the value of the exponential function of the estimated residuals has to be considered.

conservative policy choices.<sup>12</sup> The effects of decoupling on calf prices are the smallest in Germany and the greatest is France.

## 5.2. Dynamic analysis

As mentioned above, *ceteris paribus* interpretations of the cointegration relationships in multivariate systems can be misleading due to the complex and comprehensive dynamics of such models (Lütkepohl and Reimers, 1992). A number of approaches can be used to study the dynamic behaviour of linear multivariate cointegration models, that is, the effects of a shock to the prices considered. We discuss three closely related concepts<sup>13</sup>: impulse response functions (IRF), persistence profiles (PP) and generalised impulse response functions (GIRF). What these measures have in common is that they track the dissemination of a disequilibrium shock<sup>14</sup> in one period to the price system over time. They are particularly useful for gaining insight into the dynamic behaviour of multivariate systems. The measures are functions of the cointegrating vector  $\beta$  and of  $B_o$ , the recursive sum of parameters of the VAR representation of the VECM (and hence of the parameter matrices  $\hat{\Pi}$  and  $\hat{\Gamma}_i$ ).

These approaches offer information on the relative magnitudes (which can be interpreted as semi-elasticities), the time paths, and whether the responses are transitory or permanent. If the variable responding to the shock returns to zero, then the effect of the shock is transitory, otherwise permanent since it is pushed to a new equilibrium value. The measures differ in some aspects. First, they either focus on a variable specific shock or on a system wide shock. Second, some methods are based on the variance–covariance matrix  $\Omega = PP'$  of the model residuals without further transformation. Other

12 At first glance, the hypothetical equilibrium price of ES seems very high as does the margin between the German and Spanish prices. As Table 11 shows, the estimated deviations from equilibrium, e.g. for ES–NL, lie in the range between −0.43 and 0.73. Hence, observed prices might well have been EUR 300/head for example since  $\text{EUR } 434 \cdot e^{-0.3} = \text{EUR } 322$ .

13 Ben-Kaabia and Gil (2007) give a detailed account for the dynamic analysis of multivariate non-linear cointegration models with one cointegration relationship. We are indebted to two anonymous referees for pointing out further approaches for multivariate linear models, such as forecast error variance decomposition (Lütkepohl and Reimers, 1992) and structural VECM (e.g. Amisano and Giannini, 1997).

14 Usually, it has the size of one standard error of the variable regarded as its source (*unit shock*).

methods orthogonalise this matrix using the Cholesky decomposition  $\Omega = PP'$  where  $P$  is a  $n \times r$  lower triangular matrix.

Lütkepohl and Reimers (1992) recommend the use of IRF,  $\psi_{eq,m}^{\circ}(o)$ , in cointegrated systems. IRF track the dissemination of a shock to the  $m$ th,  $m \in \{1, \dots, n\}$ , price equation into the price system. They are usually estimated by using orthogonalised shocks as:

$$\psi_{eq,m}^{\circ}(o) = \beta' B_o P e_m, \quad o = 0, 1, 2, \dots, \quad (4)$$

where  $e_m$  is an  $n$ -dimensional selection vector for the  $m$ th equation shocked.

Lee and Pesaran (1993) suggest PP,  $h_{eq}(o)$ , as an alternative in order to measure the difference between the conditional variances of the  $o$ -step and the  $(o-1)$ -step-ahead forecasts scaled by a suitable matrix  $G$ . They are computed as:

$$\begin{aligned} h_{eq}(o) &= GH_{eq}(o)G = G(\beta' B_o P)(\beta' B_o P)' G = G\beta' B_o P P' B_o' \beta G \\ &= G\beta' B_o \Omega B_o' \beta G, \end{aligned} \quad (5)$$

where  $H_{eq}(o)$  denotes the unscaled PP of an  $o$ -step-ahead forecast of a shock to the whole system of equations (Pesaran and Shin, 1996).

Pesaran and Shin (1998) generalise the impulse response approach to GIRF,  $\psi_{eq,m}^g(o)$ , which are calculated as:

$$\psi_{eq,m}^g(o) = \sigma_{mm}^{-1/2} \beta' B_o \Omega e_m. \quad (6)$$

Equations (4)–(6) illustrate the close relationships of the three measures. IRF and GIRF, which are identical only for  $m = 1$  (Pesaran and Shin, 1998), quantify the effect of variable specific shocks, that is, shocks in one of the price series. PP measure system wide shocks to all prices (see, e.g. Pesaran and Smith, 1998). IRF are not suitable for multivariate linear cointegration models since they suffer from the composition problem (Koop *et al.*, 1996), that is, the resulting functions are not unique because they depend on the ordering of the variables and, hence, the implied orthogonalisation of  $\Omega$ . Both PP and GIRF are robust to the ordering of the variables in the VECM and are thus superior measures in the given context.

The following figures which were calculated based on the VAR representation illustrate the benefits of these measures for interpretation. The dissemination patterns of system wide and variable specific shocks in the cointegration relationships are displayed in Figures 2 and 3, respectively. In Figure 2, the time paths of responses of the three long-run price equilibria to a system wide shock appear to be very similar. After overshooting in the first week after the shock (i.e. temporarily increased disequilibrium), the PP converge rapidly to zero. The French and the Spanish relationships overshoot the most, but also show the steepest decline afterwards. Within 4 weeks, more

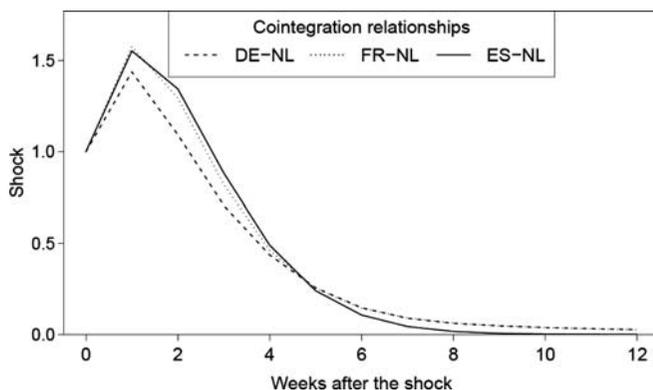


Fig. 2. Persistence profiles.

than half of the initial shock is absorbed by each relationship. After 10 weeks, the shock is almost completely absorbed into the market system.

A similar picture is found for the responses of the price equilibria to variable specific shocks as depicted in Figure 3. Shocks to the two largest markets, Germany and France, lead to disequilibria of high magnitude and longest duration. The initial negative equilibrium errors are quickly reduced within the first 5 weeks. However, these movements are followed by moderate over-shooting which is slowly corrected within the next 4–5 months. A shock to the Dutch market which corresponds to the system wide shock in Figure 2 is adjusted for most quickly.

In Figure 4 we show the responses of the four prices to a shock to each of the prices, as calculated in equation (6) without pre-multiplying by  $\beta'$ . In the first 5 weeks after a shock, the Dutch price typically responds the strongest. However, except for a shock originating in Spain, within a few weeks the Dutch price decreases by more than half of its initial response. The Dutch price shows the by far largest semi-elasticity of a domestic price shock amounting to more than 7 per cent while in general the responses both to domestic and foreign shocks range between 1 and 2 per cent. For most cases, stable values are reached within 10 weeks. The permanent effects of the shocks on the prices are in the magnitude of around 1 per cent, i.e. the new equilibrium prices are roughly 1 per cent higher than before the shock. In the medium run, i.e. after 6–8 months, the German price consistently shows the strongest response to a shock.

Figures 3 and 4 illustrate the different effects which shocks have on equilibrium relations and prices, respectively: the shocks are transitory for the equilibrium relations since the response functions converge to zero after some time (Figure 3), that is, disequilibria are adjusted for and the equilibria are restored. In contrast, the price series attain new equilibrium values after a shock, as shown by the permanent increases in Figure 4.

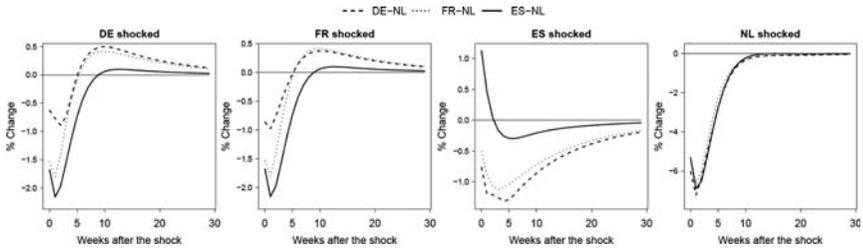


Fig. 3. Generalised impulse response functions of the cointegration relations.

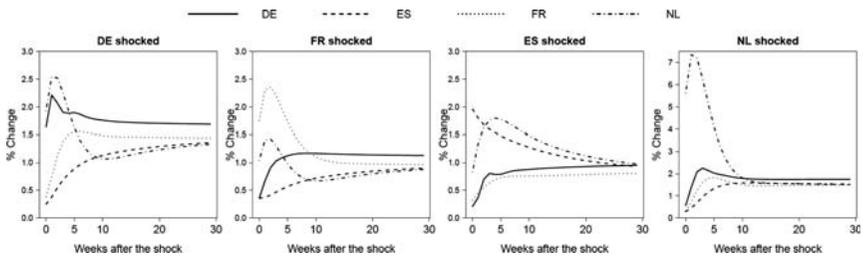


Fig. 4. Generalised impulse response functions of the price series.

All three aspects of the dynamic analysis confirm the general picture of the close interrelationships of the four European calf markets studied. We leave the analysis of price discovery for further research. An approach for studying this issue could be the permanent–transitory (P–T) decomposition (e.g. Gonzalo and Ng, 2001). This method splits the shocks and the price series into long-run trends and short-run cyclical components and can straightforwardly be calculated using the estimated loading matrices and cointegration matrices.<sup>15</sup> Garrat *et al.* (2006) suggest a similar decomposition which is applicable to models including exogenous variables which additionally splits the permanent component into a deterministic and a stochastic part. The P–T decomposition could yield insights into whether the common factor (i.e. the permanent shock) driving the price system originates from only one national market or from several sources and whether this factor was subject to structural change due to the policy reform or the BT crisis.

## 6. Conclusions

Following the 2003 reforms of the EU's CAP, decoupling of support payments was implemented differently by the member states. Since 2003, however, European cattle markets were not only subject to structural change induced by the changes in the policy framework, but also to a major animal health

<sup>15</sup> We are grateful to an anonymous referee for pointing our attention towards the relevant literature.

crisis brought on by the infection of Central European cattle herds with the BT disease. In August 2007, a large-scale outbreak occurred, resulting in extensive restrictions on animal movements between some member states.

We empirically explore how these two major exogenous shocks impacted the long-run transmission of price signals among four major European calf markets. We assess price interdependencies of calf markets using weekly price data of young male calves of Germany, France, the Netherlands and Spain ranging from 2003 to 2009. A recently developed range unit-root test, which is robust to structural breaks, and a multivariate Vector Error Correction Model with exogenous variables are used for this purpose.

We discuss the notions of market integration and price transmission. The former term is seen as a long-run measure, while the latter concept is a gradual measure of both a long-run and a short-run dimension. We find strong evidence that markets are integrated. Price transmission in the long run is found to be complete in two price relationships and impacted significantly by the EU's 2003 agricultural policy reforms. Price transmission in the short run is found to be fast. We assess the price dynamics of the markets by generalised IRF and PP. Both approaches illustrate the fast absorption of shocks either into national prices or into the entire price system. The outbreak of the BT disease played a significant role in the European beef market. We conclude that the four calf markets studied are closely interconnected and find strong evidence that they belong to a common European market.

Two counterfactual scenarios reveal the depressing effect of decoupling on calf prices. They show that national policy choices towards decoupling lowered the expected domestic market equilibrium prices. Due to the well-established interconnectedness of these national markets, the policy effects spill over to other markets.

Our results yield two policy implications. First, the strong interconnectedness of calf markets provides an interesting case against member state specific policy reforms within a common market. The decision of the European agricultural ministers to allow for deviations from a general decoupling proposal leads to discernable price effects. The cost of allowing member states to pursue different policy implementations must be carefully weighed against national interest concessions. Second, the strong price impact attributed to trade restrictions in response to the outbreak of the BT disease illustrates spill over effects to most markets in the EU. The massive shifts in trade flows in the aftermath of the disease suggest the need for a coordinated European action, as was soon established after the first recorded outbreaks.

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