LOCAL POLYNOMIAL FITTING AND SPATIAL PRICE RELATIONSHIPS: PRICE TRANSMISSION IN EU PORK MARKETS

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Summary

We study price transmission processes within EU pork markets after the implementation of the EU single market in 1993. We compare results derived from nonparametric regressions with those obtained using alternative nonlinear threshold models. Both techniques support the hypothesis that prices are transmitted across spatially separate EU pig markets and provide evidence for asymmetric price adjustments. They also suggest the existence of a range of price differentials where equilibrating price adjustments are less intense. Nonparametric techniques often suggest a higher degree of price transmission than that implied by threshold models.

Keywords: price transmission, pork markets, nonparametric methods.

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1. Introduction

The analysis of spatial price relationships has traditionally sought to shed light on spatial aspects of the functioning and interaction of localised markets. Various concepts such as spatial arbitrage, market integration and market efficiency have been used to describe spatial price linkages. Spatial price arbitrage is an equilibrium concept according to which, in a well-functioning market, transactions between spatially dispersed agents will ensure that price shocks occurring in one market evoke responses in other markets. As a result, prices of a homogeneous good in two separate locations will differ by, at most, the cost of moving the commodity from the cheapest market to the most expensive one. As Fackler and Goodwin (2001) note, the arbitrage condition is equivalent to the weak version of the law of one price (LOP). Several authors have studied price transmission processes within this framework (e.g. Ardeni, 1989). Others have framed their analyses in the context of market integration (Zanias, 1993; Baulch, 1997). According to Fackler and Goodwin (2001), market integration is usually a measure of the degree of price transmission between spatially separate markets, while the concept of efficiency is often used to denote a situation in which no opportunities for arbitrage have been left unexploited by arbitrageurs.

Our paper studies the spatial arbitrage condition within the EU pork market from the mid-1990s to the mid-2000s. A number of applied analyses on spatial price transmission have focused their attention on food markets (Goodwin and Piggott, 2001; Zanias, 1993). While most previous studies have been based on U.S. (Goodwin and Piggott, 2001) or international data (Dries and Unnevehr, 1990), European food markets have received less attention (Zanias, 1993; Gordon, Hobbs and Kerr, 1993; Sanjuán and Gil, 2001 are a few exceptions). Additionally, to our knowledge, no published study has addressed EU spatial food price relationships after the implementation of the Single Market in 1993, which involved the removal of barriers to trade within the EU.

The analysis of spatial price relationships within the EU pork market is considered economically relevant for at least three reasons. First, pork is a rather homogeneous good relative to other meats such as lamb. Additionally, EU policymakers have intervened less in pig markets relative to other markets such as beef or milk. These factors should provide an optimal environment for smooth price transmission processes across space. Second, an understanding of spatial price linkages is important because the structure of the pig industry has changed at an unprecedented rate since the 1990s, resulting in higher levels of concentration. These structural changes could have had an influence on the degree to which price shocks are regionally transmitted. Third, the EU meat sector has been seriously affected by major veterinary crises throughout the period of study (the BSE crisis, the swine fever epidemic, and the foot and mouth disease outbreak). These animal disease episodes have had important impacts on both supply and demand for meat products and have triggered various policy interventions such as trade bans or the removal of animals from the market. These episodes are thus very likely to have had an impact on price transmission processes. Although Sanjuán and Gil (2001) studied cointegration relationships between prices taken at different EU pork markets in the period that goes from the end of the 1980s to the mid 1990s, the changes experienced by the sector afterwards, as well as the recently introduced methodological refinements, make it interesting to update the analyses in order to better understand the current workings of EU pig markets.

In order to study the issue of spatial price transmission empirically, a series of analytical methods have been devised. Given the usual nonstationary nature of price data, recent contributions have underlined the need to use econometric techniques adequate for dealing with nonstationary and cointegrated data. A current issue is the possibly nonlinear nature of spatial price relationships that has been often attributed to a lack of perfect arbitrage resulting from transactions costs and uncertainty. Nonlinearities appear whenever price dynamics differ across regimes (e.g. when asymmetries characterise price adjustments). Nonlinear price responses are symmetric when a shock to a certain price elicits the same response to other prices regardless of the direction of the shock, or asymmetric otherwise. Previous research has identified several causes of asymmetries that may be relevant to the EU pork market (see section 2). We test for the presence of asymmetries in spatial price relationships. In this regard, our analysis represents a contribution to the literature, as no previous published study has made an allowance for asymmetries or for nonlinearities in regional food price transmission processes within the EU. Several econometric procedures have been proposed to capture nonlinear price relationships. Recently, Chavas and Metha (2004) suggested an extended error correction model that allows price dynamics to differ across regimes. While these authors treat regime switching as exogenous, more general models of asymmetry such as threshold autoregressive (TAR) models (Obstfeld and Taylor, 1997; Goodwin and Grennes, 1998; Goodwin and Piggott, 2001), or threshold vector error correction (TVECM) models (Lo and Zivot, 2001; Goodwin and Piggott, 2001) incorporate this issue as endogenous. All these parametric approaches have in common the fact that they model nonlinearities by using a limited number of variables and their associated parameters. Parametric approaches to modelling price relationships require assumptions about the true nature of price behaviour that may be too restrictive or unrealistic. In threshold models, for example, two regimes are usually specified - one that pertains to price relationships under trade and another that holds when price differences are too small for profitable trade to occur.

The transition from one regime to another is assumed to be sharp and discontinuous, involving price differentials that motivate individuals to undertake arbitrage activities and/or adjust prices that are common across economic agents. This assumption may be valid if transaction costs and uncertainties are homogeneous across different individuals, but may be too restrictive otherwise. Models smoothed using the approach of Teräsvirta (1994) partially overcome this limitation by allowing for gradual adjustments between regimes.[1] However, in that they are parametric, they still have the potential for specification biases as a result of inappropriate parametric assumptions.

Contrary to parametric models, nonparametric techniques such as local polynomial modelling (Fan and Gijbels, 1996) do not require any assumption about the functional form characterising price behaviour. They are especially useful for situations in which a suitable parametric form of the regression surface is unknown. In that they are data-driven methods, it is the data that inform and determine the shape of the relationship among the variables. Nonparametric models have the advantage of essentially 'nesting' parametric alternatives. For example, the nonlinear pattern of adjustment implied by a threshold model can be compared to the patterns implied by nonparametric models. To the extent that the implied patterns of adjustment differ substantially across techniques, one may question the appropriateness of the parametric threshold model. It is also important to point out that our approach is robust to bi-directional patterns of trade in that no restrictions (or assumptions) are imposed with regard to the symmetry of responses or relative sizes of transactions costs from moving in one direction

[1] Smooth transition regressions may be seen as a generalisation of other parametric approaches used to model nonlinearities in the data such as the TAR models or the nonlinear error correction models introduced by Escribano (1986) (see Escribano and Mira, 2002; or Teräsvirta and Eliassen, 2001).

relative to another. Our methods are also robust to the possibility of asymmetric adjustments that occur for other reasons as well as to misspecifications due to structural breaks.

To date, the use of nonparametric techniques to study nonlinear aspects of price transmission has been very limited (see Mancuso, Goodwin and Grennes, 2003; Barrett and Li, 2002; or Serra, Goodwin and Mancuso, 2005 for a few exceptions). In this article we assess price transmission relationships among EU pork markets by employing nonlinear methods of analysis. To our knowledge, no previous study has applied these techniques to a consideration of spatial price transmission within EU food markets. To do so, we first test for the weak version of the law of one price by using local polynomial modelling techniques. Results are then used to guide the specification of parametric threshold autoregressive models. We next compare the results derived using flexible nonparametric techniques to those obtained from the more restrictive parametric TAR models. Although both techniques yield similar qualitative results, important differences arise in the nature of price linkages.

This article is organised as follows. In the next section, the main characteristics of the EU pig sector are presented. The methodology section describes both the parametric and nonparametric techniques employed to characterise the nature of price transmission. The empirical implementation section presents a description of the data used and a discussion of the results. Concluding remarks are presented in the last section.

2. **The main characteristics of the EU pork sector**

The EU-15 occupies a prominent position in world rankings of pork production and trade. It is the world's second largest pig producer after China and is followed, at a distance, by the USA.

EU pork production capacity is consistently above 100 per cent (almost 107 per cent in 2003) [2], which explains the strong export orientation of this sector especially in some countries such as Denmark. Although there are relevant pork trade flows with non-EU member countries, the most intensive trade flows occur within the EU-15. In 2002, for example, intra-EU pork trade accounted for 2.7 million tons, while exports to and imports from countries outside EU-15 were in the order of 1.2 million and 51,000 tons respectively. Pork production represents around 8 per cent of the EU gross agricultural product. Germany, Spain, France and Denmark are the four top EU-15 pork producers. Most pork is produced under intensive farming systems, which generally reduce heterogeneity across countries. In spite of this, there exist relevant differences in average carcass weight preferences across EU member states: while some exceptions occur, north-central Europe has a preference for heavy carcasses whereas southern Europe chooses lighter deadweights.

Our analysis of pork price transmission focuses on the four leading EU pork producers: Germany, Spain, France and Denmark, which represent more than 60 per cent of net pork output. While Germany and France produce carcasses above the EU average weight (around 87-88 kg), Spain and Denmark opt for lighter animals. As is shown in Table 1, however, these differences in production do not inhibit trade flows. The four countries studied represent a conspicuous part of intra EU trade, with Germany being the largest importer and Denmark a leading exporter. There are bilateral trade flows between the four countries studied. In 1995, for example, German pork exports to Denmark, Spain and France represented almost 20 per cent of Germany's pork exports to EU countries. Additionally, Table 1 shows an increase in exports between the four countries between 1995 and 2003. With the exception of Germany, imports among the four

[2] All data quoted in this section are derived from Eurostat and refer to EU-15.

countries have also increased, and imports within the group also represent a relevant portion of total intra EU imports.

Table 1 here

Our study allows for nonlinearities thus recognising that, although prices may have a tendency to move towards a long-run equilibrium, actual movement towards this equilibrium may not occur in every period. Transaction costs or market imperfections can lead to nonlinear and asymmetric price adjustments. The economic literature has identified several theoretical reasons that may bring about asymmetries (Meyer and von Cramon-Taubadel, 2004), and that may be relevant to the understanding of price transmission processes within the EU market. First, trade flows between two markets tend to occur primarily in one direction. Denmark, for example, is a net exporter of pigmeat, with Germany being the largest EU importer. According to previous research this may cause asymmetric price adjustments (Goodwin and Piggott, 2001). Market power can be a second source of asymmetries. As Abdulai (2000) noted, if middlemen have market power, they may react more quickly to shocks involving a reduction in their marketing margins than to shocks increasing them. In the past decade the EU pig industry has undergone major structural changes. Although the family farm still plays an important role, there has been a considerable increase in the number of corporate units, which are usually vertically integrated. Structural changes have resulted in a considerable increase in concentration. While the total pig numbers increased from 121,5 to 121,95 million head during the period 1993-97, the average herd size increased from 78 to 105.8 pigs per holding. Structural changes were more accentuated in countries such as Spain or Germany, where the average herd size more than doubled, or in France or Denmark which experienced increases on the order of 80 per cent. Increased concentration may yield market power and result in nonlinearities in price transmission processes. Third, inventory management strategies or policies that insulate markets may also prevent full price transmission and motivate nonlinearities (Blinder, 1982; Wohlgenant, 1985; Kinnucan and Forker, 1987). In this regard, the relevant animal disease episodes that afflicted the EU meat sector since the mid-1990s could have caused nonlinearities in spatial price transmission processes. In any case, they are likely to have altered price transmission patterns by changing both the supply and demand for meat and triggering policy interventions such as import/export bans and the removal of animals from the market.

3. Methodology

3.1 General remarks

Price transmission studies are usually based upon empirical exercises that test predictions derived from economic theory[3] and show how quickly and to what extent price differentials (net of transactions costs) between two spatially distant markets are eliminated (Obstfeld and Taylor, 1997). Various quantitative techniques have been applied for such purpose. Early analyses, which typically used correlation and simple regression analyses, have been criticised for not recognising the usual nonstationary nature of price data (McNew and Fackler, 1997; Fackler and Goodwin, 2001, Barrett, 1996). More recent studies have addressed inferential problems associated with nonstationarity by adopting time series techniques adequate for dealing with nonstationary and cointegrated data (Goodwin and Schroeder, 1991). However, regression and

[3] Many of the spatial price relationship tests that have been proposed in the literature can be economically justified through the point-location model (Fackler and Goodwin; 2001), which represents the equilibrium conditions inherent in the Law of One Price.

cointegration-based tests suffer from a common limitation: they do not account for transaction costs (McNew and Fackler, 1997; Barrett, 1996).

Transaction costs cover the costs of all factors associated with spatial trade and arbitrage. In addition to transportation charges and freight charges, transaction costs may include risk premia, information-gathering costs, negotiation costs, spoilage, theft, or the costs of maintaining a presence in a regional market. An analogous form of transaction costs may be associated with the exchange of commodities over time – in other words, storage. In this case, the relevant comparison involves prices today and prices at some point in the future and the relevant costs are those associated with storage. In this analysis, we focus exclusively on spatial arbitrage and trade, and thus consider price linkages in the spatial dimension. As noted, the presence of transaction costs may result in nonlinear spatial price relationships (Heckscher, 1916). This occurs because price relationships may involve a combination of different "regimes," corresponding to different trade/no-trade conditions.

Recently, there have been two major developments in the methods used to allow for transaction costs: the parity bounds model (PBM) and threshold cointegration tests. The PBM compares the observed regional price differentials against exogenously predicted transaction costs. It usually distinguishes among three different price regimes: where spatial price differentials are equal, above or below transaction costs (Baulch 1997), and estimates the probability of being in each regime. PBM has however been criticised on the ground that, contrary to threshold cointegration, it is based on static comparisons and thus it does not allow price dynamics to be studied (see Fackler, 1996 for more detail). Also, the PBM requires information on transaction costs, which are usually partially or totally unobservable.

Threshold cointegration models, which we use in our analysis, do not require observations on transaction costs. They are based upon the idea that these costs create a neutral band where spatial price links are weak or even non-existent, because price differentials do not exceed transaction costs thus making spatial arbitrage unprofitable. If price differentials exceed these costs, they do not persist but are rather driven back to transaction costs through the equilibrating adjustments caused by profit-seeking activities of spatial trade. This nonlinear pattern of price adjustment is represented through a combination of different regimes. Threshold cointegration methods are based upon Tong (1978), who introduced nonlinear threshold time series models, and Tsay (1989), who developed a method to model threshold autoregressive processes. Balke and Fomby (1997) extended the threshold autoregressive models to a cointegration framework. They considered a model where there is a discontinuous adjustment to a long-run equilibrium. Specifically, the equilibrium error follows mean-reverting threshold autoregressions outside a given range, while having a unit root inside the range. To estimate these models, Balke and Fomby (1997) suggest a two-step approach whereby threshold parameters are chosen through a grid search that minimises a sum of squared errors (SSE) criterion. In this paper, we follow the proposal by Balke and Fomby (1997).

The literature review presented above reveals that analyses of price transmission have typically been based on parametric approaches. These approaches possibly make assumptions about the true nature of price transmission that are both too strong and inadequate. Because nonparametric techniques do not require any preliminary assumptions about the nature of price behaviour, we are also interested in applying these techniques to a characterisation of spatial price relationships and in comparing the results with those arising from alternative parametric

TAR models. We now present details on both the nonparametric and parametric techniques utilised in this paper.

3.2 *Local polynomial fitting*

Locally weighted regression techniques, which consist of locally approximating a polynomial regression function, have been thoroughly studied in the literature (Cleveland, 1979; Cleveland, Devlin and Grosse, 1988; Fan, 1992; or Fan and Gijbels, 1995). In our application, we use local polynomial fitting techniques to estimate a nonparametric version of a threshold autoregressive model of spatial price differentials (i.e., differences in contemporaneous prices at two different locations). These models have been widely used to assess spatial price transmission (Obstfeld and Taylor, 1997; Goodwin and Grennes, 1998; Goodwin and Piggott, 2001).

Consider a series of scatter points (X_{t-1}, Y_t) where t=1,…,n from a population (X_{-1}, Y), where $Y_t = \Delta X_t = (P_{it} - P_{jt}) - (P_{it-1} - P_{jt-1})$ represents the adjustment in regional price differentials in period t, being P_{it} and P_{jt} the prices of a certain commodity in two spatially separate markets (i and j), and $X_{t-1} = (P_{it-1} - P_{jt-1})$ is the value of the regional price differential in the previous period $t-1$. It is assumed that data are generated from a model $Y = m\left(X_{-1}\right) + \varepsilon$, where $E\left[\varepsilon\right] = 0$, and m is a smooth function differentiable at X_{-1}. The basic idea behind local fitting is to use those observations (X_{t-1}, Y_t) relatively close to a given point x_k to estimate function m at that point, i.e. $\hat{m}(x_k)$. To estimate the whole function $\hat{m}\left(X_{-1}\right)$, the process is then repeated for a number of grid values of X_{-1}. The unknown function may be approximated through a Taylor series expansion that models $m(x)$ locally by a simple polynomial model:

$m(x) \approx \sum_{j=0}^{p} \beta_j (x - x_k)^j$, where $\beta_j = \frac{m^{(j)}(x_k)}{j!}$, and $m^{(j)}$ is the jth derivative of function m. The estimator for $m(x_k)$ is therefore $\hat{m}(x_k) = 0!\hat{\beta}_0 = \hat{\beta}_0$ (Fan and Gijbels, 1996: 58). As noted by Heij et al. (2004: 292), the selection of a polynomial of order $p = 1$ is recommended in most cases and leads to the local linear regression estimator (LLRE).

The local approximation is more accurate for values of X_{t-1} that are closer to x_k compared with more distant points. This motivates the use of weighted least squares to estimate parameters β_j, in order to exclude those observations that are too distant from x_k and weight the non-excluded observations (the neighbourhood of x_k) according to their distance from x_k:

$$\min_{a,b} \sum_{t=1}^{n} \left(Y_t - \beta_0 - \beta_1 (X_{t-1} - x_k)\right)^2 K_t \left(\frac{X_{t-1} - x_k}{h_k} \right) \tag{1}$$

where h_k, the bandwidth, and K, the kernel function, are described in the following lines. The selection of the neighbourhood of x_k to be used in the estimation is done through the bandwidth h_k, which determines the maximum distance of X_{t-1} from x_k for X_{t-1} to be included in the local estimation. As Fan and Gijbels (1996, chapter 3) note, the bandwidth parameter has an important influence on the results. While a large bandwidth can cause a large modelling bias by under-parameterising the regression, an excessively small bandwidth can result in noisy estimates. We selected an optimum constant bandwidth ($h_k = h$) using the cross-validation technique. This widely used technique (Fan and Gijbels, 1996) chooses h to minimise the squared prediction error: $\sum_{t=1}^{n} \left(Y_t - \hat{Y}_t\right)^2$, where $\hat{Y}_t$ is an estimate of the regression function involving the smoothing parameter h. For each observation t, the estimate $\hat{Y}_t$ is obtained by computing the regression function without the t-th observation and predicting Y_t. The model is then validated by examining the prediction error. In our application and following previous research (Mancuso, Goodwin and Grennes, 2003), the predicted value for Y_t was obtained using the Nadaraya-Watson[4] nonparametric regression estimator (Fan and Gijbels 1996: 14-15). The minimisation process requires the computation of the squared prediction error at different bandwidth gridpoints. We searched for the bandwidth h over the range 0.1-2 standard deviations of the independent variable X_{t-1}.

After selecting the neighbourhood of x_k, the weighting of the observations included in the estimation was done through the kernel function K, which has a support contained in $[-1,1]$ and whose role is to down-weigh the contribution of those observations away from x_k. The Epanechnikov kernel was used, which is shown by Fan and Gijbels (1996, chapter 3) to be an optimal weight function. The solution to the problem in expression (1) is given by (see Fan and Gijbels 1996: 94-95):

$$\hat{m}(x_k) = \frac{S_{n,2}(x_k)T_{n,0}(x_k) - S_{n,1}(x_k)T_{n,1}(x_k)}{S_{n,2}(x_k)S_{n,0}(x_k) - S_{n,1}(x_k)^2} \qquad (2)$$

where $T_{n,l}(x_k) = \sum_{t=1}^{n} K_t\left(\frac{X_{t-1} - x_k}{h}\right)(X_{t-1} - x_k)^l Y_t$ and $S_{n,j}(x_k) = \sum_{t=1}^{n} K_t\left(\frac{X_{t-1} - x_k}{h}\right)(X_{t-1} - x_k)^j$.

[4] The Nadaraya-Watson estimator speeds up the estimation of the LLRE, since it is less computationally intensive than the estimator in (1).

Local polynomial fitting techniques are especially useful in situations for which a suitable parametric form of the regression surface is unknown. As noted by Fan (2000), nonparametric techniques allow one to search for appropriate nonlinear forms that best describe the available data, which allows one to reduce the possible modelling biases of parametric techniques. Hence, a local linear regression was fitted to our data before proceeding to the parametric approach, in order to guide the specification of the TAR model. As we discuss in section 4, the nonparametric results suggest that price relationships may be adequately characterised by a three-regime TAR, with a central band of relatively slow price adjustments and outside bands of a relatively fast adjustment.

3.3 Threshold Autoregressive Models

Our parametric estimation approach can be summarised as follows. First, we evaluated the time series properties of the data using unit root and cointegration tests. Unit root tests can be seriously distorted by structural changes in the time series (Zivot and Andrews, 1992; or Perron, 1997). Figure 1 suggests that our price series may be affected by such changes. In order to determine whether price series are nonstationary or whether the apparent nonstationarity is due to a structural break, we used Perron's (1997) sequential test.[5] Second, cointegration among prices was tested using the Johansen (1988) test.[6] In order to capture spatial price dynamics we then estimated a threshold autoregressive model as described in the following paragraphs.[7]

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[5] The truncation lag parameter is selected using the general to specific recursive method (Perron, 1997).

[6] The lag length of the vector autoregressive model was selected to ensure non-autocorrelation of the residuals and to minimise the AIC criterion. The deterministic terms of the test were selected according to the LOP.

[7] As a suggestion of an anonymous referee, a Momentum-TAR (MTAR) model was also considered. However, TAR models were found to bear more resemblance to the nonparametric regressions of price differentials than MTARs.

A simple autoregressive model (AR) of price differentials can be represented as: $Y_t = \beta X_{t-1} + e_t$, where, as noted, Y_t is the adjustment in regional price differentials in period *t*, X_{t-1} is the value of the regional price differentials in the previous period $t-1$, and e_t is a white noise error term. Usually, analyses of spatial price behaviour take the magnitude of regional price differentials $\left(X_{t-1}\right)$ as the variable that determines regime-switching (Serra and Goodwin, 2004; Mancuso, Goodwin, and Grennes, 2003). Different regimes are represented by different values of the parameter β. A three-regime TAR was estimated to allow for asymmetries in price behaviour:

$$Y_t = \begin{cases} \beta^{(1)} X_{t-1} + e_t^{(1)} & if \quad -\infty < X_{t-1} \leq c_1 \\ \beta^{(2)} X_{t-1} + e_t^{(2)} & if \quad c_1 < X_{t-1} \leq c_2 \\ \beta^{(3)} X_{t-1} + e_t^{(3)} & if \quad c_2 < X_{t-1} \leq +\infty \end{cases} \tag{3}$$

where c_1 and c_2 represent threshold parameters that define the regimes. The TAR model can be estimated using sequential iterated least squares regression in two steps (see Balke and Fomby, 1997). The aim of the first step is to estimate threshold parameters c_1 and c_2 through a grid search. The first or lower threshold was identified by searching over the space defined by the minimum and the median of the lagged price differentials, while the upper threshold was identified after searching over the range that goes from the median to the maximum value of the lagged price differentials. These searches were restricted in order to ensure an adequate number of observations in each regime. For a given pair $\left(c_1, c_2\right)$, $\beta^{(1)}$, $\beta^{(2)}$, and $\beta^{(3)}$ can be determined

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Following Fan's (2000) recommendation to use nonparametric results to enhance the parametric analysis, we have chosen to use TAR models.

through the OLS regressions of Y_t on X_{t-1} for each subsample. From this estimation, the residual sum of squares was derived giving $\hat{\sigma}(c_1,c_2)=\sum_{t=1}^{n}\hat{e}_t(c_1,c_2)^2$. The aim of the grid search is to maximise a standard *F* statistic for a linear AR against the alternative of a TAR: $F=\frac{\tilde{\sigma}^2-\hat{\sigma}^2(c_1,c_2)}{\hat{\sigma}^2(c_1,c_2)}n$, where n represents the number of observations, $\hat{\sigma}^2(c_1,c_2)$ stands for the error variance of the TAR model, being $\tilde{\sigma}^2$ the error variance of the AR model. Hence, in the second step of the process, the estimates of c_1 and c_2 were obtained as $\left(c_1,c_2\right)=\underset{c_1,c_2}{\arg\min}\quad \hat{\sigma}^2(c_1,c_2)$, which is equivalent to maximising *F*. The *F* test for the significance of the differences in parameters across regimes was performed. Because this test does not have a standard distribution, its *p*-value is determined following the method provided by Hansen (1997). If the three-regime TAR was found not to be significant against the AR model, a two-regime TAR was estimated and tested against a standard AR model.

4. Empirical Implementation

To assess price transmission processes between EU pork markets, weekly producer prices, expressed in Euro/100 kg, for Germany, Spain, France, and Denmark over the period 1994-2004 were used. Prices were obtained from the European Commission's publication *Agricultural Markets - Prices*.

Figure 1 here

Figure 1 shows that these countries' price series follow very similar patterns, implying important price transmission processes across EU markets. The graphs also suggest three main peaks for all price series in 1996, 1997, and 2001 that coincide with the three important animal disease outbreaks that affected the EU meat sector in the period studied: first, the 1996 official announcement that BSE can cause Variant Creutzfeldt-Jakob disease (vCJD), which involved an important shift in meat demand from beef to pork; second, the 1997-98 classical swine fever epidemic in the Netherlands, which involved the removal of millions of pigs from the market just as demand for pork was rising as a result of BSE concerns; and third, the foot and mouth disease in the UK, which caused a switch from beef to pork.

In light of its nonparametric nature, the local linear regression method is best interpreted by graphical representation, which favours relatively simple model specifications. We opted for pair-wise analyses. Pair-wise analyses are very common in the price transmission literature and are a natural choice for studying price linkages since arbitrage conditions, which imply integration, should hold for any pair of prices. Goodwin and Piggott (2001) suggested defining pairs of prices consisting of a central market price (P_{it}) and another market price (P_{jt}), where the central market is the largest market in terms of volume.[8] In our analysis, Germany plays this role as it is the most populated EU-15 country, the largest importer of pigmeat, a relevant exporter and has one of the highest per capita consumption levels of pork. A market with these characteristics would be expected to lead price formation. Weak exogeneity tests presented in Table 2 confirm this hypothesis.[9] For each pair of prices, a local linear regression was fitted. Results[10] are presented in Figures 2 to 4 where, for comparison, the predicted values of the TAR models are also included.

[8] Prices are in levels. The analysis was replicated using prices in logs and the results were very similar.

[9] Other studies have also provided evidence of the relevance of consumer markets (or net importer markets) in price formation processes relative to producer markets (Serra and Goodwin, 2004).

[10] As explained in the methodology section, the bandwidth used in the local linear regression technique is searched between 0.1 and 2 standard deviations of the independent variable using the cross-validation technique. The

Figures 2 to 4 here

Nonparametric results suggest that deviations from long-run equilibrium relationships tend to be arbitraged away in a nonlinear fashion. Specifically, they suggest the existence of a central neutral band or area of price differentials of relatively slow price adjustments and outside bands of quicker price movements. The fact that out-of-band adjustments are not symmetric suggests the presence of asymmetries in the process of price adjustment. Within each regime or band, price dynamics seem to follow a linear pattern. This suggests that price relationships may be adequately characterised by a three-regime TAR, which we estimated.[11]

Table 2 here

Now shifting to the parametric analysis, Table 2 presents test statistics that evaluate the time series properties of the data. The sequential modality of the Perron (1997) test that allows for both a change in the intercept and the slope[12] suggests that the null hypothesis of a unit root cannot be rejected for any series. It is interesting to note that all selected breakpoints coincide with the classical swine fever outbreak in 1997 that involved the removal of millions of pigs

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optimum result of the grid search is $h = 1.1,\ 1.5,\ 1.9$ for the Spain-Germany, France-Germany, and Denmark-Germany models respectively. It should be noted here that the corrected Akaike information criteria (Hurvich and Simonoff, 1998) was also used as an alternative method for bandwidth selection and the results derived were very similar.

[11] As a suggestion of an anonymous referee, we also estimated a TVECM for each pair of prices. Results, which are available upon request, do not substantially differ from the ones derived from TAR models.

[12] Results for the other modalities of the test are not presented, but are available upon request. These results do not differ however from the ones offered here.

from the market at a time when demand was strongly growing. Johansen tests provide evidence of stationary linear long-run relationships among all pairs of prices.[13]

Table 3 here

Results derived from the estimation of the threshold autoregressive models are presented in Table 3. While a three-regime TAR was statistically significant against an AR model for Germany-Denmark and Germany-France pairs of prices, price relationships between Germany and Spain were best captured by a two-regime TAR. Hence, asymmetries characterise price transmission processes between Germany and France and Denmark, while a symmetric relationship best represents the Germany-Spain price linkage. In-band parameter estimates are not statistically different from zero, which is consistent with the existence of transaction costs that cause price adjustments to take place only after price differentials exceed a certain minimum amount. Out-of-band parameters are all negative and significantly different from zero, suggesting that price differentials exceeding threshold values are arbitraged away. The three-regime TAR models suggest that asymmetries grant a certain advantage to Germany over Denmark and France; while negative price differentials are quickly adjusted, positive price differentials are corrected more slowly. In contrast, the two-regime TAR shows that price transmission processes leave Spain on equal terms with Germany: the out-of-band adjustment has the same speed independent of whether price differentials are positive or negative. The advantage of Germany over France and Denmark but not over Spain, might be explained both by the greater physical distance between Spain and Germany, which could reduce trade flows, and

[13] As a suggestion of an anonymous referee we tested for structural changes in the cointegrated autoregressive models following Hansen and Johansen (1999). Results, which are available upon request, suggest stability of parameter estimates throughout the period studied.

by the fact that both Germany and Spain are the leading EU pork producers and thus can compete under more similar conditions.

We find transaction cost bands to be the largest for the Germany-Denmark model. While Denmark is a leading EU pork exporter, Germany is the primary importer. Hence, prices in Germany probably carry a significant transaction cost charge not present in Danish prices. Thus, transaction cost bands are expected to be wider between an exporter and an importing country, than between two importing markets. Consistent with this hypothesis, transaction costs between France (which, as Germany, has a negative balance in the intra-EU pork trade) and Germany are considerably below the transaction cost band derived in the Germany-Denmark model. Spain has a positive balance in the intra-EU pork trade, being thus a net exporter. The transaction cost band for the Germany-Spain model is unexpectedly small. This could be explained by a less intensive commercial flow between Germany and Spain relative to the trade between Denmark and Germany, which may reduce the adequacy of the interpretation of the thresholds as transaction costs band (Goodwin, Grennes and Craig, 2002).

Direct comparison between parametric and nonparametric techniques (Figures 2-4) suggests that both techniques support the hypothesis that prices are transmitted across spatially separate EU pig markets. Additionally, both models suggest that there exists a range of price differentials where equilibrating price adjustments may be less intense, which is compatible with transaction costs. In spite of the similarities between the two models, important differences arise. First, because local linear regression techniques do not assume homogeneous transaction costs across individuals, the transition from one regime to another is allowed to be smooth, which contrasts with the sharp and discontinuous transitions implied by the parametric techniques. A second difference is apparent between the parametric within-band price behaviour and the analogous values predicted by the nonparametric techniques. Where TAR models suggest a market in equilibrium, local polynomial fitting shows that a price adjustment takes place. Furthermore, this adjustment can be relatively quick, as is the case with the Germany-France model. Hence, the LLRE implies that markets are more strongly interconnected than what one would conclude from simple observation of the TAR model. This is compatible with Mancuso, Goodwin and Grennes (2003) and Serra, Goodwin and Mancuso (2005).

Third, nonparametric techniques suggest that TAR models, in that they are estimated with a limited number of regimes, may have difficulty in capturing the true nature of price relationships. Though the true nature of the price transmission is unknown, nonparametric techniques, in that they are more flexible and do not carry the potential for specification biases as a result of an inappropriate parametric assumption, are expected to represent true price linkages better. In the Germany-Spain model, for example, big positive price differentials above €15, accelerate the speed of price adjustment. This acceleration is not captured by the parametric method, suggesting that another regime might be necessary if the TAR is to correctly represent true price relationships. However, a three-regime TAR for this pair of markets was estimated and rejected against a linear AR. Following the same argument and as a general rule, for big positive price differentials the speed of price adjustment suggested by parametric models is slower than that derived from the LLR. Conversely, for negative differentials, the slope of the TAR regression is steeper than (or coincides with) the nonparametric one.

5. Concluding Remarks

The economic literature on price transmission has argued that spatial price adjustments may only take place when regional price differentials exceed a certain minimum amount. These frictions reflect the presence of transaction costs, imperfect information and other barriers to arbitrage. Threshold parametric models have been widely used to capture nonlinear price adjustments. We argue that these techniques may involve too restrictive or unrealistic assumptions about the true nature of price behaviour and thus that nonparametric techniques such as local polynomial fitting may offer significant advantages. Additionally, nonparametric methods are especially useful in those situations in which a suitable parametric form of the regression surface is unknown. The use of nonparametric techniques to guide the specification of the parametric model may reduce the likelihood of model biases.

We used nonparametric methods to analyse price transmission within EU pork markets for the period 1994 to 2004. We used weekly, country-level price series for Germany, Spain, France, and Denmark, representing the four leading EU pork producers and traders. Results obtained with nonparametric techniques (LLR) and more restrictive parametric threshold models (TAR) were compared. Both methodologies suggest that EU pork markets are interrelated in that price shocks are spatially transmitted. However, LLR techniques often suggest a higher degree of price transmission than that implied by TAR models. Specifically, while TAR models support the existence of a band of price differentials within which no adjustment takes place, LLR results imply price adjustments even within thresholds, albeit at different rates. Hence, according to the LLR, markets are more strongly interconnected either through information or trade flows. Also, TAR models seem to have difficulty in capturing the true speed of price changes for out-of-band price differentials. Whereas an increase in out-of-band price differentials changes the slope of the nonparametric regression to make transmission processes quicker, TAR models assume the speed of adjustment is constant.

Both methodologies suggest that price transmission processes between Germany and Denmark and France are asymmetric and grant a certain advantage to Germany. They also imply that price relationships between the two leading EU pork producers (Spain and Germany) are symmetric and leave these two countries more on equal terms. Whereas negative price differentials between Germany and Denmark or France are quickly corrected, and positive price differentials are arbitraged away at a slower path, the out-of-band adjustment for the Germany-Spain model has the same speed independently of the sign of the price differentials.

These results might be explained by the larger physical distance between Spain and Germany and by the traditionally less intensive trade flow relative to those between the latter country and Denmark or, to a lesser extent, France. Moreover, both Germany and Spain are leading pork producers, which allow them to compete more on equal terms. Another interesting result, from our perspective, is that transaction costs bands are wider when we look at price transmission between the largest importing and exporting countries. This indicates that in the importing country, pig prices include larger transaction costs than in the exporting country.

Obviously, the interpretation of our results has to be restricted to the pig sector and the sample period analysed. In this sense, further research is needed. Apart from applying new methodological refinements, which, with no doubt, will appear in coming years, this study should be extended to other meat sectors in Europe with two main objectives. The first one would be to try and replicate our findings relating distances, trade flows, transaction costs and spatial price transmission. Second, as different meat sectors are supported with different degrees of market intervention, comparing results from new studies could provide valuable information about the potential impact of policy measures on spatial price transmission.

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[A few references have been removed as you can check 56 references is rather a lot! Is there any scope for culling a few of them?]

Table 1. Trade flows (in current euros, thousands)

Exports from	year	Imports by				
		Germany	**Denmark**	**Spain**	**France**	**EU-15**
Germany	1995	0	8,132	5,396	44,044	299,375
	2003	0	65,934	19,993	73,561	1,182,498
Denmark	1995	424,194	0	6,920	200,825	1,376,279
	2003	550,157	0	7,746	74,762	1,701,310
Spain	1995	56,774	5,863	0	88,486	315,415
	2003	162,909	21,625	0	320,067	966,883
France	1995	123,601	14,428	32,683	0	634,245
	2003	68,709	11,712	38,508	0	689,104
EU-15	1995	2,289,515	70,272	147,068	787,222	6,498,309
	2003	2,002,046	159,141	196,640	798,552	8,258,635

Table 2. Time series properties of the data

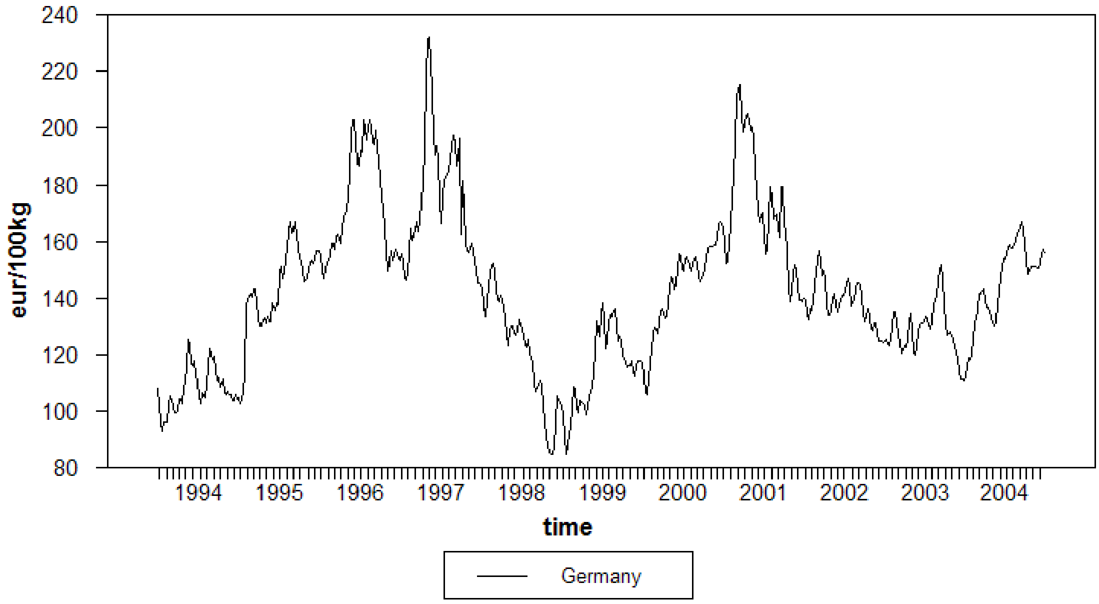
Unit Root Tests			
Price	Perron (10% significance level)	Break date	
Spain	-4.24 (-4.82)	1997:09:01	
Germany	-4.61 (-4.82)	1997:09:15	
Denmark	-3.23 (-4.82)	1997:05:15	
France	-3.91 (-4.82)	1997:09:29	
Johansen Tests			
Model	λ max, r=0 (critical value at 10%)	λ max, r=1 (critical value at 10%)	Weak exogeneity test (p-value) H_0: P_{jt} is weakly exogenous
Spain – Germany	33.56 (10.29)	7.23 (7.50)	19.46 (0.00)
Denmark – Germany	75.88 (10.29)	7.30 (7.50)	56.93 (0.00)
France – Germany	26.15 (10.29)	7.37 (7.50)	12.23 (0.00)

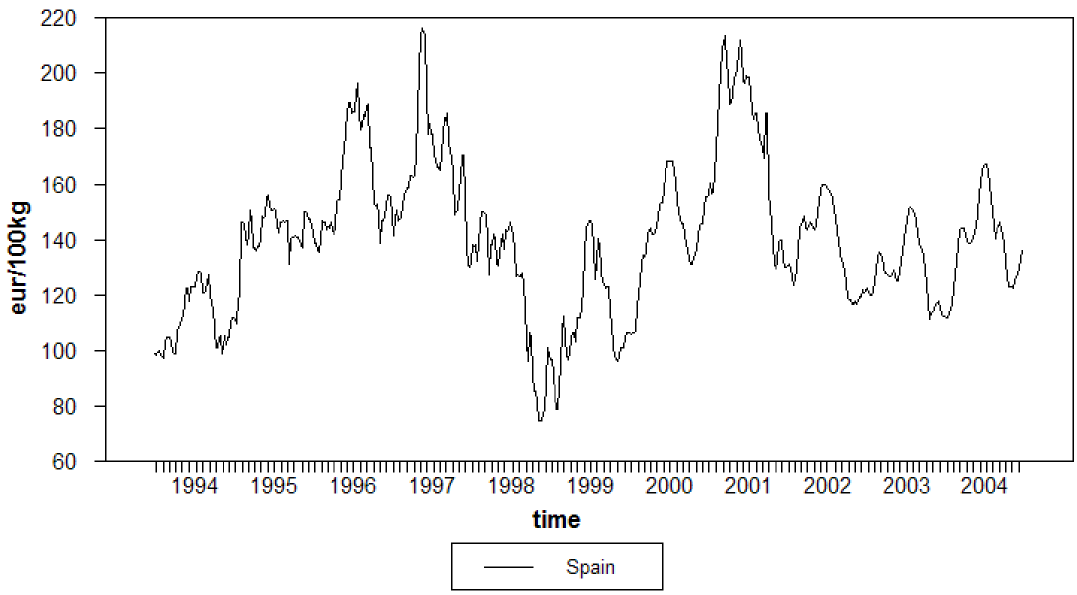
Table 3. TAR model parameter estimates

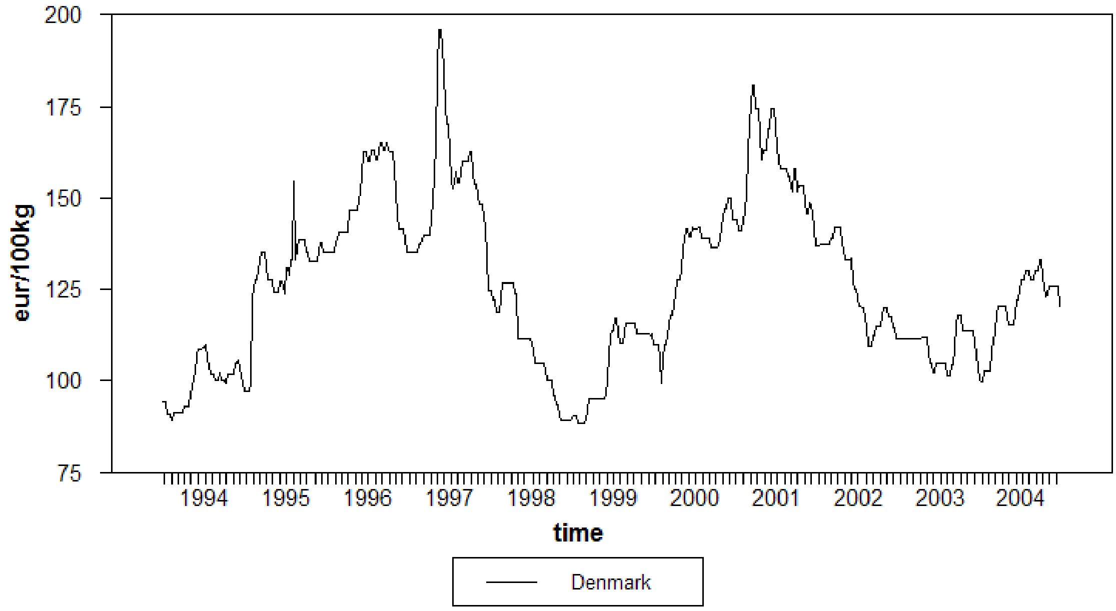
Markets	Thresholds and F-test			TAR parameters		
	Lower threshold c_1	**Upper threshold** c_2	**F-test (*p*-value)**	**First regime** $\beta^{(1)}$ **(std error)**	**Second regime** $\beta^{(2)}$ **(std error)**	**Third regime** $\beta^{(3)}$ **(std error)**
Germany-Denmark	2.41	33.3	18.30* (0.02)	-0.50* (0.16)	-0.01 (0.01)	-0.08* (0.02)
Germany-Spain	4.77		2.46* (0.00)	-0.10* (0.02)	0.14 (0.15)	-0.10* (0.02)
Germany-France	-0.08	11.40	20.24* (0.01)	-0.36* (0.07)	0.02 (0.04)	-0.11* (0.02)

Note: An asterisk (*) denote statistical significance at the 5 per cent significance level.

FIGURE 1. Price series Germany, Spain, Denmark and France






200
175
150
125
100
75
eur/100kg
1994
1995
1996
1997
1998
1999
2000
2001
2002
2003
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time
Denmark

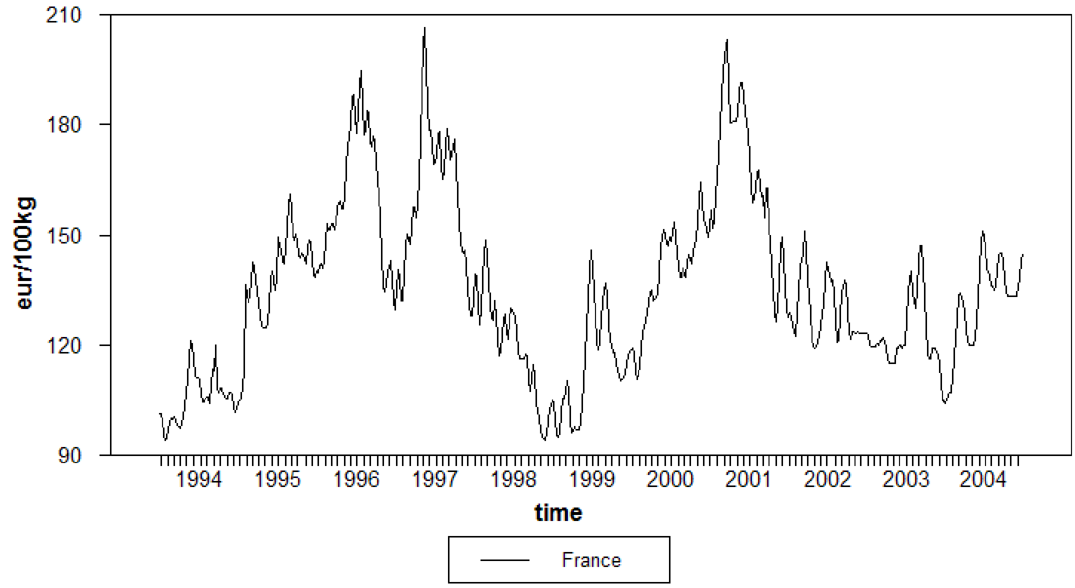
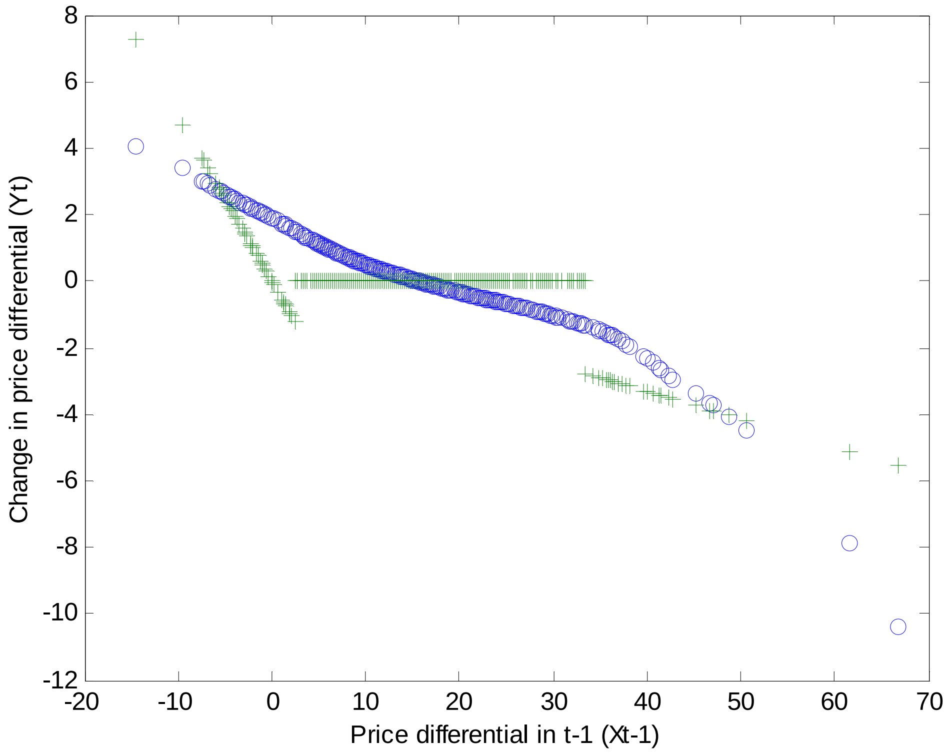

210
180
150
120
90
eur/100kg
1994
1995
1996
1997
1998
1999
2000
2001
2002
2003
2004
time
France

FIGURE 2. Nonparametric and TAR model: Germany-Denmark

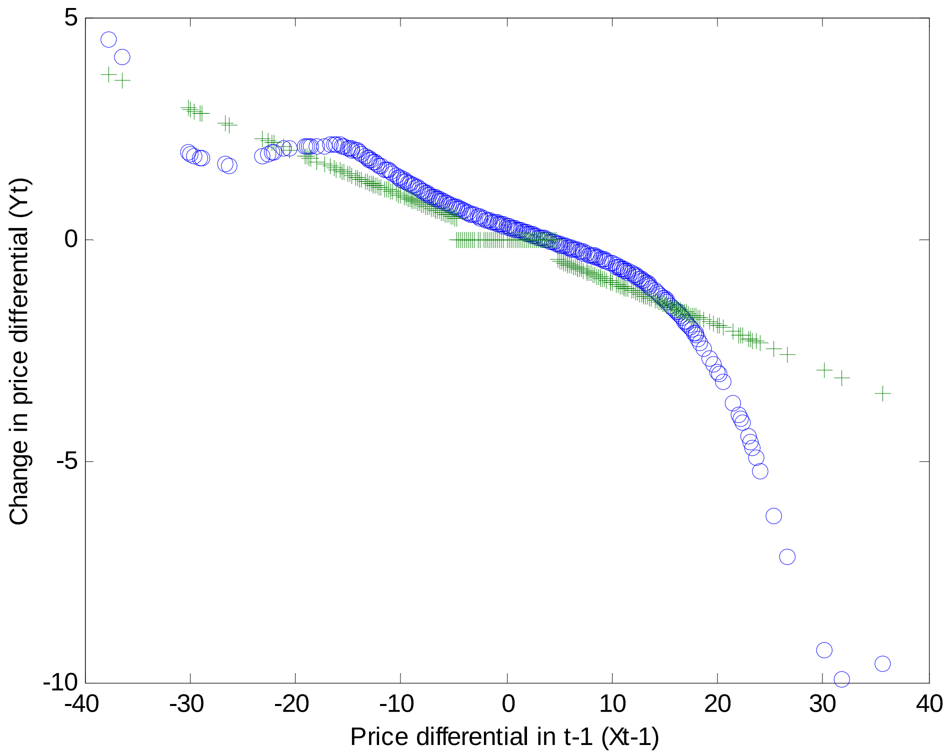


where:

○ represents the LLRE model

+ represents the TAR model

FIGURE 3. Nonparametric and TAR model: Germany-Spain

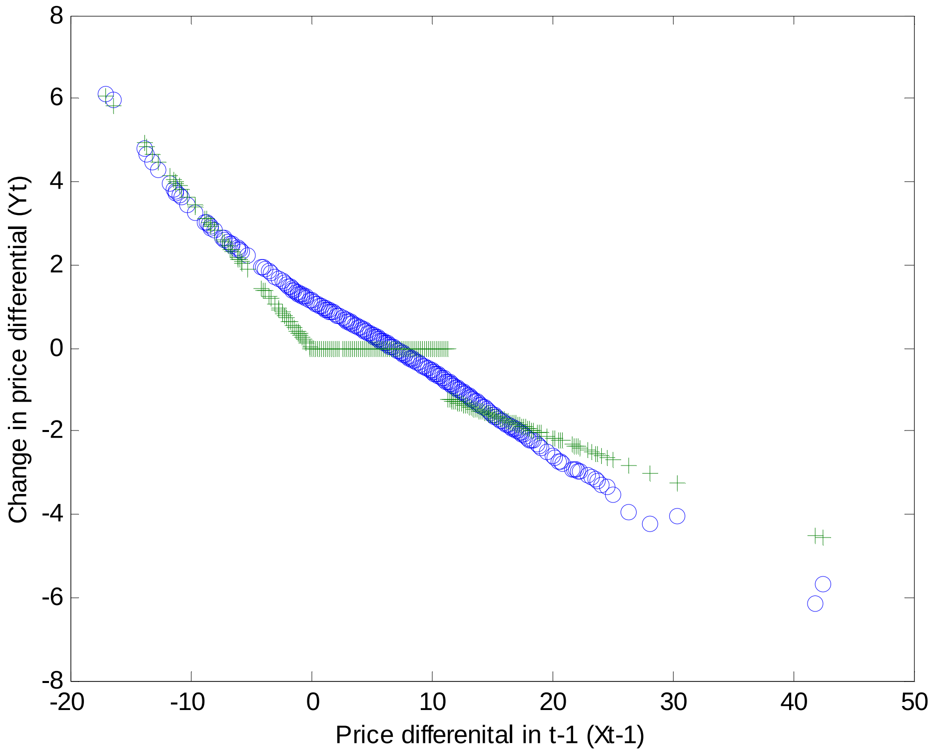


where:

○ represents the LLRE model

+ represents the TAR model

FIGURE 4. Nonparametric and TAR model: Germany-France



where:

○ represents the LLRE model

+ represents the TAR model