TESTING COLLUSIVE PRICES IN A MULTIMARKET CONTEXT: THE EUROPEAN CASE OF VITAMIN C*

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ABSTRACT

In this paper, we suggest a method to test price-fixing agreements. Prices fixed to multiple shipments are decomposed into a set of destination market effects and time effects in order to allow us to perform an analysis of residuals. We examine the pricing behavior of vitamin C in the European destination markets of German exports. We explore two different periods: January 1991 to August 1995 and September 1995 to September 2001. Empirical results on the first period, which are consistent with our knowledge obtained from firms' confessions about illegal agreements, contrast notably with those obtained on the more recent period.

Keywords: Collusion; International markets; Vitamin C

JEL classification: D43, L12, L41, L65

RESUMEN

En este trabajo proponemos un método para contrastar la presencia de prácticas no competitivas en precios. Cada uno de los precios fijados para diferentes mercados es descompuesto en un efecto fijo por destino y un efecto temporal con objeto de analizar los residuos. Los precios de la vitamina C en los mercados de destino de las exportaciones alemanas son examinados en dos periodos: de enero de 1991 a agosto de 1995 y de septiembre de 1995 a septiembre de 2001. Los resultados para el primer periodo, los cuales son consistentes con nuestro conocimiento obtenido de las confesiones de las empresas participantes en los acuerdos ilegales, contrastan notablemente con los obtenidos para un período más reciente.

Palabras clave: Colusión; Mercados internacionales; Vitamina C

Clasificación JEL: D43, L12, L41, L65

1. Introduction

Unfortunately, price-fixing agreements are a frequent feature in oligopolistic market structures. It is obvious that sanctions are not enough to dissuade these illegal practices. This is not surprising because it can yield considerable profits for participating firms and because there is no straightforward way it can be identified by antitrust agencies. Social welfare is likely to suffer notable losses in cases in which there are elements like the importance of the product for consumers, the large size of the geographic markets where participants operate, or maintaining illegal practices over a number of years. The very well known collusion between the primary manufacturers of vitamins is a good example displaying those elements. This is the reason why, at the end of 2001, the European Commission imposed record fines against eight pharmaceutical companies (bearing in mind that an important part of the agreement took place in the European economic area).[1] The damaging effects of the firms' infringements could be illustrated by the following declarations made by the President of the Commission:

"This is the most damaging series of cartels the Commission has ever investigated due to the sheer range of vitamins covered which are found in a multitude of products from cereals, biscuits and drinks to animal feed, pharmaceuticals and cosmetics… The companies' collusive behavior enabled them to charge higher prices than if the full forces of competition had been at play, damaging consumers and allowing the companies to pocket illicit profits. It is particularly unacceptable that this illegal behavior concerned substances which are vital elements for nutrition and essential for normal growth and maintenance of life." (Mario Monti, November 2001)

In this case, some participants cooperated with the European Commission[2] and admitted the existence of cartel agreements quite early on, thus facilitating the investigation, after being encouraged by the possibility of obtaining a substantial reduction in the fines that would eventually be imposed. Nevertheless, identification of illegal practices is a complicated affair. The tools available for antitrust agencies are

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[1] Fines were imposed on the following companies (in order of importance): Hoffmann-La Roche (Switzerland), BASF AG (Germany), Aventis SA (France), Solvay Pharmaceuticals BV (Netherlands), Merck KgaA (Germany), Daiichi Pharmaceutical Co Ltd (Japan), Esai Co Ltd (Japan), and Takeda Chemical Industries Ltd (Japan).

[2] Price agreements take place in other important markets like USA. Therefore, participants also cooperated with the United States Department of Justice Antitrust Division's Grand Jury.

poor. The aim of this paper is to provide a methodology that can be used to help identify the presence of non-competitive pricing behavior in a simple fashion. We suggest an empirical approach which will allow us to test the hypothesis of price collusion when firms operate in multiple markets.[3] There are at least two interesting implications that characterize this approach. First, only a data set of product prices fixed to several destination markets will be necessary, which implies that the methodology has a low cost. Second, since the methodology is based on a multimarket model, the analysis is expected to be useful for application to certain contexts where illegal agreements on prices are more likely. This idea is supported by recent studies, which show that firms' contact in multiple markets may enhance abilities to collude (Paker and Röller, 1997; Gupta, 2001).

A data set regarding prices fixed by German exporters of vitamin C over the admitted period of collusion will allow us to determine whether the empirical approach is valid to help identify price-fixing agreements. We also used the suggested methodology to investigate pricing behavior after this period. In this way, we tested whether authorities' control over the later period, with regard to the defense of competition, had been successful.

The rest of this paper is organized as follows. Section 2 develops an economic framework and discusses the implications of different market structures. In section 3 we propose the econometric specification and present the implications of a framework of price collusion and for two alternative market structures. In section 4 we explore the approach using export data on vitamin C for the European market. The two different periods are examined using month-to-month data. The first covers the period 1991:1 to 1995:8, where collusion is admitted by firms, and the second involves the period 1995:9 to 2001:9, where no price coordination is expected. The final section provides the summary conclusions.

2. The economic approximation

The economic framework used in this paper is focused on a profit-maximizing firm behavior that produces homogeneous goods for sale in N separate destination

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[3] The presence of collusion could be caused by an express agreement or by the existence of conscious parallelism (or tacit collusion). Nevertheless, for the purpose of this paper there is no need to distinguish between the two terms (although it has led to some confusion in legal proceedings). For a traditional explanation of cartel theory in both cases, see Stigler (1964).

markets, which are indexed by *i*. First order conditions for profit maximization could be described by prices (in *fob* terms) obtained as the product of the marginal cost (MC_t) and a mark-up (f_{it}), as follows:

$$p_{it} = MC_t\left(\sum_i^N q_{it}, w_t\right) f_{it}, \forall i \qquad (1)$$

where $t=1,\ldots,T$ index time. The marginal cost MC_t, which is common to the N destination markets, is a function of the level of total production of the firm $\sum_i^N q_{it}$, and the input price w_t.

Assuming that destination-specific mark-up could depend on price competitiveness, we can define it as the relative price in local market terms (Hung et al., 1993; Kongsted, 1998). In this way, an approximation of the optimal price, where arguments of MC_t have been removed, can be written (in logs) as:

$$\ln p_{it} = \ln MC_t + \ln k_i + b \ln\left(\frac{p_{it}^c}{p_{it} f_i}\right) + e_{it} \qquad (2)$$

where k_i is a constant term of mark-up over marginal cost which is specific to market *i*,[4] p_{it}^c is an index of prices of local competitors (in *cif* terms), f_i represents the specific iceberg cost of shipping consumer goods to destination *i*. Thus, $\left(\frac{p_{it}^c}{p_{it} f_i}\right) = P_{\text{it}}$ is the relative price in local market *i*. The coefficient b should be interpreted as the relative price elasticity of the mark-up ($b>0$). Lastly, e_{it} is a stochastic variable which is independent, identically normally distributed with mean zero and variance s_e^2.

The involvement of the different market structures on non-random components in equation (2) is straightforward. Let us consider the single competitive market, where $p_{it}^c = p_{it} f_i$. The fulfillment of this hypothesis requires prices to be equal to marginal cost, which is common across destination markets and implies $\ln k_i = b\ln(1) = 0$.

An alternative hypothesis is the presence of collusive prices. Although this hypothesis also implies price parallelism between firms within each destination market,

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[4] A constant term of mark-up could be associated to stable aspects like the distance between the firms and the destination market or the dimension of the specific market *i*.

in this case the relative price can take a value other than one. When this occurs, then $b\ \ln\left(\frac{p_{it}^{c}}{p_{it}\ f_i}\right) = b\ \ln\left(P_i\right) \neq 0$. Furthermore, in this interesting case, firms' practices presumably involve third degree price discrimination as well as the segmentation of markets. Thus, the existence of mark-up differences across destinations are represented in equation (2) by $\ln k_i + b\ \ln\left(P_i\right)$.

Lastly, changes in the relative price are inconsistent with both hypotheses (perfect competition and price collusion). This phenomenon is consistent with an extensive range of imperfect competition market structures in which the degree of competition is directly related to the value of the relative price elasticity b. In this range of cases, firms take into account changes in relative prices. Since the optimal price becomes an endogenous variable, the total effect is obtained by:

$$\ln p_{it} = g\ \ln MC_t + g\ \ln k_i + (1-g)\ f_i + (1-g)\ln p_{it}^{c} + g e_{it} \qquad (3)$$

where γ=1/(1+β). Equation (3) shows that while b is large (or γ is low), the firm follows competitors' prices and the market is characterized by a high degree of competition. The next section presents a simplification of equation (3) that will be capable of distinguishing the market structure in which we are interested from a range of intermediate market structures and from the perfect competitive case.

3. Research design

This section centers on the research design used to test collusion in a simple fashion and its empirical implications. Research design is based on a decomposition of prices p_{it}, where explanatory variables of equation (3) are taken to be unobservable. Differences between the closely related markets will allow us to trace the effects of the idiosyncratic characteristics of destination markets and common changes (like marginal cost, which is in accordance with the central ideas of the "new empirical industrial organization" studies, surveyed in Bresnahan, 1989). Since we take the relative price to be an unobservable variable, requirements of statistical information are low and the methodology involves a notable degree of pragmatism. It is obvious that sometimes a scarcity of data at firm level makes it difficult to build an accurate index of competitors' prices for each of the destination markets.

Price decomposition could be represented as a fixed-effects regression model, as follows:

$$\ln p_{it} = a + q_t + l_i + n_{it} \qquad (4)$$

where α is a constant term, q_t are the coefficients of the time effects which capture any common movements in prices over time across all destination markets (and, obviously, excludes q_1 to avoid collinearity with the constant term), and l_i is a set of market effects (that excludes l_1). The last term, $\nu_{it,}$ is an error term. The implication of the coefficients of equation (4) in the diverse market structures could be illustrated by the economic framework described in the previous section.

First, the fulfillment of the perfect competitive hypothesis implies that firm prices are fully explained by the constant term and the set of time effects. Thus, l_i is zero for overall destination markets and the regression disturbance (ν_{it}) equals e_{it}. In this case, while the constant α captures the marginal cost that corresponds to the first period (CM_1), the time effects q_t will measure their evolution (MC_t, $t=2,..,T$). This evolution equals the common changes in price in each period.

Second, the fulfillment of the hypothesis concerning non-competitive segmented markets, where pricing behavior depends on the idiosyncratic demand schedule faced by the firm inside each market, requires l_i to have a value other than zero. While q_t captures both the evolution of common mark-up and marginal cost, l_i, ($i=2,..,N$) indicates the specific mark-up difference from the destination country indexed by $i=1$. If price coordination implies that p_{it} equals p_{it}^c,[5] then ν_{it} is distributed in an identical and independent manner and is equal to e_{it}.

In the alternative market structures relative prices come into play. Therefore, the constant term, and the fixed effects from equation (4) are insufficient to explain pricing behavior. In this case, the error term (ν_{it}) captures both a random factor and a deterministic factor. More specifically, it captures $ge_{it} + (1-g)\, lnp_{it}^c$, where γ acts as the coefficient for the linear combination of the random and a deterministic factor.

[5] In the more general case of price parallelism in destination markets, it could also include the constant gap between firms' prices.

4. Data and results

In this section, we conduct an empirical analysis of the pricing behavior in the European Union markets of vitamin C, which is, together with vitamins A and E, one of most important products sold by drug firms (Table 1). A first stage of the analysis is based on OLS estimation from equation (4). To do so, we use unit values as a proxy variable of export prices of the vitamin. Unit value indices are calculated from data collected from the COMEXT database (published by *Eurostat*).[6] We have chosen to study the value indices obtained from German exports, which will represent the price dynamics of two of the four firms involved in the illegal agreements (BASF AG and Merck KgaA).[7] Furthermore, German shipments provide a steady, important volume of vitamin exports to several destination markets on a monthly basis. The study focuses on the main destination markets of German vitamin C exports in the European Union. In particular, we will obtain specific results for France, Belgium and Luxembourg, Netherlands, Italy, United Kingdom, Spain and Austria. The export data covers both the whole period of successful collusion in the European market, following the information offered by the *Official Journal of the European Communities* (Commission Decision of 21 November 2001, Case COMP/E-1/37.512),[8] that is to say, 1991:1 to 1995:8[9] and a more recent period between 1995:9 and 2001:9. From the results obtained by the investigation carried out by the Commission, we expect the illegal agreements to have been finished in this latter period.[10] We will thus be able to compare results for both periods.

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[6] Database follows the *Integrated Tariff of the European Communities* classification. Vitamin C is classified with code number 293627.

[7] The remaining firms implied in the vitamin C agreements belong to Switzerland (Hoffmann-La Roche) and Japan (Takeda Chemical Industries Ltd).

[8] As the *Official Journal* indicates the importance and the duration of the agreements is not necessarily the same for all participants and destination countries.

[9] German firms are subjected in destination markets of exports to frequent variations in the exchange rates throughout of this period, which induces automatically changes in competitiveness. Then, without the existence of price agreements, the relative prices would change regularly following the models based in the pricing-to-market behaviour (Krugman, 1987; Dornbusch, 1987).

[10] Although the precise stage at which the agreement on vitamin C was withdrawn is not documented, the Commission's decision is based on this date.

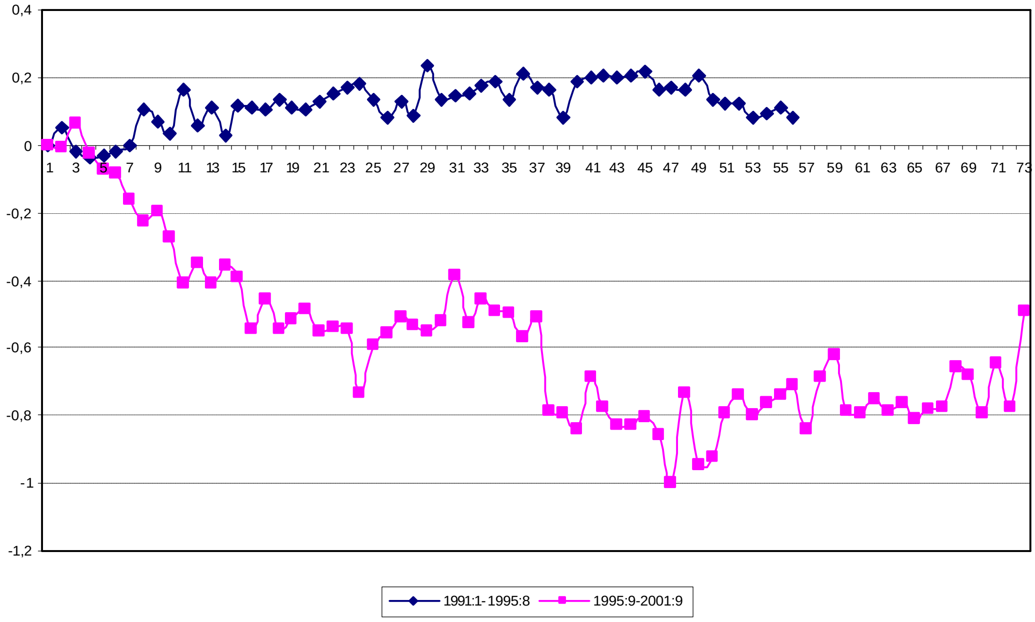
TABLE 1. Market shares of sales by vitamin types in the EU market

Product	1994	1996	1998
Vitamin C	27.79	20.84	15.24
Vitamin E	23.18	27.81	31.74
Vitamin A	16.97	18.30	18.98
Beta Carotene	6.81	8.45	9.83
Canthaxanthin	6.01	6.34	6.45
Vitamin B2	5.06	4.88	4.36
Biotin (H)	4.32	3.96	2.98
Pantothenates/calpan (B5)	3.83	4.08	4.51
Vitamin B1	2.18	1.58	1.91
Vitamin D3	1.98	2.54	2.61
Vitamin B6	1.87	1.21	1.39

Source: Percentages calculated from data published in the *Official Journal of the European Communities* (10/1/2003)

Estimated time effects are represented in Figure 1. We can observe that the common price evolution of vitamin C differs clearly between the two periods. While the common component of prices was quite stable over the collusion period, it decreased dramatically during the first two years of the later period. This finding appears as an immediate consequence of ending the agreements on vitamin C.

FIGURE 1. Estimated time effects for the two periods.



In a second stage, we studied the residuals of regression for both periods with the purpose of determining whether a deterministic component is included. If such a deterministic component is revealed, then collusion and a perfect competitive market would be rejected. Otherwise, the presence of idiosyncratic mark-up according to destination can separate these last two alternatives.

Table 2 and 3 present the analysis of residuals for the first period (1991:1 to 1995:8) based on the Box-Pierce statistic (Box and Pierce, 1970) and on the Box-Ljung statistic (Ljung and Box, 1979), respectively. We explored the statistical results with a set of degrees of freedom (*L=1,2,..10*) for each destination country. We cannot deny that, in general, the residuals follow a white noise process. When we attempt to apply the Box-Pierce statistic, we cannot reject independence of residuals over time for sixty-five out of seventy cases at the 1-percent level. When the Box-Ljung statistic is applied, the hypothesis is not rejected in sixty-two out of seventy cases. The only clear exception corresponds to Belgium-Luxembourg as a destination where, in view of results, we can conclude that illegal agreements did not take place at all during the period under consideration. Therefore, the evidence obtained is, in general, consequent with the presence of price parallelism over this period. However, a white noise process is a necessary but insufficient condition to support price coordination.

Rejection of perfect competition, in favor of non-competitive prices, is corroborated by significant differences in the estimations of country-fixed effects. As well as the product's being homogeneous, from the results included in Table 4, we can infer that there is a constant mark-up component which depends critically on the destination country. [11] That is, for example, mark-ups obtained from sales to both Spain and Italy are significantly smaller than the mark-up obtained from sales to the French market (which is the reference country in the regression). Obviously, differences in mark-ups across destinations agree with the price collusion that was proved to have existed over this period.

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[11] In practice, differences in country effects may result from other aspects that have not been considered. For example, country effects could pick up differences in product quality across destination countries which are caused by differences in tastes or per capita incomes.

TABLE 2. White noise test for residuals based on the *Box-Pierce* statistic for the period 1991:1 to 1995:8.

Lags(L) *Destination country(i)*	1	2	3	4	5	6	7	8	9	10
France	0.085 (0.926)	0.721 (0.697)	0.915 (0.822)	3.288 (0.511)	5.322 (0.378)	5.442 (0.489)	7.705 (0.359)	7.762 (0.457)	7.811 (0.553)	7.854 (0.643)
Belgium-Luxembourg	2.894 (0.089)	9.946 (0.007)	13.594 (0.004)	14.374 (0.006)	14.774 (0.011)	16.529 (0.011)	17.108 (0.017)	18.379 (0.019)	23.475 (0.005)	25.529 (0.004)
Netherlands	3.270 (0.071)	3.3505 (0.187)	3.406 (0.333)	3.464 (0.483)	3.615 (0.606)	3.625 (0.727)	4.771 (0.688)	5.506 (0.702)	5.537 (0.785)	6.086 (0.808)
Italy	0.825 (0.364)	0.982 (0.612)	0.988 (0.804)	1.600 (0.809)	2.147 (0.828)	3.114 (0.794)	3.120 (0.874)	3.142 (0.925)	3.772 (0.926)	4.751 (0.907)
United Kingdom	5.496 (0.019)	5.506 (0.064)	5.720 (0.126)	6.918 (0.140)	9.273 (0.159)	8.583 (0.198)	8.583 (0.284)	9.684 (0.288)	9.782 (0.368)	15.013 (0.136)
Spain	0.224 (0.636)	3.811 (0.149)	4.239 (0.237)	4.300 (0.367)	5.758 (0.330)	5.760 (0.451)	6.495 (0.483)	6.542 (0.587)	12.740 (0.175)	15.976 (0.101)
Austria	0.098 (0.753)	0.484 (0.785)	2.790 (0.425)	3.766 (0.439)	8.767 (0.119)	9.032 (0.172)	9.750 (0.203)	10.062 (0.261)	10.926 (0.281)	11.528 (0.318)

Note: Statistic is distributed as the chi-squared with *L* degrees of freedom. Values in brackets correspond to significance levels. Results were obtained with *Limdep 7.0*.

TABLE 3. White noise test for residuals based on the *Box-Ljung* statistic for the period 1991:1 to 1995:8.

Lags(L) Destination Country(i)	1	2	3	4	5	6	7	8	9	10
France	0.009 (0.924)	0.774 (0.679)	0.986 (0.805)	3.634 (0.458)	5.946 (0.311)	6.085 (0.414)	8.765 (0.270)	8.833 (0.356)	8.894 (0.447)	8.948 (0.537)
Belgium-Luxembourg	3.052 (0.081)	10.626 (0.005)	14.619 (0.002)	15.488 (0.004)	15.943 (0.007)	17.979 (0.006)	18.665 (0.009)	20.201 (0.010)	26.488 (0.002)	29.079 (0.001)
Netherlands	3.448 (0.063)	3.535 (0.171)	3.595 (0.309)	3.660 (0.454)	3.832 (0.574)	3.843 (0.698)	5.200 (0.636)	6.089 (0.637)	6.126 (0.727)	6.818 (0.743)
Italy	0.871 (0.351)	1.039 (0.595)	1.045 (0.790)	1.718 (0.786)	2.351 (0.799)	3.472 (0.748)	3.479 (0.838)	3.505 (0.899)	4.283 (0.892)	5.517 (0.854)
United Kingdom	5.795 (0.016)	5.807 (0.054)	6.040 (0.110)	7.496 (0.117)	9.106 (0.105)	9.273 (0.159)	9.274 (0.234)	10.605 (0.225)	10.725 (0.295)	17.321 (0.068)
Spain	2.236 (0.627)	4.089 (0.129)	4.558 (0.207)	4.626 (0.328)	6.284 (0.280)	6.286 (0.392)	7.155 (0.413)	7.212 (0.514)	14.861 (0.095)	18.931 (0.041)
Austria	0.104 (0.741)	0.518 (0.772)	3.041 (0.385)	4.131 (0.389)	9.817 (0.081)	10.125 (0.120)	10.975 (0.140)	11.352 (0.183)	12.418 (0.191)	13.178 (0.214)

Note: Statistic is distributed as the chi-squared with *L* degrees of freedom. Values in brackets correspond to significance levels. Results were performed with *Limdep 7.0*.

TABLE 4. Estimates of constant term and country-fixed effects for the period 1991:1 to 1995:8.

Constant and destination country effect (i)	Coefficient (P-value)
α	2.341
	(0.000)
$\lambda_{\text{Belgium-Luxembourg}}$	0.013
	(0.388)
$\lambda_{\text{Netherlands}}$	0.013
	(0.371)
λ_{Italy}	-0.055
	(0.000)
$\lambda_{\text{United Kingdom}}$	-0.012
	(0.407)
λ_{Spain}	-0.535
	(0.000)
λ_{Austria}	0.022
	(0.142)

Note: The reference group is France. A positive coefficient means that export prices are on average more expensive than the reference group. The t-statistics. which are in brackets, are robust to heteroscedasticity. The estimates were performed with *Limdep 7.0*.

The residuals obtained from the most recent period were also analyzed using the values of the Box-Pierce and the Box-Ljung statistics. Table 5 and 6 includes results for these statistics, respectively. As we can see, results contrast notably with those obtained over the admitted price coordination period. We now reject, at the 1 per cent level, a white noise process for residuals in all destination countries in a very clear way. It can therefore be inferred that residuals capture a deterministic component and, consequently, time and destination effects are insufficient to explain the German pricing behavior. The presence of a deterministic component supports the presence of non-cooperative strategies on prices in an imperfect competition framework.

5. Concluding remarks

The main purpose of this paper was to provide a simple and useful method with which to investigate the presence of illegal price agreements in international markets. The hypothesis of price collusion has been stated throughout our analysis. We divided the other market structures into two alternative hypotheses, namely, the perfectly competitive market hypothesis, and a wide range of market structures where there are changes in competitiveness which affect optimal prices. We suggest a simple empirical model in which each of these alternatives can be chosen with little statistical information being required. The approach is based on the estimation of a two-factor fixed-effects model and on the analysis of residuals.

The European market for vitamin C provides an excellent framework in which to explore the usefulness of an empirical approach and to obtain evidence about price behavior in recent years in scenarios where we expect competitive behavior to have occurred. To achieve this, we use a variable proxy for the German export prices over two different periods. First, a period of general agreements of price collusion is admitted by participating firms (from January 1991 to August 1995). Second, there is another later period where information reported by firms indicates that there was no price collusion (from September 1995 to September 2001). Empirical results obtained about the first period were consistent with the fulfillment of the non-competitive price hypothesis for most of the destination markets. More specifically, we found a great deal of evidence for price parallelism within international markets and discriminatory pricing across destinations. These empirical results are radically different to those obtained from the last period. We infer that, in the most recent period, the mark-up of the German firms was reacting to changes in international competitiveness. This finding coincides with the withdrawal of the vitamin C cartel. We hope that the empirical approach suggested in this paper helps to identify future illegal pricing behaviors and to hinder their growth.

TABLE 5. White noise test for residuals based on the *Box-Pierce* statistic for the period 1995:9 to 2001:9.

Lags(L) Destination country(i)	1	2	3	4	5	6	7	8	9	10
France	30.801 (0.000)	48.479 (0.000)	64.413 (0.000)	79.405 (0.000)	90.940 (0.000)	101.989 (0.000)	112.450 (0.000)	121.859 (0.000)	132.291 (0.000)	141.840 (0.000)
Belgium-Luxembourg	22.255 (0.000)	32.385 (0.000)	34.822 (0.000)	35.780 (0.000)	38.170 (0.000)	41.164 (0.000)	46.240 (0.000)	46.968 (0.000)	48.017 (0.000)	48.375 (0.000)
Netherlands	32.591 (0.000)	65.320 (0.000)	94.424 (0.000)	123.863 (0.000)	152.575 (0.000)	175.166 (0.000)	199.529 (0.000)	216.480 (0.000)	235.786 (0.000)	249.430 (0.000)
Italy	17.306 (0.000)	35.006 (0.000)	52.689 (0.000)	74.519 (0.000)	87.046 (0.000)	95.853 (0.000)	109.144 (0.000)	122.062 (0.000)	129.752 (0.000)	132.304 (0.000)
United Kingdom	5.886 (0.015)	11.145 (0.004)	14.563 0.002)	22.111 (0.000)	36.496 (0.000)	32.529 (0.000)	40.485 (0.000)	42.614 (0.000)	48.452 (0.000)	53.781 (0.000)
Spain	35.856 (0.000)	63.042 (0.000)	79.222 (0.000)	94.405 (0.000)	111.519 (0.000)	130.025 (0.000)	147.767 (0.000)	159.817 (0.000)	48.452 (0.000)	170.792 (0.000)
Austria	5.491 (0.019)	11.399 (0.003)	18.297 (0.000)	24.722 (0.000)	35.486 (0.000)	41.588 (0.000)	50.378 (0.000)	52.825 (0.000)	57.597 (0.000)	60.642 (0.000)

Note: Statistic is distributed as the chi-squared with *L* degrees of freedom. Values in brackets correspond to significance levels. Results were performed with *Limdep 7.0*.

TABLE 6. White noise test for residuals based on the *Box-Ljung* statistic for the period 1995:9 to 2001:9.

Lags(L) Destination country(i)	1	2	3	4	5	6	7	8	9	10
France	32.084 (0.000)	50.758 (0.000)	67.830 (0.000)	84.126 (0.000)	99.848 (0.000)	109.217 (0.000)	121.105 (0.000)	131.961 (0.000)	144.186 (0.000)	155.554 (0.000)
Belgium-Luxembourg	23.182 (0.000)	33.883 (0.000)	36.494 (0.000)	37.557 (0.000)	40.171 (0.000)	43.523 (0.000)	49.291 (0.000)	50.131 (0.000)	51.360 (0.000)	51.787 (0.000)
Netherlands	33.949 (0.000)	68.522 (0.000)	99.705 (0.000)	131.704 (0.000)	163.372 (0.000)	188.660 (0.000)	216.344 (0.000)	235.904 (0.000)	258.528 (0.000)	274.771 (0.000)
Italy	18.028 (0.000)	36.724 (0.000)	55.670 (0.000)	79.398 (0.000)	93.216 (0.000)	103.074 (0.000)	118.177 (0.000)	133.082 (0.000)	142.094 (0.000)	145.133 (0.000)
United Kingdom	6.131 (0.013)	11.687 (0.003)	15.349 0.001)	23.553 (0.000)	39.484 (0.000)	35.044 (0.000)	44.0174 (0.000)	46.473 (0.000)	53.351 (0.000)	59.659 (0.000)
Spain	37.350 (0.000)	66.067 (0.000)	83.403 (0.000)	99.907 (0.000)	118.782 (0.000)	139.498 (0.000)	159.659 (0.000)	173.563 (0.000)	53.315 (0.000)	186.527 (0.000)
Austria	5.720 (0.017)	11.961 (0.002)	19.351 (0.000)	26.335 (0.000)	57.850 (0.000)	38.207 (0.000)	55.026 (0.000)	57.850 (0.000)	63.442 (0.000)	67.067 (0.000)

Note: Statistic is distributed as the chi-squared with *L* degrees of freedom. Values in brackets correspond to significance levels. Results were performed with *Limdep 7.0.*

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