# Market Power in UK Food Retailing: Theory and Evidence from Seven Product Groups

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# Market Power in UK Food Retailing: Theory and Evidence from Seven Product Groups

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**Abstract**

Establishing the presence of market power in food chains has become an increasingly pertinent line of enquiry given the trend towards increasing concentration that has been observed in many parts of the world. This paper presents a theoretical model of price transmission in vertically related markets under imperfect competition. The model delivers a quasi-reduced form representation that is empirically tractable using readily available market data to test for the presence of market power. In particular, we show that the hypothesis of perfect competition can be rejected if shocks to the demand and supply function are significant and correctly signed in price transmission equations. Using a cointegrated vector autoregression, we find empirical results that are consistent with downstream market power in six out of seven food products investigated, supporting both the findings of the UK competition authority's recent investigation in to supermarkets and renewed calls for further scrutiny of supermarket behaviour by the UK's Office of Trading.

**Key words:** imperfect competition, Cointegrated VARs, UK food industry

**JEL**: D4, L81.

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## Introduction

As the degree of market concentration in European food retail markets has increased in recent years, concern has been expressed by many, including regulatory bodies, over the potential impact this might have on relationships between retailers and their suppliers in the food chain (Clarke *et al*, 2002). A key issue, as highlighted by the UK's Competition Commission (2000), is the extent to which retailers can exert buyer power over their suppliers and what impact this has on welfare, broadly defined. However, before welfare effects can be evaluated, it is vitally important to establish that market power actually exists and it is here that this paper seeks to make a theoretical and empirical contribution to the interpretation and understanding of vertically related markets.

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). Alternatively, industry-wide enquiries such as that undertaken by the UK's competition authorities (Competition Commission *op cit*) to gather very rich data are both time consuming and expensive. For example, the UK enquiry took 18 months to complete at a costs of some £30 million. Consequently, investigations of this sort are unlikely to be carried out every time concerns are raised over possible abuses of market power. What is needed therefore is the provision of a simple yet robust test to detect the existence of market power, which avoids the naivety of simple concentration ratios and the costs of a full regulatory enquiry. In this paper we provide such a test by devising a simple quasi-reduced form model of pricing in a vertical market that facilitates the testing of hypotheses it posits with readily available market data from seven UK food groups.

The most accessible data are prices and these can be traced along a vertical chain as food products move pass through it. The transmission of prices in such markets has received a great deal of attention since Gardner's (1975) seminal work. However, what Gardner (*op cit*) assumed was perfect competition and as McCorriston *et al* (2001) show, price transmission is greatly affected once we allow for imperfect competition in the chain. In other words, the pattern of prices we expect to see will be different in a world characterised by imperfect competition compared to one where perfect competition exists. We contend that this notion can allow researchers to use price data supplemented by appropriate marketing cost and other data to establish the presence of market power in a vertically related market.

Price and marketing data provide good indicators of behaviour in markets. In a perfectly competitive world, the difference (or spread) between two prices at different marketing levels can be attributed solely to marketing costs. If market power exists then the spread will not behave in this predictable fashion since price setting by the sector with market power will be reflected in the mark up that the sector can earn, and so affect the spread.

Hence, as we show in section 2, where market power exists market shocks have a differential impact at each stage in the marketing chain and thus determine the behaviour of the spread in addition to marketing costs. In effect, shocks to the underlying supply and demand functions are mediated through market power parameters and thus give rise to predictable effects on the spread. In the absence of market power, the effect of shocks is common at all 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-stage vertical market that explicitly allows for shocks to both the demand and supply functions of the product under investigation. Our aim is not to measure the extent of market power but to develop an empirical test for its presence. Moreover, given that the impact of shocks appear with definite sign in the theoretical model of the spread, the basis for reliable inference regarding market power is strengthened accordingly. Our approach is applied to data from seven food groups in the UK food industry. The empirical test rejects the null of perfect competition in all but one case. Furthermore, coefficients are signed according to the predictions in the theoretical model in the overwhelming majority of cases.

The paper is structured as follows. In Section 1 we outline the theoretical model that underpins our conceptualisation of a vertically related market. The econometric techniques employed are discussed in Section 2 while Section 3 describes the data. The results of the testing procedure are outlined in Section 4 and we offer some concluding thoughts and caveats in Section 5.

**1. Theory**

In this section, we outline a simple framework that delivers a formal test of market versus perfect competition that we use to motivate the empirical analysis. The demand function for the processed product is given by:

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

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

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

where $A$ is the quantity of the agricultural raw material and $N$ is the exogenous shifter in the farm supply equation.

In accordance with the findings of the Competition Commission (*op. cit.*) the source of market power in the food chain is given to be at the retail level. 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 which is assumed to equal 1. This assumption corresponds closely to the construction of the data in the vertical market chain used in the empirical analysis that follows. Constant returns to scale are assumed. 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):

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

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

with domestic supply being given by:

$$S = Q + N$$

where N is the level of exports which are exogenously determined. From this we can rewrite (4) as:

$$R - \frac{\mathrm{q}}{b} Q = M + P + \mathrm{mg}Q \tag{4'}$$

where q and m as average output and input conjectural elasticities respectively, such that with n firms in the industry $q = (\Sigma_i [\partial Q/\partial Q_i][Q_i/Q])/n$ and $m = (\Sigma_i [\partial A/\partial A_i][A_i/A])/n$. These parameters can be interpreted as an index of market power with $q = m = 0$ representing competitive behaviour and $q = m = 1$ representing collusive behaviour. $M$ 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 costs function of the form:

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

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:

$$Q = \frac{(h - by - bk) + cX - bzE - bgN}{(1+q) + bg(1+m)} \tag{6}$$

$$R = \frac{h + [(1+q) + bg(1+m)][(1-b)(y + k + gN) + (1-bz)E + cX)]}{(1+q) + bg(1+m)} \tag{7}$$

$$P = \frac{g[h - by + cX - bzE] - g[b - ((1+q) + bg(1+m))(k+N)]}{(1+q) + bg(1+m)} \tag{8}$$

To derive the retail-farm spread, use (7) and (8) to give

$$R - P = \frac{h(q/b + gm) + (1+bg)(y+zE) + (q/b + gm)cX - (q + bgm)(k+gN)}{(1+q) + bg(1+m)} \tag{9}$$

Note that if neither oligopoly nor oligopsony power matters in determining the retail-farm price spread (i.e. $q = m = 0$), then equation (9) reduces to:

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

i.e. the source of the retail-farm price margin 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 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 either oligopoly and/or oligopsony power in the food sector is important, then they will influence the margin between retail and farm prices. In other words, each shifter will affect the two prices differentially and thus the margin between the prices will change.

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

$$R = b_0 + b_1 P + b_2 M + b_3 X + b_4 N \tag{11}$$

Hence the test for the existence of market power is whether the coefficients on these variables in the retail-farm spread equation are statistically significant. Specifically, rejection of the null hypothesis,

$$H_0 : b_3 = b_4 = 0$$

implies market power. Furthermore, equation (9) unambiguously signs the effect of the shifters in the presence of market power. Whereas shocks to the demand shifter widens the margin, supply-side shocks narrow it, hence if market the shifters are significant in the margin equation, theory predicts that $b_3 > 0$ and $b_4 < 0$ in (11). In the following empirical section, we test these propositions using data for seven commonly purchased product groups in the UK.

## 2. 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 eight product groups it is assumed that the data may be approximated by a VAR($p$) model,

$$\mathbf{x}_t = \Phi_1 \mathbf{x}_{t-1} + \Phi_2 \mathbf{x}_{t-2} + \ldots + \Phi_p \mathbf{x}_{t-p} + \Psi \mathbf{D}_t + \varepsilon_t \qquad (12)$$

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 constants and centered seasonals and each $\Phi_i$ $(i = 1, \ldots, p)$ and $\Psi$ are $(k \times k)$ and $(k \times d)$ matrices of coefficients to be estimated using a ($t$ = 1, . . .$T$) sample of data. $\varepsilon_t$ is a $(k \times 1)$ vector of i.i.d. disturbances with zero mean and non-diagonal covariance matrix, $\Sigma$.

Equation (12) represents an unrestricted reduced form representation 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 information criteria and vector-based diagnostics. The preferred lag length is thus the most parsimonious model that is free of residual correlation at the 5% significance level.

The presence of cointegration is detected by estimating (1) in its error correction representation using Johansens's (1988) maximum likelihood procedure,

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

Attention focuses on the ( $k \times \mathrm{r}$ ) matrix of co-integrating vectors, comprising $\beta$, 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, $\alpha$, the elements of which load deviations from equilibrium (*i.e.* $\boldsymbol{\beta}' \mathbf{x}_{t-k}$ ) into $\Delta \mathbf{x}_t$, for correction. The $\Gamma_i$ 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. The number of cointegrating relations, corresponding to the rank of $\beta$ in (12), is evaluated by Johansen's Trace ($\eta_r$) and Maximal Eigenvalue ($\xi_r$) test statistics (Johansen, 1988). The $\eta_r$ statistic tests the null that there are at least $r$ cointegrating relationships ($0 \leq r < n$) and the $\xi_r$ evaluates the null that there are $r$ against the alternative that there are at most $r+1$ such relationships. While the $\eta_r$ 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. In the empirical analysis that follows we also report both asymptotic and the degree-of-freedom-adjusted test statistics of Cheung and Lai (1993).

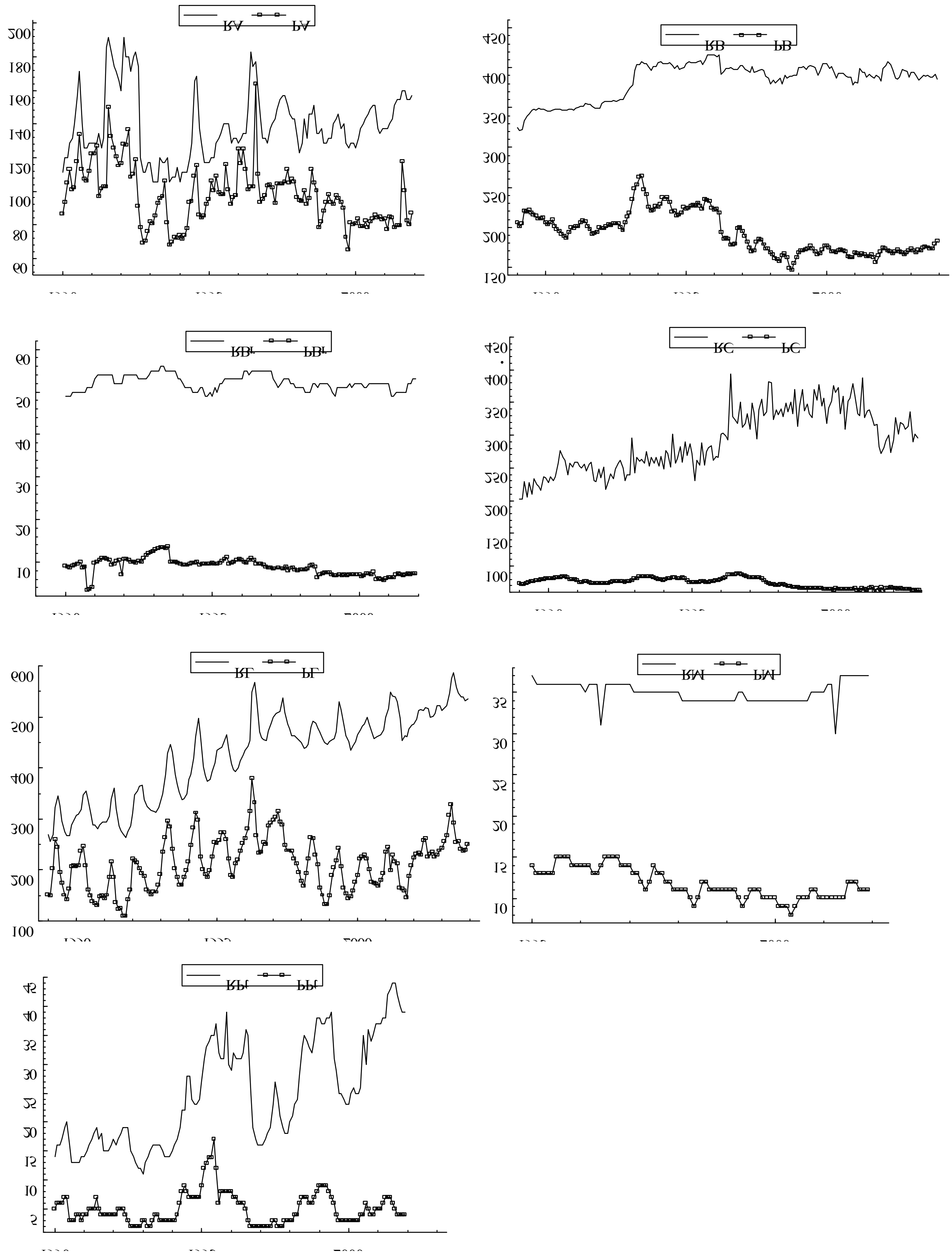
Where a single cointegrating relationship is detected, formal testing is undertaken to investigate whether market power is implied. Following from section 2, if the vertical market for a product is perfectly competitive, retail and producer prices may be expected to form a cointegrated relationship with at most marketing costs. Where retail market power is present, the 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 we can offer some additional evidence on market power by comparing the estimated signs of the shifters in the cointegrating relation with that predicted by the theoretical model.

## 3. Data[2]

In this paper we analyse the nominal monthly prices of seven UK food products, namely: apples (A); beef (B); bread (Br); chicken (C); lamb (L); milk (M) and potatoes (Pt) at retail (R) and producer (P) levels. In addition, each price model includes three industry-level 'shifters' representing proxies for marketing costs and shocks to the demand and supply functions. Where possible, retail and producer product prices are expressed in prices per standard unit (pence/kg of carcass weight for all meats; pence/pint for liquid milk, pence/lb for potatoes, and apples are an index [1987=100] of prices in pence/lb). For bread, price series are expressed in natural logs (of a standard sliced loaf and bread wheat respectively) and thus differ from the other prices in that there is no common unit of measurement. While this is inevitable given the product's transformation between retail and producer levels, it does have implications for the underlying functional form of the pricing relation, which was assumed to be linear in Section 2. Hence, bread does not sit as neatly in the theoretical framework as the other products analysed in this study. The price series are illustrated in Figure 1.

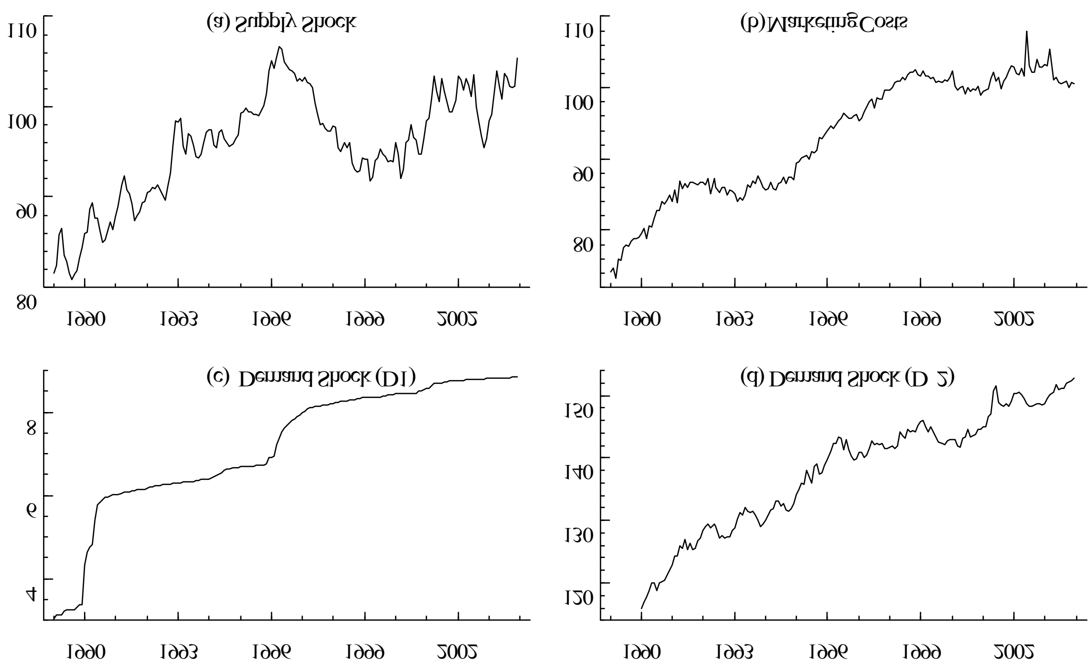
[2] Details and sources of data series used are given in Appendix 1. All statistical analysis is undertaken in PCGIVE 10.0 Hendry and Doornik (2001). Data and results are available upon request.

Figure 1: Product Price Series at Retail and Producer Levels



As Figure 1 illustrates, there is considerable variation in the price series between products and across marketing levels, although a tendency to diverge over time is a common feature, with the possible exception of bread.[3] While growth in the price spread is not in itself indicative of market power (marketing costs may account for it), it is necessary given the strong trend-like behaviour of the shifters, which are plotted in Figure 2.

Figure 2: Shifters




Referring to Figure 2 it is evident that all shifters display the tendency to grow over time. As noted in section 1, measures of product-specific marketing costs are not available in the UK and thus we use an index of unit wage cost index for manufacturing

[3] Time series plots of the spreads themselves (not shown in the interests of brevity) clearly demonstrate this tendency, even for bread.

industries (M), on the grounds that such costs are typically thought to represent some 70% of food manufacturing costs (Wholgenant, 2001). In order to incorporate the impact of farm-level production costs, the supply shifter (S) represents a price index of all goods and services purchased on UK farms. Demand-side shocks are proxied by two measures: for meat products we use the (natural logarithm of the) cumulative count of articles regarding the health and safety of food published in four broadsheet newspapers (D1) and the food retail price index (D2) for non-meat products. Application of the Augmented Dickey-Fuller test indicates that all prices and shifters is integrated of order one in levels and stationary in first differences. ADF test statistics are reported in Appendix 2.

## 4. Results

Having established the non-stationarity of the data, equation (13) is estimated for each of the seven product groups sequentially for $k = 13$ to 1. Since there is no consensus on the best criterion to use to determine lag length, three commonly applied measures are used here, namely the information criteria developed by Shartwz, Hannan-Quinn and Akaike (SBC, HQC and AIC respectively) and vector diagnostic tests for residual autocorrelation, heteroscedasticity and normality. The SBC tends to select the most parsimonious model and the AIC the least with the HQC selecting a lag length that is generally common to one of the other two, in roughly equal measure. In only one case (milk) is the lag length selected by the three information criteria unanimous. The vector tests for residual autocorrelation and heteroscedasticity tend to select models with longer lag lengths and hence concur with the AIC in most cases. To determine the preferred lag length, a consensus view is taken, although this usually conforms to the most parsimonious model in which the null of no residual correlation cannot be rejected at 5% significance. In many cases, test statistics reject the null of (residual) normality emphasizing that care should be exercised in interpreting results. Notwithstanding this caveat, the selected models are unrestricted reduced forms and represent the baseline models against which parameter restrictions are evaluated.

As a first step, the cointegrating rank is evaluated in the selected specification for each product group. Table 1 reports the results from the cointegration analysis using the Trace ($\eta_r$) and maximal Eigenvalue ($\xi_r$) tests in asymptotic ($\infty$) and finite sample (*T-mp*) forms (Cheung and Lai, 1993). Overall, the evidence points firmly to the presence of a single cointegrating vector in all product groups. Evaluating hypotheses at the 5% significance level, the null of no cointegration is rejected in 14 out of 16 tests using asymptotic critical values and on 10 out of 14 occasions using degree-of-freedom-adjusted critical values. Confining inference to the more stringent (degree of freedom adjusted) tests, every product has at least one statistic rejecting the null of no cointegration at the 5% level. Evidence for two cointegrating vectors is confined to $\eta_r(\infty)$ statistics which rejects at 5% for chicken and lamb. No finite sample statistics reject the null of multiple cointegrating vectors at this level of significance.

On the basis of the results in Table 1 and plots of cointegrating residuals (not shown), we proceed on the assumption that a single cointegrating vector is present for each product group. Coefficients of the cointegrating vectors along with their asymptotic standard errors are reported in Table 2.

Table 1: Asymptotic ( $\infty$ ) and Finite Sample Test Statistics for Cointegration

| **Product** | **Rank** | **Trace** $h_r(\infty)$ | **Maximal Eigenvalue** $x_r(\infty)$ | **Trace** $h_r(T\text{-}mp)$ | **Maximal Eigenvalue** $x_r(T\text{-}mp)$ |
|---|---|---|---|---|---|
| Apples | 0 | 83.77 [0.002]** | 36.88 [0.018]* | 77.87 [0.009]** | 34.28 [0.041]** |
| | 1 | 46.89 [0.060] | 23.28 [0.165] | 43.59 [0.118] | 21.64 [0.247] |
| | 2 | 23.62 [0.224] | 16.36 [0.213] | 21.95 [0.311] | 15.20 [0.286] |
| | 3 | 7.26 [0.554] | 4.66 [0.782] | 6.75 [0.613] | 4.33 [0.819] |
| | 4 | 2.60 [0.107] | 2.60 [0.107] | 2.42 [0.120] | 2.42 [0.120] |
| Beef | 0 | 78.75 [0.007]** | 40.89 [0.004]** | 71.18 [0.037]* | 36.96 [0.017]* |
| | 1 | 37.86 [0.312] | 23.90 [0.140] | 34.22 [0.495] | 21.60 [0.250] |
| | 2 | 13.96 [0.843] | 7.29 [0.932] | 12.62 [0.905] | 6.59 [0.959] |
| | 3 | 6.67 [0.622] | 5.11 [0.729] | 6.03 [0.695] | 4.62 [0.787] |
| | 4 | 1.56 [0.211] | 1.56 [0.211] | 1.41 [0.235] | 1.41 [0.235] |
| Bread | 0 | 79.27 [0.006]** | 32.69 [0.066] | 73.69 [0.022]* | 30.39 [0.124] |
| | 1 | 46.58 [0.064] | 27.31 [0.051] | 43.30 [0.125] | 25.39 [0.092] |
| | 2 | 19.27 [0.485] | 12.62 [0.501] | 17.91 [0.583] | 11.73 [0.586] |
| | 3 | 6.64 [0.625] | 4.77 [0.769] | 6.18 [0.679] | 4.43 [0.807] |
| | 4 | 1.88 [0.171] | 1.88 [0.171] | 1.74 [0.187] | 1.74 [0.187] |
| Chicken | 0 | 85.85 [0.001]** | 35.92 [0.024]* | 76.84 [0.011]* | 32.15 [0.077] |
| | 1 | 49.93 [0.030]* | 26.21 [0.072] | 44.69 [0.095] | 23.46 [0.158] |
| | 2 | 23.72 [0.219] | 14.84 [0.313] | 21.24 [0.353] | 13.28 [0.441] |
| | 3 | 8.89 [0.383] | 6.24 [0.590] | 7.96 [0.477] | 5.59 [0.671] |
| | 4 | 2.65 [0.104] | 2.65 [0.104] | 2.37 [0.124] | 2.37 [0.124] |
| Lamb | 0 | 82.11 [0.003]** | 34.23 [0.042]* | 75.15 [0.016]* | 31.32 [0.097] |
| | 1 | 47.88 [0.048]* | 25.79 [0.082] | 43.83 [0.113] | 23.61 [0.152] |
| | 2 | 22.09 [0.303] | 15.68 [0.254] | 20.22 [0.419] | 14.35 [0.351] |
| | 3 | 6.41 [0.651] | 5.25 [0.712] | 5.87 [0.713] | 4.81 [0.765] |
| | 4 | 1.16 [0.281] | 1.16 [0.281] | 1.06 [0.302] | 1.06 [0.302] |
| Milk | 0 | 103.04 [0.000]** | 61.83 [0.000]** | 96.83 [0.000]** | 58.11 [0.000]** |
| | 1 | 41.20 [0.183] | 20.87 [0.294] | 38.72 [0.275] | 19.61 [0.381] |
| | 2 | 20.33 [0.411] | 11.38 [0.619] | 19.11 [0.496] | 10.70 [0.684] |
| | 3 | 8.95 [0.377] | 8.53 [0.335] | 8.41 [0.430] | 8.02 [0.385] |
| | 4 | 0.42 [0.517] | 0.42 [0.517] | 0.39 [0.530] | 0.39 [0.530] |
| Potatoes | 0 | 67.89 [0.069] | 39.08 [0.008]** | 60.67 [0.216] | 34.92 [0.033]* |
| | 1 | 28.81 [0.777] | 13.35 [0.857] | 25.75 [0.894] | 11.93 [0.925] |
| | 2 | 15.47 [0.754] | 10.93 [0.662] | 13.82 [0.850] | 9.77 [0.767] |
| | 3 | 4.53 [0.852] | 3.43 [0.904] | 4.05 [0.893] | 3.07 [0.932] |
| | 4 | 1.10 [0.295] | 1.10 [0.295] | 0.98 [0.322] | 0.98 [0.322] |

** denotes significance at 1%; * at 5% and $p$-values are in parentheses. Asymptotic ( $\infty$ ) critical values are those of Osterwald-Lenum (1992) and finite sample (degree of free dom) adjusted test statistics are those of Cheung and Lai (1993) where the correction is ( $T - mp$ ) where $T$ is sample size and m is number of endogenous variables and $p$ is the lag length in the VAR.

Table 2: The Cointegrating Vectors
(normalised on retail prices)

| Product | Producer prices ($b_1$) | Marketing costs ($b_2$) | Demand shifter ($b_3$) | Supply shifter ($b_4$) |
|---|---|---|---|---|
| Apples | 1.94** (0.23) | -6.42** (2.2) | 8.07** (2.21) | -3.73** (1.33) |
| Beef | 2.02** (0.23) | 6.15** (1.44) | 18.5* (7.39) | -3.19** (0.88) |
| Bread | 0.273** (0.048) | 0.016** (0.005) | -0.012* (0.005) | 0.004 (0.003) |
| Chicken | 10.38** (1.55) | 12.31** (3.04) | 30.3 (16.24) | -11.79** (1.93) |
| Lamb | 3.95** (0.62) | -7.19 (6.55) | 148.03** (42.12) | -29.01** (5.73) |
| Milk | 0.55** (0.13) | 0.06 (0.08) | 0.10 (0.10) | -0.13* (0.05) |
| Potatoes | 0.49 (0.32) | -2.02** (0.54) | 3.24** (0.32) | -1.67** (0.53) |

Figures in bracket are asymptotic standard errors; ** denotes significance at the 1% and *denotes significance at the 5% level.

As noted in section 2, the theoretical model signs the coefficients of the long run relationship in the presence of market power, namely, $b_1 > 0$, $b_2 > 0$, $b_3 > 0$ and $b_4 < 0$. Although inference in cointegrated VARs is best undertaken using formal likelihood ratio tests rather than coefficient standard errors (see below), a number of the results in Table 2 are worthy of note: first, price transmission coefficients ($b_1$) are positive in all cases and statistically significant at the 5% level for all products except for potatoes; second, marketing costs, as proxied by labour costs in manufacturing, ($b_2$) are positive in four cases, significantly so in three; third, the demand shifter coefficient ($b_3$) is significantly positive in the cointegrating relations of six out of seven products; and fourth, the coefficient on the supply shifter is significantly negative in six out of seven products.

These results suggest that in the main, the shifters play an important role in the long run determination of prices, and enter the cointegrating relations with signs that are consistent with the use of retail market power. To investigate this issue further, we perform a second set of tests to evaluate the validity of excluding the shifters from the cointegrating vectors. The results from evaluating these exclusion restrictions using likelihood ratio statistics are reported in Table 3. The first two columns test the individual significance of each shifter in each cointegrating vector and thus perform the same role as the standard errors in Table 2. The performance of the $c^2$ tests of Table 3 is known to be superior to the use of asymptotic standard errors, however in this case both yield very similar results. The final column evaluates the hypothesis that both shifters are jointly zero. As described in section 2, both shifters are statistically significant in the presence of market power, so the joint hypothesis in Table 3 explicitly tests this.

Table 3: Tests for Market Power

| Product | $H_0 : b_3 = 0$ | $H_0 : b_4 = 0$ | $H_0 : b_3 = b_4 = 0$ |
|---|---|---|---|
| Apple | 6.38 [0.01]* | 3.86 [0.05]* | 7.04 [0.03]* |
| Beef | 4.06 [0.04]* | 10.76 [0.00]** | 11.12 [0.00]** |
| Bread | 6.49 [0.01]* | 0.26 [0.610] | 5.66 [0.06] |
| Chicken | 4.12 [0.04]* | 0.47 [0.49] | 26.48 [0.00]** |
| Lamb | 4.69 [0.03]* | 8.34 [0.00]** | 15.50 [0.00]** |
| Milk | 0.66 [0.42] | 5.83 [0.02]* | 7.71 [0.02]* |
| Potatoes | 18.14 [0.00]** | 16.1 [0.00]** | 18.30 [0.00]** |

Figures in bracket are asymptotic *p*-values; ** denotes significance at the1% and *denotes significance at the 5% level

The null, which corresponds to perfect competition, is rejected for all products except bread at the 5% level. Similar likelihood ratio tests for the significance of the shifters individually reject in 11 out of 14 cases. Overall, the behaviour of prices in the majority of products considered here are consistent with the use of market power.

## 5. Conclusion

In this paper we have attempted to devise a simple yet robust means of testing for the presence of market power. By constructing a quasi-reduced form model of a vertically related food market, we can establish a simple hypothesis that the null of perfect competition can be rejected if the shifters from the supply and demand equations are significant and correctly signed. In framing this approach we are able to move away from the naivety of simple measures of concentration, and although the results from our statistical tests are far less authoritative than the findings of a regulatory inquiry they are relatively quick and costless to conduct. Indeed, out tests are better thought of as forming part of an preliminary assessment prior to any such authoritative investigation.

Drawing on data from seven food products in the UK food industry we show that in all but one case, we reject the hypothesis of perfect competition, implying that for these food products at least, the market is characterised by imperfect competition. Bread is the exception and something of an anomaly: although it rejects the perfectly competitive null at the 6% level, the shifters are perversely signed (albeit insignificantly so in the case of the supply shifter). Whether this reflects that bread is sold to supermarkets by a concentrated bakery sector with a degree of countervailing power that suppliers of the other products do not command, or simply that the data used do not sit neatly in the theoretical framework, is impossible to assess.

As always, conclusions, particularly those based on statistical tests from market-level data, are subject to caveat. Whilst care has been taken to select products appropriate to the theoretical framework and use reliable data from official sources, there are number of issues that should be borne in mind. First and foremost is the quality of the proxies used, particularly the measure of marketing cost. Whilst labour costs commonly represent the single most important component of total costs, it is nevertheless an industry-wide measure, which may or may not be representative of the actual costs of transforming individual products at the farm gate into the consumer product. Indeed, in two of the eight products studied (apples and potatoes) the labour cost proxy entered the pricing relationship with a significantly negative coefficient, contrary to the prediction of the theoretical model. Also, the theoretical model itself is predicated on a number of simplifying assumptions, (e.g. constant proportions, conjectural variations) whose empirical veracity in the cases studied is difficult to determine. However, notwithstanding these and other limitations the results point firmly to the rejection of perfectly competitive pricing behaviour in the majority of products analysed. 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.

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## Appendix Table 1: Data Definitions and Sources

| Label | Variable | Units | Sample | Obs | Area | Comments | Data Source |
|---|---|---|---|---|---|---|---|
| **RA** | Retail apple | Index of pence/lb (1987=100) | 1990.1 – 2001.12 | 144 | UK | Desert apples only | Employment Gazette/Labour Market Trends |
| **PA** | Producer apple | Index of pence/lb (1987=100) | 1990.1 – 2001.12 | 144 | UK | Exclude direct subsidies | Department of Food, Environment and Rural Affairs |
| **RB** | Retail beef price | Pence/kg carcass weight equivalent | 1989.1 – 2003.12 | 168 | GB | Converted in to c.w.e. by MLC | Meat and Livestock Commission |
| **PB** | Producer beef price | Pence/kg carcass weight | 1989.1 – 2003.12 | 168 | GB | MLC sample average | Meat and Livestock Commission |
| **RBr** | Retail bread price | ln(pence/800g loaf) | 1990.1 – 2001.12 | 144 | UK | Standard white sliced | Employment Gazette/Labour Market Trends |
| **PBr** | Producer bread price | ln(£/ton) | 1990.1 – 2001.12 | 144 | UK | Bread wheat | Department of Food, Environment and Rural Affairs |
| **RC** | Retail chicken price | Pence/kg carcass weight | 1989.1 – 2002.12 | 156 | GB | Uncooked whole birds including frozen <1.81 kg | National Food Survey/Expenditure and Food Survey |
| **PC** | Producer chicken price | Pence/kg carcass weight | 1989.1 – 2002.12 | 156 | E&W | Birds <2.27 kg | National Farmers Union |
| **RL** | Retail lamb price | Pence/Kg carcass weight equivalent | 1989.1 – 2003.12 | 168 | GB | Converted in to c.w.e. by MLC | Meat and Livestock Commission |
| **PL** | Producer lamb price | Pence/kg carcass weight | 1989.1 – 2003.12 | 168 | GB | MLC sample average | Meat and Livestock Commission |
| **RM** | Retail milk price | Pence/pint | 1995.1 – 2001.12 | 84 | UK | Semi skimmed only | Employment Gazette/Labour Market Trends |
| **PM** | Producer milk price | Pence/pint | 1995.1 – 2001.12 | 84 | UK | Average all milk | Department of Food, Environment and Rural Affairs |
| **RPt** | Retail potato price | Pence/lb | 1990.1 – 2001.12 | 144 | UK | Old white, sold loose | Employment Gazette/Labour Market Trends |
| **PPt** | Producer potato price | Pence/lb | 1990.1 – 2001.12 | 144 | UK | Average all potatoes (including processor sales) | Department of Food, Environment and Rural Affairs |
| **D1** | Meat demand shock | Ln(cumulative count of newspaper 'food scare' articles) | 1985.1 – 2003.12 | 216 | UK | Articles appearing in Times, Sunday Times, Guardian and Observer about health and safety of food. | Euro-PA Associates, Northhampton. |
| **D2** | Non-meat demand shock | Food Retail Price Index (1987=100) | 1987.1 – 2003.12 | 192 | UK | Includes all food items in RPI | Office of National Statistics |
| **S** | Farm Supply Shock | Index of farm input prices (1997=100) | 1989.1 – 2003.12 | 168 | UK | Includes all Goods and services currently consumed on UK farms | Department of Food, Environment and Rural Affairs |
| **M** | Marketing shock | Index (2000=100) of seasonally adjusted unit wage costs in manufacturing | 1989.1 – 2003.12 | 168 | UK | Index of average unit wage cost in UK manufacturing. | Office of National Statistics |

**Appendix Table 2: ADF Test Statistics**

| Variable | Levels | | First-difference | |
|---|---|---|---|---|
| | ADF | Lag | ADF | Lag |
| **RA** | -2.67 | 0 | -10.88** | 0 |
| **PA** | -2.38 | 4 | -6.94** | 9 |
| **RB** | -1.88 | 0 | -12.70** | 0 |
| **PB** | -2.49 | 1 | -8.41** | 0 |
| **RBr** | -2.74 | 0 | -7.77** | 1 |
| **PBr** | -2.91 | 1 | -9.49** | 1 |
| **RC** | -1.52 | 3 | -11.10** | 2 |
| **PC** | -2.38 | 4 | -4.13** | 3 |
| **RL** | -1.83 | 6 | -7.22** | 6 |
| **PL** | -1.50 | 6 | -8.27** | 5 |
| **RP** | -1.67 | 0 | -11.20** | 0 |
| **PP** | -2.24 | 8 | -6.68** | 5 |
| **RM** | -1.14 | 3 | -7.78** | 2 |
| **PM** | -2.11 | 13 | -7.29** | 1 |
| **RPt** | -2.12 | 0 | -11.30** | 0 |
| **PPt** | -2.61 | 2 | -8.16** | 1 |
| **RE** | -2.75 | 1 | -9.46** | 1 |
| **PE** | -2.86 | 5 | -2.97* | 4 |
| **S** | -2.61 | 12 | -3.06* | 10 |
| **D1** | -1.93 | 3 | -3.18* | 3 |
| **D2** | -2.37 | 0 | -12.02** | 0 |
| **M** | -1.26 | 9 | -3.92** | 7 |

Lag length is selected on basis of the information criteria (see main text for details). Regressions include constant, trend and seasonals (if appropriate) in the levels; constant (and seasonals) only in first differences. 95% (*) and 99% (**) critical values are -3.45 and -2.88 respectively.