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Buyer Power in U.K. Food Retailing: A 'First-Pass' Test

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Buyer Power in U.K. Food Retailing: A 'First-Pass' Test*

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Abstract

The potential existence of buyer power in U.K. food retailing has attracted the scrutiny of the U.K.'s anti-trust authorities, culminating in the second of two comprehensive regulatory inquiries in recent years. Such inquiries are authoritative but correspondingly time-consuming and costly. Moreover, detection of buyer power has been dogged by the paucity of reliable evidence of its existence. In this paper, we present a simple theoretical model of oligopsony which delivers quasi-reduced form retailer-producer pricing equations with which the null of perfect competition can be tested using readily available market data. Using a cointegrated vector autoregression, we find empirical results that show the null of perfect competition can be rejected in seven of the nine food products investigated. Though not conclusive on the existence of buyer power, the proposed test offers a means via which the behaviour of the retail-producer price spread is consistent with it. At the very least, it can corroborate the concerns of the anti-trust authorities as to whether buyer power is potentially one source of concern.

KEYWORDS: buyer power, cointegrated VARs, U.K. food industry

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1. INTRODUCTION

In common with many national retail food markets in Europe, the rising degree of market concentration in the UK food sector has been a cause of concern to both consumer groups and food producers in recent years. By 2006, the four leading food retailers in the UK had a combined share of the grocery market of around 75 per cent, with the largest of these accounting for around one-third of all food sales (Office of Fair Trading, 2007). The issue has also aroused the attention of the UK's principal anti-trust authority, the Competition Commission, which has undertaken two statutory inquiries into food retailing in the last decade (Competition Commission, 2000, 2008). A key motivation underlying their scrutiny of the supermarkets was:

> ' . . . [the] public perception of . . . an apparent disparity between farm-gate and retail prices . . . which is seen as evidence by some that grocery multiples were profiting from the crisis in the farming industry'. Competition Commission (2000), vol.1, p.3

Statutory inquiries are expensive in terms of time and resources and are thus not undertaken without good grounds for doing so. This paper offers one possible approach based on a 'first-filter' test of price data that may be used as part of the preliminary analyses into the presence of buyer power in such markets. Contingent on assumptions relating to functional form and technology, we reject the null hypothesis of perfect competition in seven out of nine specific food groups investigated. While not conclusive that buyer power is the primary cause of widening margins between retail and farm price spreads, as a first-pass test, it suggests that buyer power is a potential candidate among others.

The paper is structured as follows. In Section 2, we provide some background material to the UK Competition Commission's concerns about buyer power exercised by dominant food retailers and the motivation for our testing procedure as a 'filter.'[1] In Section 3, we outline the theoretical model that underpins our conceptualisation of a vertically-related market. The model is by no means intended as a detailed description of the UK food chain, but it does serve as a useful device for characterising how prices are transmitted in such a market, albeit in simplified form. It also forms the basis for determining the appropriate econometric approach and the interpretation of the key variables used to identify the existence of oligopsony power. Section 4 describes the data that are used in the testing procedure while Section 5 shows how the test for oligopsony power

[1] The issue of countervailing power in vertical markets, while of growing interest in the academic literature (see, for example, Connor *et al.* (1996), Chen (2003)) and to policy-makers, is not examined here.

can be implemented using tractable techniques of time series analysis. The results are outlined in Section 6 and we offer some concluding comments in Section 7.

2. BACKGROUND

(a) Concerns of Buyer Power by UK Food Retailers

A key issue highlighted in the 2000 report was the extent to which retailers can exert buyer power over their suppliers and the potential impact this has on consumer choice and competition in the food chain (Competition Commission 2000). The belief that buyer power existed and was potentially being abused had been one of the primary reasons for instigating the report. However, collating evidence of buyer power during the investigation had not been easy, not least because of the large number of ways that it may be applied.[2] The report concluded that while there was only limited potential for abuse of seller power with respect to consumers, there were grounds for significant concern regarding food retailers' relationships with suppliers, highlighting 27 oligopsonistic practices that specifically gave cause for concern.[3] Despite the subsequent imposition of a Supermarket Code of Practice in October 2001 effectively outlawing such practices, concerns over buyer power remain and were not allayed by the findings in interim reports on the Code of Conduct in 2004 and 2005. Such concerns formed the basis for the Office of Fair Trading's recent decision to refer the supermarkets to a further Competition Commission inquiry (Competition Commission, 2008).

Concerns were most cogently illustrated by the nature of trading between retailers and suppliers of "fresh" food products in that '[g]enerally, suppliers of fresh produce appear to be most dependent on their largest main party customers [big supermarkets] for their sales' (Competition Commission, 2000, 11.15, p232) and '. . . most suppliers of fresh fruit and vegetables meat and poultry . . . appear to concentrate on trade with a limited number of suppliers (often four or less)' (Competition Commission, 2000, 11.8 p.231). Indicative figures from the food industry underline this reliance with some 75%[4] of total UK output of apples and 80%[5] of total UK fresh potato output being sold to the supermarkets. Around 65% of liquid milk sales are accounted for by the main food retailers (KPMG, 2002).

[2] Buyer power can affect almost all aspects of the retailer-supplier contractual arrangement including the timing, form and level of the negotiated payment as well as shifting risks between the parties (see GfK (2007) for examples from the UK food industry).

[3] These practices primarily related to the retailers' interactions with suppliers rather than food manufacturers. See Table 2.14, pp.140-143 of the Competition Commission's 2000 report for details.

[4] English Apples and Pears Limited, personal communication

[5] Yakovleva and Flynn (2004)

With respect to meat products, the data are more indirect in that they relate to consumption of meat via the retail sector as a whole rather than the supermarkets alone, though given their share of consumer markets, the figures are informative of the likely dominance in the procurement market. With this caveat in mind, the data show that 85% of beef is consumed via the retail sector, with the corresponding figures for pork and lamb being 81% and 90% respectively.[6]

Undertaking regulatory inquiries is time consuming and expensive and establishing detailed empirical evidence of the existence of buyer power is problematic especially as it can occur in many different forms (Dobson, 2005). The Competition Commission, however, felt it gathered enough evidence to show that buyer power existed. Given the range of different practices that can characterise retailer-supplier relationships and that may be the mechanism via which buyer power is exerted, to what extent though can these findings be foreshadowed using less expensive means based on available market data? In this paper we offer an approach that provides a simple, inexpensive first-filter test for exploring the UK food retail sector that relies on price data and acts as a precursor to potentially more detailed analysis. This test fits as mid-way between two different means of addressing the issue of buyer power. On the one hand, indicative measures often rely on anecdotal accounts, small-scale surveys of the parties involved or at a more representative level, summary measures of concentration. Relating simple measures of concentration to the existence of selling power has long been recognised as of limited value and the same is true for buying power (Clarke *et al*, 2002). For example, the high level of concentration evident in the UK food retailing sector, coupled with the high profits they report, is not necessarily indicative of the exploitation of buyer power. On the other hand, there is a spectrum of econometric approaches that may be employed to detect buyer power that encompass a wide range of challenges including accessing data, the level of dis-aggregation and so on.

Where estimation is based upon price data alone, such as in orthodox price transmission studies (e.g. London Economics, 2004), the veracity of anti-trust inference is undermined by the reduced-form nature of the price regressions employed (Hoehn *et al.* 1999, p.113). Although structural econometric models address this issue of 'measurement without theory' directly, they are often confounded by data limitations and methodological shortcomings relating to market definition and the validity of the behavioural assumptions employed (Baker and Bresnahan, 1992). In these circumstances, a simple test derived from economic theory detecting the potential existence of buyer power offers some appeal (see, for example, Raper, Love and Shumway, 2007), and it is in this regard that this paper seeks to make a contribution. Specifically, we provide one

[6] Meat and Livestock Commission, personal communication.

such possible test by devising a simple restricted quasi-reduced form model of price formation at retailer and supplier levels in which rejection of the null hypothesis of perfect competition can be readily tested using widely available market-level data. While the approach does not aim to derive an explicit measure of buyer power, it does provide grounds for further testing for its potential existence. In doing so, it emphasises the test's 'path-finder' role and is suggestive of the desirability of further scrutiny such as that undertaken by regulatory authorities.

There are two aspects to this ‘first-pass’ test. First, what is the nature of the specific concern(s) being addressed by the anti-trust authority? If there is concern over buyer power, then whatever the nature of the specific contractual arrangements between retailers and suppliers, the spread between retail and farm-level prices should be increasing, as is implied by the quote in the Introduction. Second, with this as background, given some limiting but widely-used assumptions relating to functional form and the technology characterising the vertical food chain for a given commodity, a basis for testing can be established. The null hypothesis is that, in the presence of exogenous shocks, if the retail food sector is perfectly competitive, all changes in the retail-producer price spread are accounted for by marketing costs. Rejection of the null hypothesis suggests that the behaviour of the retail-producer price spread is consistent with a number of possible causes including the presence of buyer power.

Of course, there are other factors that could cause the retail-food price spread to behave as it does but the simple test proposed here suggests that the anti-trust authorities have ‘good’ reason to pursue an investigation. In terms of the limiting assumptions, the role of technology (fixed versus variable proportions) may be an issue and can affect the price spread in the face of exogenous shocks, though fixed proportions is likely to be a ‘reasonable’ assumption for many product groups.[7] Moreover, it is well-known in the industrial organisation literature that functional form may be an issue in identifying the exercise of buyer power. For example, with constant elasticity demand and supply functions, even with the existence of buyer power, the gap between the retail-farm price spread remain constant.[8] However, our approach here is to start from the premise that widening margins have been already been identified as existing and thus our approach is to determine whether buyer power can be considered as one possible explanation of this and thus rejecting the null hypothesis would suggest more in-

[7] McCorriston *et al.* (1998) also show that the behaviour of the retail-farm price spread is only marginally affected by the fixed versus variable proportions assumption and that buyer/seller power is likely to dominate the issue of technology. See also Sexton and Lavoie (2001) for a general critique of the variable proportions assumption in characterising links in the vertical chain.

[8] See McCorriston *et al.* (1998) for the oligopoly case and Weldegebriel (2004) for the oligopsony case.

depth consideration by regulatory authorities is warranted. In identifying buyer power as a possible cause, we are not stating that it is the *only* possible cause merely that its potential presence would warrant further, more detailed investigation.

In sum, the model here should be interpreted as a 'first-pass' test and used to complement investigations initiated by regulatory authorities. It confirms the possibility that the exercise of buyer power is a candidate in characterising the behaviour of the retail-producer price spread. It is relatively inexpensive in terms of time and data requirements and familiar to applied economists who focus on the behaviour of prices in related markets. It does not, however, prove the existence of buyer power nor, if buyer power does exist, detail the extent of buyer power. Nevertheless, as a first pass test, it is both potentially informative and useful.

(b) Related Literature

In terms of the academic literature, the test proposed here lies between two related fields in the industrial organisation literature. At one end is the estimation of structural models in the context of the new empirical industrial organisation literature. Bresnahan (1989) provides an overview. The key feature of this methodology is the use of exogenous shocks (such as exogenous shifts in the demand or supply functions) in order to identify the presence of buyer/seller power more generally. From this, one can retrieve a measure of the aggregate conjectures representing the degree of buyer/seller power in a specific market. In the approach followed here, we also employ exogenous shocks as a means to detect the potential for buyer power. Extensions of this methodology to the case of buyer power between the downstream food sector and producers have been explored in the agricultural economics literature. Just and Chern (1980) were amongst the first to develop this methodology for identifying buyer power with reference to the US tomato industry.

At the other end is the theoretical and empirical literature on the incidence of policy changes (such as tax changes) or other shocks since the incidence of taxes may differ in the presence of buyer/seller power. There is a substantive theoretical literature on the issue of incidence on the presence of buyer/seller power (see, for example, Bulow and Pfleiderer (1983) as an early example). McCorriston *et al*. (1998) extend this analysis in the context of the traditional retail-farm spread model while Weldegebriel (2004) further extends this theoretical framework to explore the role of oligopsony and the spread between retail and farm level prices. From the empirical side, Feuerstein (2002) and Delipalla and O'Donnell (2001), represent recent examples using the role of exogenous shifters to detect the relationship between seller power and incidence.

The approach followed here relates to these theoretical and empirical strategies in that we exploit the presence of exogenous shocks in order to identify the presence of buyer power based on a theoretical model of the incidence of shocks on both upstream and downstream prices. As we explain below, the detection of buyer power simply depends on how these shocks affect *both* sets of prices. While the simplicity of the approach does not allow us to retrieve an empirical estimate of the degree of buyer power, the trade-off does circumvent some of the obstacles inherent in the estimation of structural econometric modelling and the difficulties associated with the interpretation of estimated conjectures.

More specifically, in the framework we present, the difference (or spread) between prices at different marketing levels can be attributed solely to marketing costs under competitive conditions. In other words, shocks impact on prices at each marketing level equally. If buyer power exists then the spread between retail and producer supply prices behaves differently since price setting by the sector with buyer power will be reflected in the mark down that the firms can earn, and so affects the spread. Hence, as we show in section 2, where buyer power exists, with albeit limiting but widely-used assumptions, market shocks have a differential impact at each stage in the marketing chain and thus determine the behaviour of the spread between prices at different vertical levels in addition to marketing costs. In effect, shocks to the underlying supply and demand functions are mediated through buyer power parameters and thus give rise to predictable effects on the spread. In the absence of buyer power, the effect of shocks is common at all vertical market levels so that the spread is simply determined by marketing costs.

In what follows, we develop a model of price transmission in a two-level (i.e. retail and farm-gate) vertical market that explicitly allows for shocks in both the demand and supply functions for a food product. The theoretical framework delivers an equation for the determination of the price spread in which the impact of these shocks appears with definite sign in the presence of oligopsony power. This provides the theoretical basis for a simple empirical test of the presence or otherwise of perfect competition. We apply our approach to data from nine basic food groups (such as apples, eggs and beef) in the UK food industry. Results strongly point to the presence of a single relationship between the retail and producer prices for the majority of products, with the empirical test rejecting the null of perfect competition in seven out of nine of products at conventional levels of statistical significance. Furthermore, coefficients on the exogenous shifters are signed according to the predictions in the theoretical model in thirteen out of eighteen cases. Overall, the results suggest that the spread between producer and retailer prices is not consistent with perfectly competitive behaviour and thus might be caused by, at least as a candidate amongst other factors, the existence of oligopsony power in UK food retailing, a finding that is consistent with the conclusions of the Competition Commission (2000) investigation. Whilst by no means a substitute for legalistic scrutiny of accounts and contracts in anti-trust cases, our statistical test offers a complementary indicator of anti-competitive behaviour that may be easily applied in other similar analyses.

3. THEORETICAL MODEL

In this section, we outline a simple framework that delivers a formal test of perfect competition that we use to motivate the empirical analysis. The model can be readily adapted to account for oligopoly power and the co-existence of oligopoly and oligopsony power.[9] The model is static and ignores repeated interaction between downstream firms and suppliers over time, a simplification that allows for ease of interpretation of the shifters as a ‘first-pass’ test.. The demand function for the processed product is given by:

$$Q = h(R, D) \tag{1}$$

where R is the retail price of the good under consideration and D is a general demand shifter. The supply function of the agricultural raw material is given by (in inverse form):

$$P = k(A, S) \tag{2}$$

where A is the quantity of the agricultural raw material supplied to retailers by farmers and then resold by retailers to consumers as Q and S is the exogenous shifter in the farm supply equation.

In accordance with the findings of the Competition Commission (*op. cit.*), the source of power in the food chain is given to be at the retail level in the form of buyer power. For a representative retail firm, the profit function is given by:

$$\pi_i = R(Q)Q_i - P(A)A_i - C_i(Q_i) \tag{3}$$

where C_i is other costs and, assuming a fixed proportions technology, $Q_i = A_i / a$ where a is the input-output coefficient. This assumption corresponds closely to

[9] We limit the discussion here to the oligopsony case as this was the issue of immediate concern to the UK Competition Commission. Including oligopoly will not change the nature of the tests outlined below; if the concerns related to oligopoly, the methodology still applies. If both oligopoly and oligopsony exist then the methodology is still applicable though it cannot distinguish between the two. However, since the methodology aims at a first pass test, this is not an immediate concern here.

the construction of the data in the vertical market chain used in the empirical analysis that follows.[10] Constant returns to scale in distribution are assumed. The theoretical set-up assumes a static game and has no dynamic links on the basis that contracts between retailers and suppliers are assumed to be negotiated every year and the commodities for which we are concerned are essentially perishable except at very short time intervals. The first-order condition for profit maximisation is given by:

$$R + Q_i \frac{\partial R}{\partial Q}\frac{\partial Q}{\partial Q_i} = \frac{\partial C_i}{\partial Q_i} + aP + aA_i \frac{\partial P}{\partial A}\frac{\partial A}{\partial A_i} \tag{4}$$

In order to get an explicit solution, consider linear functional forms for equations (1) and (2) and assume $a = 1$ (which is consistent with the construction of the data series):

$$Q = h - bR + cD \tag{1'}$$

$$P = k + gA \tag{2'}$$

with supply being given by:

$$A = Q + S$$

where S is the exogenous supply shifter. From this we can rewrite (4) as:

$$R = M + P + \mu g Q \tag{4'}$$

where μ is the aggregate input conjectural elasticity, such that with n firms in the retail sector, $\mu = (\Sigma_i\ [\partial A/\partial A_i][A_i/A])/n$. This parameter can be interpreted as an index of buyer power with $\mu = 0$ representing competitive behaviour and $\mu = 1$ representing monopsony behaviour. While μ is the measure of buyer power, as noted above, we do not aim to derive an explicit value for this parameter, but test only for its existence. M is a composite variable that represents all other costs that affect the retail-farm price margin.

To allow for changes in costs, we assume a linear marketing cost function of the form:

[10] This technology is likely to be most appropriate for many of the products we include in the empirical analysis.

$$M = y + zE \tag{5}$$

where y is a constant and zE represents the costs of inputs from the marketing sector (for example, wages). Using (1'), (2'), (4') and (5), we can derive an explicit solution for the endogenous variables:

$$Q = \frac{(h - by - bk) + cD - bzE - bgS}{1 + bg(1 + \mu)} \tag{6}$$

$$R = \frac{h + [1 + bg(1 + \mu)][(1 - b)(y + k + gS) + (1 - bz)E + cD)]}{1 + bg(1 + \mu)} \tag{7}$$

$$P = \frac{g[h - by + cD - bzE] - g[b - (1 + bg(1 + \mu))(k + S)]}{1 + bg(1 + \mu)} \tag{8}$$

To derive the spread between retail and producer prices, use (7) and (8) to give:

$$R - P = \frac{hg\mu + (1 + bg)(y + zE) + g\mu cD - bg\mu(k + gS)}{1 + bg(1 + \mu)} \tag{9}$$

Note that if oligopsony power does not matter in determining the retail-producer price spread (i.e. $\mu = 0$), then equation (9) reduces to:

$$R - P = y + zE = M \tag{10}$$

i.e. the source of the retail-producer price spread in a perfectly competitive industry is due to changes in marketing costs only. In this case, the exogenous shifters relating to the retail and agricultural supply functions play no role in determining the spread. This is not to say that they do not affect each price individually, but rather that because they affect retail and producer prices equally in a perfectly competitive industry they play no role in determining the relative gap between the prices at each stage of the food chain. Correspondingly, if oligopsony power in the food sector is important, each shifter affects the two prices differentially and thus the margin between the prices changes. In particular, in the presence of buyer power the demand shifter will be unambiguously positive and the supply shifter unambiguously negative. As such the demand shifter causes the margin to widen whereas the supply shifter will cause it to narrow. Intuitively, a rightward shift in the demand function will raise both the retail and farm-gate prices; but from (7) and (8), the changes to each of these prices vary in the presence of μ such that the spread widens as indicated by equation (9). Similarly, an exogenous shift in the supply function has a different relative effect on retail and farm level prices in the presence of buyer power with the 'net' effect on the spread being negative as indicated in equation (9).[11]

Equations (7)-(9) form the basis of our econometric modelling. Consider, first of all, equation (9) that relates to the retail-producer spread. Note that if buyer power does characterise the UK food sector, then the supply and demand shifters should enter our econometric model of the margin between retail and producer prices. Writing the margin equation in unrestricted form (i.e. in terms of prices) gives an empirical testable equation:

$$R = \beta_0 + \beta_1 P + \beta_2 M + \beta_3 D + \beta_4 S \tag{11}$$

The expected signs for the betas relate to the reduced form expressions for the determination of the retail-farm spread as reported in equations (9) and (10). Specifically, $\beta_1 > 0$, and $\beta_2 > 0$ irrespective of the degree of retail competition. The test for the rejection of perfect competition is whether the coefficients on the remaining variables in the retail-producer spread equation are statistically significant. Specifically, rejection of the (perfectly competitive) null hypothesis:

$$H_0 : \beta_3 = \beta_4 = 0$$

implies that perfectly competitive pricing is not congruent with the data.[12] Furthermore, equation (9) unambiguously signs the effect of the shifters. Whereas shocks to the demand shifter (which shift the demand curve to the right) widen the margin, supply-side shocks (which shift the supply curve to the left) narrow it, hence if oligopsony power is exercised (or is at least one of the potential candidate factors for the widening spread), the shifters are significant in the spread equation with signs such that $\beta_3 > 0$ and $\beta_4 < 0$ in (11).[13] In the empirical section, we test this proposition using data for nine product groups.

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[11] As with all studies in empirical industrial organisation, the issue of functional form matters. Given the focus here on the role of the exogenous shifters on the retail-farm spread, if we had constant elasticity inverse supply and demand functions, the spread would be constant. In this case, buyer/seller power may determine the magnitude of the size of the (static) spread but would not be consistent with exogenous shifters contributing to a widening spread.

[12] In principle, we only require one of these shocks to be significant to point to rejection of perfect competition, though the evidence will be 'stronger' if both shocks are. In the results presented below, in most cases, both shocks matter in determining the rejection of competition.

[13] With constant elasticity functional forms, the mark-up/mark-down will not change in the presence of buyer power. However, if we cannot reject the null hypothesis, then buyer power will not characterise the behaviour of the retail-producer price spread whatever the functional form.

4. DATA FROM THE FOOD INDUSTRY[14]

We apply our test method to assess whether we can reject perfect competition in UK food retailing using widely available market level data on prices of nine products, namely: apples (A); beef (B); bread (Br); chicken (C); Eggs (E) lamb (L); milk (M); pork (K) and potatoes (Pt) at retail (R) and producer (P) levels. The prices are deflated by the 'all products' Retail Price Index (1987=100) and expressed in terms of a standard unit that is comparable at both levels of the food chain (such as: pence/dozen for eggs; pence/kg of carcass weight equivalent for the meat products and pence/pint for liquid milk). The sample for each product begins in January 1990 and runs until October 2001 (giving 130 monthly observations) the date at which the Competition Commission's Code of Conduct came into force.[15] The price series are plotted in Figure 1 and details of the construction of the data are summarized in Appendix 1.

We use fresh products as these are subject to the smallest degree of processing by the post-farm gate chain and thus potentially provide a clearer correspondence between theory and data. As highlighted in section 2, it is also in the fresh food sector where asymmetry in bargaining is most likely to be revealed, since this is where small suppliers and large buyers co-exist most visibly. Clearly, however, this correspondence between theory and reality is not perfect in a number of respects. For example, prices represent the weighted average across a category, rather than a single product (eggs includes branded and non-branded sales, albeit of a standard size); the product sold at retail may not be identical to that sold at the farm-gate (it is the 'all milk' price that is recorded for producers but semi-skimmed milk at retail).

While meat products, which arguably undergo some of the most extensive processing in the sample of products in this study, have been adjusted by the Meat and Livestock Commission who convert them to a carcass weight equivalent, this is not so for bread which uses the price of a standard loaf at retail but the price of bread wheat to proxy for the producer price. Furthermore, retail prices are weighted across outlets and so include sales from independent retailers as well as supermarket chains, although the latter do dominate retail sales in the UK. Whilst these and other measurement issues potentially weaken the quality of the data for the purpose at hand, we merely point out here that they represent the best estimates that are currently available at the national level. We return to the

[14] Details of data series used and sources are given in Appendix 1. All statistical analysis is undertaken in *PCGIVE 12.,0* Doornik and Hendry (2007) except the stationarity testing for which *Eviews 6.0*, Quantitative Micro Software (2007). Data and detailed results are available upon request.
[15] Milk price data begins in January 1995.

important issue of data quality in light of the empirical results obtained in the following section.

As Figure 1 illustrates, there are some product-specific idiosyncrasies in the price series, such as the seasonal price fluctuations for lamb (and to a lesser extent, milk) and the disparity between retail and producer price volatility in chicken. Interestingly, one feature that is common to all products is the steady decline in prices at the producer level. While some retail prices also decline, many do not. Interestingly, even in cases where retail price decline is evident, it does not appear to be as rapid as the decline observed at the producer level. This tendency for retail and producer prices to diverge over time gives rise to a widening in the price spread, a feature that is common to all products analysed here, with the exception of milk.[16] While growth in the price spread is not in itself indicative of buyer power (since marketing costs may account for the observed behaviour), it is however noteworthy if only because growing spreads appear to be the norm over the sample period. Nevertheless, concerns about the impact of potential buyer power as investigated by the Competition Commission related to the existence of widening spreads between retail and farm prices in a number of commodity sectors.

The key issue that we address formally in the following section is whether the movement in the price spreads of these products can be attributed solely to marketing costs or whether it is also correlated with supply and demand shocks, as predicted by the theoretical model in the presence of buyer power. As noted in section 1, measures of product-specific marketing costs are not available in the UK and thus given the importance of labour costs in food retailing we use an index (base year 2000) of real average earnings in the UK service sector (M) to proxy for these costs. While an undeniably crude measure, this proxy does appear to explain much of the behaviour in the price spreads, as we report in the following section. Since buyer power is mediated through market shocks in the theoretical model, we attempt to capture such shocks using the following proxies in the empirical model. Specifically, to incorporate the impact of farm-level production costs the supply shifter (S) represents a real price index (base year 1997) of all goods and services purchased on UK farms. Demand-side shocks are proxied by one of two measures depending on the market at hand. For meats and animal based products we use an index of media activity relating to the health and safety of food (D1). Specifically, this index represents the natural log of the cumulative count of articles relating to the health and safety of food published in four national broadsheet newspapers. This index is dominated by articles relating

[16] Plots of the spread over time (not shown in the interests of brevity) more clearly demonstrate this tendency.

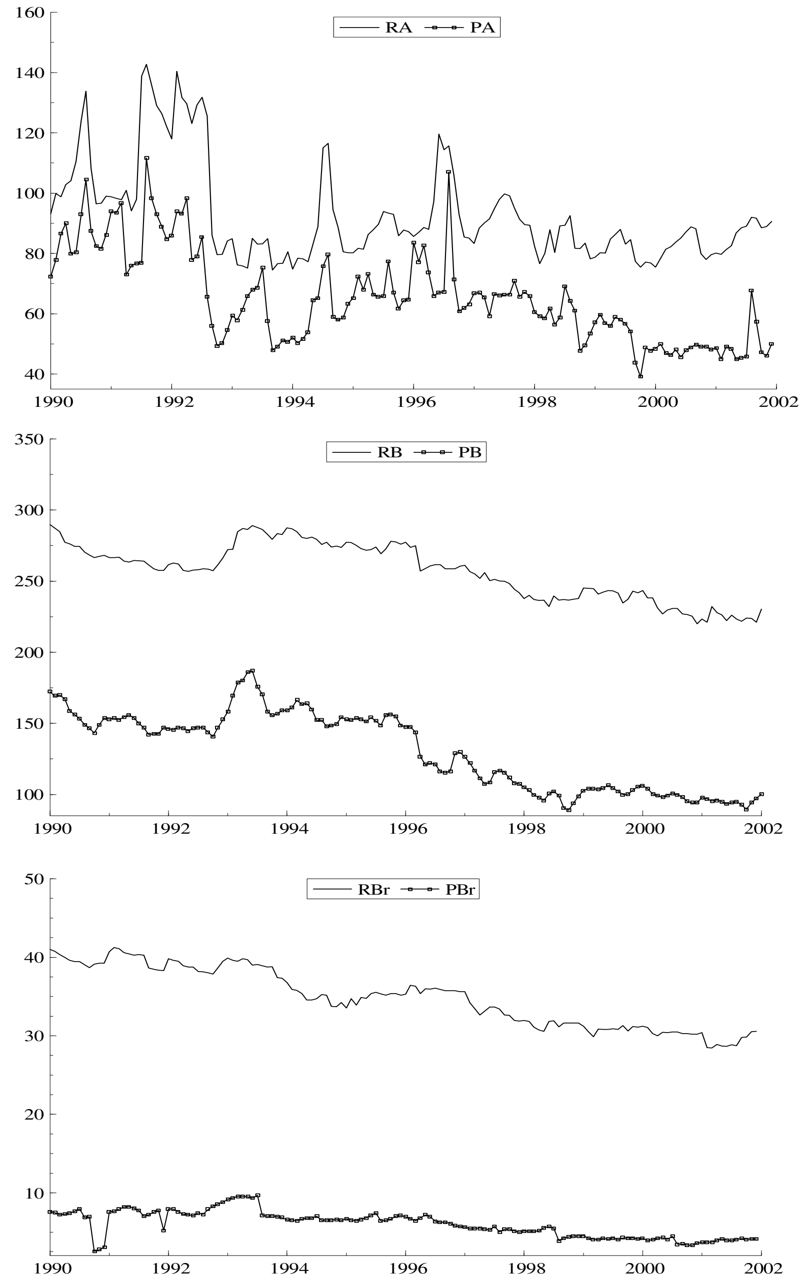
to BSE and can act as a shifter for all the meat products studied here.[17] On the basis that such articles rarely relate to non-meat products, we have therefore used the food retail price index (D2) for these products on the basis that this represents a general demand shifter affecting the food retailing sector as a whole.[18] The individual fresh products reported here represent such a small weight in the food retail price index that the movement in the latter can be a reasonable proxy for a general demand shifter for these non-meat products.[19] Details about the construction of the proxy variables are given in Appendix 1 and the time series are plotted in Figure 2 for each measure.

[17] It should be noted that as the number of stories increases, this is a positive increase in the value of the shifter. This will cause the demand for non-beef meats to increase as consumers substitute away from beef and thus the prices of these meats will increase, *ceteris paribus*. However, with a positive increase in the shifter the price for beef should fall as this represents a leftward shift in the demand curve for beef and hence a price fall, *ceteris paribus*.

[18] Ideally, the measure of income would be preferred as a demand shifter but this is not available at the same monthly frequency that we need to tie in with the price data.

[19] One should also recall that we are not totally reliant on this shifter in determining buyer power since the significance of the supply shifter also matters.

Figure 1: Real Product Prices at Retail and Producer Levels[20]



Source: The data sources for these series are listed in the Appendix (Table 1).

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[20] The quantity units for the vertical axes vary from commodity to commodity (e.g. per kg or per lb) but are all expressed in pence per unit.

Figure 1: (Continued)

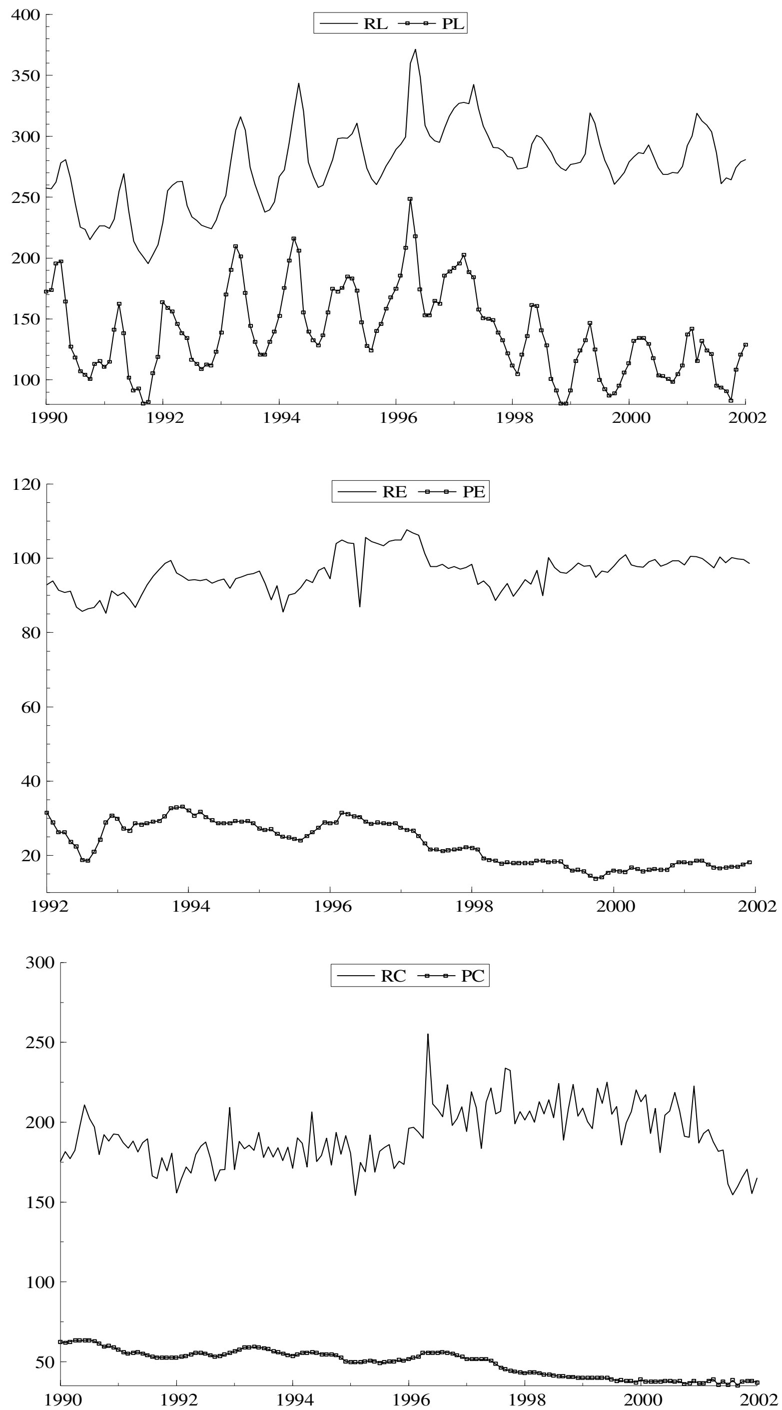
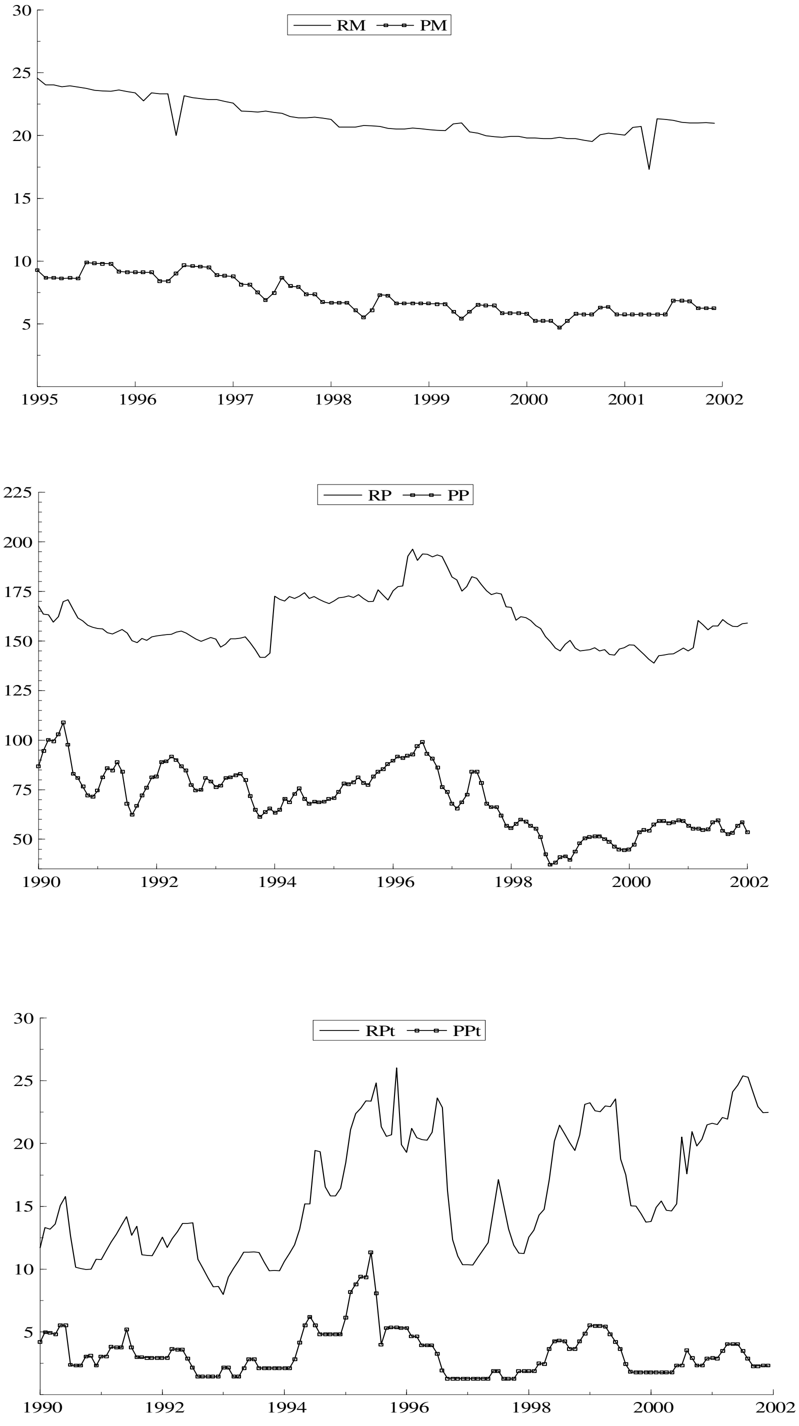


Figure 1: (Continued)



5. EMPIRICAL METHOD

To allow for the possibility that retail and producer prices of each product group are non-stationary and cointegrated, we couch the empirical analysis in a vector autoregressive (VAR) framework. For each of the nine product groups it is assumed that the data may be approximated by a VAR(*p*) model,

$$\mathbf{x}_t = \mathbf{\Phi}_1 \mathbf{x}_{t-1} + \mathbf{\Phi}_2 \mathbf{x}_{t-2} + \ldots + \mathbf{\Phi}_p \mathbf{x}_{t-p} + \mathbf{\Psi} \mathbf{D}_t + \boldsymbol{\varepsilon}_t \tag{12}$$

where $\mathbf{x}_t$ is a $(k \times 1)$ vector of jointly determined I(1) variables, $\mathbf{D}_t$ is a $(d \times 1)$ vector of deterministic terms (constants, trends and centred seasonals) and each $\mathbf{\Phi}_i$ $(i = 1,\ldots, p)$ and $\mathbf{\Psi}$ are $(k \times k)$ and $(k \times d)$ matrices of coefficients to be estimated using a (t = 1, . . .T) sample of data. $\boldsymbol{\varepsilon}_t$ is a $(k \times 1)$ vector of n.i.d. disturbances with zero mean and non-diagonal covariance matrix, $\mathbf{\Sigma}$.

Equation (12) represents an unrestricted reduced form of the variables in $\mathbf{x}_t$ comprising retail and producer prices, a measure of marketing costs and the supply and demand shifters. Given the monthly frequency of the data, lag length (p) of the VAR is determined for each product group in step-wise fashion $(p = 13,12,\ldots,1)$ using standard vector-based diagnostics, so that the preferred specification is the most parsimonious model that is free of residual correlation at the 5% significance level.[21] Prior to the cointegration analysis each preferred model is checked for parameter constancy using vector-based recursive Chow tests.

The presence of a price transmission relationship between retailer and producer is indicated by the detection of cointegration among the variables in $\mathbf{x}_t$. Rearranging (12) into its error correction form,

$$\Delta\mathbf{x}_t = \boldsymbol{\alpha\beta}'\mathbf{x}_{t-p} + \sum_{i=1}^{p-1} \mathbf{\Gamma}_i \Delta\mathbf{x}_{t-i} + \mathbf{\Psi}\mathbf{D}_t + \boldsymbol{\varepsilon}_t \tag{13}$$

we test for cointegration using Johansens's (1988) maximum likelihood procedure in which attention focuses on the $(k \times \mathrm{r})$ matrix of co-integrating vectors,

[21] Although commonly applied in VAR analyses, information criteria (such as AIC, SBC and HQC) tended to select (overly parsimonious) models characterised by residual autocorrelation in this study, a feature most likely to reflect the large number of parameters required for a general model involving monthly data. Given the object is to adopt the most parsimonious model comprising white noise errors, lag length selection was based upon diagnostic tests directly, rather than the likelihood-based information criteria.

comprising β, that quantify the 'long-run' (or equilibrium) relationships between the variables in the system and the $(k \times \mathrm{r})$ matrix of error correction coefficients, α, the elements of which load deviations from equilibrium (*i.e.* $\boldsymbol{\beta}'\mathbf{x}_{t-p}$) into $\Delta\mathbf{x}_t$, for correction. The Γ_i coefficients in (13) estimate the short-run effect of shocks on $\Delta\mathbf{x}_t$, and thereby allow the short and long-run responses to differ.

Since there may exist up to $k-1$ cointegrating relations among the k variables in $\mathbf{x}_t$, the precise number is evaluated by Johansen's Trace (η_r) and Maximal Eigenvalue (ξ_r) test statistics (Johansen, 1988). The η_r statistic tests the null that there are at least r cointegrating relationships ($0 \le r < k$) and ξ_r evaluates the null that there are r against the alternative that there are at most $r+1$ such relationships. While the η_r test is generally preferable because it is robust to residual non-normality and delivers a sequentially consistent test procedure, it is standard practice to report both test statistics (Hariss and Sollis, 2005, p.123).

The specification of the deterministic terms in (13) play a pivotal role in cointegration inference, not least because the distributions of η_r and ξ_r depend on these terms and how they enter the model. In the empirical analysis that follows, we estimate (13) with unrestricted constant (but without linear trend) to allow for drift in any of the non-stationary variables in $\mathbf{x}_t$, a property that characterises some of the series we investigate. Linear trend terms are excluded on the grounds that while they allow for quadratic and/or trend stationary behaviour in $\mathbf{x}_t$ (depending on whether the data is I(1) or I(0) respectively) neither behaviour is a plausible representation of our data, as we explore in the following section.[22]

Where a single cointegrating relationship is detected between retail and producer prices, formal testing of the significance of the supply and demands shocks is undertaken to investigate whether buyer power is present. Following from section 2, if the vertical market for a product is perfectly competitive, retail and producer prices may be expected to form a cointegrated relationship with at most marketing costs. Where retailers exert buying power, the supply and demand shifters also enter the pricing relationship. This then gives rise to a null hypothesis of perfect competition which can be evaluated empirically by a standard likelihood ratio test of the exclusion restrictions on the shifters in the cointegrating relation. In addition, given that the theoretical model signs the parameters in the pricing relation (11), we can offer some additional evidence on the possible

[22] Where seasonality is present, we augment $\mathbf{D}_t$ with centred seasonals, however these do not affect the asymptotic distributions of the test statistics. See Juselius (2006) p.139.

rejection of perfect competition by comparing the estimated signs of the shifters in the cointegrating relation with that predicted by the theoretical model.

6. RESULTS

As a first step in the descriptive analysis, we analyse the time series properties of the data to determine the most appropriate form for the deterministic part of the VAR (such as constant, trend and seasonals). Inspection of Figures 1 and 2 suggests that the series possess the (stochastic) trends that characterize the random walk I(1) model. The sustained nature of these trends suggests that a random walk with drift may represent a better approximation for many of the series, and for lamb (and possibly milk) prices, a seasonal pattern is also apparent. Considerations of this sort suggest that the unit root I(1) null should be evaluated in a maintained model with constant, trend and where necessary, centred seasonals. While the alternative hypothesis of trend stationarity seems an unlikely outcome from an economic viewpoint, inclusion of the trend term does ensure invariance of the (unit root) test statistic to the presence of drift (see, for example, Patterson, pp.233-238).

Results from the application of the Augmented Dickey-Fuller (ADF, 1979) are reported in Table 1. The test is applied to the levels (Model 1) and first differences (Model 2) of each data series and indicate that the data are I(1) in levels and I(0) in first differences, as indeed visual inspection suggests. In cases where the data are actually (mean) stationary rather than I(1), the ADF test is known to have low statistical power to reject the unit root null (Dickey and Fuller, 1981) owing to the inclusion of (redundant) trend terms in the ADF regression. In recognition of this, we also apply the stationarity test of Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992) to the price and shifter series shown in Figures 1 and 2, results of which are also presented in Table 1. With a null hypothesis of stationarity, the KPSS test offers an appealing complement to the ADF test.[23] Referring to Table 1, the KPSS test results confirm the non-stationarity of the shifters but there is less unanimity regarding prices, where inference largely depends on whether a time trend is included in the maintained regression of the KPSS test. Specifically, where the null hypothesis is of trend stationarity (Model 3) the KPSS test rejects in favour of a unit root in ten out of eighteen cases.

[23] Note however that owing to the non- equivalence of these tests, in that the null of the ADF test and alternative hypothesis of the KPSS are not identical (see Patterson p.268) a dual testing procedure does not offer a panacea. Furthermore, differences in finite sample performance, sensitivity of the tests to nuisance parameters and lag length selection also mean that in practice contradictory results from the two testing strategies may occur. For further details see Maddala and Kim (p.128).

With a (mean) stationary null (Model 4) the unit root alternative is favoured in fifteen out of eighteen prices, RL, RP and PPt being the exceptions (despite the unit root having been favoured for two of these cases using Model 3). While these contradictory results underline the fragility of unit root testing *per se* and the finite sample similarity of the random walk and trend stationarity (even in relatively large samples such as those used here), the implausibility of trend stationarity from an economic viewpoint suggests we proceed on the basis that the random walk (possibly with drift) model is an adequate (albeit imperfect) statistical approximation of the data.[24]

To recap, vector error correction models (equation (13)) with unrestricted constant are estimated for each of the nine products in a general-to-specific strategy for k = 13 to 1, the preferred model in each case being the most parsimonious model in which the vector tests of the null of no residual correlation, homoscedasticity and parameter constancy cannot be rejected at the 5% level. The resulting models represent the baseline against which subsequent parameter restrictions are evaluated.[25] The presence of a price transmission relationship is indicated by the cointegrating rank of each baseline model using the Trace (η_r) and maximal Eigenvalue (ξ_r) tests statistics reported in Table 2.

[24] The negative trend observed in many of the price series seems unlikely to persist indefinitely. As Stein's Law states, "Things that can't last forever, don't". Locally linear trend models (see Harvey, 1989) are not considered here.

[25] Results from parameter constancy tests have been made available to referees but are not included in the interest of brevity.

Table 1: ADF and KPSS Test Statistics

Variable	ADF		KPSS	
	(1) Levels (c,t)	(2) Differences (c)	(3) Levels (c,t)	(4) Levels (c)
Prices				
RA	-2.88	-10.93**	0.10	0.63**
PA	-2.34	-8.51**	0.11	0.99**
RB	-2.21	-11.19**	0.18*	1.49**
PB	-2.46	-7.49**	0.13	1.56**
RBr	-2.85	-11.52**	0.05	1.34**
PBr	-2.92	-8.86**	0.19*	1.23**
RC	-1.35	-10.67**	0.17*	0.51*
PC	-2.65	-3.83**	0.14*	1.40**
RE	-2.58	-16.09**	0.12	0.52*
PE	-2.94	-6.96**	0.17*	1.00**
RL	-3.14	-8.67**	0.23**	0.96**
PL	-3.25	-7.78**	0.15*	0.31
RM	-1.52	-8.08**	0.28**	1.10**
PM	-1.57	-9.36**	0.20*	1.11**
RP	-1.59	-10.20**	0.19*	0.22
PP	-2.23	-8.33**	0.08	1.23**
RPt	-2.33	-11.70**	0.06	0.74**
PPt	-2.17	-10.48**	0.08	0.10
Shifters				
S	-1.91	-8.25**	0.14*	1.50**
D1	-1.94	-5.27**	0.21*	1.56**
D2	-2.38	-11.09**	0.20*	1.26**
M	-0.94	-14.12**	0.33**	1.52**

The ADF test on the variables in levels is conducted using Model (1) which include a constant and linear trend (c,t) conditioned on centred seasonals where appropriate. The ADF test is also applied to the variables expressed in first differences (Model 2) which includes constant (c) and seasonals where appropriate. Lag length is determined by serial correlation test on the residuals evaluated at the 5% level. The 5% and 1% (finite sample) critical values for the ADF test are -3.45 and -4.03 for Model (1) and -2.89 and -3.52 for Model (2). Asterisks denote rejection of the unit root null at 5% (*) and 1% (**). For the KPSS test, the correction factor for serial correlation is by Barlett Kernel based on Newey-West weights. Test statistics evaluate the null of stationarity around linear trend (Model 3) and non-zero mean (Model 4) and relate to the variables expressed in levels. The 5% and 1% (asymptotic) critical values are 0.15 and 0.21 for Model (3) and 0.46 and 0.74 for Model (4).

Figure 2: Shifters

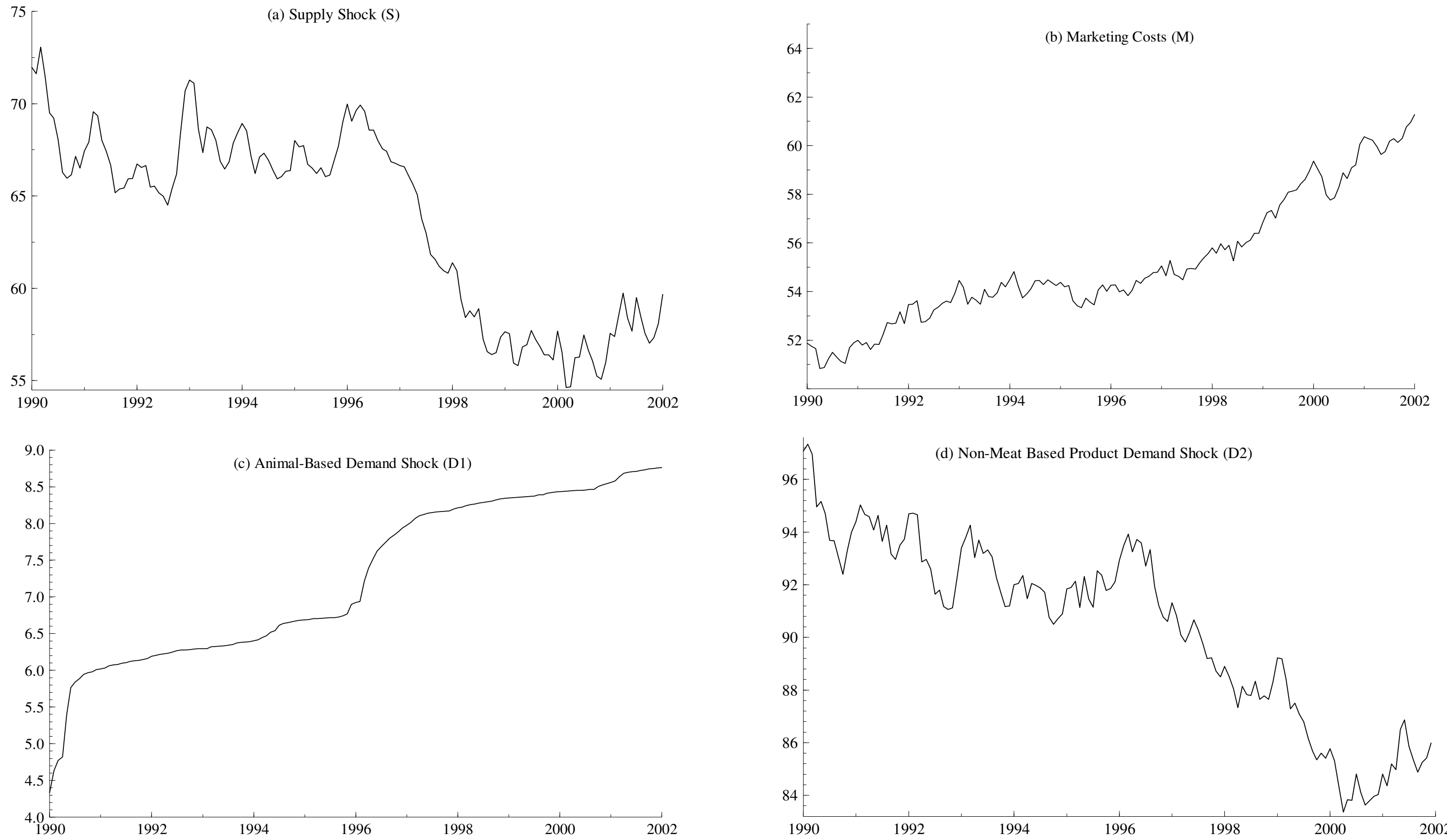


Table 2: Test Statistics for Cointegration

Product (lag length)	**Rank**	**Trace** η_r	**Maximal Eigenvalue** ξ_r	**Product** (lag length)	**Rank**	**Trace** η_r	**Maximal Eigenvalue** ξ_r
Apples (4)	0	66.59 [0.086]*	35.36 [0.029]**	Lamb (2) Seasonals	0	102.05 [0.000]***	50.95 [0.000]***
	1	31.23 [0.657]	14.76 [0.766]		1	51.10 [0.022]**	25.23 [0.097]*
	2	16.47 [0.686]	10.77 [0.677]		2	25.87 [0.136]	17.45 [0.157]
	3	5.70 [0.732]	5.58 [0.672]		3	8.42 [0.429]	8.36 [0.351]
	4	0.13 [0.722]	0.13 [0.722]		4	0.06 [0.812]	0.06 [0.812]
Beef (6)	0	83.69 [0.002]***	46.00 [0.001]***	Milk (4) Seasonals	0	131.34 [0.000]***	65.05 [0.000]***
	1	37.69 [0.320]	18.79 [0.444]		1	66.28 [0.000]***	35.34 [0.003]***
	2	18.90 [0.511]	12.04 [0.557]		2	30.95 [0.036]**	18.41 [0.118]
	3	6.86 [0.600]	5.82 [0.642]		3	12.53 [0.134]	11.40 [0.137]
	4	1.04 [0.307]	1.04 [0.307]		4	1.14 [0.286]	1.14 [0.286]
Bread (9)	0	109.81 [0.000]***	54.83 [0.000]***	Pork (4)	0	76.26 [0.013]***	28.77 [0.184]
	1	54.98 [0.008]***	33.53 [0.006]***		1	47.49 [0.153]	21.04 [0.283]
	2	21.45 [0.340]	14.24 [0.359]		2	26.45 [0.119]	16.49 [0.205]
	3	7.21 [0.560]	6.63 [0.542]		3	9.96 [0.289]	9.95 [0.220]
	4	0.58 [0.446]	0.58 [0.446]		4	0.01 [0.913]	0.01 [0.913]
Chicken (3)	0	96.02 [0.000]***	52.62 [0.000]***	Potatoes (3)	0	70.60 [0.041]**	32.54 [0.069]*
	1	43.39 [0.123]	22.16 [0.219]		1	38.06 [0.303]	16.99 [0.590]
	2	21.23 [0.354]	15.45 [0.269]		2	21.06 [0.364]	12.53 [0.510]
	3	5.78 [0.723]	5.39 [0.695]		3	8.53 [0.418]	8.26 [0.361]
	4	0.39 [0.531]	0.39 [0.531]		4	0.28 [0.600]	0.28 [0.600]
Eggs (2)	0	70.32 [0.044]**	27.77 [0.231]				
	1	42.55 [0.144]	17.97 [0.509]				
	2	24.58 [0.183]	15.49 [0.267]				
	3	9.09 [0.363]	8.97 [0.296]				
	4	0.13 [0.722]	0.13 [0.722]				

Notes: *** denotes significance at 1%; ** at 5% and * at 10%. *p*-values are in parentheses. Critical values are those of Doornik (1998). Seasonals denotes that the VAR contains monthly centred seasonal dummies.

Overall, the evidence points to the presence of a single cointegrating vector in the majority of products. Evaluating hypotheses at the 5% significance level, the null of no cointegration is rejected in fourteen out of eighteen tests. At least one test rejects the null of no cointegration for every product. Using the more reliable Trace statistic, cointegration is detected in all products at the 5% level, except apples where the *p*-value is 9%. Test statistics suggest the presence of multiple cointegration relations for bread, lamb and milk although the presence of centred seasonals in the models for the last two products is a potential explanation for the apparent distortion to the size of the tests in these cases.[26] In the absence of any obvious economic explanation for relationships other than the price transmission relation, we proceed on the assumption that a single cointegrating vector is present for each product. Finally, as an informal check on the adequacy of the long-run specification, we inspect the cointegrating residuals from each baseline model for the tell-tales of model mis-specification such as trending or structural change. Each set of residuals appear to be 'well-behaved' with zero mean and no trend (see appendix) lending casual support to the adequacy of the chosen specifications.

Table 3 reports the parameters of the cointegrating vectors normalized on retail prices obtained from each baseline model. Recall that the theoretical model presented in section 2 signs these coefficients such that, $\beta_1 > 0$ and $\beta_2 > 0$; and where buyer power exists, $\beta_3 > 0$ and $\beta_4 < 0$ (for demand increasing and supply decreasing shocks respectively). Referring to the table a number of points seem noteworthy: first, price transmission coefficients (β_1) are positive in all cases; second, marketing costs, as proxied by the index of real average earnings in services, (β_2) are positive in seven out of nine cases; third, the coefficient on the demand shifter (β_3) is correctly signed in seven out of nine cases; and fourth, the coefficient on the supply shifter is correctly signed in six out of nine cases. Overall, the results accord well with theoretical predictions. Of key interest are the results relating to the demand and supply shifters, since it is through these variables that the existence of buyer power is mediated in the theoretical model. Since the standard errors from the cointegrating relations in Table 3 are based upon approximation, we perform a set of likelihood ratio tests to evaluate the statistical significance of these coefficients, results of which are contained in Table 4.

[26] Although the asymptotic distributions of these test statistics are invariant to centred seasonals, the small sample performance is currently unknown (see Juselius, 2006, p.136). Estimation without the seasonal dummies supports the existence of a single cointegrating relationship in both lamb and milk, suggesting that size distortion is responsible for the apparent multiple cointegration findings indicated in the table. Alternative test statistics such as those proposed by Saikkonen and Lütkepohl (2000) may be beneficial however they are not available in the software used.

Table 3: The Cointegrating Vectors (normalised on retail prices)

Product	Producer prices (β_1)	Marketing costs (β_2)	Demand shifter (β_3)	Supply shifter (β_4)
Apples	1.67*** (0.22)	5.74*** (2.18)	1.41 (2.33)	-0.13 (0.99)
Beef	1.62*** (0.13)	4.46*** (0.93)	27.45*** (3.16)	-1.57*** (0.48)
Bread	3.64*** (0.69)	4.62*** (1.04)	3.59*** (0.97)	-0.82** (0.32)
Chicken	6.13*** (1.01)	-5.13*** (1.80)	20.36*** (4.85)	-5.13*** (1.30)
Eggs	1.29*** (0.36)	0.33 (0.53)	11.95*** (1.65)	0.59 (0.42)
Lamb	1.25*** (0.10)	2.17 (1.61)	0.14 (4.64)	-4.91*** (1.08)
Milk	1.08*** (0.15)	0.20*** (0.04)	-0.31** (0.15)	-0.02 (0.04)
Pork	0.87** (0.43)	-8.57*** (3.29)	54.09*** (7.77)	5.82*** (1.77)
Potatoes	2.54** * (0.44)	5.91*** (0.96)	3.29*** (0.95)	0.21 (0.41)

Figures in bracket are asymptotic standard errors; *** denotes significance at 1%; ** at 5% and * at 10%. Note however that estimators of the variance of the cointegrating parameters are, strictly speaking, not defined (see Banerjee *et al.* 1993 p.61-64). While it is common practice to use approximations to calculate standard errors, we use formal likelihood ratio tests to facilitate inference regarding buyer power (see Table 3).

Each cell contains a likelihood ratio statistic and its associated asymptotic *p*-value. The first column of results evaluates the null hypothesis that both shifters are jointly insignificant, and is thus our test of buyer power. Test statistics are distributed as $\chi^2(2)$ under the null hypothesis of no buyer power (i.e. perfect competition). Results indicate that the null of perfect competition can be rejected in seven of the nine products, milk and apples being the exceptions. Tests of the individual significance of each shifter are distributed as $\chi^2(1)$ with rejection of the perfectly competitive null occurring in 10 of the 18 tests at the 5% level. Again, milk and apples are found not to reject the perfectly competitive nulls. Importantly, nine out of ten of the statistically significant coefficients are signed in accordance with buyer power in the theoretical model.

Table 4: Tests for Competition

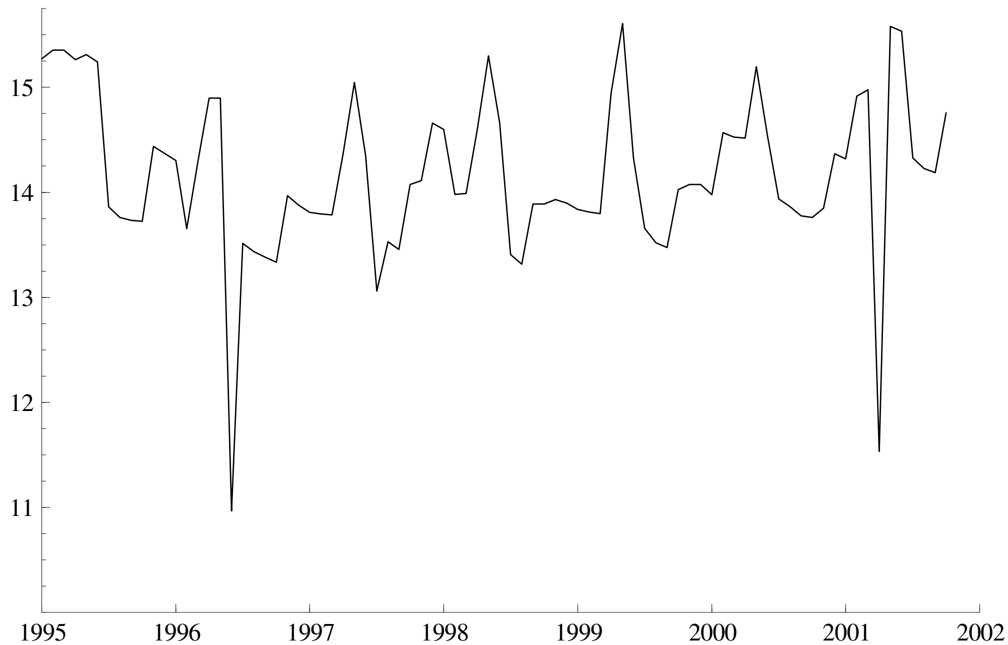
Product	$H_0:\beta_3=\beta_4=0$	$H_0:\beta_3=0$	$H_0:\beta_4=0$
Apples	0.42 [0.8111]	0.24 [0.6275]	0.01 [0.9104]
Beef	23.95 *** [0.0000]	18.30*** [0.0000]	8.38*** [0.0038]
Bread	7.19** [0.0275]	7.10*** [0.0077]	4.33** [0.0374]
Chicken	29.48*** [0.0000]	6.54** [0.0105]	29.37*** [0.0000]
Eggs	12.14 *** [0.0023]	9.77*** [0.0018]	0.70 [0.4044]
Lamb	19.76*** [0.0001]	0.00 [0.9844]	13.37*** [0.0003]
Milk	2.82 [0.2443]	2.43 [0.1194]	0.14 [0.7102]
Pork	10.14 *** [0.0063]	7.17*** [0.0074]	1.91 [0.1669]
Potatoes	15.2*** [0.0005]	6.66*** [0.0099]	0.17 [0.6764]

Figures in brackets are asymptotic *p*-values; *** shows significance at 1%; ** at 5%; * at 10%

Reflecting upon the pattern of results across products, it appears that rejection of the perfectly competitive null is spread across meat and non-meat products. Whilst any justification for the observed pattern is understandably conjectural, it is interesting that we find no evidence for the exercise of buyer power in milk, a product around which much attention, and indeed controversy has centred in recent years.[27] Most major UK retailers[28] have accepted fines amounting to £116 million imposed by the Office of Fair Trading in 2007 for price collusion with milk and dairy product processors during the early 2000s (Office of Fair Trading, 2007b). Thus if the milk price spread was being maintained by collusion rather than competition, as the regulatory authorities have found to be the case, it is little wonder that our simple test is unable to detect what amounts to relatively sophisticated strategic pricing behaviour.

[27] See, for example, House of Commons (2004).
[28] At the time of writing, Morrisons is exempt for the ruling and Tesco is disputing the judgment.

Figure 3: The Milk Retail-Producer Price Spread (pence per pint)



In this regard, note also from Figure 3 that there was no overall trend in the retail-farm spread; the "success" of the test reported for other commodities is consistent with a widening spread and the exogenous shifters indicating the potential existence of buyer power. Given the raw data, it is therefore not surprising that we cannot reject the null hypothesis in this specific case. There could of course be some other aspect of buyer/seller power that exists in this market and that the concerns about collusion between retailers and processors did not negatively impact on milk producers taken over the period for which our data applies. While similar observations or explanations cannot account for the apple results (where the spread between retail and farm level prices was rising over the period), we merely note here that although statistically insignificant, both shifters are signed in accordance with theoretical prediction in this case.

Returning to the results presented in Table 2, there are two further caveats to note. First, while the theoretical model additionally implies that $\beta_1 < 1$, this condition is seldom met in the empirical setting. This may be due to heterogeneity within product groups and other practical factors such as wastage and differences in product specification that interfere with the strict one-to-one correspondence of products as they move through the marketing 'chain' in the theoretical model. Second, the estimated coefficients on the marketing costs proxy is negative in the models for chicken and pork, and as such are at odds with their role in the theoretical model. This is likely to reflect the inadequacy of a general marketing cost variable such as labour costs and/or that we are unable to pick-up specific trends in marketing technology or costs with it in these cases. Nevertheless, the overall correspondence between theoretical prediction and empirical finding is a noteworthy feature of the analysis. To the extent that this merely confirms the potential existence of buyer power identified by the competition authorities, our findings are not new but arguably lend 'scientific' substance to the survey and witness-based evidence compiled by the competition authorities. Moreover, since the purpose of our analysis is to evaluate whether the exercise of buyer power may be detected econometrically using aggregate level data, the corroboration offered in this paper is potentially useful to those conducting analyses of this type in other countries where retailer dominance is also of public concern. However, it is important to re-iterate the caveat reported in the Introduction; while these results lend a filter to explore whether buyer power is important, the rejection of the null hypothesis of perfect competition under the assumptions employed suggest that buyer power is at least a candidate for explaining the spread between retail and farm prices. As such, rejection of the null hypothesis is supportive of further investigation but not the end of the road in terms of concluding that buyer power unambiguously characterises the links between retailers and producers.

7. CONCLUDING COMMENTS

In this paper, we have devised a simple means of testing for the presence of buyer power in vertically-related markets such as those characterising the food chain. By constructing a quasi-reduced form model of the retailer-supplier pricing equations, the null of perfect competition can be rejected if the shifters from the supply and demand equations are significant and correctly signed. In principle, the approach sits between other methods of evaluation, to which it is complementary. In particular, we are able to move away from naïve concentration-based indicators of buyer power and the practical limitations of structural econometric modelling. The approach is simple and transparent yet delivers a statistical test derived from a theoretically-consistent basis. Furthermore, the test demands relatively little in terms of data and is implemented using standard techniques of modern time-series analysis. The technique is most applicable where products undergo relatively little transformation between marketing levels and is thus particularly well-suited to the relatively unprocessed products of the food chain. In the UK at least, these are also products over which concerns of potential buyer power abuse have been most acute.

Drawing on data from a basket of nine basic products of the UK food industry, we show that in seven cases, the hypothesis of perfect competition can be firmly rejected at conventional levels of significance, implying that for these food products at least, the market is characterised by buyer power by our measure. As such, our findings corroborate the findings of Competition Commission (2000) and lend support to the recent request by the Office of Trading for further detailed scrutiny of the UK food chain by the UK's competition authorities. Of course, we cannot interpret our results as being conclusive of the use of buyer power in UK food retailing. Among many important caveats are that the test is predicated on simplifying assumptions, the data subject to measurement problems and the procedures prone to statistical error. However, the methods we employ are both familiar to applied economists and readily implemented, and deliver what we may call a 'first pass' test, that when used in combination with other evidential indicators, can be useful in contributing to uncovering the existence of buyer power in the vertical food chain.

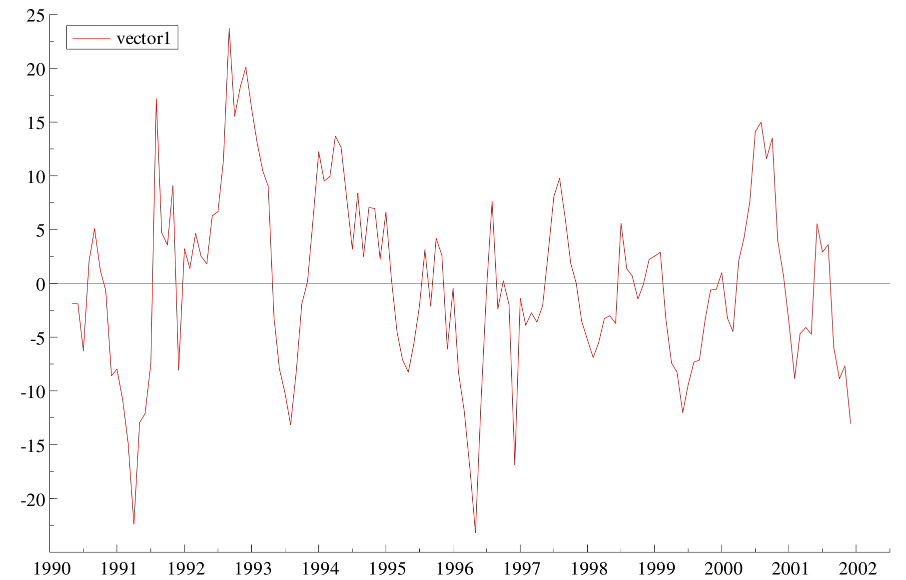
Appendix Table 1: Data Definitions and Sources[29]

Label	Variable	Units	Area	Comments	Source
RA	Retail apple	Index of pence/lb (1987=100)	UK	Desert apples only	Employment Gazette/Labour Market Trends
PA	Producer apple	Index of pence/lb (1987=100)	UK	Exclude direct subsidies	Department of Food, Environment and Rural Affairs
RB	Retail beef price	Pence/kg carcass weight equivalent	GB	Converted to c.w.e by MLC	Meat and Livestock Commission
PB	Producer beef price	Pence/kg carcass weight	GB	Sample of auction & abattoir average	Meat and Livestock Commission
RBr	Retail bread price	ln(pence/800g loaf)	UK	Standard 800g white sliced loaf	Employment Gazette/Labour Market Trends
PBr	Producer bread price	ln(£/ton)	UK	Bread wheat	Department of Food, Environment and Rural Affairs
RC	Retail chicken price	Pence/kg carcass weight	GB	Uncooked whole birds including frozen <1.81 kg	National Food Survey/Expenditure and Food Survey
PC	Producer chicken price	Pence/kg carcass weight	E&W	Birds <2.27 kg	National Farmers Union
RE	Retail egg price	Pence/dozen		Eggs of size 2	
PE	Producer egg price	Pence/dozen		Eggs of size 3	
RL	Retail lamb price	Pence/Kg carcass weight equivalent	GB	Converted in to c.w.e. by MLC	Meat and Livestock Commission
PL	Producer lamb price	Pence/kg carcass weight	GB	Sample of auction & abattoir average	Meat and Livestock Commission
RM	Retail milk price	Pence/pint	UK	Semi skimmed only	Employment Gazette/Labour Market Trends
PM	Producer milk price	Pence/pint	UK	Average all milk	Department of Food, Environment and Rural Affairs
RPt	Retail potato price	Pence/lb	UK	Old white, sold loose	Employment Gazette/Labour Market Trends
PPt	Producer potato price	Pence/lb	UK	Average all potatoes (including processor sales)	Department of Food, Environment and Rural Affairs
S	Farm supply shock	Index of farm input prices (1997=100)	UK	Includes all Goods and services currently consumed on UK farms	Department of Food, Environment and Rural Affairs
D1	Meat demand shock	Ln (cumulative count of newspaper 'food scare' articles)	UK	Articles appearing in Times, Sunday Times, Guardian and Observer.	Euro-PA Associates, Northhampton.
D2	Non-meat demand shock	Food Retail Price Index (1987=100)	UK	Includes all food items in RPI	Office of National Statistics
M	Marketing shock	Index (2000=100) of average earnings in GB service sector, including bonuses.	GB	Series known as RLNMT by ONS. Seasonally adjusted	Office of National Statistics

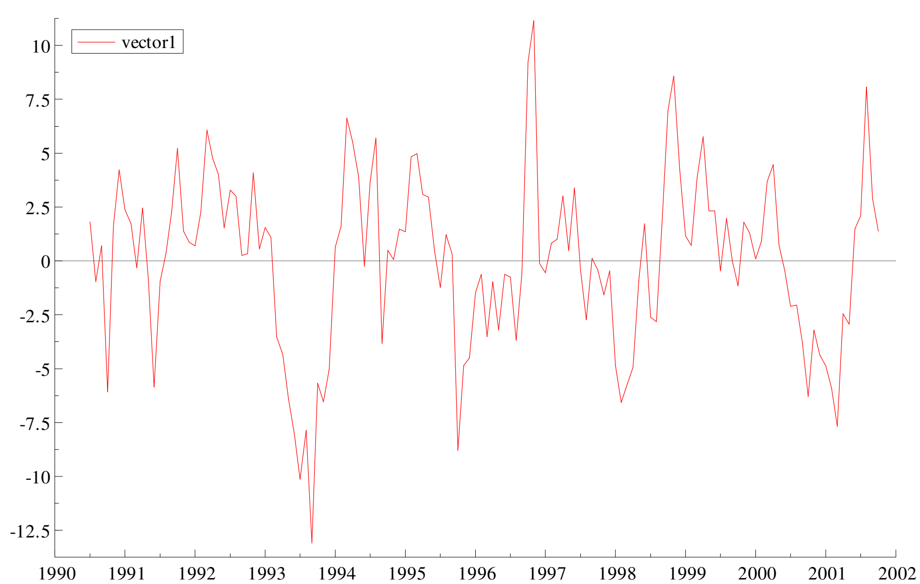
[29] Data available from 1990.1 to 2001.10 (130 observations) except eggs (1992.1 – 2001.10; 118 observations) and milk (1995.1 – 2001.10; 82 observations). All monetary variables expressed in real terms using the (all products) RPI.

Appendix Figure 1: Residuals of Cointegrating Vectors

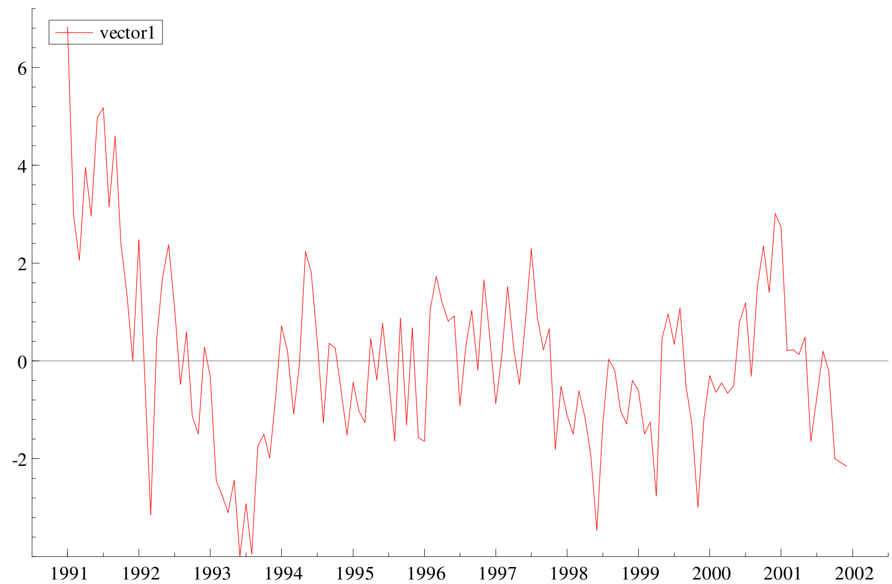
Apples



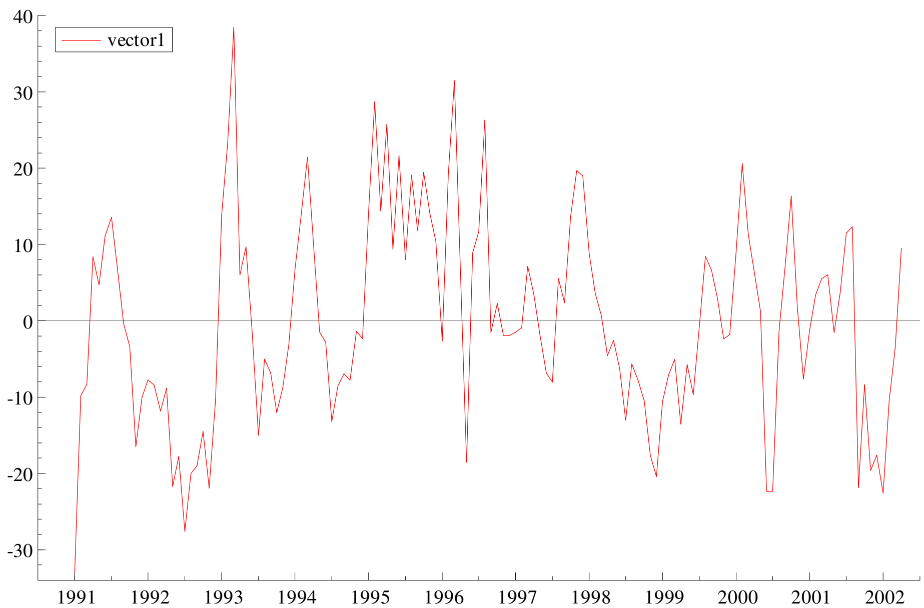
Beef



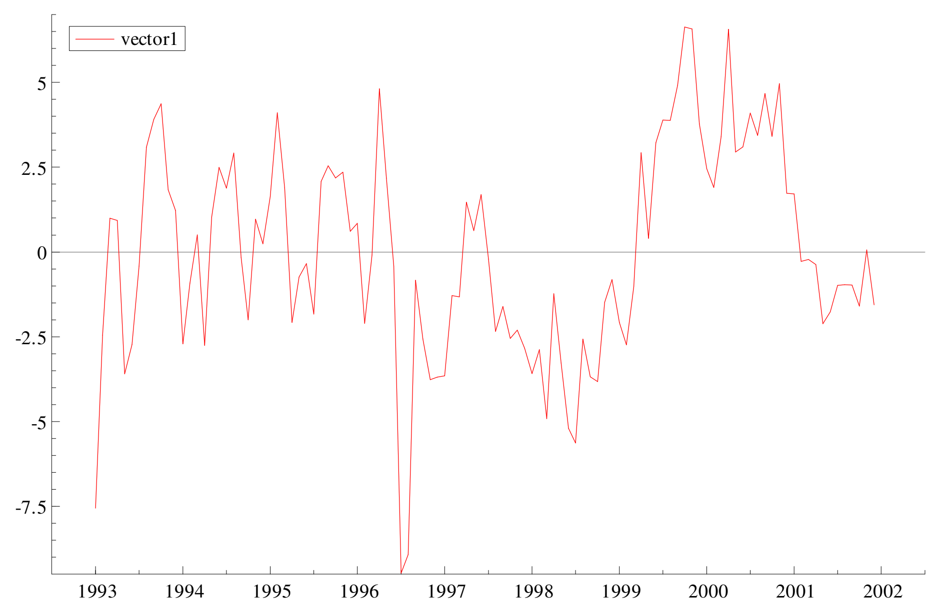
Bread



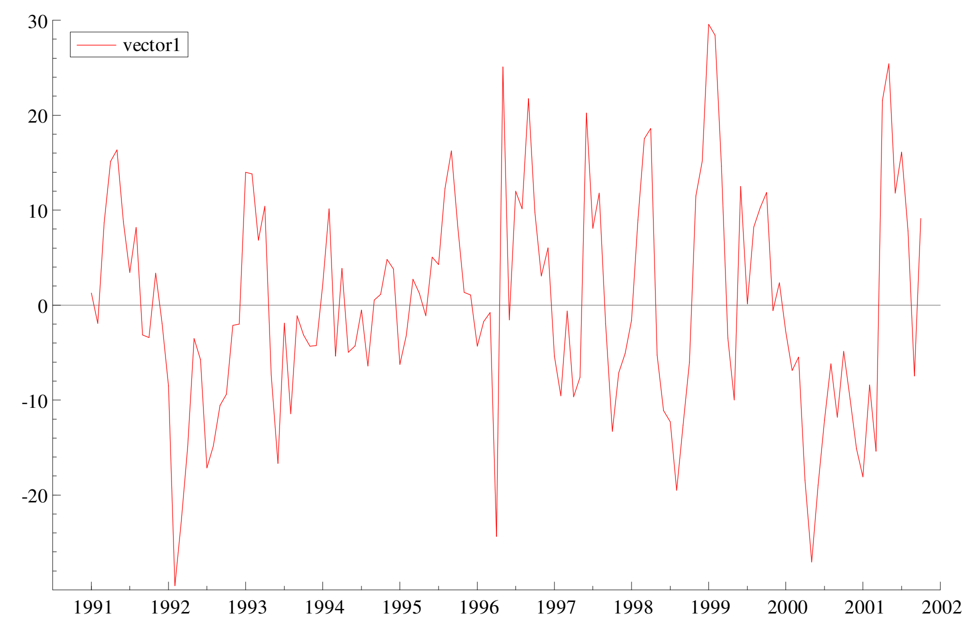
Chicken



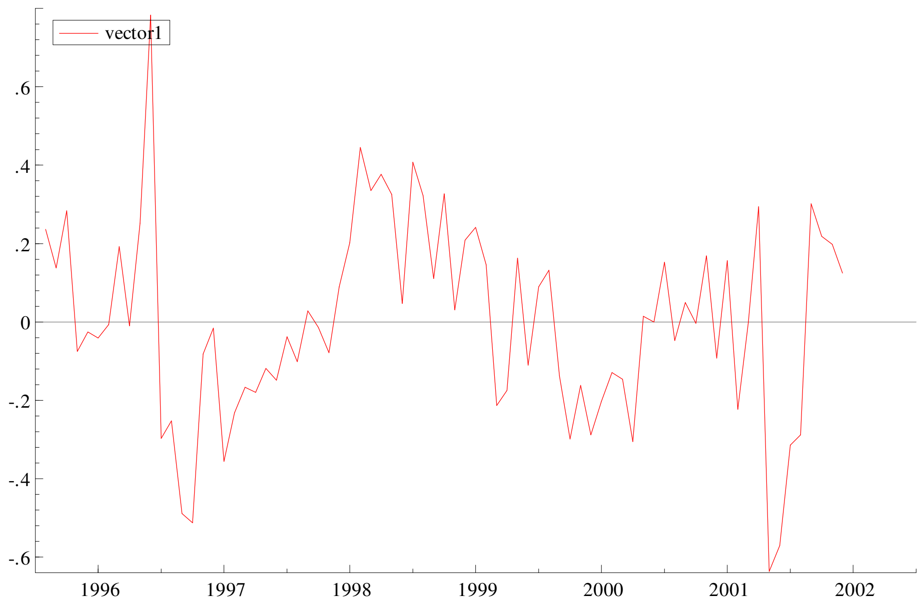
Eggs


vector1
5
2.5
0
-2.5
-5
-7.5
1993
1994
1995
1996
1997
1998
1999
2000
2001
2002

Lamb


vector1
30
20
10
0
-10
-20
1991
1992
1993
1994
1995
1996
1997
1998
1999
2000
2001
2002

Milk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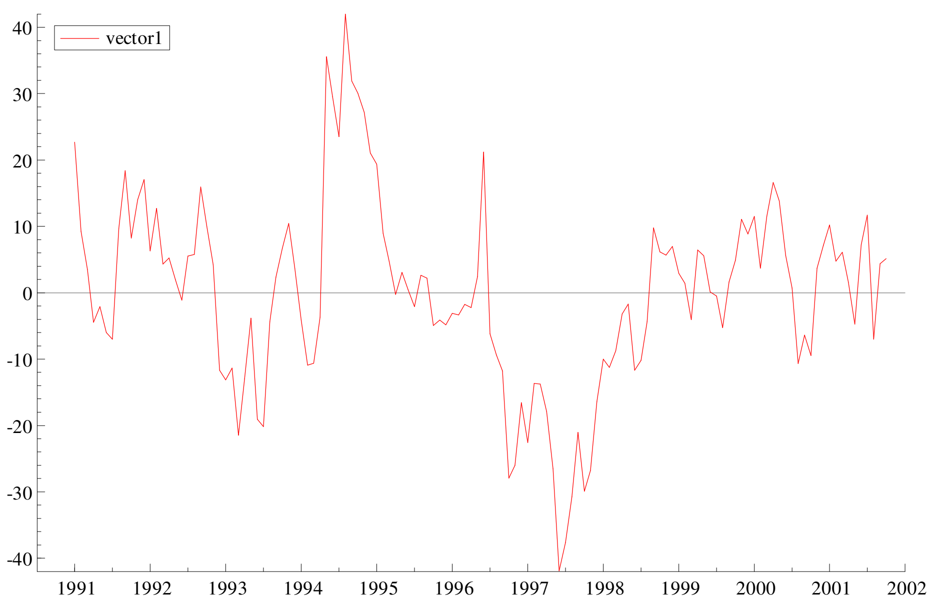
.4
.2
0
-.2
-.4
-.6
1996
1997
1998
1999
2000
2001
2002

Pork


vector1
40
30
20
10
0
-10
-20
-30
-40
1991
1992
1993
1994
1995
1996
1997
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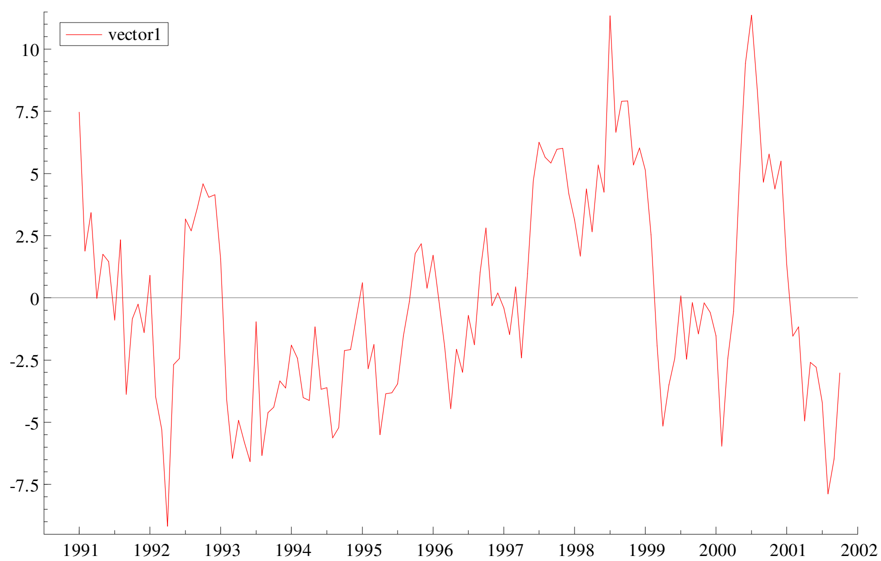
1999
2000
2001
2002

Potatoes



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